TOO MUCH FINANCE FOR WHOM?
RETHINKING THE CREDIT–OUTPUT NEXUS

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ABSTRACT

TOO MUCH FINANCE FOR WHOM?
RETHINKING THE CREDIT–OUTPUT NEXUS

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This dissertation asks whether a single coefficient or a universal threshold can describe the relationship between credit and economic performance, and concludes that neither can. Two linked investigations are conducted. The first estimates the association between the pace of credit expansion and the sectoral composition of value added in an unbalanced annual panel covering 1992 to 2019 and up to 186 countries, comparing pre-crisis, post-crisis and full-period results and, in a separate partition, low-, middle- and high-income economies. Because the credit variables are integrated of order one while the sectoral value-added shares are stationary, credit enters in first differences, so the estimated nonlinearities lie in credit-growth rather than credit-depth space. Estimation uses fixed effects with Driscoll–Kraay standard errors and two-step System GMM, with functional form selected sequentially and turning points tested formally. The second builds a matched quarterly panel for Türkiye, pairing total cash credit with value added across eleven activity groups from 2005 to 2026, and estimates a cross-sectionally augmented error-correction model. Sectoral responses differ qualitatively, not only in magnitude. In the post-crisis decade faster credit growth accompanies a falling industrial share and a rising service

share, while the industrial maximum shifts leftward as income rises, falling below zero in high-income economies. In Türkiye a negative average short-run coefficient conceals an inverted U peaking near 7.23 billion Turkish liras, and sectoral thresholds differ by more than an order of magnitude. Credit policy should therefore address the destination, timing and horizon of lending alongside its volume.

Keywords: financial development, credit growth, sectoral value added, nonlinearity, structural transformation

ÖZ

KİMİN İÇİN ÇOK FAZLA FİNANS?
KREDİ-ÇIKTI BAĞLANTISINI YENİDEN DÜŞÜNMEK

HUYUGÜZEL TÜRK, Nurbanu
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Bu tez, kredi ile ekonomik performans arasındaki ilişkinin tek bir katsayı ya da evrensel bir eşikle tanımlanıp tanımlanamayacağını sormakta ve her ikisinin de mümkün olmadığı sonucuna varmaktadır. Birbirine bağlı iki inceleme yürütülmektedir. Birincisi, kredi genişlemesinin hızı ile katma değerin sektörel bileşimi arasındaki ilişkiyi 1992–2019 dönemini ve 186 ülkeye kadar geniş bir örneklemi kapsayan dengesiz yıllık bir panelde tahmin etmekte, kriz öncesi, kriz sonrası ve tüm dönem sonuçlarını karşılaştırmakta, ayrıca ekonomileri düşük, orta ve yüksek gelirli gruplara ayıran ikinci bir bölümleme üzerinden de karşılaştırma yapmaktadır. Kredi değişkenleri birinci dereceden bütünleşik, sektörel katma değer payları ise durağan olduğundan kredi modellere birinci farklarıyla girmekte, dolayısıyla tahmin edilen doğrusal olmayışlar kredi derinliği değil kredi büyümesi uzayında konumlanmaktadır. Tahminler, Driscoll–Kraay standart hatalarıyla sabit etkiler ve iki aşamalı Sistem GMM ile yürütülmekte, fonksiyonel biçim sırayla seçilmekte ve dönüm noktaları biçimsel olarak sınanmaktadır. İkinci inceleme, 2005–2026 döneminde on bir faaliyet grubunda toplam nakdî krediyi katma değerle eşleştiren çeyreklik bir Türkiye paneli kurmakta ve yatay kesitle genişletilmiş bir

hata düzeltme modeli tahmin etmektedir. Sektörel tepkiler yalnızca büyüklük bakımından değil niteliksel olarak da ayrışmaktadır. Kriz sonrası on yılda daha hızlı kredi artışı gerileyen bir sanayi payına ve yükselen bir hizmet payına eşlik etmekte, sanayideki maksimum ise gelir yükseldikçe sola kayarak yüksek gelirli ekonomilerde sıfırın altına düşmektedir. Türkiye'de negatif bir ortalama kısa dönem katsayısı, yaklaşık 7,23 milyar Türk lirasında tepe yapan ters U biçimli bir ilişkiyi gizlemekte ve sektörel eşikler bir büyüklük mertebesinden fazla farklılaşmaktadır. Bu nedenle kredi politikası, hacmin yanı sıra kredinin gittiği alanı, zamanlamayı ve ufku da dikkate almalıdır.

Anahtar Kelimeler: finansal gelişme, kredi büyümesi, sektörel katma değer, doğrusal olmayış, yapısal dönüşüm

To my elementary family, who are the reason for my existence

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LIST OF ABBREVIATIONS

Abbreviation	: Full Form
ADF	: Augmented Dickey–Fuller
AIC	: Akaike Information Criterion
AR	: Autoregressive
ARDL	: Autoregressive Distributed Lag
BRSA	: Banking Regulation and Supervision Agency of Türkiye
CADF	: Cross-Sectionally Augmented Dickey–Fuller
CCE	: Common Correlated Effects
CD	: Cross-Sectional Dependence (Pesaran Test Statistic)
CIPS	: Cross-Sectionally Augmented Im–Pesaran–Shin
COVID-19	: Coronavirus Disease 2019
CPIA	: Country Policy and Institutional Assessment
CS-ARDL	: Cross-Sectionally Augmented Autoregressive Distributed Lag
CS-DL	: Cross-Sectionally Augmented Distributed Lag
CS-ECM	: Cross-Sectionally Augmented Error Correction Model
CSD	: Cross-Sectional Dependence
DCCE	: Dynamic Common Correlated Effects
ECT	: Error Correction Term
EU	: European Union
FAO	: Food and Agriculture Organization of the United Nations
FE	: Fixed Effects
GDP	: Gross Domestic Product
GFD	: Global Financial Development Database
GMM	: Generalised Method of Moments
GNI	: Gross National Income
IDA	: International Development Association
IMF	: International Monetary Fund
IPS	: Im–Pesaran–Shin

ISIC	:	International Standard Industrial Classification of All Economic Activities
IV	:	Instrumental Variable
MG	:	Mean Group
NACE	:	Nomenclature statistique des Activités économiques dans la Communauté Européenne
OECD	:	Organisation for Economic Co-operation and Development
PMG	:	Pooled Mean Group
SBIC	:	Schwarz Bayesian Information Criterion
TFP	:	Total Factor Productivity
TL	:	Turkish Lira
TP	:	Turning Point
TurkStat	:	Turkish Statistical Institute
WDI	:	World Development Indicators

CHAPTER 1

INTRODUCTION

The relationship between financial development and economic performance has occupied economists for more than a century, and it retains its prominence because the answer bears directly on how societies organise the allocation of capital. Schumpeter (1911/1934) placed the banker at the centre of the development process, arguing that credit is what allows the entrepreneur to command resources before those resources have been earned, and the modern literature has elaborated that intuition into a set of functions that financial systems perform: mobilising savings, reducing information and transaction costs, monitoring investment, diversifying risk and directing funds towards their highest-return uses (Levine, 2005). Cross-country evidence assembled over three decades broadly supported this account. King and Levine (1993) reported that measures of financial depth predict subsequent growth, capital accumulation and productivity improvement, Levine and Zervos (1998) found that banking and stock market development each carry independent explanatory power, Rajan and Zingales (1998) demonstrated that industries dependent on external finance grow disproportionately faster in financially developed economies, Beck et al. (2000) identified total factor productivity growth rather than physical capital accumulation as the more robust channel through which financial intermediation reaches real activity, and Levine et al. (2000) showed that the exogenous component of financial intermediary development, isolated through differences in legal origin, is positively associated with long-run growth. Read collectively, these contributions established financial development as a determinant of long-run economic performance rather than as a passive reflection of it.

That body of work was widely interpreted as describing a monotonic relationship, in which deeper financial systems deliver better real outcomes at every level of

development, and the interpretation shaped both academic practice and several decades of financial liberalisation policy. The global financial crisis of 2008 exposed the fragility of that presumption. Rapid credit expansion in the years preceding the crisis coincided with the misallocation of resources, the accumulation of leverage, the growth of speculative positions and eventually with severe contractions in real activity, prompting a reorientation of the research agenda from whether finance promotes growth towards the conditions, the magnitudes and the horizons over which it does so.

The literature that emerged from this reorientation is usually organised around the 'too much finance' hypothesis, which holds that the marginal contribution of financial development is not constant across levels of financial deepening. Cecchetti and Kharroubi (2012) reported that the growth contribution of financial sector size becomes negative beyond a point, Arcand et al. (2015) located a threshold near the point at which private credit approaches the size of the economy, Law and Singh (2014) reached a comparable conclusion using a dynamic panel threshold estimator, and Samargandi et al. (2015) documented an inverted U-shaped pattern in a sample of middle-income economies. Several mechanisms have been proposed to account for the reversal. Once financing constraints have been relaxed, additional lending may flow towards consumption, property or the refinancing of existing obligations rather than towards productive investment, which makes the composition of credit consequential in its own right, since enterprise lending is associated with growth while household lending is not (Beck et al., 2012). Cecchetti and Kharroubi (2015) identify a resource channel, arguing that a rapidly expanding financial sector draws skilled labour away from other activities and favours projects that are easy to collateralise over those that are more productive, with research-intensive industries and firms holding few tangible assets suffering the largest losses. Continued lending to weak borrowers may delay restructuring and depress industry-level productivity, as Caballero et al. (2008) documented for Japan, while at the macroeconomic level rapid credit and leverage growth raise the probability of systemic crises (Schularick & Taylor, 2012). The thresholds themselves, however, prove sensitive to samples, periods and estimators, and Botev et al. (2019) find no universal tipping point beyond which financial development is invariably harmful, suggesting that the

relationship is nonlinear and context-dependent rather than governed by a single global parameter.

A second complication has attracted considerably less attention, and it turns out to be decisive for the analysis presented here. Almost the whole of the nonlinear literature treats financial development as a stock, measured by the ratio of outstanding credit to GDP, yet credit ratios are highly persistent series whose empirical properties rarely satisfy the conditions under which level-on-level regressions can be interpreted. Where two trending series are regressed on one another, statistically significant coefficients may reflect common trends rather than economic relationships, a hazard identified in the classic treatment of spurious regression by Granger and Newbold (1974). Second-generation panel unit-root procedures that allow for cross-sectional dependence, in particular the cross-sectionally augmented test of Pesaran (2007), provide the appropriate diagnostic in macroeconomic panels where common shocks are pervasive. Applying that test to the data assembled for this dissertation yields an unambiguous result: the sectoral value-added shares of agriculture, industry and services are stationary in levels, whereas private credit by deposit money banks and credit to agriculture, forestry and fishing are integrated of order one. The credit variables consequently enter every estimated equation in first differences, and the parameter of interest ceases to describe financial depth, describing instead the pace of credit expansion.

This substitution changes the object of analysis rather than merely the units in which it is expressed, and a substantial body of work indicates that stocks and flows of credit behave differently. Loayza and Rancière (2006) confronted the apparent contradiction between a growth literature reporting positive effects of financial depth and a crisis literature reporting that credit aggregates are among the best predictors of banking crises, concluding from an encompassing model of long- and short-run effects that a positive long-run relationship between intermediation and output growth coexists with a largely negative short-run relationship linked to indicators of financial fragility. Beck et al. (2014) reach a compatible conclusion by separating the intermediation function of financial systems from their overall size, finding that intermediation activities raise growth and reduce volatility over long horizons while

expansion along other dimensions has no long-run real effect. The most direct evidence comes from Ductor and Grechyna (2015), who show that the growth effect of financial development depends on the rate of private credit expansion relative to real output growth and turns negative where credit outpaces the real economy, and from Bezemer et al. (2016), who distinguish credit stocks from credit flows and report positive growth effects for flows to non-financial business but not for mortgage and other asset-market lending.

Evidence from long-run macro-financial history reinforces the case for treating the pace of credit expansion as an object of interest in its own right. Schularick and Taylor (2012) established credit growth as a leading indicator of financial crises, Mendoza and Terrones (2008) documented the macroeconomic and firm-level regularities that accompany credit booms, and Jordà et al. (2013) showed that more credit-intensive expansions tend to be followed by deeper recessions and slower recoveries. Borio et al. (2016) extend the argument into productivity, finding that credit booms tend to reallocate labour towards sectors with lower productivity growth – a temporarily inflated construction sector being the leading example – and that the resulting damage is far greater where a crisis follows. Müller and Verner (2024) document a parallel asymmetry in the allocation of credit itself, showing that booms disproportionately favour non-tradable activities and that such expansions are systematically associated with subsequent slowdowns. One interpretive consequence follows immediately and governs the whole of this dissertation. A turning point recovered from a specification in first differences is located in credit-growth space, expressed as a rate of change, and it is therefore not comparable with the thresholds of eighty to one hundred per cent of GDP reported by Arcand et al. (2015), Law and Singh (2014) or Cecchetti and Kharroubi (2012). A finding that the marginal contribution of credit turns negative beyond some rate of expansion speaks to the speed at which sectors can absorb new lending, not to the level of financial depth at which an economy becomes overbanked, and conflating the two would misstate both the result and its policy content.

Nonlinearity and the choice of margin are nonetheless only part of the difficulty. Conventional finance–growth studies aggregate on the financial side, and an

economy-wide credit measure pools lending to activities that differ in technology, investment horizon, collateral structure, working-capital requirement and capacity to generate value added. Credit extended for agricultural modernisation, industrial machinery, information technology, construction, financial intermediation or household consumption cannot plausibly be expected to produce equivalent real effects, and the point is not resolved by moving from levels to growth rates, since an aggregate rate of credit expansion averages over destinations exactly as an aggregate stock does. The allocation literature has made this concrete from several directions, with Beck et al. (2012) showing that the cross-country association between credit and growth is driven mainly by enterprise lending, Mian et al. (2017) finding that expansions in household debt predict medium-term slowdowns, Bahadir and Gumus (2016) demonstrating that the cyclical signature previously attributed to aggregate credit is considerably stronger for household than for business borrowing, and Bai et al. (2018) showing that a shift in the composition of bank credit towards more productive borrowers raised aggregate industry productivity following state banking deregulation in the United States. Hsu et al. (2014) add a structural dimension, reporting that equity-market development supports innovation in externally dependent and technology-intensive industries whereas credit-market development does not, which implies that a predominantly bank-based system may finance conventional capital accumulation effectively while remaining a weaker instrument for innovation-intensive activity.

Aggregation on the output side introduces a symmetrical problem. Total output combines sectors that respond differently to a common financial shock, so expansion in one activity may offset contraction in another and leave the aggregate largely unchanged, while shifts in the composition of production remain invisible within a single growth rate. An aggregate specification cannot therefore identify which sectors benefit from faster credit expansion, which encounter diminishing returns and which may already have exhausted their capacity to absorb additional lending productively, nor can it establish whether an observed aggregate nonlinearity originates within sectors or arises from a reallocation of credit between them. This dissertation addresses the problem by taking sectoral value added as the dependent variable, and a clarification of measurement is required at the outset, since the cross-

country models use sectoral value added expressed as a share of GDP. A change in such a share indicates a change in the relative weight of an activity rather than in its absolute output, so a negative coefficient implies that a sector grew more slowly than the rest of the economy rather than that it contracted, an interpretation the structural transformation literature treats as a normal feature of development rather than as evidence of sectoral failure.

The structural transformation literature also explains why the distinction matters analytically. Growth involves not only an increase in total output but a reallocation of labour, capital and production from agriculture towards industry and subsequently towards services (Herrendorf et al., 2014; McMillan & Rodrik, 2011), and financial development may shape that process by altering the relative financing conditions facing each sector. Capital-intensive industry typically depends on long-horizon external funding, agriculture requires instruments suited to seasonal cycles, biological lags and covariate risk, and many modern services rest on intangible assets, human capital and innovation rather than on collateralisable plant. A financial system well adapted to one phase of structural change need not serve the next equally well, and the phase itself has become harder to traverse, since Rodrik (2016) documents that the hump-shaped relationship between industrialisation and income has shifted downwards and closer to the origin, so that economies now exhaust their industrialisation opportunities sooner and at lower income levels.

Heterogeneity across countries compounds these considerations. Converting credit into productive value added depends on institutional quality, financial inclusion, technological capacity, human capital, infrastructure and the effectiveness of supervision, and the binding constraint differs by stage of development. Greenwood and Jovanovic (1990) modelled the joint evolution of financial structure and income distribution, Rioja and Valev (2004a) argued that the effect of financial development is uncertain at very low levels, large at intermediate levels and positive but smaller at high levels, and Demetriades and Law (2006) reached a compatible conclusion through an institutional lens, reporting that finance exerts its strongest influence in middle-income economies and its weakest in low-income ones. Aizenman et al. (2015) extend the argument to the sectoral level, documenting substantial sectoral

and regional heterogeneity in the effect of financial depth on output growth, including negative effects in several sectors. Sahay et al. (2015) add the consideration that bears most directly on a differenced specification, cautioning that the speed of financial deepening carries risks of its own once it outpaces supervisory and regulatory capacity.

Temporal context matters for the same reason. The three decades covered by the cross-country analysis encompass substantial changes in financial regulation, in the destination of bank lending, in the technological environment and in the organisation of production, and the years following 2008 brought extensive regulatory reform, unconventional monetary policy, altered bank risk appetite and a growing policy interest in the composition rather than merely the volume of credit. Rousseau and Wachtel (2011) document a weakening – and in some specifications an apparent disappearance – of the estimated effect of financial deepening on growth in more recent samples, which suggests that estimating a single coefficient across a long window may average over regimes in which finance operated differently.

Four propositions therefore need to be examined jointly rather than in isolation. The relationship between credit and real activity may be nonlinear, its direction and intensity changing as expansion accelerates. The margin at which finance is measured is consequential, since the accumulated stock of credit and its rate of expansion are distinct objects with distinct empirical properties. The composition of credit and of output is decisive, because aggregate indicators conceal sector-specific dynamics that may differ in sign as well as in magnitude. And the influence of finance is conditional on context, varying across stages of development, historical periods and institutional settings. The central claim advanced in this dissertation follows from taking all four seriously: no single coefficient, and no universal credit threshold, can characterise the relationship between financial development and economic performance, which must instead be evaluated by sector, by stage of development, by period and by time horizon.

The empirical work is organised around a set of connected questions. Does the pace of credit expansion bear a stable relationship to sectoral value added, or does its

marginal contribution change as expansion accelerates? Are the linear and nonlinear patterns observed in agriculture, industry and services common across sectors or specific to each? Did the relationship change after the global financial crisis, and if so in which direction and for which activities? Do economies at different income levels display different functional forms and different turning points? Do the estimated relationships survive the introduction of dynamic persistence and the instrumentation of potentially endogenous credit? Turning to the national evidence, how do changes in sectoral credit relate to short-run output growth within a single economy, what long-run equilibrium relationship links sectoral credit levels to sectoral activity, and do sectors differ systematically in their capacity to absorb additional credit productively? A final question runs across the two settings and follows from the integration properties of the data, asking whether the pace of credit expansion and the level of financial depth are separable margins that carry different implications for real activity.

The dissertation addresses these questions through two linked empirical investigations. The first provides broad international coverage, and the second supplies national depth. This pairing combines the comparative reach of a cross-country panel with the institutional and sectoral resolution available within a single economy, allowing the international analysis to establish whether systematic differences exist across major sectors, income groups and financial environments while the Türkiye analysis examines sectoral heterogeneity at finer resolution and separates short-run adjustment from long-run equilibrium. Results that appear in both settings carry wider significance, whereas divergences point towards the influence of national institutions, policy design and sectoral composition.

The international analysis employs an unbalanced annual panel spanning 1992 to 2019 and covering economies from every World Bank income category (World Bank, 2026b), with up to 186 countries in the industrial and service models and 109 countries in the agricultural models, where coverage is limited by the availability of sector-specific credit data. The sample ends in 2019 so that the exceptional and unevenly distributed disruptions of the COVID-19 pandemic remain outside the estimation window, since those years combined simultaneous shocks to production,

credit supply, public expenditure and the composition of output. The full period is supplemented by separate estimations for 1992–2007 and 2010–2019, allowing the pre-crisis expansion to be compared with the decade that followed, while 2008 and 2009 are excluded from the subperiods because they correspond to the most acute phase of the crisis and to the immediate policy response and cannot be assigned to either regime.

Economic performance is decomposed into three sectors following the International Standard Industrial Classification (United Nations, 2008), with the dependent variables defined as the natural logarithms of value added in agriculture including forestry and fishing, in industry including construction and in services, each expressed as a share of GDP. Financial development is measured according to availability and sectoral relevance. The agricultural models use credit to agriculture, forestry and fishing normalised by GDP, which links financing directly to the activity being financed, whereas the industrial and service models use private credit by deposit money banks relative to GDP, since sector-specific credit data are not available for a broad international sample. All credit measures enter in first differences of their logarithms for the reasons set out above. The specifications additionally include secondary school enrolment, gross capital formation, government consumption, inflation and trade openness, capturing human capital, physical investment, the scale of public activity, macroeconomic stability and external integration, with secondary school enrolment likewise entering in first differences because its logarithmic level series also contains a unit root.

The estimation strategy responds to the diagnostic properties of the panel, which exhibits cross-sectional dependence, slope heterogeneity, heteroskedasticity, serial correlation and mixed orders of integration. Baseline models are accordingly estimated with country and year fixed effects and Driscoll–Kraay standard errors, which remain reliable under heteroskedasticity, serial correlation and general forms of dependence across countries (Driscoll & Kraay, 1998). Two-step System GMM estimates following Blundell and Bond (1998), with the finite-sample correction of Windmeijer (2005), address output persistence, reverse causality and omitted-variable endogeneity, and their validity is assessed through the Arellano–Bond tests

for serial correlation (Arellano & Bond, 1991) together with the Hansen, Sargan and Difference-in-Hansen statistics (Hansen, 1982; Sargan, 1958), with explicit attention to the loss of test power that accompanies instrument proliferation (Roodman, 2009). Functional form is selected from the data rather than imposed, with linear, quadratic, cubic and quartic specifications estimated in sequence and retained only where the additional term is statistically significant and improves fit. A significant squared term does not by itself establish a genuine turning point, so quadratic specifications are subjected to the test proposed by Lind and Mehlum (2010), which requires the slope to be positive at the lower bound of the data and negative at the upper bound with the turning point located inside the observed interval, while cubic specifications are evaluated through the roots and the discriminant of their first derivative so that a curve is described as N-shaped or inverted-N-shaped only where two real turning points can actually be identified.

A preliminary aggregate analysis motivates the sectoral design rather than serving as a finding in its own right. Estimating GDP per capita growth on the first difference of log private credit produces a consistently concave relationship, with a negative and highly significant squared term in every period, and the turning point calculations locate the growth-maximising rate of credit expansion at −0.023 in logarithmic terms for 1992–2007 and at −0.276 for the full period, corresponding to contractions of roughly two and twenty-four per cent in the credit-to-GDP ratio. Within-country explanatory power nonetheless remains modest, ranging from 0.059 to 0.122, which is itself informative, indicating that a single economy-wide credit measure captures only a fraction of the mechanisms through which finance may reach real activity and that disaggregation is required to identify them.

The sectoral results confirm that expectation, since the three sectors describe a pattern that no aggregate specification could recover. Agriculture displays no stable association between the pace of credit expansion and the agricultural share of value added in either subperiod, and the only specification in which the credit terms attain significance – the full-period quadratic estimated with Driscoll–Kraay standard errors – proves fragile on inspection. The Lind and Mehlum (2010) test rejects monotonicity at the ten per cent level and turning point calculations place the

maximum at 0.502, equivalent to annual agricultural credit growth of approximately sixty-five per cent, yet the curvature reverses sign once macroeconomic controls are introduced and the Fieller (1954) interval for the turning point is wide enough to accommodate both negative values and implausibly rapid expansions. The defensible reading is that agricultural credit growth exhibits at most a weak and imprecisely located nonlinearity, which accords with the argument that the effectiveness of agricultural finance depends on the institutional conditions of rural credit markets – transaction costs, collateral institutions, insurance provision and exposure to covariate risk – rather than on lending volumes alone (Conning & Udry, 2007).

Industry is the only sector in which both estimators produce statistically significant credit coefficients in all three periods, and the shape of the relationship changes across financial regimes in a way that carries substantive content. In the pre-crisis window the static estimator returns a weakly positive linear coefficient that does not survive the addition of controls, while the dynamic estimator identifies a significant negative cubic term implying a local maximum at 0.510, or annual private credit growth of about sixty-seven per cent, a rate located just below the ninety-ninth percentile of the pre-crisis distribution. The post-crisis decade produces the clearest results in the periodical analysis, with a cubic specification whose coefficients remain stable across the baseline and extended models and whose turning points at −1.642 and 0.246 are firmly identified by a significant discriminant. Since the upper maximum corresponds to credit growth of roughly twenty-eight per cent and lies between the ninety-fifth and ninety-ninth percentiles, the interval over which the marginal association is negative encompasses almost the entire mass of the data, so faster credit expansion accompanied a lower industrial share across the range in which countries ordinarily operate. Over the full period both estimators agree on the broad shape, the static model locating a local maximum at 0.216 and the dynamic model at 0.156, corresponding to credit growth of approximately twenty-four and seventeen per cent respectively.

Services move in the opposite direction, and the contrast is the central empirical finding of the periodical analysis. Neither the pre-crisis window nor the full sample yields significant credit coefficients under either estimator, but in the post-crisis

decade the linear term is positive and highly significant in both the baseline and the extended specification, establishing that faster private credit growth accompanied a higher service share of value added. The eventual reversal implied by the fitted cubic lies at 0.405, beyond the ninety-ninth percentile of the distribution, so the relationship is positive and rising over virtually the whole of the observed range and its curvature cannot be characterised with confidence. What matters is that the same regressor, the same denominator and the same decade produce a negative coefficient in industry and a positive coefficient in services, an asymmetry that cannot be attributed to scaling and points instead towards allocation, with post-crisis lending plausibly directed disproportionately towards property, consumption and service activities rather than towards industrial capacity. The evidence assembled by Beck et al. (2012) and Bezemer et al. (2016) on the differential real effects of credit by borrower type supports this reading, although testing it directly would require disaggregated credit data that the cross-country sample does not contain.

Partitioning the sample by income level confirms that no single finance–growth curve applies across stages of development. Agricultural credit growth registers only in low-income economies, where a weakly supported inverted U places the maximum at 0.414 and the great majority of observations lie on its upward-sloping segment, implying that faster agricultural credit growth tends to accompany higher agricultural value added across the range in which these economies actually operate, an outcome consistent with the persistence of binding constraints in smallholder finance (World Bank, 2007). Industry registers in all three groups but with different functional forms and, more importantly, with the estimated maximum shifting steadily leftward as income rises. In low-income countries a cubic profile places the working range of the data on a declining middle segment, in middle-income countries the fitted curve is approximately concave over the region where observations are dense, with the static estimator locating the maximum at 0.309 and a dynamic quadratic specification placing it at 0.026 – almost exactly at the median rate of private credit growth in that group – and in high-income countries the maximum lies at −0.224, below zero, so that essentially the whole range of positive credit growth falls on the declining branch. The middle-income evidence provides the strongest support in this study for the ‘too much finance’ proposition, and it does

so in the sector for which that proposition has the clearest theoretical rationale, since manufacturing remains a substantial share of output in these economies while supervisory frameworks are still maturing (Cecchetti & Kharroubi, 2015; Sahay et al., 2015). Services, by contrast, respond only weakly and only in middle-income countries, where a positive linear coefficient loses significance once controls are added, and nothing in those estimates indicates a threshold beyond which further credit growth becomes detrimental.

The second empirical investigation turns to Türkiye, where the policy relevance of credit allocation is unusually visible. The sample period encompasses rapid financial deepening, repeated credit cycles, the expansion of publicly supported guarantee schemes, episodes of macroeconomic instability, a pronounced credit response to the pandemic and a growing reliance on selective lending and macroprudential instruments, which together make the country an informative case for examining whether the destination and scale of credit matter as much as its volume. The analysis constructs a quarterly panel of eleven sectors running from the first quarter of 2005 to the first quarter of 2026, yielding 902 observations. Sectoral total cash loans are drawn from the Banking Regulation and Supervision Agency (BRSA, 2026) and comprise both performing and non-performing receivables recorded by deposit, participation, development and investment banks, so that the measure captures the full exposure the banking system has committed to each activity rather than only its currently performing component. Because the credit series records outstanding stocks at monthly reporting dates, the final month of each quarter is taken as that quarter's value, a rule consistent with the accounting logic of a balance-sheet variable.

Sectoral value added is taken from the production-based national accounts of the Turkish Statistical Institute (TurkStat, 2026) and matched to the credit data through a concordance that assigns the supervisory credit categories, reported on NACE Rev. 2.1, to the NACE Rev. 2 activity groups used in the national accounts (Eurostat, 2008, 2025), with care taken to avoid overlapping categories, which required subtracting manufacturing from the broader industrial aggregate in order to construct a series for industry excluding manufacturing.

This design disaggregates both sides of the relationship, pairing the credit each activity receives with the value added that activity generates, which distinguishes the Türkiye chapter from the cross-country analysis and removes much of the aggregation problem that constrains conventional finance–growth models. It also permits both margins of financial development to be estimated within a single framework, which the cross-country chapter cannot do. The Turkish panel exhibits very strong cross-sectional dependence, since sectors share a common monetary, financial and macroeconomic environment and are linked through production chains and common demand conditions, while panel Granger-causality tests indicate bidirectional predictive feedback between sectoral credit and output. These features render a static fixed-effects specification inadequate and make System GMM unsuitable in this application, given that the time dimension substantially exceeds the number of sectors, that instrument proliferation would follow and that the common factors driving the dependence would remain unmodelled. The analysis therefore adopts a cross-sectionally augmented autoregressive distributed lag model in error-correction form, following the correlated effects approach of Pesaran (2006) and its dynamic heterogeneous extension by Chudik and Pesaran (2015), estimated by Mean Group so that sector-specific slopes are preserved. Including two lags of the cross-sectional averages removes residual cross-sectional dependence in the linear, quadratic and cubic specifications, and a negative and significant error-correction term in each confirms that deviations from long-run equilibrium are progressively corrected. The framework separates the short-run component, in which credit enters through changes in the logarithmic credit terms, from the long-run component, in which credit enters in levels, so the results speak to credit growth and credit depth respectively.

The Turkish results demonstrate the limits of any economy-wide summary. In the linear specification the Mean Group short-run coefficient is negative and significant, indicating that larger quarterly increases in sectoral credit are on average associated with weaker contemporaneous output growth, an average driven principally by agriculture, by trade, transport, accommodation and food services and by public and social services. The long-run Mean Group coefficient is positive but insignificant, and that insignificance conceals rather than summarises the sectoral evidence, since

industry excluding manufacturing, agriculture, the trade-related group and other services all return positive and significant long-run coefficients. The quadratic specification then shows the negative linear result to be an artefact of aggregation, since the short-run Mean Group relationship is inverted U-shaped, with credit changes supporting quarterly output growth up to a turning point of approximately 7.23 billion Turkish liras and their contribution turning negative beyond it. The sectoral turning points behind that average differ by more than an order of magnitude, running from roughly 2.46 billion liras in information and communication and 4.94 billion in agriculture through 8.68 billion in public and social services to 40.23 billion in the trade, transport and accommodation group, which indicates that the capacity to absorb credit varies substantially across activities within a single national economy. Long-run patterns are more heterogeneous still, with financial and insurance activities displaying an inverted U-shape whose maximum lies near 26.07 billion liras – a result that speaks directly to the ‘too much finance’ debate, since the sector in question is the financial sector itself – while agriculture and information and communication display U-shaped profiles in which credit appears costly at low levels and productive only once it reaches sufficient scale to finance modernisation, fixed investment or technological upgrading.

The cubic specification refines rather than overturns this account. At the Mean Group level the third-order term is insignificant, so the panel average remains best described by the single-threshold adjustment identified in the quadratic model, but several sectors display long-run trajectories that a single turning point cannot represent. Real estate follows a clear N-shaped path with turning points near 1.36 and 13.84 billion liras, a configuration consistent with the collateral feedback mechanism formalised by Kiyotaki and Moore (1997), in which durable assets serving simultaneously as factors of production and as collateral transmit and amplify shocks, though the terminal positive segment should not be read as benign given the association between credit-driven property booms and asset-price inflation. Manufacturing, the trade-related group and public and social services display inverted-N configurations in which credit is associated with weaker outcomes at low levels, supports activity over an intermediate range and becomes negatively associated again at very high levels, a pattern compatible in manufacturing with the

possibility that very large exposures sustain less efficient structures and delay adjustment (Caballero et al., 2008). Agriculture, industry excluding manufacturing and construction require the most careful treatment, since their coefficient sign sequences resemble N-shaped or inverted-N-shaped curves while the discriminants of their derivatives are negative, so the fitted functions possess inflection points rather than two genuine turning points and describe changes in the momentum of the relationship rather than reversals in its direction.

Read together, the two investigations support a common conclusion. The relationship between credit and real activity depends on the sector receiving the funds, on the pace at which credit expands relative to that sector's capacity to absorb it, on the stage of development, on the prevailing financial regime and on the horizon over which the effect is measured. Finance may relax binding constraints, fund working capital, enable technological upgrading and expand productive capacity, yet where expansion outruns absorptive capacity it may instead raise debt-service burdens, sustain low-productivity activity, encourage speculative positions or interact unfavourably with weak institutions. A single sector may accordingly display a negative short-run coefficient, a beneficial range below a quadratic threshold and a further nonlinear long-run trajectory without any contradiction among the three, since they describe different margins and different horizons of the same adjustment process.

The dissertation makes six contributions. The first is conceptual, joining the finance–growth and structural transformation literatures by treating financial development not merely as a determinant of aggregate growth but as a force capable of reshaping the sectoral composition of production. This reframing shifts the question from whether credit is beneficial in general towards how its benefits and costs are distributed across productive activities, and the empirical results give the reframing content, since faster credit expansion accompanied a rising service share and a falling industrial share in the same decade and in the same sample.

The second contribution concerns measurement, and it follows from taking the integration properties of the data seriously rather than treating them as a technicality. Because private credit and agricultural credit are integrated of order one while the

sectoral value-added shares are stationary, the credit variables enter in first differences and the estimated nonlinearities are located in credit-growth space rather than in credit-depth space. The distinction between the stock of credit and its rate of expansion is well established in the crisis and business-cycle literature, yet it has not been carried systematically into sectoral analyses of value added, where the two margins are frequently treated as interchangeable. The Türkiye chapter goes further by recovering both margins within one specification, retaining sectoral credit levels in the long-run component while modelling credit growth in the short-run component, which allows the divergence between them to be observed directly rather than inferred.

The third contribution is empirical. Agriculture, industry and services are examined within a common framework that allows relationships to differ across pre-crisis, post-crisis and full-sample periods as well as across income groups, and the Türkiye analysis adds resolution by matching eleven credit–output pairs at quarterly frequency over a period that includes the principal shifts in the country's financial and policy environment. Combining international breadth with national depth permits patterns common to many economies to be distinguished from those specific to particular institutional and policy configurations.

The fourth contribution is methodological. Functional form is selected sequentially rather than assumed, turning points are recovered from marginal effects rather than read off coefficient signs, quadratic profiles are subjected to formal slope tests and interpreted only where the maximum lies within the supported range of the data, and cubic profiles are evaluated through the roots and discriminant of the derivative so that an inflection point is not mistaken for a pair of turning points. Static estimates are compared with dynamic ones throughout, and the Turkish analysis addresses common factors, sectoral heterogeneity, bidirectional feedback and the separation of horizons explicitly. Locating each estimated turning point within the empirical distribution of the regressor is a further element of this discipline, since a statistically significant higher-order term whose turning points lie in the extreme tail describes a relationship that is, for practical purposes, monotonic over the range that countries actually experience.

The fifth contribution concerns the geometry of the nonlinearity rather than its existence. The results confirm that the relationship between credit and real activity departs from linearity, which is by now a familiar conclusion, but they also indicate that the departure does not take the form the ‘too much finance’ literature has made canonical. That literature is conventionally represented by an inverted U in which the contribution of finance rises, peaks at an interior threshold and declines thereafter (Arcand et al., 2015; Cecchetti & Kharroubi, 2012; Law & Singh, 2014; Samargandi et al., 2015). The aggregate estimates reported here are concave in every period examined, yet their maximum lies at or below zero, so the ascending branch of the fitted curve falls entirely within the range of credit contraction and the configuration cannot be read as one in which credit expansion supports growth up to a threshold and impedes it beyond. Disaggregation indicates why an aggregate specification may return a shape of this kind. Sectors differ not merely in the location of their turning points but in the order of the polynomial their data support and in the sign of the association itself, since industry requires a cubic form in all three periods of the international panel and carries a negative marginal association across almost the entire observed range after the crisis, whereas services in that same decade display a positive first-order effect whose reversal lies beyond the ninety-ninth percentile of credit growth. The selected functional form also varies across income groups within a single sector, and the Turkish evidence is more differentiated still, yielding an N-shaped long-run path in real estate, inverted-N configurations in manufacturing, in the trade-related group and in public and social services, and cubic functions whose derivatives possess no real roots in agriculture, in industry excluding manufacturing and in construction, where the correct description is an inflection point marking a change in momentum rather than a pair of regime reversals. Curvature may even reverse within one sector across horizons, since agriculture and information and communication combine a short-run inverted U with a U-shaped long-run profile. Two implications follow. An aggregate turning point is a compositional average of sectoral relationships differing in shape as well as in sign, so it need not correspond to the behaviour of any individual sector and may, as in the present estimates, be displaced outside the region of the data that carries economic interest. The appropriate object of inference is accordingly the shape of the relationship at a specified horizon rather than the location of a single economy-wide threshold, and

recovering that shape requires the functional-form discipline described above, since a cubic sign sequence read directly off the coefficients would have produced N-shaped and inverted-N-shaped labels for three Turkish sectors in which no such structure exists.

The sixth contribution is directed at policy. Aggregate credit targets and uniform lending rules presuppose that sectors convert finance into value added at broadly similar rates, and the evidence assembled here indicates that they do not. A rate of expansion that is beneficial in one activity may exceed the absorptive capacity of another, and a negligible economy-wide association may coexist with strong positive, negative or nonlinear relationships at the sectoral level. Effective financial policy should therefore attend not only to how much credit is extended but to where it goes – which sectors receive it, at what maturity, for what purpose and against what productive capacity – a conclusion that carries particular weight in economies where allocation is shaped substantially by guarantee schemes, state banks, targeted lending incentives and macroprudential instruments. Policies aimed simply at expanding credit may encounter diminishing returns in the absence of project selection, productivity-enhancing investment and institutional capacity, while restrictive rules anchored in aggregate indicators may constrain sectors whose productive financing needs remain unmet. Estimated turning points are not regulatory limits. They mark the ranges within which the marginal contribution of additional lending changes and within which the quality of allocation warrants closer scrutiny.

The argument advanced throughout is accordingly that financial development should be evaluated by the capacity of the financial system to support productive activity rather than by the size or the speed of credit creation. Volume continues to matter, particularly where financing constraints bind, but it must be assessed alongside composition, timing, institutional environment and the use to which borrowed funds are put. Examining how these dimensions interact across sectors, income levels, historical periods and time horizons yields a more granular account of the finance–growth relationship and a firmer empirical basis for sector-aware financial policy.

The remainder of the dissertation develops this argument in stages. Chapter 2 reviews the theoretical and empirical literature on the conventional finance–growth relationship, the emergence of the 'too much finance' hypothesis, the distinction between credit stocks and credit flows, credit composition and allocation, sectoral transmission, structural transformation, income-level heterogeneity and the Turkish financial environment. Chapter 3 presents the international dataset, the variable definitions and the descriptive evidence, together with the preliminary aggregate analysis that motivates the sectoral design, and Chapter 4 sets out the diagnostic testing sequence and the estimation strategy. Chapter 5 reports the periodical and income-level results for agriculture, industry and services, comparing the pre-crisis, post-crisis and full-period estimates before turning to the low-, middle- and high-income partitions. Chapter 6 constructs the matched Turkish sectoral panel, describing the sectoral credit and value-added series and the concordance that links them, Chapter 7 develops the cross-sectionally augmented error-correction framework together with the diagnostic sequence that motivates it, and Chapter 8 presents the short-run and long-run results in linear, quadratic and cubic form. Chapter 9 draws the two investigations together, considers their implications for theory and for policy, sets out the limitations of the analysis and identifies directions for further research.

CHAPTER 2

LITERATURE REVIEW

2.1. Financial Development and Economic Performance: Theoretical Foundations

The relationship between financial development and economic performance has occupied economists for more than a century, and the modern debate still rests on the intuition that intermediaries do something more than record real activity. Schumpeter assigned banks an active role in innovation, arguing that they identify entrepreneurs, evaluate projects and create the purchasing power that redirects resources toward new combinations (Schumpeter, 1911/1934). Later work translated this intuition into a general theory of intermediation, in which financial institutions mobilise dispersed savings, produce information about borrowers, monitor the use of funds, ease transactions, diversify risk and allocate capital, with the growth consequences of these functions depending on how well the system performs them rather than on how large it becomes (Levine, 2005).

Patrick (1966) supplied the causal vocabulary that still organises much of the empirical literature, distinguishing ‘supply-leading’ finance, in which institutions and instruments emerge ahead of the demand generated by real-sector expansion, from ‘demand-following’ finance, in which output, investment and commercial activity raise the demand for financial services and the system responds. He also anticipated that the balance between the two shifts along the development path, with supply-leading effects mattering most in early stages and demand-following responses becoming more prominent as economies mature. That framework remains directly relevant to a study comparing countries at different income levels and examining the short-run feedback between sectoral credit and output in Türkiye, since it implies that

the direction of the relationship is itself a parameter that may vary rather than a maintained assumption.

Formal growth models clarify the channels through which intermediation could affect long-run performance. Bencivenga and Smith (1991) show that intermediaries reduce households' need to hold liquid but unproductive assets, channelling a larger share of savings into productive investment. Greenwood and Jovanovic (1990) construct a model in which financial development and growth are jointly determined, since participation in more sophisticated financial arrangements improves risk sharing and allocation while the capacity to bear participation costs rises with income, generating feedback between the two processes and stage-dependent effects. Neither framework implies that every increment of credit is productive, and both locate the benefit in improved selection, monitoring, diversification or allocation rather than in the volume of claims as such.

The financial liberalisation literature introduces a complementary consideration. McKinnon (1973) and Shaw (1973) argued that administered interest rates, directed lending and related forms of financial repression depress saving, restrict intermediation and distort investment, so that liberalisation could raise growth by increasing loanable funds and improving their allocation. Subsequent experience suggests that removing quantitative restrictions is not sufficient when regulatory oversight, information infrastructure, contract enforcement and supervisory capacity remain weak, indicating that the institutional setting largely determines whether a larger financial system performs its functions or merely expands the volume of claims.

Financial development is a multidimensional concept encompassing the depth, access, efficiency and stability of institutions and markets (Čihák et al., 2012). This dissertation engages only with the quantitative dimension and does so through credit-based indicators. Private credit extended by financial intermediaries to the private sector, scaled by GDP, is the conventional proxy for banking-sector depth in the empirical finance–growth literature (Beck et al., 2000; Levine, 2005; Čihák et al., 2012), and the cross-country analysis adopts it alongside agricultural credit relative

to GDP for the agricultural equations. The Türkiye analysis measures sectoral financial development by the stock of total cash loans extended to each activity branch. What the term ‘financial development’ denotes here is therefore the scale and evolution of credit intermediation, not an assessment of the system’s quality, efficiency, inclusiveness or stability.

That restriction to a quantitative proxy carries an implication that the empirical chapters make explicit. Theory speaks predominantly about the level of intermediation, so that the natural empirical counterpart of an economy with a deeper financial system is a stock, measured at a point in time. The counterpart of an economy in which credit is expanding is a flow, measured as a rate of change, and the two objects need not carry the same sign or the same interpretation. Because the credit series used in the cross-country panel are integrated of order one, they enter the estimated equations in first differences, so the coefficients recovered in Chapters 4 and 5 describe the association between credit growth and sectoral outcomes rather than between credit depth and those outcomes. Section 2.4 develops this distinction at length, since it governs how the results reported later can legitimately be compared with the thresholds reported in the ‘too much finance’ literature.

Taken together, the theoretical and empirical literature supports four propositions that frame the analysis. Finance can influence real activity through capital accumulation, productivity growth, risk diversification and support for innovation (Schumpeter, 1911/1934; Levine, 2005). The relationship is plausibly bidirectional, since financial deepening may stimulate real activity while growth raises the demand for financial services (Patrick, 1966; Greenwood & Jovanovic, 1990). Its strength and its channels are likely to vary with the level of development and institutional quality (Rioja & Valev, 2004a, 2004b; Law et al., 2013).

The consequences of credit expansion depend not only on how much credit is available but on how it is distributed across borrowers, sectors and uses (Wurgler, 2000; Beck et al., 2012; Cecchetti & Kharroubi, 2015). These propositions motivate the move beyond a linear finance–growth specification toward the sectoral, nonlinear and context-dependent framework developed in this dissertation.

2.2. Early Empirical Evidence and the Direction of Causality

The modern empirical literature was shaped by cross-country studies that documented a robust positive association between financial development and subsequent growth. King and Levine (1993), using data on 80 countries between 1960 and 1989, found that several indicators of financial development were strongly related to real per capita income growth, physical capital accumulation and improvements in the efficiency with which capital is used. Their results lent empirical weight to the Schumpeterian view and established private credit and comparable balance-sheet measures as standard regressors in growth analysis. It is worth noting that this design regresses average growth over a long horizon on the initial level of financial depth, so its estimand is a level effect and not a response to credit expansion.

Later contributions addressed whether finance merely responds to growth. Levine et al. (2000) combined cross-sectional instrumental-variable methods with dynamic panel estimators and found that the exogenous component of financial intermediary development was positively and robustly associated with long-run growth, while stressing that their evidence did not exclude causality running from real activity to finance but showed that reverse causation alone could not account for the observed association. Beck et al. (2000) reached a complementary conclusion by decomposing growth into its sources, reporting that the finance–growth link operates mainly through total factor productivity, with estimates for saving and capital accumulation proving considerably less stable.

Evidence from equity markets widened the analysis beyond banking. Levine and Zervos (1998) showed that stock-market liquidity and banking development are each strongly associated with subsequent growth, capital accumulation and productivity improvements, with each entering independently of the other. That banks and markets are complements rather than substitutes is a cautious reading of this pattern rather than a hypothesis the study tests directly. This insight later proved important for research separating financial structure from financial depth and for studies suggesting that bank-based and market-based finance support innovation through different mechanisms.

Industry-level evidence offered a particularly influential identification strategy. Rajan and Zingales (1998) asked whether industries that depend more heavily on external finance for technological reasons grow disproportionately faster in countries with more developed financial systems, and found that they do. Because the dependence measure was constructed from the financing behaviour of mature firms in the United States and then applied to industries elsewhere, the design substantially reduced the risk that domestic industry growth was driving the country-level measure of financial development. The study therefore did more than confirm a correlation, since it identified a mechanism in which finance matters most where production technologies require external funding, and it established that the effect of financial development is heterogeneous across activities.

Related industry studies refine this conclusion. Cetorelli and Gambera (2001) found that the effect of banking-market concentration on industrial growth depends on the degree to which firms rely on external finance. Fisman and Love (2003) showed that trade credit can partially substitute for formal intermediation, with industries more dependent on supplier finance growing relatively faster in countries where financial institutions are less developed. Kroszner et al. (2007) demonstrated that banking crises impose disproportionate costs on industries that depend heavily on external finance, hold fewer tangible assets or consist largely of young firms. Read together, these studies indicate that any assessment of the role of finance must take account of production technology, asset structure, firm age, the availability of financing substitutes and exposure to financial shocks.

The direction of causality remained contested. Calderón and Liu (2003) found that financial development generally drives growth while also documenting bidirectional causality, with the supply-leading channel more pronounced in developing countries. Rousseau and Wachtel (2011), in contrast, reported that the cross-country finance–growth relationship weakened markedly in more recent samples, attributing part of the attenuation to the rising incidence of financial crises. These findings do not overturn the earlier literature so much as bound it, suggesting that an average positive association estimated over one period need not persist as financial systems, regulation, credit composition and crisis risk evolve.

The early consensus is best stated precisely. The evidence indicates that well-functioning financial systems are associated with faster growth and stronger productivity performance, and that this association is not merely an artefact of reverse causation. It does not establish that all forms of credit have equivalent effects, that the marginal effect is constant, or that larger financial systems are uniformly more efficient. Nor does it settle whether the relevant margin is the accumulated stock of credit or its rate of expansion, a question that only became prominent after the global financial crisis and that Section 2.4 takes up directly.

2.3. Nonlinearity and the 'Too Much Finance' Debate

Early empirical work generally modelled the effect of financial development as linear, implicitly assuming that an additional unit of financial depth delivers the same marginal benefit irrespective of the prevailing level. The global financial crisis exposed the limits of that assumption. Rapid credit expansion, rising leverage, resilient asset prices and a growing financial sector preceded severe contractions in output and employment across a range of economies, shifting the research question from whether finance supports growth toward the conditions under which it does so and the point at which further expansion may become counterproductive.

Cecchetti and Kharroubi (2012) reported that financial development supports growth only up to a point, beyond which it becomes a drag on productivity, and additionally found that rapid growth of the financial sector is associated with weaker productivity performance. Arcand et al. (2015) identified a comparable nonlinearity, estimating that the marginal contribution of private credit turns negative at high credit-to-GDP ratios. Law and Singh (2014), applying a dynamic threshold approach, found that financial development raises growth below an estimated threshold and depresses it above. Samargandi et al. (2015) documented an inverted-U relationship in a sample of middle-income countries, while Sahay et al. (2015) concluded that further deepening still offers substantial gains in many emerging economies but that the pace and extent of financial expansion involve a trade-off between growth and stability.

The thresholds reported in this literature cluster loosely around 80 to 100 per cent of GDP, and they are best treated as sample-specific estimates rather than as regulatory

limits. Their location depends on the proxy for financial development, the country coverage, the sample period, the treatment of endogeneity, the assumed functional form and the conditioning set. A threshold estimated from aggregate private credit in an international panel cannot be transferred directly to a particular sector, income group or national banking system, and even within one dataset the position and precision of a turning point may shift when a quadratic specification is replaced by a threshold model or when banking and market finance are modelled separately.

Contrary evidence therefore carries real weight. Botev et al. (2019) estimated several nonlinear threshold specifications across developing, emerging and advanced economies without finding a robust tipping point beyond which financial development becomes systematically harmful, emphasising instead the complementarity between bank-based and market-based finance. Their results caution against treating the inverted U as a law. The defensible conclusion is that the marginal effect of finance depends on its level and its composition, and that the shape and location of any nonlinearity remain empirical questions rather than settled parameters.

Several mechanisms could generate diminishing or negative marginal effects. Once the most binding financing constraints have been relaxed, additional credit may fund projects with progressively lower expected returns, shifting the allocative margin as the system expands. A rapidly growing financial sector may also draw skilled labour away from activities that contribute more directly to technological progress, a reallocation that Cecchetti and Kharroubi (2015) argue penalises research-intensive industries and projects whose returns are difficult to pledge. Excessive credit may further sustain unproductive firms through repeated refinancing, as Caballero et al. (2008) document for Japan, where subsidised lending to impaired borrowers delayed restructuring and depressed investment and employment among healthier competitors.

A second mechanism runs through financial vulnerability. Schularick and Taylor (2012), using data on 14 countries from 1870 to 2008, established that credit growth is a powerful predictor of financial crises. Credit expansion may raise output in the

short term while simultaneously increasing the probability and severity of a subsequent downturn, so that the average coefficient in a conventional growth regression can conflate a stimulatory impulse with a later correction. This intertemporal pattern is one reason why separating short-run from long-run effects is central to the Türkiye analysis presented in Chapter 7.

A third mechanism concerns the destination of credit. Beck et al. (2012) found that enterprise credit is associated with growth whereas household credit is not. Mian et al. (2017) showed that increases in the household debt-to-GDP ratio predict lower medium-term output growth and higher unemployment. Müller and Verner (2024) demonstrated that credit growth concentrated in non-tradable sectors tends to precede growth slowdowns and financial crises, while credit growth to tradable sectors is associated with sustained increases in output and productivity without elevated crisis risk. These results suggest that what appears as a problem of ‘too much finance’ may in part be a problem of misallocated finance, a theme developed in Section 2.5.

Nonlinearity may also arise because recipients differ in their capacity to absorb credit. At low levels of intermediation, modest amounts of financing support high-return investments that would otherwise not be undertaken. As credit expands, the pool of remaining productive projects thins, debt-service obligations accumulate and lending standards may loosen, so that the marginal benefit declines even in the absence of a crisis. A turning point need not indicate that credit has become excessive in any normative sense; it may instead indicate that prevailing institutional, technological and sectoral conditions can no longer convert additional credit into proportionate increases in value added.

Two implications for empirical practice follow. A statistically significant squared or cubic term is not by itself evidence of an economically meaningful nonlinearity, since the estimated turning point must lie within the observed data range and the slopes on either side must be consistent with the proposed shape. Higher-order terms should also not be added mechanically, so that the choice of functional form requires joint testing, comparison of explanatory power and consistency with the support of the regressor. The empirical chapters follow this procedure by estimating linear,

quadratic, cubic and quartic specifications, selecting among them systematically and interpreting turning points only when they fall inside the sample.

2.4. From Credit Depth to Credit Growth: Measurement, Integration and Interpretation

The literature reviewed so far treats financial development predominantly as a stock, measured by the ratio of outstanding credit to GDP. That convention is convenient but not harmless, because the empirical properties of credit ratios rarely satisfy the assumptions under which level-on-level regressions can be interpreted. Credit-to-GDP series are highly persistent and typically behave as integrated processes, so that regressions relating one trending series to another may produce statistically significant coefficients that reflect common trends rather than economic relationships, a hazard identified in the classic analysis of spurious regression by Granger and Newbold (1974). Second-generation panel unit-root procedures that allow for cross-sectional dependence, notably the cross-sectionally augmented test proposed by Pesaran (2007), provide the appropriate diagnostic in macroeconomic panels where common shocks are pervasive. Where such tests indicate that credit variables are integrated of order one while the output variables are stationary, this dissertation resolves the resulting mixed-integration problem by entering the credit series in first differences. That step is a modelling decision taken here rather than a prescription contained in either source, and its consequence is that the parameter of interest ceases to describe financial depth and instead describes credit growth.

This is not a technical footnote but a change in the object of analysis, and a substantial body of work indicates that the two objects behave differently. Loayza and Rancière (2006) confronted the apparent contradiction between the growth literature, which finds positive effects of financial depth, and the crisis literature, which finds that credit aggregates are among the best predictors of banking crises. Estimating an encompassing model of long- and short-run effects with the pooled mean group estimator of Pesaran et al. (1999), they concluded that a positive long-run relationship between financial intermediation and output growth coexists with a largely negative short-run relationship, and they linked the short-run estimates to

measures of financial fragility. The distinction they draw between trend and transitory components of credit aggregates is conceptually related to the distinction between a level specification and a differenced one, although the correspondence drawn here is an interpretation rather than a claim advanced by those authors. It suggests that a negative coefficient on credit growth is neither anomalous nor inconsistent with a positive long-run association between credit depth and output.

Beck et al. (2014) reach a compatible conclusion from a different angle by separating the intermediation function of financial systems from their overall size. Drawing on 77 countries between 1980 and 2007, they find that intermediation activities raise growth and reduce volatility in the long run, that expansion of the financial sector along other dimensions has no long-run effect on real outcomes, and that over shorter horizons a large financial sector stimulates growth in high-income countries at the cost of higher volatility. Their results underline that horizon and concept interact, so that what finance does over a decade and what it does over a few quarters are separate empirical questions requiring separate specifications.

The most direct evidence that credit growth carries information distinct from credit depth comes from Ductor and Grechyna (2015), who examine the interdependence between financial development and real-sector output in 101 developed and developing countries between 1970 and 2010. They show that the effect of financial development on growth depends on the growth of private credit relative to real output growth, and that the effect becomes negative when private credit expands rapidly without a corresponding expansion of real output. Their finding is important for the present study in two respects. It establishes that the relevant conditioning variable can be a rate of change rather than a level, and it supplies a mechanism – divergence between financial and real expansion – that generates nonlinearity in credit growth space rather than in credit depth space.

Bezemer et al. (2016) sharpen the point by distinguishing explicitly between stocks and flows of credit. Using newly assembled data on 46 economies between 1990 and 2011, they show that financial development since 1990 was driven mainly by growth in credit to real estate and other asset markets, which carries a negative growth

coefficient, and they find positive growth effects for credit flows to non-financial business but not for mortgage and other asset-market credit flows. By accounting simultaneously for the composition of credit stocks and for the effect of credit flows, they account for the insignificant or negative growth effects of financial development reported in more recent samples. Their design demonstrates that the stock–flow distinction and the composition question are not competing explanations but complementary ones, since the growth effect of a credit flow depends on where that flow is directed.

Evidence from long-run macro-financial history reinforces the case for treating credit growth as an object of interest in its own right. Schularick and Taylor (2012) established credit growth as a leading indicator of financial crises, and Jordà et al. (2013), using the same historical panel of 14 advanced economies, documented that financial-crisis recessions are costlier than ordinary ones and that more credit-intensive expansions tend to be followed by deeper recessions and slower recoveries. Their use of local projection methods, conditioning on a broad set of macroeconomic controls, indicates that past credit accumulation carries predictive content for output, investment, lending, interest rates and inflation. Where these studies measure the intensity of an expansion by the change in credit rather than by its level, they estimate precisely the kind of parameter that a differenced specification recovers.

The productivity literature adds a sectoral dimension that bears directly on the design of this dissertation. Borio et al. (2016), examining more than twenty advanced economies over four decades, find that credit booms tend to undermine productivity growth by inducing labour reallocation toward sectors with lower productivity growth, with a temporarily inflated construction sector serving as the leading example, and that the damage caused by reallocations during a boom is far larger when a crisis follows. Their evidence implies that rapid credit expansion may change the composition of output as much as its level, which is exactly the outcome captured by the sectoral value-added shares used as dependent variables in the cross-country chapter. Mian et al. (2020) reach a related conclusion through a different identification strategy, showing that credit-supply expansion raises non-tradable employment and non-tradable prices with limited effects on tradable employment,

and that such expansions amplify the business cycle and lead to more severe recessions.

Three interpretive consequences follow for the empirical chapters. The first concerns the meaning of estimated thresholds. When the credit variable enters in first differences, any turning point recovered from a quadratic or higher-order specification is located in credit growth space, expressed as a rate of change, and it is therefore not comparable with the 80 to 100 per cent of GDP thresholds reported by Arcand et al. (2015), Law and Singh (2014) or Cecchetti and Kharroubi (2012). A finding that the marginal effect of credit growth turns negative beyond some rate of expansion speaks to the pace at which sectors can absorb new lending, not to the level of financial depth at which an economy becomes overbanked. Conflating the two would misstate both the finding and its policy content, and the results reported in Chapters 5 and 6 are interpreted accordingly.

The second concerns what differencing removes. First differencing eliminates the persistent, country-specific component of financial depth, so the estimates identify how variation in the pace of credit expansion relates to sectoral outcomes and are silent about the long-run level relationship that much of the earlier literature sought to measure. This is a genuine restriction, though it is also the restriction that makes inference valid given the integration properties of the series, and the evidence assembled by Loayza and Rancière (2006) and Beck et al. (2014) suggests that the short-run and long-run parameters would differ even if both could be estimated consistently. The Türkiye chapter partially compensates for this limitation, since the cross-sectionally augmented error-correction framework retains sectoral credit levels in the long-run component while modelling credit growth in the short-run component, allowing both margins to be examined within one specification.

The third concerns the sign expected on credit growth. Because rapid credit expansion may raise debt-service burdens, finance projects of declining quality, shift activity toward collateral-intensive and non-tradable production and increase the probability of a subsequent correction, a negative or non-monotonic coefficient on credit growth is consistent with a wide range of theoretical priors and with a positive

long-run role for financial intermediation. Evidence of this kind should not be read as showing that credit is harmful; it should be read as showing that the rate of credit expansion and the level of financial development are distinct margins with distinct effects, which is the position this dissertation adopts throughout.

2.5. Credit Composition, Allocation and Macroeconomic Risk

Post-crisis research increasingly separates the size of the financial system from the composition of its assets. Aggregate private credit pools lending to non-financial firms, households, property developers, financial institutions and public or quasi-public entities, and these borrowers differ in their purposes, project horizons, collateral, exposure to international competition and contribution to productive capacity. Two economies with identical ratios of private credit to GDP may therefore present very different growth prospects and very different financial risks, and two economies with identical rates of credit expansion may do so as well.

The distinction between firm and household credit is the most extensively documented. Credit to firms finances machinery, working capital, research, export capacity and organisational change, whereas household borrowing relaxes consumption constraints and supports housing investment but may raise leverage without adding productive capacity. Beck et al. (2012) show that the positive cross-country association between credit and growth is driven mainly by enterprise lending. Mian et al. (2017) find that expansions in household debt predict medium-term slowdowns, consistent with credit-supply shocks, optimistic expectations and eventual balance-sheet adjustment. Bahadir and Gumus (2016) extend the analysis to emerging markets, showing that the business-cycle patterns previously attributed to aggregate credit are considerably stronger for household credit than for business credit, which implies that the composition of a credit expansion matters for its cyclical signature and not only for its long-run growth contribution.

Sectoral allocation introduces a further and, for this dissertation, central distinction. Tradable activities face international competition and typically employ credit to expand capacity, adopt technology or enter foreign markets, whereas non-tradable

activities, particularly those tied to land and property, can generate powerful collateral feedback loops. Müller and Verner (2024) show that credit booms disproportionately favour non-tradable sectors and that such expansions are systematically associated with subsequent slowdowns and crises, while Mian et al. (2020) document the same asymmetry in employment and prices. None of this implies that lending to non-tradable activities is unproductive, since construction, transport, retail, health and other domestic services supply essential capital and services. It implies rather that the macroeconomic consequences of a credit expansion depend on where the incremental lending is directed.

Allocation also operates at the level of individual borrowers, and the evidence here is more encouraging about what credit can achieve when it is well directed. Bai et al. (2018) find that the deregulation of state banking markets in the United States raised the relative employment and capital growth of more productive local firms, with the effect concentrated among young firms and driven by a shift in the composition of local bank credit supply toward more productive borrowers, translating into economically important gains in aggregate industry productivity through the reallocation of labour. Read alongside the finding of Borio et al. (2016) that credit booms can reallocate labour toward lower-productivity sectors, the two results describe the conditions under which credit expansion raises or lowers aggregate productivity, since what matters is whether the marginal loan reaches a more productive or a less productive user.

Financial structure conditions this selection process. Hsu et al. (2014) show that equity-market development and credit-market development affect technological innovation differently, since equity-market development supports innovation in industries that are more dependent on external finance and more technology-intensive whereas credit-market development does not. The reasoning adopted here — that equity finance suits risky projects involving intangible assets because investors share in the upside, while collateral-based bank lending favours tangible assets and predictable cash flows — is an interpretation of that pattern rather than a proposition the authors test directly. A predominantly bank-based system may therefore support conventional capital accumulation effectively while remaining a

weaker instrument for innovation-intensive activity. That asymmetry is particularly relevant to the service sector, where information technology, professional expertise, brands, software and human capital account for much of the productive base.

The allocation problem does not end at loan origination. Weak banks may roll over exposures to avoid recognising losses, politically motivated lending may prioritise employment or short-term activity over project quality, and borrowers may use new credit to refinance existing obligations rather than to fund new investment. Credit may also migrate across sectors through intra-group transactions, supplier finance or property collateral, so that the sector recorded as the borrower is an informative but imperfect indicator of the activity ultimately financed. This measurement limitation warrants caution in causal interpretation, though it does not undermine the value of sectoral analysis, since sectoral credit data reveal the distribution of banking-system exposures that aggregate figures conceal.

The allocation literature accordingly reframes what financial development means. A larger credit stock, or a faster rate of credit expansion, is not an objective in itself but a potential resource whose consequences depend on borrower selection, loan purpose, maturity, collateral and the productive capacity of the recipient. This perspective connects the ‘too much finance’ debate to sectoral analysis, since negative marginal effects observed at high levels of aggregate credit may partly reflect a compositional shift in lending, while sector-specific turning points may reveal that particular activities have exhausted their absorptive capacity even when aggregate credit remains moderate.

2.6. Why an Aggregate Framework Is Insufficient

Most finance–growth studies aggregate on both sides of the regression. Financial development is proxied by total private credit or by a composite index, while economic performance is measured by GDP growth, per capita income growth or total factor productivity. These measures are appropriate for macroeconomic assessment, but aggregation on both sides obscures the mechanisms through which finance affects the real economy, and the problem is not resolved by moving from

levels to growth rates, since an aggregate credit growth rate averages over destinations just as an aggregate credit stock does.

On the financial side, a uniform increase in total credit may represent agricultural input financing, manufacturing investment, construction lending, working capital for trade, mortgage expansion or the refinancing of existing obligations, and the output response should differ across these uses. On the production side, aggregate GDP combines activities with different technologies, adjustment speeds, financing requirements and exposure to shocks, so that expansion in one sector may offset contraction in another and leave aggregate output largely unchanged. Substantial aggregate growth may also coincide with declining value-added shares in particular sectors as the structure of the economy evolves.

The distinction is especially consequential in the first empirical chapter, where the dependent variables are the value-added shares of agriculture, industry and services. A change in a sector's share of GDP does not necessarily correspond to a change in its absolute output, since a share may fall because the sector contracts, because other sectors expand more rapidly or because relative prices move. The variable should therefore be understood as an indicator of sectoral composition and of the relative weight of each activity, which is precisely why it is informative here, as it permits an assessment of whether credit growth is associated with shifts in the structure of production rather than only with aggregate expansion.

The second empirical chapter addresses a different aspect of the same problem by matching sectoral total cash loans to sectoral value added in Türkiye across eleven activity groups. This design reduces the mismatch between an economy-wide credit measure and a highly aggregated output measure, while permitting a separation of short-run dynamics from long-run equilibrium relations and an assessment of whether credit expansion affects sectors differently.

The disaggregated approach rests on established foundations. Rajan and Zingales (1998) established that dependence on external finance varies systematically across industries, Kroszner et al. (2007) showed that industry characteristics condition

vulnerability to financial crises, and Müller and Verner (2024) demonstrated that macroeconomic outcomes depend on whether credit growth supports tradable or non-tradable activity. Each of these findings implies that a single aggregate coefficient averages over effects that may be heterogeneous in magnitude and even opposed in sign.

Aggregation may also conceal nonlinearity. If credit growth exhibits diminishing returns in construction, a positive linear effect in manufacturing and no apparent relationship in services, the aggregate association could appear weakly positive, flat or nonlinear depending on sectoral weights and the phase of the credit cycle. A turning point estimated from aggregate data may therefore reflect a compositional shift in lending, a change in the structure of output or genuine diminishing returns within sectors. Sectoral analysis cannot resolve every identification problem, but it does separate mechanisms that aggregate data necessarily conflate.

The policy implications differ accordingly. Aggregate credit targets presuppose that the marginal value of lending is broadly similar across recipients, so that a common rule can be efficient. The evidence surveyed here suggests otherwise, since credit expansion may be highly productive in sectors facing binding financing constraints while adding little in sectors that are already leveraged, demand-constrained or operating with excess capacity. Sector-specific evidence is therefore necessary for evaluating both the quantity of finance and the pattern of its allocation.

2.7. Sectoral Transmission of Financial Development

2.7.1. Agriculture

Agriculture presents a distinctive financing problem. Production is seasonal, returns are delayed and biological and climatic risks are substantial, while many producers lack conventional collateral or formal financial records. Information costs are high in rural areas and small loan sizes raise the unit cost of screening and monitoring, so that restricted supply, elevated transaction costs and reluctance to borrow against pledged assets frequently coexist. Financial development may therefore make a large

difference where credit is scarce, though converting credit into value added requires institutions capable of managing these sector-specific risks.

Microeconomic evidence establishes that credit constraints bind. Feder et al. (1990) modelled Chinese agriculture in disequilibrium and showed that restricted credit access and productivity are linked when borrowing constraints prevent households from choosing efficient input combinations. Guirkinger and Boucher (2008), using farm-level data from Peru, found that collateral requirements and asymmetric information generate quantity rationing alongside transaction-cost and risk rationing, with measurable output losses among constrained producers. Petrick (2004) examined how access to subsidised credit shaped investment behaviour among Polish farms subject to rationing, and considered whether public funds were diverted away from productive investment. These studies illustrate both sides of the relationship, since inadequate financing depresses productivity while additional lending does not guarantee that funds will be devoted to durable productive assets.

Cross-country evidence on agricultural credit growth is more recent and more directly relevant to the specification adopted here. Seven and Tumen (2020) find that agricultural credits raise agricultural productivity, estimating that a doubling of agricultural credit is associated with an increase in agricultural productivity of roughly four to five per cent. Their results also indicate that the channel differs along the development path, operating mainly through the agricultural component of GDP in developing countries and through agricultural labour productivity in developed countries, which suggests that the same expansion of agricultural lending may show up as output growth in one setting and as productivity growth in another. This heterogeneity is one reason why the cross-country chapter estimates income-group specifications separately rather than imposing a common coefficient.

Evidence on rural finance more broadly is nuanced. Binswanger and Khandker (1995) found that the expansion of formal rural finance in India raised rural non-farm employment, output and wages, while its direct effect on agricultural output was comparatively modest. This pattern indicates that the effects of rural credit may extend well beyond the sector to which the loan is nominally attributed, since

agricultural households combine production, consumption and non-farm activity, so borrowed funds may smooth consumption, finance education, support small enterprises or refinance debt rather than raising agricultural value added.

The productivity of agricultural credit further depends on complementary conditions. Mechanisation, irrigation, improved seeds, storage, transport, extension services and access to output markets all shape the return on borrowed funds, and where these are absent additional credit may provide little more than short-term liquidity. Public investment and institutional quality can therefore alter the effectiveness of financing, which helps explain why empirical results differ across periods and income groups even when the credit measure is held constant.

Structural transformation adds a final complication that is specific to the dependent variable used in the cross-country chapter. As economies develop, the share of agriculture in GDP typically declines even when agricultural productivity and absolute output rise. Where credit raises farm productivity, it may release labour to industry and services and thereby reduce the agricultural share of value added. A negative association between agricultural credit growth and the agricultural share of GDP should accordingly not be read immediately as evidence that finance harms agricultural performance, since it may instead reflect faster expansion elsewhere, relative price movements or a credit-financed transformation of the production structure. Interpreting the coefficient correctly requires combining the agricultural finance literature with the structural change literature discussed in Section 2.8.

2.7.2. Industry and Manufacturing

Industrial production has long been central to finance research because many industrial activities require substantial initial outlays, imported equipment, inventories and extended development periods, and because internal cash flow is often insufficient to cover these costs in young and rapidly growing firms. The external-finance-dependence framework of Rajan and Zingales (1998) is therefore closely associated with industrial growth, implying that industries facing a larger technological financing gap benefit disproportionately where banks and capital

markets are accessible, provided that intermediaries direct funds toward promising projects.

Financial development affects industry through both investment and productivity. Improved screening and contract enforcement allow firms to adopt more advanced technologies, expand capacity and reorganise production. Love (2003) shows that financial development reduces the sensitivity of investment to internal funds, indicating weaker financing constraints and improved capital allocation. These benefits are especially material for small and medium-sized enterprises with limited access to bond and equity markets, though established firms with substantial collateral may nonetheless dominate bank lending in ways that constrain entry and innovation.

The composition of industrial assets shapes the financial channel. Conventional manufacturing and infrastructure-intensive activities generally involve machinery, buildings and receivables that can be pledged, whereas innovative manufacturing depends more heavily on research, software, organisational capital and uncertain future cash flows. Hsu et al. (2014) find that equity-market development particularly supports innovation in high-technology industries dependent on external finance, while credit-market development does not display the same effect, indicating that the institutional structure of finance matters alongside its scale.

Industrial credit can also produce adverse effects. Readily available lending may sustain low-productivity firms, finance excess capacity or encourage investment justified by collateral rather than by expected productivity. Caballero et al. (2008) demonstrate that continued lending to impaired firms depressed the performance of healthier competitors, and Cecchetti and Kharroubi (2015) argue that finance gravitates toward projects that are easier to collateralise, diverting resources from research-intensive activity. A large industrial credit stock may therefore combine productive investment with defensive refinancing, and the same holds for a rapid rate of industrial credit expansion.

Evidence on reallocation clarifies when industrial credit raises aggregate productivity. Bai et al. (2018) show that a shift in the composition of local credit

supply toward more productive firms generates economically important gains in industry productivity, operating largely through the reallocation of labour. Borio et al. (2016) show the converse, since credit booms that direct resources toward temporarily inflated construction activity reallocate labour toward sectors with lower productivity growth and leave lasting damage when a crisis follows. Taken together, these findings suggest that the productivity consequences of industrial credit growth depend on the selection mechanism operating within the banking system rather than on the volume of lending alone.

Heterogeneity within industry is substantial. Mining, utilities, manufacturing and construction differ in technology, market structure, trade exposure and reliance on collateral. Construction is particularly sensitive to domestic demand, land prices, interest rates and collateral cycles, whereas manufacturing faces greater international competition and may generate stronger learning and export spillovers. The international industry definition, which includes construction, averages across these differences, and the Türkiye chapter improves on this by separating manufacturing, construction and industry excluding manufacturing, thereby aligning activity boundaries more closely with the underlying credit data.

The industrial literature accordingly supports two expectations. Finance can raise industrial growth substantially where firms face binding external financing constraints and hold productive investment opportunities, and it can also generate diminishing or negative marginal effects where credit sustains unproductive capacity. Which outcome dominates is likely to depend on financial structure, collateral, the intensity of innovation and the proportion of credit devoted to new investment rather than to refinancing, making industry a plausible setting for both strong positive effects and pronounced nonlinearity.

2.7.3. Services, Intangible Assets and Innovation

The service sector is exceptionally heterogeneous, encompassing trade, transport, accommodation, information and communication, finance, real estate, professional services, public administration, education, health and personal services. These

activities differ markedly in capital intensity, tradability, regulation, skill requirements and reliance on intangible assets, so that aggregating them into a single sector facilitates international comparison while potentially concealing offsetting effects within the aggregate. The more disaggregated Türkiye analysis is therefore particularly useful for understanding how finance operates within specific service branches.

The structural transformation literature indicates that services expand differently at different stages of development. Eichengreen and Gupta (2013) identify two waves of service-sector growth, an initial wave occurring at lower income levels and a second wave at higher income levels, with the first associated with traditional activities linked to urbanisation and domestic demand and the second with modern business, communication and knowledge-intensive services. This distinction implies that the relationship between finance and services may vary systematically across income groups, since lending that supports retail and transport in a lower-middle-income economy is not equivalent to financing information technology or professional services in a high-income economy.

Many modern services depend on assets that are difficult to pledge, including human capital, software, data, organisational knowledge and reputation, none of which lenders can readily recover on default. Banks that prioritise physical collateral may therefore underfund innovative service firms even where aggregate credit is abundant, with equity finance, venture capital, guarantees and relationship lending offering partial remedies. The evidence of Hsu et al. (2014) on finance and innovation is directly relevant here, since it demonstrates that the structure of a financial system conditions its capacity to support activities with uncertain and intangible returns.

Other service activities are heavily collateralised. Real estate and parts of construction can absorb large volumes of credit because property security is straightforward to value, generating a feedback loop in which rising credit supports property prices, higher valuations raise collateral values and further borrowing becomes feasible. Such lending may stimulate activity temporarily while raising

leverage and the risk of a subsequent correction, and the distinction drawn by Müller and Verner (2024) between tradable and non-tradable credit growth is especially pertinent, since non-tradable credit booms are systematically associated with later slowdowns and crises. The parallel finding of Mian et al. (2020) that credit-supply expansions raise non-tradable employment and prices while leaving tradable employment largely unaffected reinforces the same conclusion from an independent identification strategy.

Financial and insurance activities pose a further interpretive difficulty, since credit extended to the financial sector may expand balance sheets without financing non-financial production, while the sector simultaneously provides services to all other activities. Its value added may grow through intermediation, fees and asset-market transactions, so that a larger financial share does not necessarily indicate stronger support for the real economy. The resource-crowding mechanism described by Cecchetti and Kharroubi (2015) is particularly relevant, as rapid financial-sector growth may draw skilled labour away from productive and innovative activities.

Public administration, education, health and other social services depend primarily on public funding and regulation rather than on conventional bank lending. Credit may still influence private providers, infrastructure and working capital, but government expenditure, demographic requirements and institutional arrangements are likely to dominate. A small credit coefficient in these activities should therefore not be interpreted as evidence that finance is irrelevant, only that bank credit is not the principal determinant of their value added.

The service-sector literature consequently predicts no uniform sign or functional form. Some activities face severe constraints on the financing of intangible assets, others benefit from abundant property-backed credit and others are governed largely by public provision. The three-sector international analysis identifies the broad differences between services and the other main sectors, while the Türkiye analysis examines whether these aggregate findings mask distinct short- and long-run relationships within individual service branches.

2.8. Structural Transformation and Income-Level Heterogeneity

Structural transformation, understood as the reallocation of economic activity among agriculture, manufacturing and services, is a constitutive feature of growth rather than a by-product of it (Herrendorf et al., 2014). As productivity, income and demand evolve, labour and capital move between sectors, sectoral value-added shares adjust and the sources of aggregate productivity growth change. Finance influences this process by accelerating or delaying it and by shaping its direction through the relative cost of investment across activities.

McMillan and Rodrik (2011) document large productivity differences across sectors in developing economies and show that moving labour from low- to high-productivity activities can raise aggregate growth substantially. They also show that structural change is not invariably growth-enhancing, since labour may move toward lower-productivity sectors and thereby reduce aggregate productivity growth. A financial system that channels credit into expanding but low-productivity activities may therefore facilitate reallocation without improving long-run performance, a possibility that the sectoral evidence of Borio et al. (2016) renders concrete, since they identify credit booms as a cause of precisely this kind of adverse reallocation.

The role of finance is expected to change along the development path. In low-income economies, access to formal finance remains limited, agriculture employs a large share of the labour force, small firms predominate and informal arrangements retain importance, so that a moderate expansion of formal credit may relax some constraints while weak institutions and infrastructure limit the productivity response. In lower-middle-income economies, urbanisation and industrialisation raise demand for working capital, housing, transport and manufacturing investment, and financial deepening may support structural change even as rapid adjustment channels credit into construction and other non-tradable activities.

Upper-middle-income economies face different constraints. Financial access is broader, but productivity gains depend increasingly on technological progress, export upgrading and a transition from investment-driven to innovation-driven growth, and

collateral-based lending is a less effective instrument for intangible investment than for conventional capital accumulation. High-income economies generally possess developed financial systems and face fewer classical credit constraints, while remaining exposed to risks arising from household debt, property markets, complex financial services and credit saturation, so that the marginal effect of additional bank credit tends to be smaller or less stable.

Greenwood and Jovanovic (1990) provide the theoretical basis for stage dependence, Rioja and Valev (2004a) show empirically that the growth effect of financial development differs across low, intermediate and high regions of financial development, Rioja and Valev (2004b) find in a separate study that finance operates mainly through productivity growth in more developed economies and through capital accumulation in less developed ones, and Samargandi et al. (2015) demonstrate the importance of nonlinear effects within middle-income countries. Seven and Tumen (2020) add a sectoral instance of the same principle, since the channel through which agricultural credit operates differs between developing and developed economies. Pooled estimates may therefore conceal relationships that are specific to particular groups, and income classification functions as an analytical device for testing whether the influence of financial development differs across structural and institutional contexts rather than as mere description.

Income groups nonetheless remain broad categories. Countries within the same group differ in legal systems, financial inclusion, export structure, political institutions and banking regulation, and the purpose of classification is not to assert homogeneity but to relax the assumption that a single global coefficient applies equally to economies at different stages of development. Country and time fixed effects absorb other forms of heterogeneity, while separate income-group estimation allows the functional form itself to vary, which matters when the object of interest is the shape of a nonlinear relationship rather than a single slope.

The outcome variable conditions interpretation in this setting as well. A decline in the agricultural share of value added may signal successful structural transformation, whereas an increase in the services share may reflect either the expansion of modern

knowledge-intensive activity or the growth of low-productivity informal services. Financial development may shift sectoral shares without affecting aggregate growth in a corresponding way, so this dissertation treats sectoral value-added shares as outcomes of structural composition and interprets them alongside both the finance–growth and the structural transformation literature.

The pattern of nonlinearity may itself vary by income level. Where access constraints are severe, the rising portion of the relationship should be more prominent within the observed range, whereas in economies with substantial financial depth, evidence of diminishing or negative marginal effects should be more visible. This remains a hypothesis rather than a prediction, since weak institutions may generate misallocation even at low levels of credit while well-regulated advanced systems may intermediate deep finance effectively, and the empirical analysis tests whether the data support different shapes or different turning points across groups.

2.9. Time Variation, Financial Crises and the Post-2008 Environment

The finance–growth relationship is unlikely to be stable over long horizons. Since the early 1990s, regulation, monetary policy, banking technology, capital mobility, financial inclusion and credit composition have all changed substantially, so a single coefficient estimated over several decades may average across structurally distinct regimes and obscure changes in the way credit affects production.

The global financial crisis constitutes a natural dividing line. The preceding period featured financial liberalisation, cross-border banking, rising leverage and rapid growth of property-related credit in many economies. The subsequent period brought bank recapitalisation, tighter prudential regulation, unconventional monetary policy, altered risk appetite and a stronger emphasis on financial stability, changes that affected not only the volume of credit but bank capital, lending standards, borrower composition and the role of market finance.

Several strands of evidence indicate that the timing of measurement matters. Rousseau and Wachtel (2011) report that the finance–growth relationship weakened

in later samples and attribute part of this weakening to the increased frequency of crises. Schularick and Taylor (2012) identify credit growth as a strong predictor of crises, and Jordà et al. (2013) show that credit-intensive expansions are followed by deeper recessions and slower recoveries. Mian et al. (2017) find that surges in household debt precede medium-term weakness. Sahay et al. (2015) add that the pace of financial deepening, distinct from its level, carries implications for stability. Credit may therefore be positively associated with contemporaneous activity and negatively associated with future growth, producing results that vary systematically with the horizon under examination and reinforcing the interpretive point made in Section 2.4.

The cross-country analysis accordingly distinguishes 1992–2007 from 2010–2019 and also examines the full 1992–2019 period. The sub-period division does not assert that all countries experienced the crisis identically or that 2010 marks a universal institutional break; it provides a transparent comparison between the pre-crisis environment and the subsequent decade while excluding 2008 and 2009, whose exceptional disruptions could otherwise dominate the estimates. Full-sample estimation remains necessary for identifying long-horizon patterns and for exploiting the wider variation in credit growth that the complete period provides.

Period comparison is particularly important for nonlinear specifications. A relationship that appears quadratic over the full sample may reflect two distinct linear regimes, a change in credit composition or a shift in sectoral weights, whereas shorter sub-periods may lack sufficient variation to locate turning points precisely. Reporting period-specific alongside full-sample estimates therefore helps distinguish temporal instability from genuine nonlinearity, and the empirical chapters follow this practice throughout.

2.10. Evidence from Türkiye

2.10.1. Aggregate Finance–Growth Evidence

The Turkish literature on finance and growth reproduces the international debate over causality and stability while adding features specific to the national experience.

Türkiye underwent financial liberalisation, experienced repeated macroeconomic instability and banking crises, implemented substantial regulatory reform after 2001 and subsequently saw rapid credit expansion accompanied by policy-driven lending cycles. This trajectory suggests that a single linear relationship estimated over a short window is unlikely to characterise the credit–output relationship adequately.

Early research focused on aggregate indicators of financial development and aggregate output. Aslan and Küçükaksoy (2006) examined the direction of causality between financial development and growth and reported evidence of causality running from financial development to growth, consistent with the supply-leading hypothesis. Aslan and Korap (2006) analysed the same relationship for Türkiye using time-series methods, contributing to a body of work in which conclusions vary with the proxy, the period and the specification. Ceylan and Durkaya (2010), using quarterly data, identified a long-run relationship between domestic credit and growth with causality running from growth to credit. Taken together, these early studies already point in opposing directions, indicating that the balance between supply-leading and demand-following mechanisms is itself an empirical question in the Turkish case rather than a settled feature of it.

Other contributions point toward a more prominent supply-leading channel. Göçer et al. (2015), allowing for multiple structural breaks, found that increases in banking-sector credit exert positive effects on employment and economic growth. Divergence of this kind across studies is expected given differences in sample period, data frequency, credit measure, treatment of crises and econometric approach, and it reinforces the case for a specification that permits the relationship to differ across sectors and horizons.

Two cautious conclusions emerge from the Turkish aggregate evidence. A strong association between credit and output is well established, though the estimated direction of causality differs across studies. Structural breaks and credit composition matter materially for the results. Both conclusions favour a methodology that accommodates feedback, common shocks and heterogeneous adjustment speeds rather than one assuming a unidirectional and contemporaneous effect.

2.10.2. Sectoral Credit and Output Evidence

A smaller literature disaggregates Turkish credit by sector. Koç (2015) analysed bank loans extended to ten sectors between 1999 and 2011 using panel methods and found a positive long-run relationship between sectoral credit and economic growth that was not present in the short run. This distinction is directly relevant to the present study, since the effects of credit may take time to materialise where investment, production and repayment decisions adjust slowly.

Sever and Han (2015) examined sectoral credit relative to sectoral GDP in agriculture, industry and services between 2002 and 2012. They found cointegration for the aggregate, industrial and service credit–income pairs but not for agriculture, and their causality analysis indicated a stronger influence running from real activity to finance, alongside a notable finance-to-output channel in agriculture and an output-to-credit channel in services. Their evidence is important precisely because it challenges the presumption that a single causal direction applies uniformly across sectors.

Kandemir et al. (2018) analysed the relationship between participation-bank credit extended to ten sectors and both GDP and exports, finding causality running from several sectoral credit types to macroeconomic outcomes while reporting substantial variation across sectors. Sarı and Konukman (2023), using quarterly data from 2007 to 2021, identified a positive bidirectional relationship between sectoral credit concentration and economic growth. Collectively these studies establish that the distribution of lending across sectors is empirically consequential, though their conclusions depend on the sector classification adopted and on the outcome variable examined.

The existing sectoral literature nonetheless leaves several gaps that motivate the present design. Many studies retain aggregate GDP as the outcome even when the credit data are disaggregated, producing a mismatch between a sectoral regressor and an economy-wide dependent variable. Some cover relatively short periods, work with a small number of broad sectors or adopt pooled-coefficient methods that do not

explicitly accommodate cross-sectional dependence. The distinction between short-run and long-run relationships is often left implicit, and systematic testing for nonlinearity is uncommon.

The Türkiye chapter addresses these gaps by matching sectoral total cash loans reported by the Banking Regulation and Supervision Agency to sectoral value added published by the Turkish Statistical Institute across eleven activity groups from 2005Q1 to 2026Q1. The matching follows NACE-based activity boundaries and separates manufacturing, construction, industry excluding manufacturing and several service categories, so that the analysis can assess whether credit supports production within the sector to which it is allocated rather than inferring a relationship between sectoral lending and aggregate output.

The panel structure raises a further methodological requirement. Turkish sectors are exposed to common influences including monetary policy, exchange-rate movements, domestic demand, inflation, financial regulation and crisis shocks, so their errors are unlikely to be independent across units. Cross-sectionally augmented estimation permits these common factors to be absorbed while retaining sector-specific responses, and the panel causality evidence reported in Chapter 7 supports a framework in which sectoral credit and output adjust jointly rather than sequentially.

2.10.3. Policy-Directed Credit and the Credit Guarantee Fund

Credit allocation in Türkiye is shaped not only by private banking decisions but also by state-owned banks, guarantee schemes, regulatory measures and targeted programmes. Such instruments can relax collateral constraints and support employment during downturns, though their longer-term effectiveness depends on whether the assisted firms convert additional finance into sustainable production and investment.

Evidence from the Credit Guarantee Fund illustrates this trade-off. Akçiğit et al. (2021) find that guaranteed loans raised firm employment and sales and reduced default risk relative to comparable firms, while also increasing indebtedness, with

effects varying by firm size and sector. That variation is central to any evaluation of policy-directed credit, since a single guarantee instrument can produce divergent outcomes when firms and sectors differ in financing constraints, demand conditions, technology and capacity to expand.

Bağır and Seven (2021), using a large sample of Turkish firms, find that the relationship between debt growth and productivity growth strengthens as financial constraints become more binding. This suggests that relaxing severe constraints can improve firm performance while implying that the productivity gain depends on the borrower's initial financial position. Credit extended to one sector may also affect suppliers and customers in others, so sectoral estimates may capture both direct effects and transmission through input–output linkages.

Policy evaluation should accordingly separate access from allocation. Guarantee schemes are well suited to addressing collateral and information failures faced by productive firms that cannot obtain finance on commercial terms, and less well suited to situations in which they refinance ongoing losses, sustain excess capacity or generate lending driven by policy incentives rather than project quality. Estimated turning points should therefore be understood as indicators of a changing marginal effect rather than as prescriptive lending limits.

2.11. Research Gap and Positioning of the Dissertation

The literature reviewed in this chapter establishes that financial development can raise economic performance without doing so uniformly. Finance and growth may influence one another, the effect may weaken or reverse, and the composition and destination of credit condition the outcome. Sectors differ in their dependence on external finance, their collateral structures, their adjustment speeds and their capacity to absorb investment productively, while income level, institutional quality and historical period shape all of these relationships. A further and less frequently acknowledged distinction concerns the margin at which finance is measured, since the accumulated stock of credit and its rate of expansion are separate objects with separate empirical properties and, on the evidence assembled in Section 2.4, separate effects.

Four gaps remain. Most nonlinear research relies on aggregate measures of both finance and output, acknowledging potential turning points without establishing whether nonlinearity originates within individual sectors or arises from shifts in the composition of credit and output. Many sectoral studies concentrate on manufacturing, relate sectoral credit to aggregate output or impose linearity, leaving agriculture and the heterogeneous service branches comparatively unexamined. Integrated designs that combine cross-country and single-country evidence, allowing global regularities to be compared with detailed national sectoral dynamics, are rare. And although the distinction between credit levels and credit growth is well established in the crisis and business-cycle literature, it has not been carried systematically into sectoral analyses of value added, where the two margins are often treated interchangeably.

The first empirical investigation addresses the international gap by examining agriculture, industry including construction and services between 1992 and 2019, using separate financial indicators where the data permit, comparing pre-crisis, post-crisis and full-sample periods and contrasting World Bank income groups. Because the credit series are integrated of order one, they enter in first differences, so the estimates describe how the pace of credit expansion relates to sectoral composition. Linear, quadratic, cubic and quartic forms are tested rather than assumed, and fixed-effects estimation with Driscoll–Kraay standard errors is compared with two-step System GMM in order to assess robustness to cross-sectional dependence, serial correlation, dynamics and potential endogeneity.

The outcome variables in the international analysis are sectoral value-added shares of GDP, which situates the study at the intersection of the finance–growth and structural transformation literatures. No coefficient is treated as an effect on absolute sectoral output. The question is instead whether credit growth is associated with changes in the relative weight of agriculture, industry and services, and whether that association differs across financial environments and stages of development, a question given additional force by the evidence that credit booms can reallocate resources toward lower-productivity activities (Borio et al., 2016; Müller & Verner, 2024). The second empirical investigation addresses the allocation and horizon gaps using quarterly

Turkish data, constructing a matched panel of sectoral total cash loans and sectoral value added across eleven activities from 2005Q1 to 2026Q1. It tests for cross-sectional dependence and heterogeneous integration, examines causal feedback and estimates cross-sectionally augmented distributed lag and error-correction specifications that separate short-run growth effects from long-run equilibrium relationships. Linear, quadratic, cubic and quartic forms are again tested rather than assumed. The value of this design for the present argument is that it retains credit levels in the long-run component while modelling credit growth in the short-run component, so the two margins that the cross-country chapter cannot separate are estimated jointly within a single framework.

Combining the two investigations provides a form of triangulation. The cross-country panel identifies broad patterns across sectors, periods and income groups, while the Türkiye panel establishes whether comparable patterns survive when credit and output are matched at a finer sectoral resolution. Results appearing in both settings carry wider significance, whereas divergences point toward the influence of national institutions, policy, financial structure and sectoral composition.

The contribution is therefore not a single replacement for the conventional finance–growth coefficient. It is an argument, supported by evidence, that the relationship must be evaluated along several dimensions simultaneously: the pace and level of credit, its destination, the producing sector, the stage of development, the historical period and the time horizon. This position is consistent with the existing literature while extending it, since sectoral heterogeneity, nonlinearity and the distinction between credit stocks and credit flows are treated as interrelated empirical problems rather than as separate ones.

The policy implication is correspondingly specific. Financial development should be assessed by the capacity of the financial system to support productive activity rather than by the volume or the speed of credit creation. An aggregate expansion may conceal underfunding in some sectors and excessive leverage in others, so effective policy must attend to allocation, maturity, borrower quality, collateral arrangements and the complementary institutions that allow finance to generate value added.

Estimated turning points are not targets. They mark ranges in which the marginal relationship changes and in which the quality of additional lending merits closer analysis.

CHAPTER 3

CROSS-COUNTRY DATA AND EXPLORATORY ANALYSIS

3.1. Panel Data Construction and Sample Selection

Cross-country differences in data coverage and reporting practices require a deliberate construction strategy, and this study therefore assembles a consistent and comparable panel dataset that prevents the exceptional economic disruptions caused by COVID-19 from affecting the results. The analysis draws on annual data covering 1992 to 2019, a window selected for variable availability and for its capacity to isolate relationships formed within a pre-pandemic economic environment. Alongside the full-sample analysis, separate estimations are conducted for the periods 1992–2007 and 2010–2019, allowing the study to determine whether the relationships of interest vary before and after the 2008 global financial crisis while preventing the abnormal conditions of the crisis from skewing the results. The years 2008 and 2009 are excluded from these subperiods, reflecting the severity of the disruptions recorded during that interval, which simultaneously impacted financial conditions, credit, investment and economic activity. These subperiod estimations serve a further purpose, assessing the stability of the relationships across pre-crisis and post-crisis environments.

Country selection follows the World Bank's official income classification, a widely recognised and standardised framework for classifying economies, which keeps the composition of the sample objective and transparent rather than the product of arbitrary choices. Including all four income groups – low-income, lower-middle-income, upper-middle-income and high-income – the dataset captures countries at markedly different stages of economic development, thereby enabling meaningful cross-country comparison. The dissertation examines advanced economies but

extends equally to transition and developing countries, allowing a thorough analysis of how the independent variables perform across diverse economic and social contexts. This breadth may reveal distinct patterns between low-income and high-income economies, adding depth to the research and contributing to the literature.

Only countries with reliable and comprehensive data are retained, while those with significant data gaps are excluded, reducing bias and improving the accuracy of the econometric analysis. Anchoring the sample in the World Bank's classification additionally ensures that the findings remain comparable with other empirical studies, strengthening both the methodological rigour and the external validity of the research.

Countries are sorted into income groups using the World Bank's 2026 framework, which ranks economies by GNI per capita calculated through the Atlas method, a standardised approach that smooths out exchange rate fluctuations and makes comparisons across nations easier. Under this system, low-income countries have a GNI per capita of $1,135 or less in 2024 values, lower-middle-income countries range from $1,136 to $4,495, upper-middle-income countries fall between $4,496 and $13,935 and high-income countries exceed $13,935. This classification aligns with international standards and facilitates comparison with extensive empirical research. Because the Atlas method mitigates market exchange-rate volatility and inflation, it offers a stable basis for economic analysis, and these advantages make it the most reliable classification system available and the one used throughout this study.

Consistent with the World Bank's income classification, the sample is designed to represent all income groups, ensuring that the dataset captures a wide range of economic development levels and regions. Retaining within each income category only those countries with reliable and sufficient data, the study maintains methodological robustness while remaining aligned with existing research. Table 1 shows the countries included in the industry and services sectors, categorised according to the World Bank's income groups.

Table 1. World Bank Income Classification: Country List for the Industry and Services Sectors

Low income	Lower-middle income	Upper-middle income	High income	Not Classified
Afghanistan	Angola	Albania	Antigua and Barbuda	Ethiopia
Burkina Faso	Bangladesh	Algeria	Aruba	Venezuela, RB
Burundi	Benin	Argentina	Australia	
Central African Republic	Bhutan	Armenia	Austria	
Chad	Bolivia	Azerbaijan	Bahamas, The	
Congo, Dem. Rep.	Cambodia	Belarus	Bahrain	
Eritrea	Cameroon	Belize	Barbados	
Gambia, The	Comoros	Bosnia and Herzegovina	Belgium	
Guinea-Bissau	Congo, Rep.	Botswana	Brunei Darussalam	
Liberia	Cote d'Ivoire	Brazil	Bulgaria	
Madagascar	Djibouti	Cabo Verde	Canada	
Malawi	Egypt, Arab Rep.	China	Chile	
Mali	Eswatini	Colombia	Costa Rica	
Mozambique	Ghana	Dominica	Croatia	
Niger	Guinea	Dominican Republic	Cyprus	
Rwanda	Haiti	Ecuador	Czechia	
Sierra Leone	Honduras	El Salvador	Denmark	
South Sudan	India	Equatorial Guinea	Estonia	
Sudan	Jordan	Fiji	Finland	
Syrian Arab Republic	Kenya	Gabon	France	
Togo	Kyrgyz Republic	Georgia	Germany	

Table 1. (continued)

Uganda	Lao PDR	Grenada	Greece	
Yemen, Rep.	Lebanon	Guatemala	Guyana	
	Lesotho	Indonesia	Hong Kong SAR, China	
	Mauritania	Iran, Islamic Rep.	Hungary	
	Micronesia, Fed. Sts.	Iraq	Iceland	
	Morocco	Jamaica	Ireland	
	Myanmar	Kazakhstan	Israel	
	Namibia	Kosovo	Italy	
	Nepal	Libya	Japan	
	Nicaragua	Malaysia	Korea, Rep.	
	Nigeria	Maldives	Kuwait	
	Pakistan	Mauritius	Latvia	
	Papua New Guinea	Mexico	Lithuania	
	Philippines	Moldova	Luxembourg	
	Sao Tome and Principe	Mongolia	Macao SAR, China	
	Senegal	Montenegro	Malta	
	Solomon Islands	North Macedonia	Netherlands	
	Sri Lanka	Paraguay	New Zealand	
	Tajikistan	Peru	Norway	
	Tanzania	Samoa	Oman	
	Timor-Leste	Serbia	Panama	
	Tunisia	South Africa	Poland	
	Uzbekistan	St. Lucia	Portugal	
	Vanuatu	St. Vincent and the Grenadines	Qatar	
	Viet Nam	Suriname	Romania	
	West Bank and Gaza	Thailand	Russian Federation	

Table 1. (continued)

	Zambia	Tonga	San Marino	
	Zimbabwe	Türkiye	Saudi Arabia	
		Ukraine	Seychelles	
			Singapore	
			Slovak Republic	
			Slovenia	
			Spain	
			St. Kitts and Nevis	
			Sweden	
			Switzerland	
			Trinidad and Tobago	
			United Arab Emirates	
			United Kingdom	
			United States	
			Uruguay	

The geographical coverage of the agricultural sector analysis is somewhat narrower than that of the industry and services sectors. Data on credit to agriculture, forestry and fishing exist for 131 countries and territories, yet the effective sample size decreases once these data are matched with the relevant dependent variable needed for the analysis, leaving 109 countries in the agricultural-sector estimations. This decline results mainly from variations in data availability across countries and years, and especially from the lack of overlapping observations between agricultural credit and agricultural value added.

Table 2 shows the countries included in the agricultural-sector sample, categorised by the World Bank's income classification, thereby giving a clearer picture of the income distribution and the cross-country composition of the final sample.

Table 2. World Bank Income Classification: Country List for the Agriculture Sector

Low income	Lower-middle income	Upper-middle income	High income	Not Classified
Afghanistan	Angola	Albania	Antigua and Barbuda	Ethiopia
Burkina Faso	Bangladesh	Argentina	Aruba	
Burundi	Benin	Armenia	Australia	
Guinea-Bissau	Bhutan	Azerbaijan	Austria	
Liberia	Cambodia	Belarus	Bahrain	
Malawi	Ghana	Belize	Barbados	
Mali	Honduras	Bosnia and Herzegovina	Belgium	
Mozambique	India	Botswana	Bulgaria	
Niger	Jordan	Brazil	Canada	
Rwanda	Kenya	Cabo Verde	Costa Rica	
Sudan	Lebanon	Dominica	Cyprus	
Syrian Arab Republic	Lesotho	Dominican Republic	Czechia	
Togo	Morocco	Ecuador	Estonia	
Uganda	Namibia	El Salvador	Finland	
	Nepal	Gabon	France	
	Nicaragua	Georgia	Germany	
	Nigeria	Grenada	Greece	
	Pakistan	Guatemala	Guyana	
	Papua New Guinea	Indonesia	Hungary	
	Philippines	Iraq	Ireland	
	Senegal	Jamaica	Israel	
	Sri Lanka	Kazakhstan	Italy	
	Tajikistan	Malaysia	Japan	
	Timor-Leste	Maldives	New Zealand	
	Tunisia	Mexico	Oman	
	Vanuatu	Peru	Panama	
	Viet Nam	Serbia	Qatar	

Table 2. (continued)

	Zambia	Suriname	Russian Federation	
	Zimbabwe	Thailand	Seychelles	
		Ukraine	Singapore	
			Spain	
			Switzerland	
			Trinidad and Tobago	
			United Arab Emirates	
			Uruguay	

3.2. Variable Specifications and Measurement

3.2.1. Sectoral Value-Added Variables and ISIC Framework

The main goal is to examine how the quantitative measures of financial development impact the economic growth of the key sectors – agriculture including forestry and fisheries, industry including construction, and services – while accounting for other relevant factors. This focus aligns with the traditional three-sector classification widely used in the structural transformation literature (Herrendorf et al., 2014), where these sectors represent the primary phases of economic development, with labour transitioning from resource-based activities to manufacturing and ultimately to services (McMillan & Rodrik, 2011). Including forestry and fishing within agriculture, and construction within industry, adheres to the International Standard Industrial Classification (ISIC, Rev. 4), which categorises these activities accordingly. Services are the most dynamic part of modern economies and often constitute the largest GDP share in advanced nations (Herrendorf et al., 2014). Adopting this three-part classification ensures consistency with international standards and offers a solid foundation for analysing productivity across different development stages.

Most of the literature on financial development and growth has traditionally focused on either aggregate GDP growth or per capita income growth (Levine, 2005). While

these indicators offer a broad view of economic progress, they assume that the advantages of financial development are evenly distributed throughout the economy, an assumption increasingly questioned in recent studies. Part of the literature has consequently shifted toward total factor productivity (TFP) as an alternative performance metric (Beck et al., 2000), while others have examined the long-run relationship between financial depth and output through panel unit-root and cointegration methods (Christopoulos & Tsionas, 2004). More sector-specific analyses have in turn evaluated the overall output in sectors such as agriculture, industry and services in order to explore the relationship between financial development and sectoral performance (Tabassum et al., 2025). Although these methods yield valuable insights, they often overlook the fact that the impact of finance can differ across sectors depending on factors such as capital intensity, exposure to international trade and technological change. Capital-intensive industries, for instance, might benefit more from financial deepening than subsistence-based agriculture, and service sectors may respond differently again, given their dependence on human capital and innovation.

The dependent variables in this dissertation are the GDP shares of sectoral value added, defined as the percentage of GDP attributed to agriculture including forestry and fishing, industry including construction, and services. This choice enables a disaggregated analysis of economic performance across the three main sectors, moving beyond traditional growth metrics towards a more detailed and policy-relevant understanding of how financial development impacts the economy. Focusing on sector-specific value-added shares of GDP reveals the mechanisms underlying the role of finance in growth and provides targeted insights that help policymakers tailor financial reforms to their economic structures. It also strengthens the analytical depth of the study, aligning it with a growing body of research that emphasises the importance of disaggregated methods for exploring the finance–growth relationship. The data for these variables were obtained from the World Bank's World Development Indicators (WDI) for the years 1991–2019 (World Bank, 2026c).

The sectoral composition of economic activity is illustrated through three value-added indicators, each expressed as a percentage of gross domestic product. The

variable ‘Agriculture, forestry, and fishing, value added (% of GDP)’, abbreviated as AVASGDP, measures the contribution of these sectors to overall economic output and indicates their relative importance within the economy. ‘Industry (including construction), value added (% of GDP)’, referred to as IVASGDP, captures the share of manufacturing, mining, utilities and construction, reflecting the weight of the industrial sector in the economy, while ‘Services, value added (% of GDP)’, known as SVASGDP, represents the contribution of the service sector to GDP. Using these percentages standardises cross-country comparisons by minimising differences in overall economic size and emphasising variations in sectoral output across countries and over time. AVASGDP, IVASGDP and SVASGDP accordingly serve as the key sector-specific outcome variables in the empirical models for agriculture, industry and services respectively.

The sectoral decomposition adopted here rests on the International Standard Industrial Classification of All Economic Activities (ISIC, Rev. 4), developed by the United Nations, a globally accepted framework that categorises economic activities and underpins national accounts, productivity metrics and databases such as those of the World Bank and the OECD (United Nations, 2008).

According to ISIC Rev. 4, the categories of agriculture, forestry and fishing are assigned to codes 01–03, where code 01 pertains to crop and animal production, hunting and related services, code 02 covers forestry and logging and code 03 pertains to fishing and aquaculture. These activities form the core of the primary economic sector, which is mainly resource-dependent and traditionally labour-intensive. Grouping them together underscores their shared reliance on natural resources and their significance in early economic growth.

Industry, including construction, is classified by ISIC codes 05–43 and thus covers the entire secondary sector. Codes 05–09 pertain to mining and quarrying, codes 10–33 include manufacturing activities such as food processing and advanced machinery, code 35 relates to electricity, gas, steam and air conditioning supply, codes 36–39 encompass water supply, sewerage, waste management and remediation, and codes 41–43 are specifically for construction. The classification of

construction within industry reflects its role in transforming materials into durable capital goods and its contribution to productive capacity. Construction is additionally essential for infrastructure development, making it a key part of the industrial sector in ISIC.

Services, finally, include ISIC codes 45–99 and cover the tertiary sector, spanning a wide range of activities: codes 45–47 for wholesale and retail trade plus motor vehicle repairs, 49–53 for transportation and storage, 55–56 for accommodation and food services, 58–63 for information and communication, 64–66 for finance and insurance, 68 for real estate, 69–82 for professional, scientific, technical and administrative services, 84 for public administration and defence, 85 for education, 86–88 for health and social work and 90–99 for arts, entertainment, recreation, other services and household employment. The wide scope of this category highlights the growing significance of services in modern economies, where they constitute the majority of value added and employment in high-income nations.

By explicitly aligning the dependent variables with ISIC codes 01–03 for agriculture, forestry and fishing, 05–43 for industry including construction and 45–99 for services, this study ensures full consistency with international statistical standards. This alignment offers three main benefits: (i) it ensures that the results are comparable across countries and over time, as all official data sources adopt ISIC as their framework, (ii) it aligns with the historical and theoretical concept of structural transformation, in which labour shifts from primary to secondary and then to tertiary sectors, and (iii) it makes the analysis directly comparable to existing empirical research and to policy reports by organisations such as the World Bank, the IMF and the OECD. The sectoral focus of this dissertation is consequently not arbitrary but grounded in international statistical conventions and in the economic literature on structural change.

3.2.2. Proxies for Financial Development

Empirical research on finance and growth has traditionally debated which proxies best represent the quantitative aspect of financial development, and various

indicators have been suggested, each reflecting different facets of the system such as financial depth, intermediation and market activity. The most commonly used measures are summarised below.

The most prominent proxy is private credit by deposit money banks relative to GDP, which measures the volume of banking-system loans to the private sector as a share of national income. This indicator has been central to the empirical approach introduced by King and Levine (1993) and further refined by Rajan and Zingales (1998) and Levine et al. (2000). Its strength lies in excluding government and public enterprise credit, thereby focusing on resources directed toward potentially productive private investments. Numerous cross-country studies report a strong positive association between private credit and long-term economic growth (Levine, 2005; Levine et al., 2000), and King and Levine (1993) trace that association to the proximate sources of growth, finding that greater financial depth predicts faster physical capital accumulation and larger improvements in the efficiency with which capital is employed.

Another branch of the literature examines liquid liabilities such as M2 or M3 relative to GDP, often termed ‘financial depth’. Pioneers such as Goldsmith (1969), McKinnon (1973) and Shaw (1973) relied primarily on this measure, interpreting it as a reflection of the size of financial intermediation compared to the real economy. Despite their popularity in early growth research, liquid liabilities are criticised for failing to distinguish clearly between savings aimed at productive investment and monetary expansion driven by inflationary finance, since the aggregate records the monetary liabilities of the financial system irrespective of the use to which the underlying funds are eventually put. The choice of proxy therefore shapes the conclusions drawn, a concern carrying particular weight in a literature where even the direction of the relationship remains contested, with evidence indicating that financial development generally leads output while a reverse channel also operates (Calderón & Liu, 2003).

A related indicator is total domestic credit as a percentage of GDP, which reflects credit extended by banks and non-bank financial institutions to both private and

public borrowers. Although this measure covers a broad spectrum of financial activities, scholars such as Levine (2005) highlight its main limitation in growth studies, namely that it includes government lending. Public borrowing frequently funds consumption or deficit spending rather than productive investments, so relying solely on broad domestic credit may mask the true connection between financial deepening and actual productive capacity.

In addition to banking indicators, researchers often use stock market metrics such as market capitalisation-to-GDP ratios, turnover ratios and trading volume, indicators compiled and compared systematically across countries by Demirgüç-Kunt and Levine (1996). Levine and Zervos (1998) show that equity market growth independently and significantly impacts economic growth, even when banking factors are considered. Such metrics reflect the liquidity and risk-sharing roles of capital markets, providing a different view of financial development. In many developing and emerging economies, however, sparse and illiquid stock exchanges limit the effectiveness of these exchanges as reliable cross-country indicators.

Recent work has developed composite indices of financial development that encompass multiple aspects of the financial system, with the World Bank's benchmarking framework (Čihák et al., 2012) and the IMF's Financial Development Index (Svirydzenka, 2016) as notable examples. These indices measure financial depth, access and efficiency across banking and capital markets. While offering a comprehensive view of financial systems, they are limited by shorter time series and less extensive country coverage, posing challenges for long-term panel analysis.

Among the various quantitative indicators of financial development, the ratio of private credit from deposit money banks to GDP has become the primary proxy used in empirical research, a choice supported by theoretical reasoning as much as by extensive empirical evidence from influential studies. King and Levine (1993) first identified the advantages of this measure in cross-country growth regressions, showing that higher private credit ratios are associated with faster growth, greater capital accumulation and improved productivity. Unlike broader measures such as

M2/GDP, private credit directly reflects the role of the banking sector in channelling resources to the private sector rather than to government spending or inflationary finance. This distinction is key, since Rajan and Zingales (1998) find that private-sector financing promotes innovation and entrepreneurship, while public-sector lending may not lead to productive investment. Levine et al. (2000) provided further evidence through a panel data analysis, establishing not only a strong correlation but also a causal relationship between private credit and long-term economic growth, with increases in private lending preceding improvements in GDP growth. Similarly, Beck et al. (2000) demonstrate that this indicator outperforms alternatives such as M2/GDP or total domestic credit in explaining cross-country differences in growth. Indicators such as liquid liabilities (M2/GDP) or total domestic credit encompass flows to governments and monetary aggregates that need not correspond to intermediated lending to productive users, which weakens their explanatory power for the mechanisms through which finance affects growth. Stock market measures, although useful for developed economies, are less reliable in countries with thin or illiquid equity markets (Levine & Zervos, 1998). The literature therefore largely agrees that private credit from deposit money banks, as a percentage of GDP, is the most dependable single proxy for the quantitative aspect of financial development, its empirical robustness, its focus on productive private-sector activity and its availability across long time periods and diverse countries making it the preferred choice in cross-country growth research and in panel analyses covering both developed and developing economies.

The literature clearly differentiates between private credit by deposit money banks to GDP and a broader measure that also includes other financial institutions, likewise expressed as a percentage. Although both metrics aim to assess financial intermediation levels, scholars highlight their distinct analytical significance. The first is the standard measure, focusing solely on bank-based private-sector credit, which is directly related to capital formation, entrepreneurship and innovation. The second, being more inclusive, also accounts for lending by non-bank entities and sometimes for public credit, which can weaken its explanatory power and yield more varied results. Private credit extended by deposit money banks, as a percentage of GDP, is consequently considered a robust and reliable indicator of financial

development, especially in long-term panel studies involving both developed and developing countries.

Building on existing finance–growth research, this study uses private credit from deposit money banks, measured as a percentage of GDP and labelled PC, as the main proxy for financial development in the industry and services sectors. The analysis would ideally focus on credit specifically directed to industrial and service activities. Comprehensive sector-specific credit data are unavailable across many countries and years, however, so a consistent panel dataset of that kind cannot be constructed. Private credit by deposit money banks, relative to GDP, is used as an alternative, being a common and effective indicator of financial development and of the level of formal banking system support for the private sector. This approach allows a reasonable approximation of the availability of private-sector financing for industry and services, ensuring comparability across countries and sufficient temporal coverage.

A more specialised measure of financial development exists for the agricultural sector, where credit to agriculture, forestry and fishing – denoted as AC in this study – is used instead of the broader private-credit indicator. This choice enables the analysis to focus on financial resources specifically allocated to agricultural and related activities, including forestry and fishing, rather than relying on an overall private credit measure that might mainly reflect financing conditions in other sectors. This differentiation is crucial, because the financing structure, production cycle, risk profile and credit needs of agricultural activities can vary significantly from those in industry and services. Using a sector-specific agricultural credit measure thus enhances the alignment between the financial development indicator and the sectoral value-added variable in the agricultural models.

The series for credit to agriculture, forestry and fishing is reported annually from 1991 to 2024, using 2015 as the base year, with these data sourced from the Food and Agriculture Organisation of the United Nations (Food and Agriculture Organization of the United Nations, 2025). Private credit by deposit money banks, as a percentage of GDP, was obtained in contrast from the World Bank’s Global

Financial Development Database (World Bank, 2026a) for countries with available data, this series covering the period from 1992 to 2019 at an annual frequency. Although the agricultural credit data extend beyond the final period studied, the analysis focuses on the overlapping timeframe, ensuring consistency across the sectoral models and allowing a meaningful comparison of the results.

Several steps were necessary before incorporating the agricultural credit variable into the empirical analysis. The AC series was initially presented as an absolute monetary value, making it non-comparable across countries with different sizes, and it was not in the same relative format as private credit by deposit money banks, which is reported as a percentage of GDP. Agricultural credit was therefore normalised by GDP, specifically by dividing the AC series by GDP in constant 2015 US dollars, in order to create a comparable indicator of sectoral financial development. This approach also aligns with the 2015 base-year structure of the agricultural credit data.

A unit conversion was performed before calculating the agricultural credit-to-GDP ratio. The original AC variable is in millions of US dollars, while GDP in constant 2015 US dollars is given in individual dollar units, so the AC variable was multiplied by 1,000,000 and then divided by GDP in 2015 US dollars in order to align both series on the same scale. The resulting ratio was multiplied by 100 to express it as a percentage of GDP, making it comparable to the private credit indicator used in the industry and services models. The final variable, called ACGDP, was constructed as follows:

$$ACGDP = [(AC \times 1{,}000{,}000) / \text{GDP in constant 2015 US dollars}] \times 100 \quad (1)$$

This transformation minimises distortions caused by variations in country size and monetary units, enabling a consistent comparison of agricultural credit intensity across countries and over time. ACGDP accordingly serves as the sector-specific financial development proxy in the agricultural-sector estimations, while PC is used for the industry and services sectors. The choice of these different indicators is driven by the availability of internationally comparable data and aims to align the

financial development measure as precisely as possible with the specific economic sector in each analysis.

3.2.3. Macroeconomic Control Variables

Obtaining unbiased estimates of how financial development influences sectoral growth makes it essential to account for the key determinants of growth. Failing to include these factors could mix the effects of finance with other crucial factors such as capital accumulation, macroeconomic stability, fiscal policy and human capital development. The discussion that follows explains the theoretical reasons and the empirical evidence for incorporating each control variable.

Investment, measured as gross capital formation relative to GDP, is among the most reliable indicators of long-term economic performance. Early regressions across countries by Levine and Renelt (1992) showed that the investment-to-output ratio remains a consistently significant variable under rigorous sensitivity tests. Similarly, De Long and Summers (1991, 1992) found that equipment investment significantly influences productivity growth, especially in manufacturing. Within neoclassical and augmented Solow models, including investment accounts for the effects of physical capital accumulation rather than attributing these effects solely to financial development (Mankiw et al., 1992). Gross capital formation relative to GDP, abbreviated as GCF, is used in this study.

Price level stability is a key factor influencing growth outcomes. Fischer (1993) showed that high, volatile inflation hampers both investment and productivity, thereby slowing economic growth, and later studies confirmed that inflation crises often lead to sustained declines in growth performance (Bruno & Easterly, 1998). Khan and Senhadji (2001) found threshold effects in this relationship, indicating that moderate inflation might be harmless while sustained rates above single digits are harmful. Controlling for inflation ensures that the negative effects of macroeconomic instability are not mistakenly attributed to the financial system. Inflation is abbreviated as INF in this analysis. The size and composition of government spending significantly influence economic incentives. Barro (1991) found a negative

link between the proportion of government consumption and long-term growth, primarily due to crowding-out effects and distortionary taxes. Devarajan et al. (1996) added that the effectiveness of government expenditure depends heavily on how it is allocated between current expenses and capital investments. In OECD countries, Kneller et al. (1999) observed that 'productive' expenditures such as infrastructure and education promote growth, whereas unproductive current spending tends to hinder it. Controlling for government consumption helps prevent biases in the financial development coefficients caused by fiscal policy. Government consumption is represented in this study by the variable GC.

Human capital is widely acknowledged as a key factor in economic development. Augmented Solow models (Mankiw et al., 1992) highlight the importance of accumulating human capital, or educational achievement, for reaching steady-state income levels. Barro and Lee (2013) created a comprehensive global dataset on educational attainment that continues to support research on growth, while Hanushek and Woessmann (2008, 2012) showed that educational quality, measured by cognitive test scores, has an even stronger association with growth than enrollment rates. Secondary school enrollment nonetheless remains a common indicator of human capital development, especially in large cross-country analyses, and the study therefore uses secondary school enrollment (SE) as a proxy for human capital.

External orientation significantly influences growth trajectories. Frankel and Romer (1999) used geography-based instrumental variables to provide strong evidence that trade openness causally impacts income, and Alcalá and Ciccone (2004) showed that openness boosts productivity through specialisation and technology diffusion. Case studies also indicate that episodes of trade liberalisation lead to faster growth and increased investment (Wacziarg & Welch, 2008). Controlling for openness helps ensure that the financial variables do not simply represent the pro-growth characteristics of the external environment. In this study, the trade openness variable, TO, was calculated by summing 'imports of goods and services (% of GDP)' and 'exports of goods and services (% of GDP)'.

Institutional quality influences economic growth both directly and by affecting how effectively financial resources are mobilised and distributed. Issues such as weak

transparency, limited accountability and widespread corruption can distort lending decisions, divert credit to politically connected or low-productivity activities, raise uncertainty and undermine the link between financial development and productive investment. Mauro (1995) shows that corruption hampers investment and growth, while Demetriades and Law (2006) find that financial development has a greater impact on economic performance when the financial system operates within strong institutional frameworks. Similarly, Law et al. (2013) suggest that the growth benefits of financial development depend on reaching a certain level of institutional quality. The CPIA rating for Transparency, Accountability, and Corruption in the Public Sector was consequently included as an institutional control variable, called GOV, which ranges from 1 for weak institutions to 6 for strong institutions and evaluates aspects such as officials' accountability, access to reliable public information, state capture and resource integrity, with the data sourced from the World Bank's World Development Indicators database (World Bank, 2026c). In the empirical models, GOV was used to control for the independent impact of governance quality on sectoral economic performance, and it was also interacted with the relevant financial indicators to explore whether governance influences financial development across sectors. The interaction specifications provided no consistent evidence that public-sector transparency, accountability and corruption conditions moderate the relationship between credit and sectoral value-added shares. Neither the linear credit–GOV interaction nor the interaction involving the quadratic credit term attained statistical significance in the agricultural, industrial and services models, a result holding under fixed effects with Driscoll–Kraay standard errors and under two-step System GMM alike. The direct GOV coefficient did reach significance at the 10% level in some fixed-effects specifications, yet these associations proved sector-specific and were not reproduced by the corresponding System GMM estimates, suggesting that they may reflect the identifying assumptions of a particular estimator rather than a stable underlying relationship. The findings therefore do not support a robust conditional role for GOV in the credit–sectoral output relationship. Given this absence of consistent evidence across estimators and functional forms, GOV is excluded from the primary set of control variables and the complete results are reported in Appendix rather than in the main regression tables, a choice that prioritises simplicity and comparability rather than

presuming governance to be unimportant. The absence of significance may also reflect limitations in the CPIA data, which are available primarily for IDA-eligible countries from the mid-2000s onwards, reducing the sample size and the power of the interaction tests.

These controls serve as the main channels through which economies build capital, stabilise macroeconomic conditions, allocate public resources, develop human capital and connect with the global economy, and empirical research confirms each as a reliable long-term growth indicator. Incorporating them into the econometric model allows a thorough assessment of whether financial development independently influences the sectoral value-added shares of GDP, beyond simply representing these core factors.Table 3 details these variables with their codes, definitions and sources.

Table 3. Variables, Definitions and Sources (World Bank WDI, World Bank GFD and FAO)

Variable Code	Variable Name	Long Definition	Source
NV.AGR.TOTL.ZS	Agriculture, forestry, and fishing, value added (% of GDP)	Agriculture, forestry, and fishing encompass ISIC Rev. 4 divisions 01–03, including crop cultivation, livestock rearing, forestry, fishing, and related activities that utilise natural resources. Value added indicates the sector's economic contribution by deducting intermediate consumption from total output. This measure is expressed as a percentage of GDP, highlighting the sector's significance within the broader economy.	World Development Indicators (WDI)

Table 3. (continued)

NV.IND.TOTL.ZS	Industry (including construction), value added (% of GDP)	The industry sector, which includes construction, encompasses ISIC Rev. 4 divisions 05–43. It covers activities such as mining, manufacturing, construction, electricity, gas, and water services. Value added is determined by subtracting intermediate consumption from total output. This metric is presented as a percentage of GDP to highlight the sector's share of the overall economy.	World Development Indicators (WDI)
NV.SRV.TOTL.ZS	Services, value added (% of GDP)	Services encompass ISIC Rev. 4 divisions 45–99, including trade, transport, accommodation, finance, real estate, public administration, education, health, and other social and personal services. Value added is determined by subtracting intermediate consumption from total output, and it is expressed as a percentage of GDP to illustrate the sector's relative impact on the economy.	World Development Indicators (WDI)

Table 3. (continued)

GFDD.DI.01	Private credit by deposit money banks to GDP (%)	This represents the proportion of domestic private sector credit extended by deposit-taking banks relative to GDP, serving as an indicator of financial intermediation.	Global Financial Development (GFD)
23068	Credit to Agriculture, Forestry and Fishing	Credit to agriculture encompasses loans and advances provided by banks to farmers, rural households, agricultural cooperatives, and related businesses. These data are mainly sourced from central bank monetary and financial data and are categorised based on ISIC Rev. 4, with credit to agriculture, forestry, and fishing reported separately from credit to manufacturing and services.	Food and Agriculture Organisation of the United Nations (FAO)
NE.GDI.TOTL.ZS	Gross capital formation (% of GDP)	Gross capital formation, previously known as gross domestic investment, includes spending on new fixed assets and net changes in inventories, both expressed as a percentage of GDP.	World Development Indicators (WDI)

Table 3. (continued)

NE.CON.GOVT.ZS	General government final consumption expenditure (% of GDP)	General government final consumption expenditure, previously known as government consumption expenditure, includes all current government spending on goods and services, expressed as a percentage of the country's GDP.	World Development Indicators (WDI)
NE.IMP.GNFS.ZS	Imports of goods and services (% of GDP)	Imports of goods and services indicate the proportion of GDP that results from all goods and market services received from the rest of the world.	World Development Indicators (WDI)
NE.EXP.GNFS.ZS	Exports of goods and services (% of GDP)	Exports of goods and services indicate the total value of all goods and market services supplied to foreign countries, expressed as a portion of GDP.	World Development Indicators (WDI)
SE.SEC.ENRR	School enrollment, secondary (% gross)	The gross enrollment ratio is the proportion of total students enrolled, of any age, compared to the population of the age group designated for secondary education.	World Development Indicators (WDI)

Table 3. (continued)

FP.CPI.TOTL.ZG	Inflation, consumer prices (annual %)	Inflation is gauged by the yearly increase in the consumer price index, which indicates how much the typical consumer's cost for a basket of goods and services has changed.	World Development Indicators (WDI)
IQ.CPA.TRAN.XQ	CPIA transparency, accountability, and corruption in the public sector rating (1=low to 6=high)	The CPIA evaluates a country's policy and institutional framework to determine its support for sustainable growth, poverty alleviation, and effective use of aid. It provides an overall score, along with scores for sixteen specific criteria, divided into four categories: economic management, structural policies, social inclusion and equity, and public sector management. The last category specifically examines transparency, accountability, and corruption, emphasising oversight, access to public information, the influence of vested interests, and integrity in the handling of public resources.	World Development Indicators (WDI)

Table 3. (continued)

NY.GDP.MKTP.KD	GDP (constant 2015 US$)	Gross domestic product (GDP) represents the total income generated from producing goods and services within a specific economy during a specified period. It is quoted at constant prices, adjusted to negate the effects of inflation over time, with 2015 as the reference year. The GDP figure is presented in United States dollars.	World Development Indicators (WDI)
NY.GDP.PCAP.KD.ZG	GDP per capita growth (annual %)	Annual GDP per capita growth rate calculated in constant local currency. Data are expressed in constant 2015 U.S. dollars.	World Development Indicators (WDI)

Notes: GOV was tested but excluded from the reported specifications, as explained in Section 3.2.3.

3.2.4. Descriptive Statistics and Correlations

Before the estimation strategy is applied, the variables defined in the preceding subsections warrant a descriptive examination, which establishes the scale and dispersion of each series and exposes the unconditional associations that the econometric models will subsequently subject to controls. Each variable is described in the form in which it enters the estimations rather than in the form in which it is originally published, since the time-series properties of the data dictate that the two are not always the same. The Pesaran (2007) cross-sectionally augmented panel unit-root test rejects the unit-root null for the three sectoral value-added shares and for gross capital formation, government consumption, trade openness and inflation, all of which therefore enter in levels, whereas the null cannot be rejected for private credit, agricultural credit and secondary school enrollment, which consequently enter

as first differences of their logarithms, as the panel unit-root results reported in Table 13 of the methodology chapter document variable by variable. Table 4 reports the abbreviations used throughout the remainder of this dissertation, providing the key that links the variable names appearing in the estimation tables to the definitions, database codes and sources listed in Table 3.

Table 4. Variable Abbreviations Used in the Empirical Analysis

Name	Varlab
AVASGDP	Agriculture, forestry, and fishing, value added (% of GDP)
IVASGDP	Industry (including construction), value added (% of GDP)
SVASGDP	Services, value added (% of GDP)
PC	Private credit by deposit money banks to GDP (%)
TO	Trade openness (Imports plus Exports over GDP)
ACGDP	Agricultural credit over GDP
SE	School enrollment, secondary (% gross)
GC	General government final consumption expenditure (% of GDP)
GCF	Gross capital formation (% of GDP)
INF	Inflation, consumer prices (annual %)
GDPPG	GDP per capita growth (annual %)

Table 5 reports the number of observations, the mean, the standard deviation and the extreme values of each variable as it enters the analysis.

Table 5. Descriptive Statistics of the Variables

Summary statistics for the variables used in the empirical analysis

Variable	**Observations**	**Mean**	**Standard Deviation**	**Minimum**	**Maximum**
ln_AVASGDP	5,172	1.8744	1.3726	-4.3804	4.3700
ln_IVASGDP	5,159	3.1641	0.4662	1.1141	4.4621
ln_SVASGDP	5,042	3.9564	0.2640	2.3850	4.5660
GDPPG	5,707	2.0408	6.2708	-49.1279	140.4906

Table 5. (continued)

d_ln_PC	4,729	0.0250	0.1767	-2.2487	1.3673
d_ln_ACGDP	2,025	-0.0049	0.3976	-5.6617	4.3760
d_ln_SE	3,082	0.0174	0.0562	-0.6304	0.3992
ln_GCF	4,531	3.1074	0.3884	-1.2280	4.3410
ln_GC	4,572	2.7308	0.4391	0.7000	4.9954
INF	4,865	23.8165	374.7896	-16.8597	23,773.1300
ln_TO	4,783	4.2991	0.5723	-3.8633	6.7606

Notes: The table reports the number of observations, mean, standard deviation, minimum and maximum for each variable. The prefix ln_ denotes natural logarithms and d_ln_ first differences of natural logarithms. Observation counts differ because the panel is unbalanced. Values are rounded to four decimals.

Source: Author's calculations based on the study dataset.

The three sectoral value-added shares differ sharply in dispersion, and the ordering is itself informative about the structural position of each sector. The logarithm of the agricultural share records a standard deviation of 1.3726 across a range running from −4.3804 to 4.3700, corresponding to shares of roughly 0.01% and 79% of GDP, whereas the service share is far more tightly concentrated, with a standard deviation of 0.2640 over a range equivalent to approximately 11% and 96% of output, and industry occupies an intermediate position at 0.4662. This pattern indicates that agriculture is the sector that most sharply separates economies at different stages of development, while services account for a substantial share of output almost everywhere, a configuration consistent with the structural transformation evidence presented below in Section 3.3.1.

GDP per capita growth averages 2.0408% with a standard deviation of 6.2708, and its range from −49.1279 to 140.4906 reflects the collapses and rebounds that accompany armed conflict, commodity price shocks and financial crises. Inflation records the widest dispersion in the dataset, combining a mean of 23.8165 with a standard deviation of 374.7896 and a maximum of 23,773.13, magnitudes that reflect

the hyperinflationary episodes a panel of this geographical and temporal breadth necessarily contains.

The differenced variables describe rates of change and therefore carry a direct proportional interpretation. Private credit relative to GDP grows by 2.50% per year on average, with a standard deviation of 0.1767, so that the illustrative movement of 0.10 used in Section 3.3.2.3 represents approximately 0.57 of a standard deviation and thus a substantial though far from exceptional change in credit conditions. The observed range, extending from −2.2487 to 1.3673, coincides with the interval over which the inverted U-shape test is evaluated in Section 3.3.2.1, confirming that the test spans the full support of the data. Agricultural credit is markedly more volatile, its standard deviation of 0.3976 more than double that of private credit, while its mean of −0.0049 indicates a credit-to-GDP ratio that is broadly flat over the period. Secondary school enrollment grows by 1.74% per year with a standard deviation of only 0.0562, the smallest dispersion among the differenced series, which accords with the gradual and highly persistent nature of educational expansion. Table 6 presents Pearson correlation coefficients for each sector separately, each panel computed on a common listwise-complete sample so that the coefficients reported within a panel remain mutually comparable.

Table 6. Sector-Specific Correlation Matrices: Agriculture, Industry and Services, 1992–2019

Pearson Correlation Coefficients

Panel A. Agriculture (N = 1,321)

Variable	ln_AVASGDP	d_ln_ACGDP	d_ln_SE	ln_GCF	ln_GC	INF	ln_TO
ln_AVASGDP	1.0000						
d_ln_ACGDP	0.0518	1.0000					
d_ln_SE	0.2396*	0.0340	1.0000				
ln_GCF	-0.0484	0.0461	0.0459	1.0000			
ln_GC	-0.3194*	-0.0346	-0.0926*	0.0955*	1.0000		
INF	0.0878*	-0.0820*	-0.0249	-0.3085*	-0.2440*	1.0000	
ln_TO	-0.4839*	-0.0371	-0.1563*	0.2835*	0.2919*	-0.1026*	1.0000

Panel B. Industry (N = 2,444)

Variable	ln_IVASGDP	d_ln_PC	d_ln_SE	ln_GCF	ln_GC	INF	ln_TO
ln_IVASGDP	1.0000						
d_ln_PC	-0.0088	1.0000					
d_ln_SE	-0.0307	0.0658*	1.0000				
ln_GCF	0.2732*	0.1402*	0.0101	1.0000			
ln_GC	-0.0567*	0.0407*	-0.1286*	0.0584*	1.0000		
INF	0.0413*	-0.0984*	-0.0202	0.0306	-0.0228	1.0000	
ln_TO	-0.0863*	-0.0047	-0.1226*	0.1965*	0.2972*	-0.0310	1.0000

Panel C. Services (N = 2,426)

Variable	ln_SVASGDP	d_ln_PC	d_ln_SE	ln_GCF	ln_GC	INF	ln_TO
ln_SVASGDP	1.0000						
d_ln_PC	-0.0216	1.0000					
d_ln_SE	-0.1761*	0.0649*	1.0000				
ln_GCF	-0.0983*	0.1378*	0.0083	1.0000			
ln_GC	0.3094*	0.0443*	-0.1282*	0.0635*	1.0000		
INF	-0.0579*	-0.0989*	-0.0202	0.0308	-0.0228	1.0000	
ln_TO	0.3202*	-0.0035	-0.1227*	0.1965*	0.2975*	-0.0310	1.0000

Notes: Pearson correlation coefficients are computed on a common listwise-complete sample for each sectoral model, comprising 1,321 observations for agriculture, 2,444 for industry and 2,426 for services. The prefix d_ denotes first differences. * indicates significance at the 5% level.

Source: Author's calculations based on the study dataset.

The unconditional association between credit growth and sectoral output is weak in every sector. Agricultural credit growth correlates with the agricultural share at 0.0518, private credit growth correlates with the industrial share at −0.0088 and with the service share at −0.0216, and none of these three coefficients attains statistical significance. This pattern is precisely what the transformation of the credit variables should be expected to produce, since differencing removes the long-run co-movement through which financial depth and sectoral composition advance together

as economies develop, leaving only the year-to-year variation in credit conditions. The implication for the empirical strategy is important, because whatever relationship the sectoral models identify will be a conditional and potentially non-linear one that simple correlations cannot reveal, and the absence of a raw association accordingly establishes a demanding rather than a favourable starting point for the analysis.

The variables entering in levels continue to trace the structural transformation that Figure 1 documents. Trade openness correlates negatively with the agricultural share at −0.4839 and positively with the service share at 0.3202, while government consumption follows the same pattern at −0.3194 and 0.3094 respectively, indicating that economies with deeper external integration and larger public sectors tend to have systematically smaller agricultural and larger service sectors. Gross capital formation reverses this configuration, recording its only appreciable positive association with the industrial share at 0.2732 alongside weak negative associations with agriculture and services, which accords with the greater capital intensity of industrial production.

Secondary school enrollment growth carries a positive association with the agricultural share at 0.2396 and a negative association with the service share at −0.1761, indicating that educational expansion proceeds most rapidly in economies where agriculture still occupies a substantial place and more slowly in those where services predominate. This configuration is consistent with the catch-up dynamics of schooling, in which enrollment rises quickly from a low base and decelerates as it approaches universal coverage, and it establishes that the coefficient on this variable speaks to the pace of educational expansion rather than to the level of attainment reached.

The associations among the explanatory variables themselves remain uniformly modest, and the strongest of them reaches only 0.3085, between inflation and gross capital formation in the agricultural panel, followed by 0.2975 between trade openness and government consumption in the services panel. No pair approaches the magnitudes conventionally taken to indicate problematic multicollinearity, and the

correlations involving the differenced credit variables are smaller still, which suggests that the control variables can be included jointly without destabilising the coefficient estimates or inflating their standard errors.

3.3. Stylised Facts and Preliminary Evidence

3.3.1. Global Trends in Sectoral Structural Transformation

Changes in the contribution of the major economic sectors to gross domestic product (GDP) over time reflect the structural transformation and development progress of countries. As economies grow, the significance of agriculture typically diminishes, while the shares of industry and especially services tend to rise. Tracking the evolution of sectoral value-added shares offers an intuitive overview of each country's economic structure and highlights broad differences in developmental stages. For this purpose, the variables 'Agriculture, forestry, and fishing, value added (% of GDP)', 'Industry (including construction), value added (% of GDP)' and 'Services, value added (% of GDP)' are used to monitor sectoral shifts in GDP throughout the sample period. This analysis reveals how the relative importance of agriculture, industry and services has changed across countries and over time, providing valuable context for the empirical analysis that follows.

Figure 1 displays the outcomes of this analysis and shows how the shares of these sectors in GDP have evolved over time. It shows that the share of agriculture in GDP steadily decreases from 1992 to 2019, reflecting a gradual decline in the economic importance of the sector within the studied economies. The industrial sector meanwhile remains relatively stable, with only a slight decrease in recent years. The most significant change occurs in the services sector, which grows consistently and becomes the main component of GDP. These patterns align with the broader trend of structural transformation, in which economies shift from agriculture towards services as they develop, and the data overall indicate a slow shift of economic activity from agriculture to services, with industry remaining relatively steady. These trends are nevertheless general and may vary significantly across individual countries and income groups.

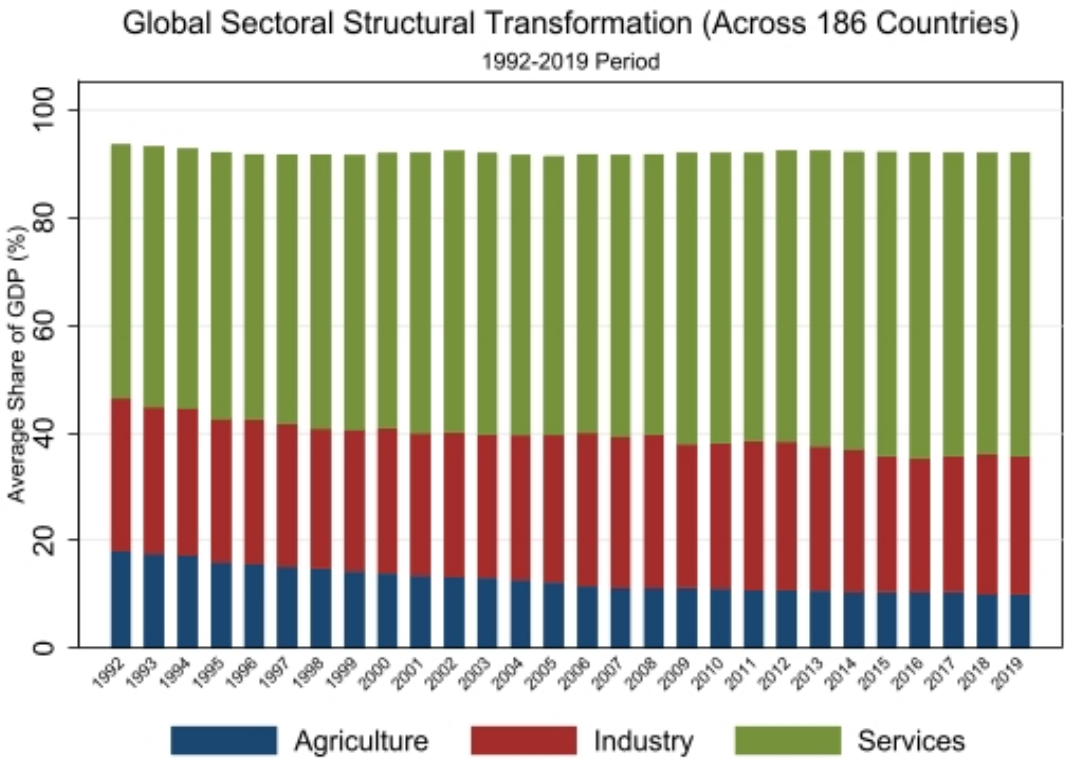


Figure 1. Sectoral Value-Added Shares in GDP: Cross-Country Averages for Agriculture, Industry and Services, 1992–2019

3.3.2. Preliminary Analysis of Aggregate Growth and Financial Development

The empirical investigation opens at the aggregate level, examining how GDP per capita growth, labelled *GDPPG*, relates to financial development across the sample countries before the analysis turns to the sectoral evidence that forms the core of this dissertation. Financial development enters the aggregate models as *d_ln_PC*, the first difference of the natural logarithm of private credit by deposit money banks expressed as a share of GDP. This transformation is dictated by the time-series properties of the data rather than adopted for convenience, since the logarithmic level series contains a unit root, as the methodology section documents, and estimation in levels would accordingly carry the risk of spurious inference. Macroeconomic control variables are subsequently added to account for the country-specific factors that shape growth, and the relationship is estimated for the entire period and for two separate subperiods, enabling an assessment of how the association varies across different global economic conditions.

GDP per capita growth requires no transformation, since it is already expressed as a rate of change and a further logarithmic conversion would serve no analytical purpose. Inflation is retained in its original rate form for a different reason, namely that a logarithmic transformation is inappropriate for a variable capable of taking zero and negative values, as it does during the deflationary and crisis episodes the analysis seeks to represent. The remaining continuous variables are converted into natural logarithms where suitable, an approach that minimises the influence of outliers, enhances comparability across countries of very different economic size and permits a proportional interpretation of the estimated coefficients. Secondary school enrollment then requires the same additional step as the credit indicators, since its logarithmic level series likewise contains a unit root and is therefore differenced, entering the models as *d_ln_SE*. *GDPPG* and inflation accordingly enter in their original form, private credit and secondary school enrollment enter as first differences of their logarithms, while gross capital formation, government consumption and trade openness are expressed in logarithmic levels.

The econometric models are estimated using country fixed effects with Driscoll–Kraay standard errors. The fixed-effects specification absorbs unobserved, time-invariant country characteristics – institutional, geographical, historical and structural traits – that could plausibly be correlated with the explanatory variables and would otherwise contaminate the estimated coefficients. Driscoll–Kraay standard errors are adopted because they remain reliable in the presence of heteroskedasticity, serial correlation and general forms of cross-sectional dependence (Driscoll & Kraay, 1998; Hoechle, 2007), a property that is particularly valuable in a global panel where countries are exposed to common shocks such as shifts in international financial conditions, synchronised business cycles and commodity-price fluctuations. All specifications additionally include year fixed effects to absorb global events and temporal factors affecting the entire sample simultaneously.

The maximum lag length entering the Driscoll–Kraay covariance estimation is not set arbitrarily but follows the rule-of-thumb bandwidth established in the relevant literature, which sets the maximum lag equal to four multiplied by T divided by one hundred, raised to the power of two-ninths, where T denotes the time dimension of

the panel (Hoechle, 2007; Newey & West, 1994). For the twenty-eight-year full sample, this calculation yields approximately 3.01 and therefore indicates an optimal lag length of three, so that all models are estimated with a maximum lag of three. Retaining the same bandwidth for the 1992–2007 and 2010–2019 estimations ensures consistent covariance estimation and preserves comparability between the full period and the two subperiods.

The Driscoll–Kraay fixed-effects estimator does not by itself resolve the endogeneity problems that pervade the finance–growth literature, including reverse causality, omitted time-varying determinants and measurement error. The estimates reported below should accordingly be read as robust conditional correlations rather than as definitive evidence of causal effects, a qualification that attaches to the fixed-effects literature as a whole rather than to this application in particular. One channel warrants explicit attention, since the credit ratio carries output in its denominator and simultaneity between the two series therefore cannot be ruled out within a single-equation framework. The year fixed effects absorb the common component of such variation, while the two-step System GMM estimates presented in Section 3.3.2.4 address the remainder directly through instrumentation, so the concern is met by the design of the analysis rather than left open. Read in these terms, the estimator provides an informative benchmark, establishing the direction, magnitude and consistency of the association between credit growth and per capita growth before the identifying assumptions are tightened.

The analysis covers 1992 to 2019 and distinguishes the subperiods 1992–2007 and 2010–2019, allowing an assessment of whether the relationship changed after the 2008 global financial crisis while preventing the exceptional conditions of the crisis itself from distorting the comparison. The years 2008 and 2009, representing the most acute phase of the crisis and its immediate aftermath, are excluded from the subperiod estimations but retained in the full-period model so that the 1992–2019 sample remains complete. Functional form is selected period by period rather than imposed uniformly: a quadratic specification serves as the baseline for 1992–2007 and for 1992–2019, whereas the 2010–2019 estimates rest on a cubic specification whose first-derivative roots are real and lie inside the observed data range. Because

the quadratic baselines imply a potential inverted U-shape, they are subsequently subjected to the formal test proposed by Lind and Mehlum (2010). Table 7 reports the baseline fixed-effects estimates with Driscoll–Kraay standard errors for all three periods.

Table 7. Fixed-Effects Estimates with Driscoll–Kraay Standard Errors: GDP per Capita Growth, Baseline Models, 1992–2007, 2010–2019 and 1992–2019

Variables	Model (1) 1992-2007 Quadratic	Model (2) 2010-2019 Cubic	Model (3) 1992-2019 Quadratic
d_ln_PC	-0.222 (0.339)	-8.513** (3.264)	-2.153 (1.442)
$d_ln_PC^2$	-4.724*** (1.158)	-17.199*** (4.427)	-3.897*** (1.037)
$d_ln_PC^3$	—	-6.317*** (1.520)	—
Year Fixed Effects	Yes	Yes	Yes
Country Fixed Effects	Yes	Yes	Yes
Observations	2,600	1,749	4,700
Number of Countries	179	183	186
Within R^2	0.059	0.122	0.066
Driscoll-Kraay Lags	3	3	3

Notes: Dependent variable: GDPPG. Fixed-effects estimates with Driscoll–Kraay robust standard errors (maximum lag = 3) in parentheses. Models (1) and (3) are quadratic and Model (2) is cubic. Country and year fixed effects are included, and year coefficients are omitted. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

The estimates in Table 7 reveal a consistently concave relationship between credit growth and per capita growth. In the 1992–2007 model, the linear term is small and statistically insignificant at −0.222 with a standard error of 0.339, while the squared term is negative and significant at the 1% level at −4.724, indicating that the marginal association weakens as credit expansion accelerates even though no average linear effect can be detected. The full-period model displays the same configuration, combining a linear coefficient of −2.153 that is not statistically significant with a squared term of −3.897 significant at the 1% level, estimated

across 4,700 observations and 186 countries. The post-crisis model differs in functional form but not in direction, since all three polynomial terms are negative and statistically significant, at −8.513 for the linear term, −17.199 for the squared term and −6.317 for the cubic term.

Two features of these baseline results deserve emphasis before the shape of the relationship is tested formally. The negative and highly significant squared term recurs in every period, suggesting that the concavity of the relationship is not an artefact of a particular window but a stable feature of the aggregate data. The within R-squared values nevertheless remain modest, ranging from 0.059 in the pre-crisis model to 0.122 in the post-crisis model, implying that aggregate credit dynamics account for a limited share of the within-country variation in per capita growth. This modest explanatory power is itself informative, indicating that a single economy-wide credit measure captures only a fraction of the mechanisms through which finance may influence real activity.

3.3.2.1. Formal Testing of the Inverted U-Shape

The sign pattern of the estimated coefficients cannot by itself establish that the relationship is genuinely inverted U-shaped. A negative squared term guarantees concavity but not that the implied maximum falls inside the observed data, so inference drawn from coefficient signs alone risks attributing a turning point to a relationship that is monotonic over the relevant range (Lind & Mehlum, 2010). The appropriate procedure, developed from the intersection-union logic of Sasabuchi (1980), requires the estimated slope to be positive at the lower bound of the data and negative at the upper bound, with the turning point located inside the observed interval. Since *d_ln_PC* is a logarithmic difference, an estimated turning point of x translates into a proportional change of exp(x) − 1 in the private credit-to-GDP ratio, which is the basis on which the percentage magnitudes reported below are obtained.

Tables 8 and 9 report the test for the quadratic baselines covering 1992–2007 and 1992–2019 respectively.

Table 8. Lind and Mehlum (2010) Inverted U-Shape Test: GDP per Capita Growth, 1992–2007

Specification: $f(x) = x^2$ Estimated Turning Point: -0.023

	Lower Bound	Upper Bound
Interval	-1.858	1.367
Slope	17.333	-13.139
t-statistic	4.046	-4.081
p-value	<0.001***	<0.001***

Overall inverse U-shape test: t = 4.046; p < 0.001***. 80% Fieller interval for the turning point: [-0.075, 0.026].

Notes: Lind and Mehlum (2010) test for an inverted U-shaped relationship between d_ln_PC and GDPPG in the quadratic fixed-effects specification with Driscoll–Kraay standard errors, 1992–2007. H_0 is a monotonic or U-shaped relationship and H_1 an inverted U-shape. The turning point, computed as $-\beta_1/(2\beta_2)$, lies within the observed interval [-1.858, 1.367], where the slope is positive and significant at the lower bound and negative and significant at the upper bound. The overall p-value of 0.000528 rejects H_0 at the 1% level. The Fieller confidence interval is constructed at the 80% level. Stars are based on unrounded p-values: *** p < 0.01, ** p < 0.05, * p < 0.10.

Table 9. Lind and Mehlum (2010) Inverted U-Shape Test: GDP per Capita Growth, 1992–2019

Specification: $f(x) = x^2$ Estimated Turning Point: -0.276

	Lower Bound	Upper Bound
Interval	-2.249	1.367
Slope	15.373	-12.809
t-statistic	2.984	-4.395
p-value	0.003***	<0.001***

Overall inverse U-shape test: t = 2.984; p = 0.003***. 80% Fieller interval for the turning point: [-0.637, -0.031].

Notes: Lind and Mehlum (2010) test for an inverted U-shaped relationship between d_ln_PC and GDPPG in the quadratic fixed-effects specification with Driscoll–Kraay standard errors, 1992–2019. H_0 is a monotonic or U-shaped relationship and H_1 an inverted U-shape. The turning point, computed as $-\beta_1/(2\beta_2)$, lies within the observed interval [-2.249, 1.367], where the slope is positive and significant at the lower bound and negative and significant at the upper bound. The overall p-value of 0.00299 rejects H_0 at the 1% level. The Fieller confidence interval is constructed at the 80% level. and excludes zero. Stars are based on unrounded p-values: *** p < 0.01, ** p < 0.05, * p < 0.10.

For the pre-crisis panel, the estimated turning point lies at a *d_ln_PC* value of −0.023, comfortably inside the observed interval running from −1.858 to 1.367. The slope is positive and significant at the lower bound at 17.333 with a t-statistic of 4.046, and negative and significant at the upper bound at −13.139 with a t-statistic of −4.081, so the overall statistic of 4.046 and its p-value of 0.0005 reject the null of a monotone or U-shaped relationship at the 1% level. Applying the conversion set out above, a turning point of −0.023 corresponds to a contraction of roughly 2.3% in the private credit-to-GDP ratio, implying that growth is maximised where credit remains broadly stable relative to output. The 80% Fieller interval, constructed to accommodate the uncertainty attaching to a ratio of estimated coefficients (Fieller, 1954), extends from −0.075 to 0.026 and therefore contains zero, indicating that the growth-maximising rate of credit expansion cannot be statistically distinguished from no expansion at all.

The full-period test conveys a stronger message. The turning point is estimated at −0.276, again inside an observed interval spanning −2.249 to 1.367, with a positive and significant slope of 15.373 at the lower bound and a negative and significant slope of −12.809 at the upper bound. The overall statistic of 2.984 with a p-value of 0.003 rejects the null at the 1% level, and the 80% Fieller interval, running from −0.637 to −0.031, excludes zero. Since a value of −0.276 in logarithms corresponds to a contraction of approximately 24% in the credit-to-GDP ratio, the estimated maximum lies well inside contraction territory, implying that the marginal association between credit expansion and per capita growth is negative across virtually the entire range of observed credit growth.

This finding warrants careful reading rather than the conclusion that shrinking financial systems promote growth. Since the credit ratio carries output in its denominator, the position of the estimated slope may partly reflect simultaneity between the two series, a possibility confronted directly in Section 3.3.2.4. The curvature documented by the test is not exposed to the same objection, however, because a denominator relationship of this kind loads growth onto the credit measure in approximately linear fashion and would therefore displace the estimated slope without readily generating a statistically significant quadratic term or a turning point

located inside the observed range. The observed interval nevertheless remains extremely wide, spanning logarithmic changes from −2.249 to 1.367 – equivalent to a contraction of about 89% and an expansion of roughly 292% – so the estimated curvature is identified partly from crisis collapses and from inflation-driven jumps in nominal credit. What the test establishes robustly is the concavity of the relationship and the location of its maximum at or below zero, both of which are consistent with the argument that the growth contribution of credit weakens and eventually reverses once financial expansion outpaces the real economy (Arcand et al., 2015; Cecchetti & Kharroubi, 2012).

3.3.2.2. The Post-Crisis Cubic Specification

The 2010–2019 panel was estimated in cubic form because, unlike the quadratic models, the first-derivative equation of the estimated function possesses real roots that fall inside the observed data range. Differentiating the cubic specification yields a quadratic in *d_ln_PC* whose reduced discriminant is given by $\beta_2^2 - 3\beta_1\beta_3$, and evaluating this expression at the coefficients reported in Table 7 returns approximately 134.5, a positive value confirming two distinct real roots. Those roots, obtained as $(-\beta_2 \pm \sqrt{(\beta_2^2 - 3\beta_1\beta_3)})$ divided by $3\beta_3$, place the turning points at *d_ln_PC* values of about −1.5195 and −0.2956, corresponding to contractions of roughly 78% and 26% in the private credit-to-GDP ratio. Because the cubic coefficient is negative, the estimated function falls below the lower turning point, rises between the two turning points and falls again beyond the upper one, so that the local minimum occurs at −1.5195 and the local maximum at −0.2956.

The economic reading of this configuration is broadly consistent with the quadratic evidence. The local maximum at −0.2956 lies remarkably close to the full-period turning point of −0.276, suggesting that the growth-maximising pace of credit expansion remained largely unchanged after the crisis and continued to sit slightly below zero, so that the additional flexibility of the cubic form alters the shape of the relationship without displacing its peak. The upward segment located at very deep contractions is more plausibly interpreted as capturing post-collapse rebound dynamics, in which output recovers while credit remains depressed, than as evidence

that severe deleveraging is itself growth-enhancing. Given that these estimates rest on a ten-year window, the shape identified for the post-crisis decade should be regarded as indicative rather than conclusive, an assessment reinforced by the sensitivity of the higher-order terms to the inclusion of controls documented below.

3.3.2.3. Estimates with Macroeconomic Controls

The baseline specifications isolate the credit variable but omit the determinants of growth discussed in Section 3.2.3, raising the possibility that the estimated polynomial absorbs variation properly attributable to investment, fiscal policy, price stability, human capital or external orientation. Table 10 therefore augments each period's baseline model with gross capital formation, government consumption, inflation, secondary school enrollment and trade openness, retaining the quadratic form for 1992–2007 and 1992–2019 and the cubic form for 2010–2019.

Table 10. Fixed-Effects Estimates with Driscoll–Kraay Standard Errors: GDP per Capita Growth, Models with Control Variables, 1992–2007, 2010–2019 and 1992–2019

Variables	Model (1) 1992-2007 Quadratic + Controls	Model (2) 2010-2019 Cubic + Controls	Model (3) 1992-2019 Quadratic + Controls
d_ln_PC	-1.788** (0.663)	-7.389*** (1.464)	-1.613** (0.738)
d_ln_PC²	-2.909** (1.195)	9.673* (4.478)	-1.575** (0.760)
d_ln_PC³	—	5.843** (1.986)	—
d_ln_SE	2.687 (2.293)	-0.319 (0.965)	2.212 (1.506)
ln_GCF	1.858*** (0.578)	0.591 (0.475)	2.507*** (0.370)
ln_GC	-2.661 (1.714)	-7.178*** (0.679)	-2.423** (0.899)

Table 10. (continued)

INF	-0.004*** (0.001)	-0.231*** (0.031)	-0.004*** (0.000)
ln_TO	2.560* (1.261)	5.615*** (0.532)	2.084*** (0.634)
Year Fixed Effects	Yes	Yes	Yes
Country Fixed Effects	Yes	Yes	Yes
Observations	988	1,062	2,257
Number of Countries	133	144	156
Within R^2	0.155	0.206	0.198
Driscoll-Kraay Lags	3	3	3

Notes: Dependent variable: GDPPG. Fixed-effects estimates with Driscoll–Kraay robust standard errors (maximum lag = 3) in parentheses. Models (1) and (3) are quadratic and Model (2) is cubic, and all three include the control variables. Country and year fixed effects are included, and year coefficients are omitted. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Adding controls sharpens rather than dissolves the credit result in the two quadratic models. The linear term now attains statistical significance in both, at −1.788 for 1992–2007 and −1.613 for 1992–2019, while the squared terms remain negative and significant at −2.909 and −1.575 respectively. Computing the implied maxima from these coefficients places them at *d_ln_PC* values of approximately −0.307 and −0.512, further into contraction territory than in the baseline models and implying a negative marginal association across the entire range of positive credit growth. Evaluating the full-period estimates at a *d_ln_PC* of 0.10 makes the magnitude concrete, indicating that an additional ten logarithmic points of credit expansion are associated with per capita growth lower by roughly 0.19 percentage points, a modest but economically meaningful difference when sustained across successive years.

The post-crisis cubic behaves very differently once controls enter, and this instability is itself an important result. The linear term strengthens to −7.389 and gains

significance at the 1% level, but the squared and cubic terms reverse sign, becoming positive at 9.673 and 5.843 respectively, which moves the turning points to approximately −1.404 and 0.300 and implies an upward-sloping segment beyond a credit expansion of about 35%. Because the estimation sample simultaneously falls from 1,749 to 1,062 observations and from 183 to 144 countries, part of this reversal may reflect a change in sample composition rather than the influence of the controls themselves. The higher-order structure of the post-crisis relationship should therefore not carry independent interpretive weight, whereas the negative and highly significant linear term, common to both specifications, remains the more reliable feature.

The control variables behave largely as the theoretical discussion in Section 3.2.3 anticipates. Gross capital formation carries a positive coefficient significant at the 1% level in the pre-crisis and full-period models, at 1.858 and 2.507 respectively, confirming that physical investment remains a primary correlate of per capita growth. Its coefficient falls to 0.591 and loses significance in the post-crisis decade, suggesting that the link between measured investment and growth may have weakened after 2010 even as credit conditions eased, although the shorter estimation window counsels caution. Inflation is negative and significant at the 1% level in every period, government consumption is negative throughout and strongly significant after the crisis at −7.178, and trade openness is positive in all three models, rising from 2.560 before the crisis to 5.615 afterwards and indicating that external integration became a more prominent correlate of growth in the later period.

Secondary school enrollment, entering in differenced logarithmic form, remains statistically insignificant in all three models, which is unsurprising given that human capital accumulation operates over horizons far longer than the annual frequency of the panel. Taken together, the control estimates raise the within R-squared to between 0.155 and 0.206 and, more importantly, establish a contrast that shapes the remainder of this section. The real-side variables retain conventional signs and stable significance across periods, whereas the higher-order structure of the credit variable shifts with the specification and the sample, suggesting that the fragility identified here is specific to the aggregate credit measure rather than a general weakness of the model.

3.3.2.4. Addressing Endogeneity: Two-Step System GMM Estimates

Neither fixed effects nor Driscoll–Kraay standard errors address the simultaneity between credit and output, and the dynamic character of growth introduces a further complication, since including a lagged dependent variable in a fixed-effects model generates the bias identified by Nickell (1981). The analysis therefore re-estimates the relationship using the two-step System GMM estimator, which combines the differenced and level equations in order to exploit lagged values as internal instruments (Arellano & Bover, 1995; Blundell & Bond, 1998), applies the Windmeijer (2005) finite-sample correction to the standard errors and collapses the instrument matrix to contain instrument proliferation (Roodman, 2009b). Estimations were carried out for all three periods, but the 1992–2007 and 2010–2019 specifications produced no significant relationships and are consequently reported in Appendix. Since instrumenting the higher-order polynomial terms proved unsatisfactory, the linear specification is adopted as the baseline in the System GMM setting, with controls added in the second model, and the full-period results appear in Table 11.

Table 11. Two-Step System GMM Estimates: GDP per Capita Growth, 1992–2019

Variables	Model (1) 1992-2019 Baseline	Model (2) 1992-2019 Controls
L.GDPPG	0.147 (0.254)	0.321*** (0.079)
d_ln_PC	12.574* (6.518)	1.291 (3.271)
d_ln_SE	—	1.056 (1.910)
ln_GCF	—	1.567*** (0.532)
ln_GC	—	-1.563*** (0.309)
INF	—	-0.0010 (0.0010)

Table 11. (continued)

ln_TO	—	0.483* (0.254)
Year Effects	Yes	Yes
Observations	4,698	2,257
Number of Countries	186	156
Number of Instruments	35	40
GMM Lag Range	2-3	2-3
Collapsed Instruments	Yes	Yes
AR(1) Test p-value	0.002	<0.001
AR(2) Test p-value	0.741	0.449
Sargan Test p-value	<0.001	0.124
Hansen Test p-value	0.125	0.136
Difference-in-Hansen p-value	0.028	0.039

Notes: Dependent variable: GDPPG. Two-step System GMM estimates with Windmeijer (2005) finite-sample-corrected robust standard errors in parentheses. Model (2) adds d_ln_SE, ln_GCF, ln_GC, INF and ln_TO. Year fixed effects are included, and year coefficients and constants are omitted. L.GDPPG and d_ln_PC enter as collapsed GMM-style instruments at lags 2–3, and the controls and year indicators as standard instruments. Diagnostic entries are p-values, and Difference-in-Hansen refers to the exogeneity test for the level-equation instruments. Stata used a generalized inverse for a singular two-step moment covariance matrix, so Difference-in-Hansen results require caution. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

The baseline System GMM model returns a positive coefficient on credit growth of 12.574 that is significant only at the 10% level, together with an insignificant coefficient on lagged growth. Its diagnostic profile is nevertheless uncomfortable, since the Arellano and Bond (1991) tests for first-order and second-order serial correlation behave as required, at 0.002 and 0.741, while the Sargan statistic rejects instrument validity decisively and the Difference-in-Hansen test rejects the exogeneity of the level-equation instruments at the 5% level. Given that the Sargan

test assumes homoskedastic errors whereas the Hansen test loses power as the instrument count rises, the divergence between the two statistics is precisely the pattern that warrants scepticism towards an over-identified specification (Roodman, 2009a, 2009b).

The specification with controls is both better identified and considerably more informative. Lagged growth now enters at 0.321 with significance at the 1% level, confirming meaningful persistence in per capita growth, and the coefficient on credit growth falls to 1.291 with a standard error of 3.271, becoming statistically indistinguishable from zero. Gross capital formation remains positive and significant at 1.567, government consumption negative and significant at −1.563 and trade openness positive at 0.483, while inflation and secondary school enrollment are not statistically significant. The Sargan and Hansen tests no longer reject at conventional levels, at 0.124 and 0.136 respectively, and second-order serial correlation is absent at 0.449, although the Difference-in-Hansen p-value of 0.039 continues to question the exogeneity of the additional level-equation moment conditions.

The contrast between Tables 10 and 11 constitutes the central lesson of the aggregate exercise. A relationship that appears robustly negative and clearly concave under fixed effects becomes statistically indistinguishable from zero once persistence and endogeneity are taken into account and the real-side determinants of growth are held constant. The corresponding subperiod estimations yielded no interpretable relationship and are reported in Appendix for completeness.

3.3.2.5. Limitations of the Aggregate Specification and the Case for Sectoral Disaggregation

The aggregate results establish a coherent but ultimately limited picture, showing a concave and predominantly negative conditional association that does not survive correction for endogeneity, and they raise the question of why an economy-wide credit measure should behave so unstably. A substantial part of the answer lies in what the measure conceals. Private credit by deposit money banks aggregates working capital for manufacturers, mortgages, consumer lending and seasonal

advances to farmers into a single ratio, treating claims with entirely different maturities, risk profiles and productive consequences as economically equivalent. Evidence that household and enterprise lending carry markedly different growth implications indicates that the composition of credit, and not merely its quantity, governs its macroeconomic consequences (Beck et al., 2012), which means that an aggregate coefficient represents a weighted average of heterogeneous underlying relationships whose weights shift with the credit cycle.

This weighting problem carries a direct econometric implication. Where sectoral effects differ in sign, the aggregate coefficient may approach zero not because finance is irrelevant but because opposing effects offset one another, and a null or unstable aggregate estimate is precisely what such cancellation would produce. Sectors differ systematically along the dimensions that govern their response to credit, including capital intensity, dependence on external finance, tradability, the length of the production cycle and the availability of collateral, and the classic finding that industries relying more heavily on external finance grow disproportionately faster in financially developed economies demonstrates that these differences are empirically consequential rather than merely conceptual (Rajan & Zingales, 1998). An aggregate specification cannot recover this structure, constraining agriculture, industry and services to share a single coefficient and a single functional form.

The limitation extends to the dependent variable as well. Per capita growth summarises the pace of expansion but reveals nothing about its composition, and the reallocation of labour and capital from primary to secondary and subsequently to tertiary activities is precisely the process through which development proceeds (Herrendorf et al., 2014; McMillan & Rodrik, 2011). An aggregate regression therefore averages across the structural transformation this dissertation seeks to examine, and it cannot indicate whether a given credit expansion supported industrial capacity, financed the expansion of services or bypassed agriculture altogether. Evidence that the finance–growth relationship varies with the stage of development compounds the difficulty, since a pooled aggregate estimate also suppresses the variation across income groups that the sample design in Section 3.1 was constructed to capture (Rioja & Valev, 2004b).

These considerations bear directly on the question motivating this study. The 'too much finance' literature is frequently framed as a debate about volume, yet the mechanisms it proposes – the diversion of credit towards asset markets, the deterioration of screening standards during expansions and the misallocation of capital and talent – are ultimately claims about allocation (Arcand et al., 2015; Beck et al., 2014; Cecchetti & Kharroubi, 2012). Testing such claims with an aggregate ratio conflates 'too much finance' with finance directed towards the wrong activities, because both generate the same aggregate signature of weakening returns to credit. Distinguishing between them requires matching credit to the activity it finances, which is the design adopted in this dissertation and set out in Sections 3.2.1 and 3.2.2.

The sectoral framework addresses each of these limitations directly. Agricultural, industrial and service value added enter as separate dependent variables, allowing the response of each sector to be estimated in its own right rather than averaged away, while the financial development indicator is matched to the sector under analysis wherever the data permit, with the differenced logarithm of credit to agriculture, forestry and fishing relative to GDP replacing the broad private-credit measure in the agricultural models. The complementary analysis of eleven matched credit–output sectors for Türkiye extends this logic further, pairing each sectoral credit series with the corresponding output series within a single institutional setting and thereby narrowing the interpretive gap that an aggregate cross-country ratio necessarily leaves open. The aggregate estimates reported in this section are best read, accordingly, as a benchmark and a motivation rather than as a finding in themselves, since they demonstrate that the relationship between credit and growth is non-linear, that its aggregate form is fragile and weakly identified, and that answering the question posed by this dissertation requires asking not how much finance an economy has but where that finance goes.

CHAPTER 4

CROSS-COUNTRY EMPIRICAL METHODOLOGY

4.1. The Panel Data Framework

Panel data analysis, which combines the cross-sectional and the time-series dimensions of observation within a single estimation framework, has become the standard instrument of comparative macroeconomic research. Cross-sectional data record the differences that separate countries, firms or individuals at one point in time, whereas time-series data trace the evolution of a single unit across successive periods. Panel structures merge these two sources of variation by following the same units over repeated intervals, thereby producing a more granular account of economic behaviour than either dimension can deliver alone (Hsiao, 2014; Wooldridge, 2010).

The principal advantage of the design lies in its capacity to attenuate omitted-variable bias, a difficulty that affects purely cross-sectional and purely temporal analyses alike. Because repeated observation of the same units permits the control of unobserved, time-invariant characteristics – geography, legal origin, institutional tradition or long-established production practices – the estimated slope coefficients are purged of confounding influences that are difficult to measure directly. Parameter estimates consequently become more consistent and more defensible as a basis for inference and for policy evaluation (Baltagi, 2021; Greene, 2018).

These advantages are conditional rather than automatic, since they hold only insofar as the underlying econometric assumptions are satisfied. Classical panel estimators rest on specific conditions governing the disturbance term and its relationship with the regressors, and, when those conditions hold, the estimators are unbiased,

consistent and efficient. Departures such as heteroskedasticity, serial correlation or cross-sectional dependence need not destroy consistency, yet they distort the estimated covariance matrix, producing standard errors that misstate sampling variability and inferences that may therefore be invalid. A systematic assessment of the structural properties of the data is thus a precondition of estimation rather than a formality, indicating that the specification should follow from diagnostic evidence rather than from convention (Arellano, 2003; Cameron & Trivedi, 2009; Hsiao, 2014).

The diagnostic sequence adopted in this chapter follows that logic and is deliberately ordered. Cross-sectional dependence is examined first, in Section 4.2, because the presence or absence of dependence across units determines which generation of panel unit-root test is admissible at the subsequent stage. The panel unit-root analysis and the associated decisions on deterministic components and lag length follow in Section 4.3, since the integration orders they establish determine the form in which the variables enter the estimated equations. Slope homogeneity, groupwise heteroskedasticity and first-order serial correlation are then assessed in Sections 4.4 to 4.6, each on the specification actually estimated. The estimation strategy that these diagnostics jointly imply – a fixed-effects estimator with Driscoll–Kraay standard errors, complemented by a two-step System GMM estimator – is developed in Sections 4.7 and 4.8.

4.2. Cross-Sectional Dependence Testing

'Cross-sectional dependence' (CSD), understood as correlation between the disturbances of different panel units, is a routine feature of cross-country macroeconomic data rather than an exceptional one. Unobserved common factors, global business cycles, trade and financial integration and spatial spillovers all transmit shocks across national borders, so the residuals of a country panel are unlikely to be mutually independent. Neglecting such dependence leaves the fixed-effects coefficients consistent under strict exogeneity but invalidates the conventional covariance estimator, thereby corrupting every test statistic derived from it (Pesaran, 2006; De Hoyos & Sarafidis, 2006).

The diagnostics were computed on regression residuals rather than on the raw series. Residuals from the within estimator are preferred at this stage because the within transformation eliminates time-invariant individual effects, leaving an error component that reflects contemporaneous correlation across units more faithfully than the unfiltered data (Pesaran, 2004; De Hoyos & Sarafidis, 2006). The tests were accordingly applied to the residuals of the agricultural value-added model (ln_AVASGDP), the industrial value-added model (ln_IVASGDP) and the services value-added model (ln_SVASGDP). These estimations rest on 1,502 observations from 95 countries for ln_AVASGDP, 2,765 observations from 158 countries for ln_IVASGDP and 2,737 observations from 157 countries for ln_SVASGDP, and all three specifications are jointly significant at the 1 per cent level, with F-statistics of 34.45, 51.56 and 68.78 respectively.

One point of sequencing deserves emphasis, since it governs the diagnostics reported throughout this chapter. Cross-sectional dependence is assessed on the level specification, in which every variable enters in the form examined at the unit-root stage, because the outcome of this test determines which class of panel unit-root test is admissible and the classification cannot therefore be conditioned on transformations that the unit-root results have yet to justify. The remaining diagnostics, reported in Sections 4.4 to 4.6, are computed on the specification actually estimated, in which the variables classified as I(1) enter in first differences, since those tests examine the residuals of the estimated equation and would be uninformative if computed on a model that is not the one taken to the data.

Four complementary statistics were computed: the Pesaran CD statistic (Pesaran, 2004, 2021), the weighted CD statistic (CDw) of Juodis and Reese (2022), the power-enhanced CDw^{+} statistic, which applies the screening-based power enhancement of Fan et al. (2015) to the weighted statistic (Juodis & Reese, 2022), and the bias-corrected CD* statistic of Pesaran and Xie (2026). Reliance on a single statistic would be unwise here, since the original CD test is known to over-reject when applied to residuals from models containing many estimated parameters, while the corrections proposed in response to that problem differ in their power against spatial and network alternatives (Juodis & Reese, 2022; Pesaran & Xie, 2026). The

nulls of the four statistics are related but not identical, and conflating them would misstate what each rejection establishes. The CD statistic and its weighted counterpart are constructed under a null of cross-sectional independence in the residuals, although Pesaran (2015) demonstrates that the CD test in fact has power only against sufficiently strong dependence, so that its effective null is better characterised as weak dependence. The bias-corrected CD* statistic is likewise derived under a null of cross-sectional independence in the idiosyncratic errors once the latent common factors have been filtered out, meaning that it targets residual network dependence rather than factor dependence as such (Pesaran & Xie, 2026). Table 12 reports the four statistics alongside the summary statistics of the underlying fixed-effects estimations.

Table 12. Cross-Sectional Dependence Test Results: Agriculture, Industry and Services, 1992–2019

Dependent Variable	Obs./ Groups	Within R^2	F-statistic	CD	CDw	CDw^+	CD*
ln_AVASGDP	1,502 / 95	0.1286	34.45*** (0.000)	7.99*** (0.000)	1.01 (0.313)	2,100.05*** (0.000)	11.29*** (0.000)
ln_IVASGDP	2,765 / 158	0.1063	51.56*** (0.000)	29.31*** (0.000)	0.87 (0.386)	8,332.45*** (0.000)	49.32*** (0.000)
ln_SVASGDP	2,737 / 157	0.1382	68.78*** (0.000)	18.31*** (0.000)	0.97 (0.330)	7,419.33*** (0.000)	13.88*** (0.000)

Notes: The table reports summary statistics for the sectoral fixed-effects (within) estimations together with residual-based cross-sectional dependence diagnostics. Obs. is the number of observations and Groups the number of cross-sectional units, and values in parentheses are p-values. CD and CDw test a null of cross-sectional independence in the residuals, which Pesaran (2015) characterises as an effective null of weak dependence, whereas CD* tests residual dependence once the latent common factors have been filtered out using four principal components. All statistics are adjusted for the unbalanced panel. *** $p < 0.01$.

Source. Author's calculations based on the study dataset.

The results point consistently towards strong dependence in all three sectoral models. For ln_AVASGDP the CD statistic is 7.99 and the bias-corrected CD* statistic is 11.29, the corresponding values being 29.31 and 49.32 for industry and 18.31 and

13.88 for services, all significant at the 1 per cent level. The power-enhanced CDw^{+} statistic reaches 2,100.05, 8,332.45 and 7,419.33 respectively, reinforcing the same conclusion through a test constructed explicitly to raise power in high-dimensional settings (Fan et al., 2015). Since these statistics are computed on panels differing in both the number of countries and the length of the observation window, their relative magnitudes should not be read as a ranking of the strength of dependence across sectors, the substantive finding being the uniform rejection of the null rather than the ordering of the rejections.

The weighted CDw statistic is the single exception, remaining insignificant in all three models with values of 1.01, 0.87 and 0.97 and p-values between 0.31 and 0.39. This divergence is not evidence of a contradiction, since the weighted statistic was designed to restore the nominal size of the CD test in panels containing many estimated parameters rather than to maximise power against every form of dependence, and it may accordingly fail to reject where the bias-corrected CD* statistic rejects decisively (Juodis & Reese, 2022; Pesaran & Xie, 2026). Read jointly, and giving weight to the statistic that is both size-correct and powerful against network alternatives, the evidence supports the presence of strong cross-sectional dependence.

The diagnostics establish that dependence exists without identifying what generates it, and the literature attributes such correlation to shared latent factors, global macroeconomic shocks, financial integration, technological diffusion or policy spillovers (Pesaran, 2006; Chudik & Pesaran, 2015). Two consequences follow for the empirical design. The first concerns stationarity testing, because first-generation panel unit-root tests assume independence across units and suffer severe size distortion when that assumption fails, which is why the cross-sectionally augmented procedure adopted in Section 4.3 is a second-generation test. The second concerns inference, since conventional covariance estimators – including cluster-robust variants that permit correlation within countries while assuming independence between them – are inconsistent in this setting, motivating the Driscoll–Kraay standard errors employed in Section 4.7 (Driscoll & Kraay, 1998; Hoechle, 2007).

4.3. Panel Unit-Root Testing, Deterministic Specification and Lag Selection

Before estimation, the time-series properties of every variable were examined in exactly the form in which the variable subsequently enters the empirical models. The sectoral value-added shares and the remaining strictly positive continuous variables were expressed in natural logarithms – ln_AVASGDP, ln_IVASGDP, ln_SVASGDP, ln_ACGDP, ln_PC, ln_SE, ln_GCF, ln_GC and ln_TO – while inflation (INF) was retained in its original annual percentage-rate form, since it takes non-positive values in a non-trivial number of country–year observations. Preserving identical transformations across the testing and the estimation stages ensures that the integration classifications refer to the same objects as the estimated equations, avoiding the inconsistency that arises when tests are conducted on one transformation and models estimated on another.

The tests cover the main estimation period, 1992–2019. Because data availability differs substantially across variables and countries, the testing sample was constructed separately for each series: observations were first restricted to the relevant period, country–year records with missing values for the series under examination were removed, the uniqueness of each country–year pair was verified and the surviving observations were ordered chronologically within each country.

4.3.1. Construction of the Testing Sample

The CIPS procedure requires that lags and differences correspond to genuinely consecutive calendar periods, since the augmented regression is defined in terms of actual temporal adjacency (Pesaran, 2007). Missing years were therefore not concealed by re-indexing observations into an artificial time sequence, a practice that would silently treat a gap of several years as a single-period lag and thereby distort the estimated persistence. Continuous observation blocks were identified separately within each country, a new block beginning whenever the current observation did not immediately follow the preceding calendar year, and the longest such block was retained for each country, the earlier block being selected whenever two blocks were of equal length.

Countries whose longest continuous block contained fewer than 15 consecutive annual observations were excluded from the test for that particular variable. The threshold reflects a compromise between retaining sufficient time-series information for the augmented regressions – which consume degrees of freedom through the lagged level, the cross-sectional averages and up to two lagged differences – and preserving broad cross-country coverage. Since both the set of eligible countries and their observation windows differ across variables, the resulting test panels are unbalanced and variable-specific, and the original calendar-year variable was retained throughout so that the lag and difference operators always span exactly one year.

4.3.2. Selection of Deterministic Components

The deterministic specification was fixed before any unit-root statistic was inspected. This ordering matters because the treatment of the deterministic component affects both the size and the power of the test: omitting a genuine trend may cause a trend-stationary series to be classified as nonstationary, whereas including a superfluous trend reduces power and biases the procedure towards non-rejection (Enders, 2015). Three complementary diagnostics were therefore used, and a trend was retained only where all three pointed in the same direction.

First, the annual change in each variable was computed as

$$\Delta y_{it} = y_{it} - y_{i,t-1} \quad (2)$$

after which an intercept-only pooled regression was estimated,

$$\Delta y_{it} = \mu + u_{it} \quad (3)$$

The null of a zero mean change, H_0: $\mu = 0$, was evaluated using Driscoll–Kraay standard errors, which remain valid under general forms of cross-sectional dependence, heteroskedasticity and serial correlation (Driscoll & Kraay, 1998). The regression was run with bandwidths of one and two years so that the inference did not rest on a single assumption about temporal dependence. This exercise serves as a

diagnostic for systematic panel-wide drift rather than as a formal unit-root or deterministic-trend test, and it was interpreted only in that limited sense.

Second, the mean annual change was computed separately for every country,

$$g_i = (1 / (T_i - 1)) \Sigma_{t=2}^{T_i} \Delta y_{it} \quad (4)$$

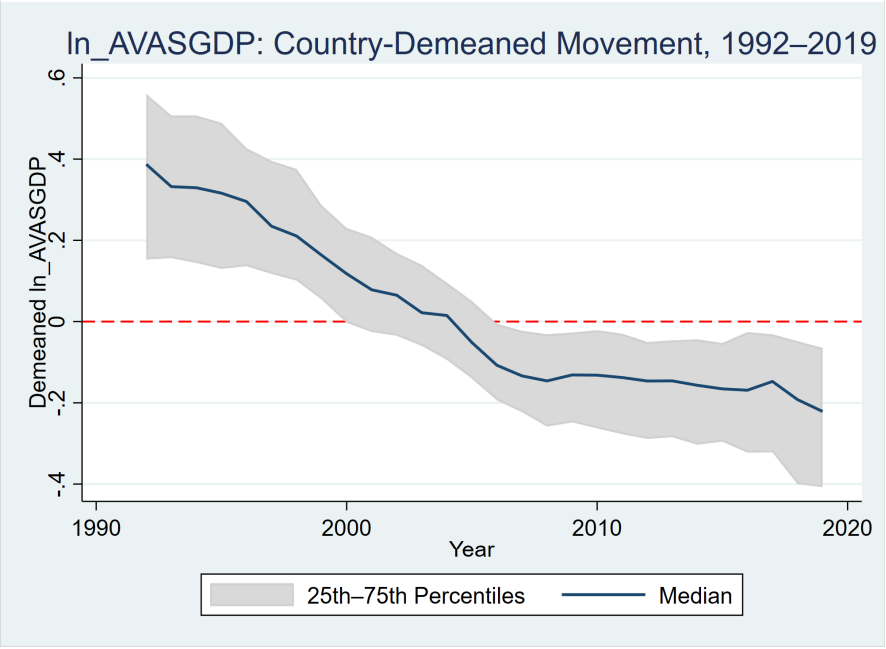
and the distribution of these country-specific averages was examined through its median and through the shares of countries recording positive and negative mean changes. Inspecting the distribution rather than the panel mean alone distinguishes a common directional movement shared by most economies from an average driven by a small number of extreme cases, a distinction that a pooled regression cannot make on its own.

Third, each variable was demeaned within countries,

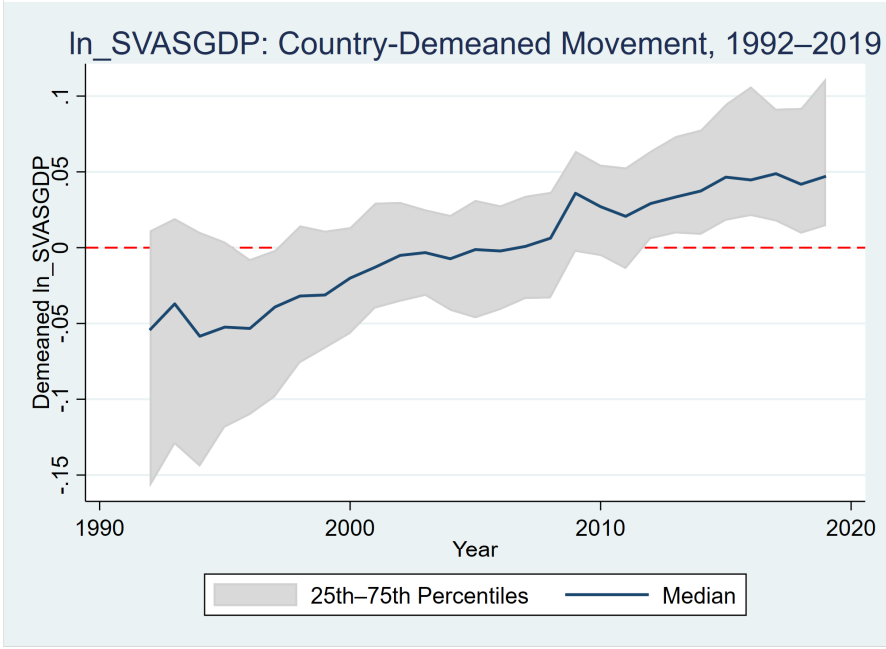
$$\tilde{y}_{it} = y_{it} - \bar{y}_i \quad (5)$$

and the cross-country median and interquartile range – the 25th and 75th percentiles – of the demeaned series were plotted year by year. Country demeaning removes persistent level differences between economies and isolates the common temporal movement of the panel, rendering a shared trend visually identifiable rather than a matter of impression. A linear trend was included when the graphical pattern, the Driscoll–Kraay inference and the distribution of country-specific changes all indicated persistent directional movement, and a constant-only specification was adopted where the evidence did not support systematic drift. Applying these criteria, an intercept and a linear trend were specified for ln_AVASGDP, ln_SVASGDP, ln_PC and ln_SE, whereas an intercept-only specification was adopted for ln_IVASGDP, ln_ACGDP, ln_GCF, ln_GC, ln_TO and INF. Notably, the deterministic specification was determined independently of the CIPS results and was not revised once those results became available, ensuring that the integration classifications reported below are not the product of specification search. The country-demeaned movements underlying these decisions are displayed in Figure 2.

Panel A. ln_AVASGDP — Log of agriculture, forestry, and fishing value added (% of GDP)

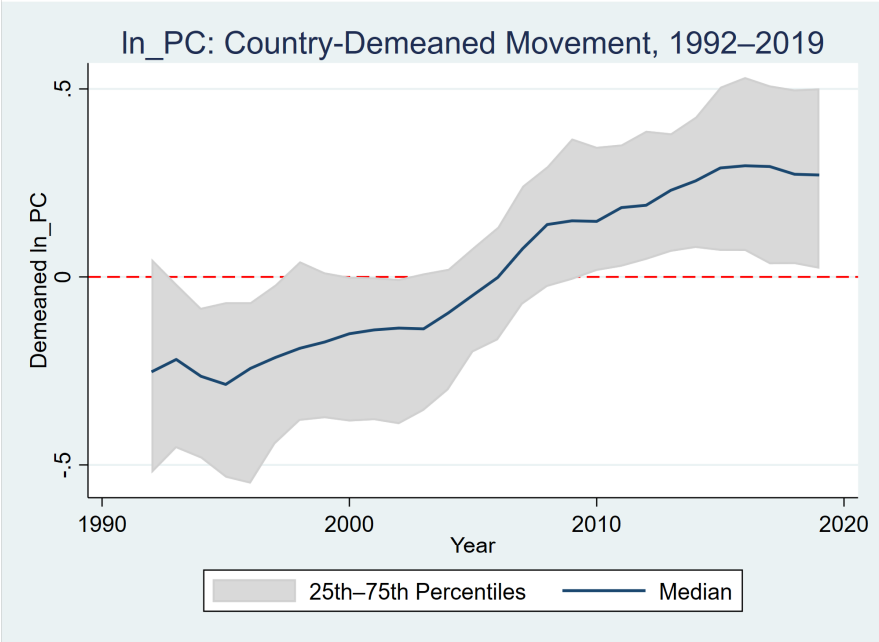


Panel B. ln_SVASGDP — Log of services value added (% of GDP)

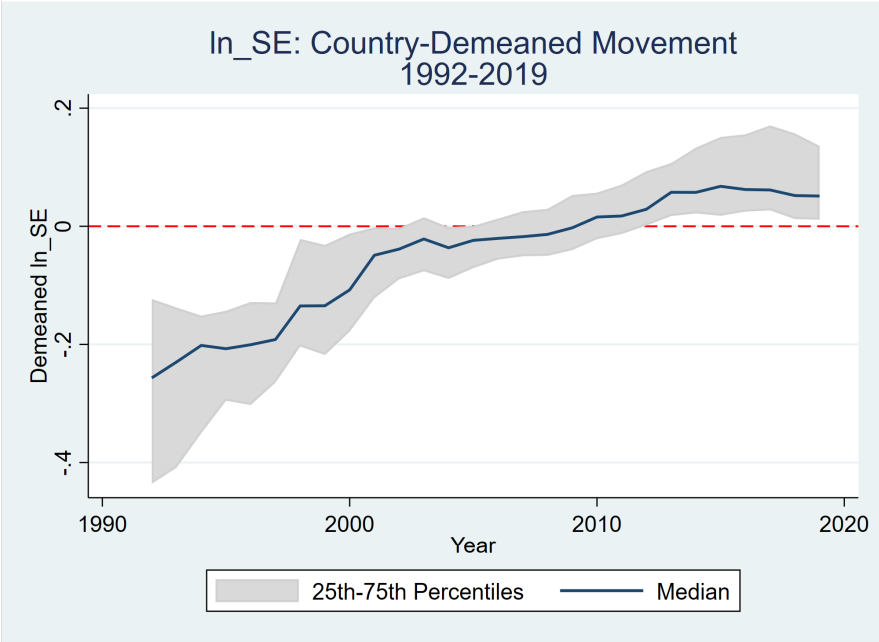


Notes: The navy line denotes the cross-country median, the shaded area represents the interquartile range (25th–75th percentiles), and the red dashed line marks zero. Samples are variable-specific longest continuous annual blocks within 1992–2019, subject to a minimum of 15 consecutive observations per country. Panels A–D correspond to variables assigned a constant-and-trend specification; Panels E–J correspond to variables assigned a constant-only specification. These plots were considered jointly with the Driscoll–Kraay tests of the mean annual change.

Panel C. ln_PC — Log of private credit relative to GDP

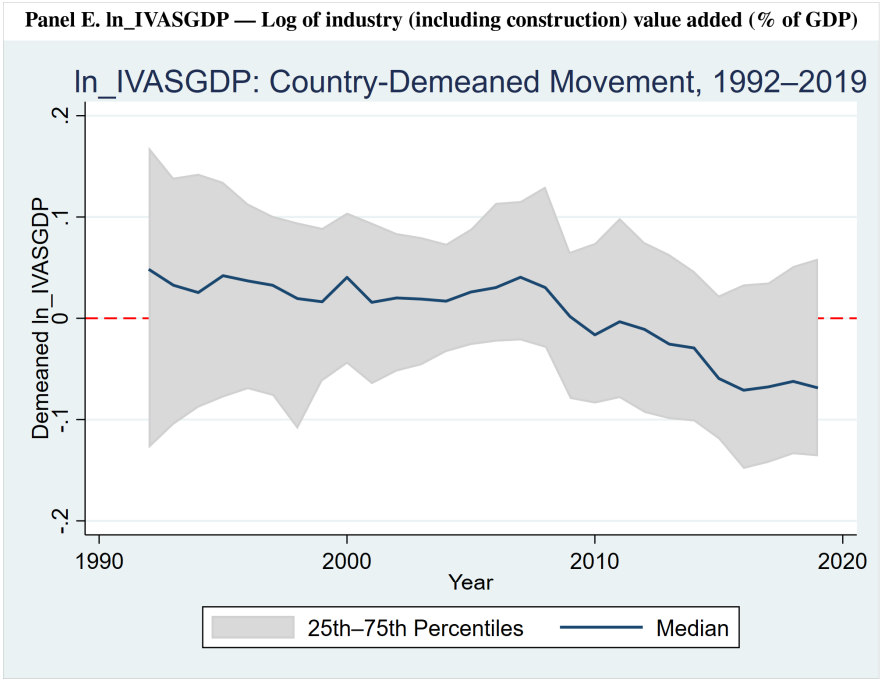


Panel D. ln_SE — Log of secondary school enrolment

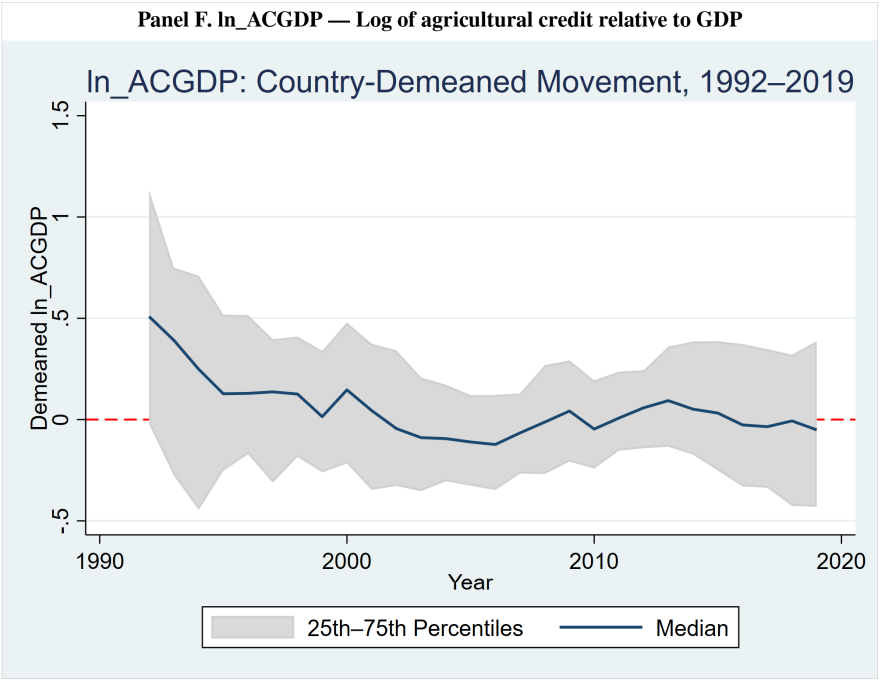


Notes: The navy line denotes the cross-country median, the shaded area represents the interquartile range (25th–75th percentiles), and the red dashed line marks zero. Samples are variable-specific longest continuous annual blocks within 1992–2019, subject to a minimum of 15 consecutive observations per country. Panels A–D correspond to variables assigned a constant-and-trend specification; Panels E–J correspond to variables assigned a constant-only specification. These plots were considered jointly with the Driscoll–Kraay tests of the mean annual change.

Panel E. ln_IVASGDP — Log of industry (including construction) value added (% of GDP)

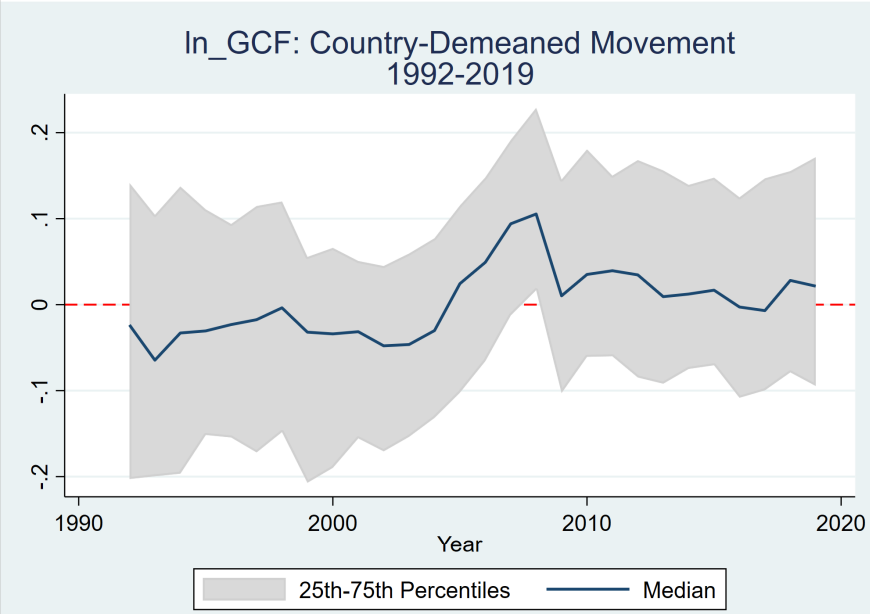


Panel F. ln_ACGDP — Log of agricultural credit relative to GDP

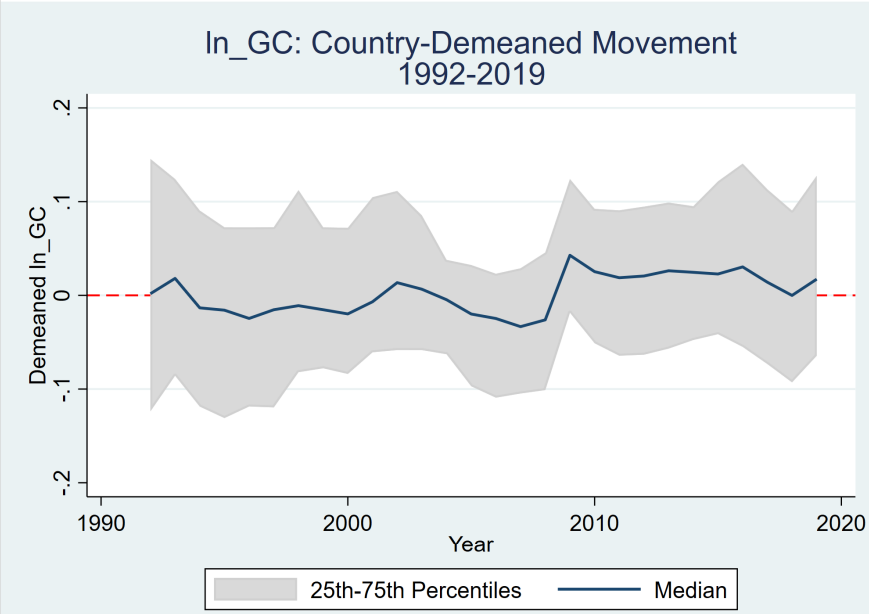


Notes: The navy line denotes the cross-country median, the shaded area represents the interquartile range (25th–75th percentiles), and the red dashed line marks zero. Samples are variable-specific longest continuous annual blocks within 1992–2019, subject to a minimum of 15 consecutive observations per country. Panels A–D correspond to variables assigned a constant-and-trend specification; Panels E–J correspond to variables assigned a constant-only specification. These plots were considered jointly with the Driscoll–Kraay tests of the mean annual change.

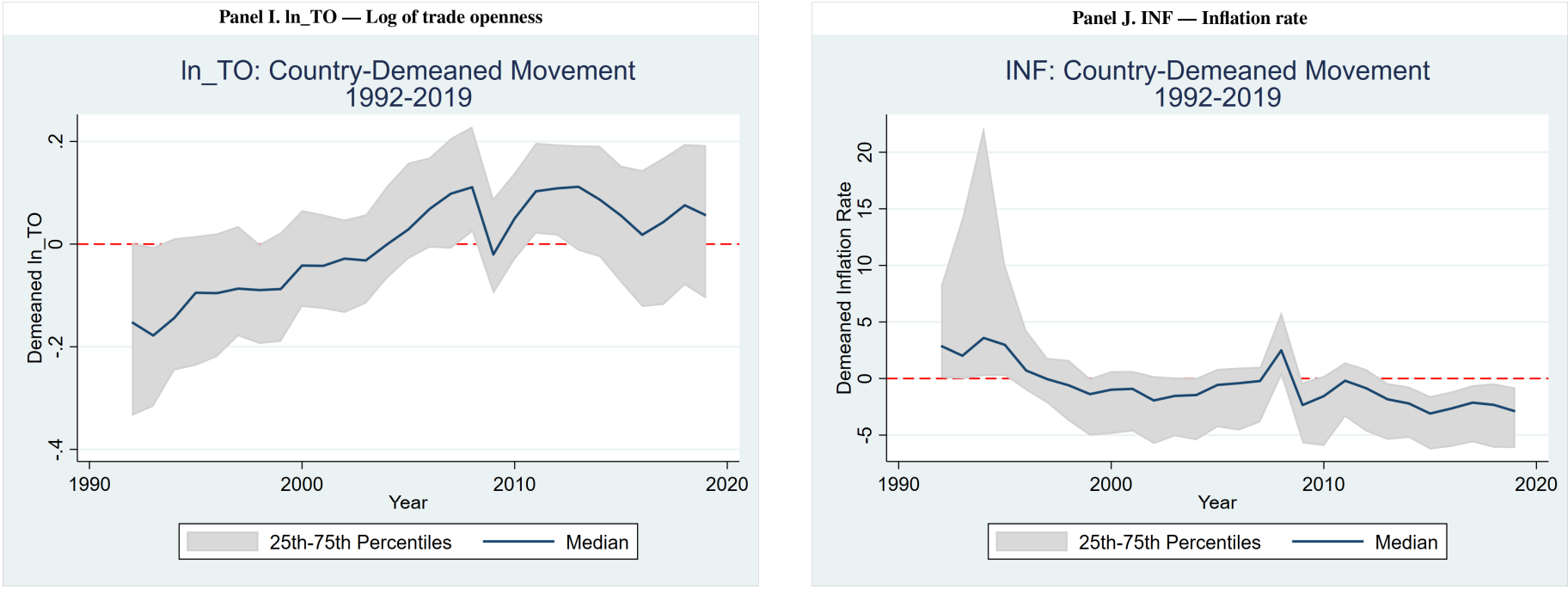
Panel G. ln_GCF — Log of gross capital formation (% of GDP)



Panel H. ln_GC — Log of government final consumption expenditure (% of GDP)



Notes: The navy line denotes the cross-country median, the shaded area represents the interquartile range (25th–75th percentiles), and the red dashed line marks zero. Samples are variable-specific longest continuous annual blocks within 1992–2019, subject to a minimum of 15 consecutive observations per country. Panels A–D correspond to variables assigned a constant-and-trend specification; Panels E–J correspond to variables assigned a constant-only specification. These plots were considered jointly with the Driscoll–Kraay tests of the mean annual change.



Notes: The navy line denotes the cross-country median, the shaded area represents the interquartile range (25th–75th percentiles), and the red dashed line marks zero. Samples are variable-specific longest continuous annual blocks within 1992–2019, subject to a minimum of 15 consecutive observations per country. Panels A–D correspond to variables assigned a constant-and-trend specification; Panels E–J correspond to variables assigned a constant-only specification. These plots were considered jointly with the Driscoll–Kraay tests of the mean annual change.

Figure 2. Country-Demeaned Movements Used in Selecting Deterministic Specifications, 1992–2019

4.3.3. Country-Specific Lag Selection Using the Bayesian Information Criterion

The lag length of the augmented unit-root regression was allowed to vary across countries. Since the persistence of macroeconomic series differs markedly between economies, imposing a common lag order would leave residual autocorrelation in some countries while consuming degrees of freedom unnecessarily in others, and either outcome would distort the country-specific t-statistics that the panel statistic subsequently averages.

For each country, candidate augmented Dickey–Fuller regressions with zero, one or two lagged first differences were estimated in the form

$$\Delta y_{it} = \alpha_i + \delta_i t + \rho_i y_{i,t-1} + \Sigma_{j=1}^{p_i} \varphi_{ij} \Delta y_{i,t-j} + \varepsilon_{it} \quad (6)$$

where $p_i \in \{0, 1, 2\}$. The trend term $\delta_i t$ was included only for those variables for which the preceding diagnostics supported a linear trend, while every specification contained an intercept.

The three candidates were compared on a common estimation sample, an observation entering the comparison only if the dependent first difference, the lagged level and both candidate lags of the first difference were available for it. Holding the sample fixed ensures that differences in the criterion reflect differences in fit rather than differences in the number or timing of the underlying observations, which would otherwise render the comparison uninformative.

For each country and candidate lag order the Bayesian information criterion was computed as

$$BIC_i(p) = -2 \ln L_{i,p} + k_{i,p} \ln n_i \quad (7)$$

where $L_{i,p}$ is the maximised likelihood, $k_{i,p}$ the number of estimated parameters and n_i the number of observations in the common estimation sample. The order minimising the criterion was selected, following the model-selection principle of Schwarz

(1978), whose penalty term increases with the sample size and therefore favours parsimony more strongly than competing criteria – an appropriate property when the country-specific series are short.

The selected lag orders were arranged in the ascending country order used by the CIPS routine implemented in Stata through the pescadf command (Lewandowski, 2006), and the length of the resulting lag vector was verified against the number of countries entering each test. For countries with shorter series in which the criterion selected two lags, additional checks confirmed that the augmented regressions remained estimable, guarding against the silent loss of units at the testing stage.

4.3.4. The Pesaran CIPS Panel Unit-Root Test

Panel stationarity was assessed with the cross-sectionally augmented IPS test of Pesaran (2007). This second-generation procedure was selected precisely because the diagnostics reported in Section 4.2 established strong dependence across countries, a condition under which first-generation tests that assume independence across units over-reject the unit-root null and yield unreliable classifications.

The country-specific CADF regression underlying the test may be written as

$$\Delta y_{it} = a_i + b_i t + \rho_i y_{i,t-1} + c_i \bar{y}_{t-1} + d_i \Delta \bar{y}_t + \Sigma_{j=1}^{pi} \varphi_{ij} \Delta y_{i,t-j} + e_{it} \quad (8)$$

where $\bar{y}_t$ denotes the cross-sectional average of the variable in the period concerned. These averages act as proxies for the unobserved common factors that generate the dependence detected earlier, filtering their influence out of the individual regressions and thereby restoring the approximate validity of the country-specific t-statistics. The trend term $b_i t$ appears only for the variables identified as trending at the deterministic-specification stage.

The individual CADF t-statistics associated with the lagged level are then averaged to construct the CIPS statistic. Because the panels are unbalanced, inference is based on the standardised Z[t-bar] statistic reported by the routine, which adjusts for the

varying time dimension across units. The null holds that every country contains a unit root, while the alternative states that a non-zero fraction of the country-specific series is stationary, so rejection is a panel-level conclusion and does not license the claim that any particular country is individually stationary.

The null was evaluated at the 10 per cent significance level. Variables for which the null was rejected in levels were classified as I(0), while for those where it could not be rejected the series was transformed to its first difference and the entire lag-selection procedure was repeated, rather than transferring the level lag orders to the differenced series, since differencing may alter the short-run dynamics of a variable materially.

For a logged variable, the transformation applied at this stage was

$$\Delta ln\, x_{it} = ln\, x_{it} - ln\, x_{i,t-1} \quad (9)$$

rather than ln Δx_{it}. The former is a log change with an approximate proportional-growth interpretation, whereas the latter is undefined whenever the change is zero or negative, a case that arises frequently in credit and investment series.

Letting $z_{it} = \Delta y_{it}$, the lag-selection regression for the first-differenced series was written as

$$\Delta z_{it} = \alpha_i + \rho_i z_{i,t-1} + \Sigma_{j=1}^{p_i} \psi_{ij} \Delta z_{i,t-j} + v_{it} \quad (10)$$

These regressions included a constant but no deterministic linear trend, since a linear trend in the level of a series corresponds to a constant drift in its first difference, which the intercept already captures.

The criterion was again evaluated over zero, one and two lags on a common estimation sample, after which the selected country-specific lag vector was supplied to the CIPS test for z_{it}, rejection at this stage – following non-rejection in levels – producing a classification of the original variable as I(1).

4.3.5. Integration Orders and Their Implications for Specification

Table 13 reports, for each variable, the deterministic specification adopted, the number of countries entering the test, the standardised statistics obtained in levels and, where required, in first differences, together with the resulting integration order.

Table 13. Pesaran (2007) CIPS Panel Unit Root Test Results: Cross-Country Panel, 1992–2019

Variable	Deterministic specification	N	Level CIPS Z[t-bar] (p-value)	First-difference CIPS Z[t-bar] (p-value)	Integration order
ln_AVASGDP	Constant and linear trend	188	−8.425 (<0.001)	—	I(0)
ln_IVASGDP	Constant	186	−3.427 (<0.001)	—	I(0)
ln_SVASGDP	Constant and linear trend	182	−1.380 (0.084)	—	I(0)
ln_ACGDP	Constant	79	0.897 (0.815)	−19.376 (<0.001)	I(1)
ln_PC	Constant and linear trend	176	−0.533 (0.297)	−31.308 (<0.001)	I(1)
ln_SE	Constant and linear trend	87	0.384 (0.650)	−16.389 (<0.001)	I(1)
ln_GCF	Constant	162	−11.751 (<0.001)	—	I(0)
ln_GC	Constant	164	−3.965 (<0.001)	—	I(0)
ln_TO	Constant	172	−4.503 (<0.001)	—	I(0)
INF	Constant	177	−25.518 (<0.001)	—	I(0)

Notes: The table reports the standardised Z[t-bar] statistic from the Pesaran (2007) cross-sectionally augmented IPS (CIPS) test, whose null is that all panel units contain a unit root and whose alternative is that at least a fraction of the units is stationary. Tests use each variable's longest continuous country-specific block within 1992–2019 and retain countries with at least 15 consecutive annual observations, so the standardised statistic is reported for these unbalanced panels. Country-specific lag orders were selected by the Bayesian information criterion from 0–2 lags on a common estimation sample. Variables nonstationary in levels were retested in first differences under a constant-only specification, and a dash indicates that no such test was required. Stata p-values of 0.000 are reported as < 0.001. Rejection is a panel-level conclusion and does not imply stationarity for every individual country.

Source. Author's calculations based on the study dataset.

Applying the 10 per cent decision rule, the unit-root null is rejected in levels for ln_AVASGDP, ln_IVASGDP, ln_SVASGDP, ln_GCF, ln_GC, ln_TO and INF, all of which are accordingly classified as I(0). Inflation yields the largest statistic in absolute value, at −25.518, a result consistent with the rapid mean reversion typically observed in annual price growth.

The null could not be rejected in levels for ln_ACGDP, ln_PC or ln_SE, which were therefore retested in first differences using newly selected country-specific lag orders. Their differences are strongly stationary, with standardised statistics of −19.376, −31.308 and −16.389 and p-values below 0.001 in every case, so all three are classified as I(1). No variable required second differencing and none was found to be I(2), which removes any need for further transformation and confirms that a single differencing operation is sufficient to render the nonstationary series usable.

These classifications govern the specification of the estimated equations. Because ln_ACGDP, ln_PC and ln_SE are integrated of order one, they cannot enter the models in levels alongside stationary regressors without exposing the estimates to the risk of spurious inference, and they are consequently included in first differences. The financial-development coefficients therefore measure the response of sectoral value added to the annual change in credit depth rather than to its level, and the results reported in the following chapters must be read in those terms, since the level of credit relative to output is no longer identified within this specification.

The specification established here also governs the diagnostics that follow. The cross-sectional dependence statistics of Section 4.2 necessarily rest on the level specification, for the reason given there, whereas the slope homogeneity, groupwise heteroskedasticity and serial correlation tests reported in Sections 4.4 to 4.6 were obtained from the residuals of the estimated equation, in which d_ln_ACGDP or d_ln_PC and d_ln_SE replace their level counterparts. Reporting residual diagnostics for a model other than the one taken to the data would leave the properties of the actual disturbance process undocumented, and the three tests were therefore recomputed on the differenced specification rather than carried over from the level results.

4.4. Slope Homogeneity Testing

Pooled panel estimators impose a common slope vector on every cross-sectional unit, and that restriction is an assumption rather than a property of the data. The test of Pesaran and Yamagata (2008) evaluates the null that the slope coefficients are identical across units against the alternative that at least one unit differs, and it is well suited to panels in which the cross-sectional dimension is large relative to the time dimension. Examining the restriction is particularly important in a sample as diverse as the present one, since economies at different levels of development, with different financial structures and different production technologies, need not respond identically to the same stimulus, and imposing homogeneity where it does not hold yields biased parameter estimates and misleading policy conclusions (Pesaran & Yamagata, 2008). The test was applied to the three sectoral fixed-effects specifications in the form established in Section 4.3, so that financial development and human capital enter in first differences while investment, government consumption, trade openness and inflation enter in levels. Both the standardised Delta statistic and its bias-adjusted counterpart are reported, the latter being the more reliable of the two in finite samples because the standardised version is subject to size distortion when the time dimension is moderate (Pesaran & Yamagata, 2008). Table 14 presents the resulting statistics.

Table 14. Pesaran and Yamagata (2008) Slope Homogeneity Test Results: Agriculture, Industry and Services, 1992–2019

Dependent Variable	Δ Statistic	p-value	Δ_adj Statistic	p-value	Conclusion
ln_AVASGDP	17.480	0.000	24.388	0.000	Slope Heterogeneity
ln_IVASGDP	18.661	0.000	26.134	0.000	Slope Heterogeneity
ln_SVASGDP	18.211	0.000	25.508	0.000	Slope Heterogeneity

Notes: The table reports the Pesaran and Yamagata (2008) slope homogeneity test, whose null is homogeneous slope coefficients across all cross-sectional units. Δ is the

original Delta statistic and Δ_adj the bias-adjusted statistic. The agricultural specification includes d_ln_ACGDP and the industrial and services specifications d_ln_PC, and all three include d_ln_SE, ln_GCF, ln_GC, ln_TO and INF. A constant was partialled out in all tests. Stata p-values of 0.000 indicate $p < 0.001$.

The null of homogeneity is rejected in every specification. In the agricultural model the Delta statistic is 17.480 and its bias-adjusted counterpart 24.388, both significant at the 1 per cent level, while the industrial model yields 18.661 and 26.134 and the services model 18.211 and 25.508. The agreement between the two versions of the statistic strengthens the finding, since the bias adjustment would be expected to weaken the rejection if the original statistic were merely oversized. Heterogeneity is thus detected in every sector rather than confined to any one of them, although the relative magnitudes of the three statistics carry no interpretation, resting as they do on samples of differing size and composition.

Substantively, the rejection means that a given change in financial development, human capital, investment, government spending, trade openness or inflation does not translate into the same change in sectoral value added everywhere. Such divergence is consistent with cross-country differences in economic structure, institutional quality, financial market depth, technological capability and policy regime, all of which may plausibly alter the marginal productivity of a unit of credit or a unit of capital (Chudik & Pesaran, 2015). Heterogeneity of this kind is precisely what an aggregate, one-size-fits-all reading of the finance–growth relationship conceals.

The implication is interpretive as much as technical, and it differs between the two estimators employed in this study. In the static fixed-effects specification the pooled coefficients are best read as an average response across the panel rather than a parameter shared by every economy, so conclusions drawn from them apply to the sample as a whole and not to any individual country. The qualification is considerably stronger for the dynamic specification of Section 4.8, since Pesaran and Smith (1995) demonstrate that pooling dynamic heterogeneous panels can yield estimates that are not merely averages but inconsistent, the bias arising because the omitted heterogeneity is correlated with the lagged dependent variable. The System

GMM results are therefore reported subject to this limitation rather than presented as free of it.

Two features of the research design mitigate the concern without eliminating it. The relationship is estimated separately by sector, by period and by income group, so that the units pooled within any single estimation are considerably more alike than the full panel, narrowing the heterogeneity that pooling must absorb. Moreover, the fixed-effects and System GMM estimates rest on different identifying assumptions, and convergence between them offers some reassurance that the reported averages are not artefacts of a single estimation strategy, although agreement between two pooled estimators cannot by itself establish that pooling is innocuous.

4.5. Heteroskedasticity Testing

Groupwise heteroskedasticity is a near-inevitable feature of cross-country panels, since the variability of the disturbance differs across economies for structural, institutional and macroeconomic reasons. Its presence leaves the fixed-effects coefficients unbiased and consistent, but it renders the conventional covariance estimator inconsistent, so that the reported standard errors, t-statistics and confidence intervals cease to be trustworthy (Wooldridge, 2010; Greene, 2018). Testing for it is therefore a step in choosing an inferential procedure rather than a check on the coefficient estimates themselves.

Following estimation of each fixed-effects model, groupwise heteroskedasticity was examined with the modified Wald test implemented in Stata as xttest3 (Baum, 2001). The test evaluates the null that every cross-sectional unit shares a common disturbance variance, H_0: $\sigma_i^2 = \sigma^2$, against the alternative that the variance of at least one unit differs, H_1: $\sigma_i^2 \neq \sigma^2$, and it is well suited to fixed-effects models estimated on unbalanced panels, which explains its standing as a routine post-estimation diagnostic in applied panel econometrics (Greene, 2018; Baum, 2001).

The diagnostics were computed on the residuals of the three estimated specifications. Agricultural value added is explained by the change in agricultural credit

(d_ln_ACGDP), the change in secondary school enrolment (d_ln_SE), gross capital formation (ln_GCF), government consumption (ln_GC), trade openness (ln_TO) and inflation (INF), while the industrial and services models retain the same macroeconomic controls and replace agricultural credit with private credit (d_ln_PC), which captures the financial development relevant to non-agricultural activity. Table 15 reports the resulting statistics.

Table 15. Modified Wald Test for Groupwise Heteroskedasticity: Agriculture, Industry and Services, 1992–2019

Dependent Variable	Model	χ^2 Statistic	Degrees of Freedom	p-value	H_0: Homoskedasticity	Conclusion
ln_AVASGDP	FE	22,193.05	93	<0.001	Rejected	Groupwise heteroskedasticity is present
ln_IVASGDP	FE	283,752.85	156	<0.001	Rejected	Groupwise heteroskedasticity is present
ln_SVASGDP	FE	9,374,967.53	155	<0.001	Rejected	Groupwise heteroskedasticity is present

Notes: The table reports the modified Wald test for groupwise heteroskedasticity (xttest3) following fixed-effects estimation, whose null is equal error variances across all cross-sectional units. The statistics follow an asymptotic chi-square distribution. The agricultural specification includes d_ln_ACGDP and the industrial and services specifications d_ln_PC, and all three include d_ln_SE, ln_GCF, ln_GC, ln_TO and INF.

The null of homoskedasticity is rejected overwhelmingly in all three models, the statistic reaching $\chi^2(93) = 22{,}193.05$ for agriculture, $\chi^2(156) = 283{,}752.85$ for industry and $\chi^2(155) = 9{,}374{,}967.53$ for services, with p-values below 0.001 throughout. Magnitudes of this order indicate a violation that is extreme rather than marginal, reflecting residual variances that differ substantially across countries within every sector. The three statistics are not comparable with one another, since they rest on different samples and carry different degrees of freedom, so no inference is drawn from the fact that one exceeds another.

The pattern is readily interpretable in terms of sample composition. The countries in the panel differ widely in development level, financial depth, institutional quality,

production technology and macroeconomic stability, and such differences plausibly generate systematic variation in the volatility of sectoral value added and hence in the variance of the regression disturbance. In panels combining advanced and developing economies, groupwise heteroskedasticity is the expected outcome rather than an anomaly requiring special explanation (Baltagi, 2021).

The consequence bears on covariance estimation alone, leaving the point estimates intact. Taken together with the dependence documented in Section 4.2, the finding also narrows the set of admissible remedies, since cluster-robust corrections address heteroskedasticity and within-country correlation while continuing to assume independence between units, an assumption the present data do not support (Hoechle, 2007).

4.6. Serial Correlation Testing

Serial correlation constitutes a further threat to conventional panel inference, as it violates one of the assumptions on which the standard covariance estimator rests. The fixed-effects estimator remains unbiased and consistent under strict exogeneity even when the disturbances are serially correlated, yet first-order autocorrelation biases the estimated standard errors and may therefore invert the conclusion of a significance test (Wooldridge, 2010; Greene, 2018). Since macroeconomic aggregates adjust slowly and absorb shocks over several periods, serial correlation is common in cross-country panels and its presence should be established rather than assumed.

Following estimation of the fixed-effects models, first-order serial correlation was examined using the Wooldridge (2010) test. This procedure is regarded as a reliable diagnostic for fixed-effects specifications because it performs well in unbalanced panels, imposes few assumptions on the structure of the error variance and requires only a single auxiliary regression on the first-differenced residuals (Drukker, 2003). It evaluates the null that the idiosyncratic disturbances are not serially correlated against the alternative that first-order autocorrelation is present, and it was applied to the residuals of the same three estimated specifications described in Section 4.5, in

which the change in agricultural credit enters the agricultural equation and the change in private credit the industrial and services equations. Table 16 presents the results.

Table 16. Wooldridge Test for First-Order Serial Correlation: Agriculture, Industry and Services, 1992–2019

Dependent Variable	Test Statistic	p-value	Null Hypothesis (H_0)	Conclusion
ln_AVASGDP	F(1, 85) = 115.560	0.0000	No first-order autocorrelation	Rejected ($p < 0.01$): First-order serial correlation detected
ln_IVASGDP	F(1, 145) = 186.542	0.0000	No first-order autocorrelation	Rejected ($p < 0.01$): First-order serial correlation detected
ln_SVASGDP	F(1, 144) = 536.012	0.0000	No first-order autocorrelation	Rejected ($p < 0.01$): First-order serial correlation detected

Notes: The table reports the Wooldridge test for first-order serial correlation in panel data, whose null is no first-order autocorrelation in the idiosyncratic disturbances. The agricultural specification includes d_ln_ACGDP and the industrial and services specifications d_ln_PC, and all three include d_ln_SE, ln_GCF, ln_GC, ln_TO and INF. Stata p-values of 0.0000 indicate $p < 0.001$.

The null of no autocorrelation is rejected in every model. The agricultural specification yields F(1, 85) = 115.560, the industrial specification F(1, 145) = 186.542 and the services specification F(1, 144) = 536.012, all significant well beyond the 1 per cent level. Shocks to sectoral value added therefore carry over from one year to the next within countries in all three sectors, though the relative size of the three statistics supports no comparison between them, computed as they are on different samples. Notably, the rejections persist even though the financial development and human capital variables enter in first differences, a pattern consistent with residual persistence that the differenced regressors do not account for, while leaving its source unidentified, since a residual-based test cannot attribute autocorrelation to any particular component of the model.

Persistence of this kind is economically intelligible rather than surprising. Sectoral value added, financial development, investment, government spending, trade

openness, educational attainment and inflation all evolve through gradual structural adjustment rather than abrupt reallocation, so a country-specific shock – whether originating in technological change, institutional reform, the business cycle or macroeconomic policy – is likely to influence sectoral performance over several consecutive periods. Annual macroeconomic data consequently tend to generate serially correlated disturbances almost by construction (Baltagi, 2021).

Two distinct implications follow. Inference must again rely on a covariance estimator that tolerates temporal dependence, which the Driscoll–Kraay procedure provides. Persistence of this magnitude also raises the possibility that the underlying relationship is dynamic rather than static, since a specification excluding the lagged dependent variable would then be misspecified rather than merely inefficient. This consideration motivates the dynamic panel estimator developed in Section 4.8, though the test result establishes only that the static specification leaves systematic temporal structure in the residuals and does not by itself demonstrate that a lagged dependent variable is the appropriate remedy.

4.7. Fixed-Effects Estimation with Driscoll–Kraay Standard Errors

The diagnostics reported in Sections 4.2 to 4.6 deliver a coherent verdict, since the three sectoral models display strong cross-sectional dependence, slope heterogeneity across countries, groupwise heteroskedasticity and first-order serial correlation. None of these findings contradicts the consistency of the fixed-effects coefficients under strict exogeneity, but together they rule out the conventional covariance estimator, and the baseline results are therefore reported with Driscoll–Kraay standard errors.

The fixed-effects estimator was retained as the baseline because it controls for unobserved, time-invariant country characteristics by exploiting only within-country variation over time, thereby limiting the omitted-variable bias that afflicts cross-sectional comparisons of economies differing in geography, legal tradition and institutional endowment (Wooldridge, 2010; Baltagi, 2021). The standard specification may be written as

$$y_{it} = \alpha_i + x'_{it}\beta + \varepsilon_{it} \quad (11)$$

where y_{it} denotes sectoral value added in country i at time t, α_i represents the unobserved country-specific effect, x_{it} is the vector of explanatory variables, β the vector of slope coefficients and ε_{it} the idiosyncratic disturbance. The individual effects are eliminated through the within transformation,

$$y_{it} - \bar{y}_i = (x_{it} - \bar{x}_i)'\beta + (\varepsilon_{it} - \bar{\varepsilon}_i) \quad (12)$$

where a bar denotes the time average for each cross-sectional unit, after which ordinary least squares is applied to the transformed variables. The transformation removes every time-invariant characteristic, whether observed or not, at the cost of discarding the cross-country variation that would otherwise identify the effects of persistent structural differences.

Conventional covariance estimation nonetheless requires homoskedastic, serially independent and cross-sectionally independent disturbances (Greene, 2018), and none of these conditions survives the diagnostic evidence assembled above. Standard errors computed under these assumptions are consequently inconsistent in the present setting, so that hypothesis tests would be unreliable and confidence intervals misleadingly narrow, regardless of how precisely the coefficients themselves are estimated.

Inference therefore relies on the covariance matrix estimator of Driscoll and Kraay (1998), which extends the heteroskedasticity- and autocorrelation-consistent framework of Newey and West (1987) to panel data by estimating the long-run covariance matrix nonparametrically from cross-sectional averages of the regression residuals. Unlike conventional robust estimators, this approach remains asymptotically valid under general forms of spatial and temporal dependence and does not require the dependence structure to be specified in advance, an important advantage when the common factors driving cross-country correlation are unobserved. Formally, the estimator is defined as

$$Var_{DK}(\hat{\beta}) = (X'X)^{-1}\,\hat{\Omega}_{DK}\,(X'X)^{-1} \quad (13)$$

Where

$$\hat{\Omega}_{DK} = \hat{\Gamma}_0 + \Sigma_{j=1}^{L} w(j, L)(\hat{\Gamma}_j + \hat{\Gamma}'_j) \quad (14)$$

$\hat{\Gamma}_j$ denotes the estimated autocovariance matrix at lag j, L is the selected bandwidth and w(j, L) represents the Bartlett kernel weights that ensure a positive semi-definite covariance estimate (Driscoll & Kraay, 1998). The estimator modifies the covariance matrix alone and leaves the fixed-effects point estimates unchanged, so the reported coefficients remain the within estimates while the inference constructed around them becomes robust, and the bandwidth was set in accordance with the standard rule relating the truncation lag to the length of the time dimension (Newey & West, 1994).

One limitation of the procedure should be acknowledged. Consistency of the Driscoll–Kraay covariance estimator is established as the time dimension grows without bound, whereas the present panel spans 28 annual observations at most, so the asymptotic justification is approximate rather than exact in this application. Simulation evidence indicates that the estimator performs well at moderate time dimensions and substantially outperforms alternatives that ignore cross-sectional dependence (Hoechle, 2007), yet the reported standard errors are best regarded as an improvement on the conventional and cluster-robust alternatives rather than as exactly sized.

The procedure is particularly well suited to a macroeconomic panel of the present kind. Countries are exposed to common shocks transmitted through global business cycles, trade integration, financial contagion, commodity prices, technological diffusion and coordinated policy responses, generating contemporaneous dependence across units, while domestic adjustment processes generate serial dependence within them. Cluster-robust estimators accommodate the second source of dependence but not the first, becoming inconsistent precisely when cross-sectional correlation is strong, whereas the Driscoll–Kraay estimator remains valid across all three departures from the classical assumptions (Hoechle, 2007). This approach nevertheless has a well-defined limit. Robust covariance estimation corrects

inference without addressing the potential endogeneity of the regressors or the dynamic persistence documented in Section 4.6, both of which affect the coefficients themselves rather than the standard errors surrounding them. A complementary dynamic estimator is therefore required, and it is set out in the following section.

4.8. Two-Step System GMM Estimation

The relationship between financial development and sectoral value added is unlikely to be recursive. Credit expansion may raise sectoral value added, yet growing sectors also generate collateral, profits and demand for external finance, so that the direction of causation runs in both directions and the regressors cannot be treated as strictly exogenous without justification (Calderón & Liu, 2003; Levine, 2005). Measurement error in credit aggregates and the omission of slowly evolving institutional characteristics compound the problem, indicating that a static estimator may recover a correlation that is only partially attributable to the effect of finance on output.

The dynamic specification employed in this study is written as

$$y_{it} = \alpha + \rho y_{i,t-1} + x'_{it}\beta + \mu_i + \varepsilon_{it} \quad (15)$$

where y_{it} denotes sectoral value added in country i in period t, $y_{i,t-1}$ its one-period lag, x_{it} the vector of explanatory variables, μ_i the unobserved country-specific effect and ε_{it} the idiosyncratic disturbance. Including the lagged dependent variable allows the model to accommodate the persistence identified by the serial correlation diagnostics, reflecting the gradual adjustment of production structures to changes in their economic environment.

Estimating this specification by conventional fixed effects would be inappropriate, since the lagged dependent variable remains correlated with the transformed disturbance even after the individual effects have been removed, producing the inconsistency documented by Nickell (1981) whenever the time dimension is short relative to the cross-sectional dimension. This dynamic panel bias is the principal reason for the widespread adoption of generalised method of moments estimators in

macroeconomic panel research (Baltagi, 2021). Difference GMM, introduced by Arellano and Bond (1991), addresses the problem by first differencing the model to eliminate the country-specific effects and by instrumenting the endogenous variables with suitably lagged levels. When the series are highly persistent, however – as macroeconomic aggregates typically are – lagged levels are weak instruments for the differenced equation, generating finite-sample bias and imprecise estimates (Blundell & Bond, 1998).

The System GMM estimator addresses this weakness by augmenting the differenced equations with an equation in levels, an approach introduced by Arellano and Bover (1995) and developed into the estimator used here by Blundell and Bond (1998). It estimates the two sets of equations simultaneously, instrumenting the differenced equations with lagged levels and the level equations with suitably lagged first differences, the latter moment conditions requiring the additional assumption that the initial deviations from the long-run level are uncorrelated with the country-specific effects. Combining the two sets of moment conditions substantially improves instrument relevance and delivers more precise estimates than Difference GMM, particularly when the variables are persistent and the time series is short. The underlying moment conditions may be expressed as

$$E[y_{i,t-s}\,\Delta\varepsilon_{it}] = 0,\ s \geq 2 \quad (16)$$

for the differenced equation and as

$$E[\Delta y_{i,t-1}\,(\mu_i + \varepsilon_{it})] = 0 \quad (17)$$

for the level equation, with analogous orthogonality conditions imposed on the predetermined and endogenous explanatory variables according to the exogeneity assumed for each. Because these instruments are constructed internally from the lag structure of the data, the estimator requires no external instrumental variables, an advantage in a cross-country setting where credible external instruments for financial development are scarce. Estimation proceeds in two steps. An initial weighting matrix is constructed under the assumption of homoskedastic disturbances, and the

residuals obtained from this first step are then used to estimate an optimal weighting matrix that accommodates heteroskedasticity and autocorrelation, so the second-step estimator is asymptotically more efficient than its one-step counterpart. Since conventional two-step standard errors are severely downward biased in finite samples, the correction proposed by Windmeijer (2005) is applied throughout, materially improving the accuracy of finite-sample inference.

The credibility of System GMM estimates depends on the validity of the instrument set and on the serial correlation properties of the differenced residuals. Instrument proliferation is a particular concern, because an instrument count that grows with the time dimension overfits the endogenous variables and weakens the Hansen test to the point of nothingness, so the instrument set was restricted deliberately rather than allowed to expand mechanically (Roodman, 2009). Instrument validity is assessed through the Hansen (1982) test of overidentifying restrictions, whose null holds that the overidentifying restrictions are valid – a joint statement about the instrument set and the specification rather than a direct test of the exogeneity of any single instrument – while the Arellano and Bond (1991) tests examine first- and second-order autocorrelation in the first-differenced residuals. First-order correlation is expected by construction, whereas failure to reject the null of no second-order correlation is necessary for the consistency of the estimator, since its presence would indicate that the lag structure of the instruments is invalid.

Several features make this estimator appropriate to the objectives of the study. It controls for unobserved country-specific heterogeneity, captures the dynamics of sectoral value added through the lagged dependent variable and mitigates the endogeneity arising from reverse causality, omitted variables and measurement error, thereby improving on static panel estimators in precisely the dimensions that the finance–growth literature identifies as problematic. It is also designed for panels combining a large cross-sectional dimension with a comparatively short time dimension, which is the configuration of most cross-country macroeconomic datasets, including the one used here (Baltagi, 2021; Roodman, 2009). These properties come at the cost of reliance on internal instruments whose validity cannot be established with certainty, and the results are interpreted with that qualification in mind.

The two-step System GMM estimator is best understood as complementary to the fixed-effects specification with Driscoll–Kraay standard errors rather than as a replacement for it. The former addresses dynamics and endogeneity while assuming that internally constructed instruments are valid; the latter imposes no such assumption but corrects inference only. Convergence between the two sets of estimates therefore constitutes meaningful evidence that the relationship between financial development and sectoral value added is not an artefact of a particular identifying assumption, whereas divergence would signal that the findings are sensitive to the treatment of endogeneity and should be reported as such.

CHAPTER 5

CROSS-COUNTRY EMPIRICAL RESULTS

5.1. Periodical Analysis

The empirical analysis is organised around three sample windows – 1992–2007, 2010–2019 and the full 1992–2019 period – in order to establish whether the association between sectoral credit and sectoral value added is stable across markedly different financial regimes. Ending the sample in 2019 keeps the exceptional disruptions of the COVID-19 pandemic outside the estimation window, since those years combined simultaneous shocks to production, credit supply, public expenditure and the sectoral composition of output, thereby raising the risk that estimated coefficients would capture short-lived crisis dynamics rather than the underlying relationship. The two sub-periods separate the pre-crisis expansion from the decade that followed the global financial crisis, while 2008 and 2009 are excluded because they correspond to the most acute phase of that crisis and to the immediate policy response, which makes them difficult to assign to either regime.

For every sector and every period, the relationship is estimated with two complementary estimators, namely fixed effects with Driscoll–Kraay standard errors (Driscoll & Kraay, 1998) and the two-step System GMM estimator of Blundell and Bond (1998) with the finite-sample correction proposed by Windmeijer (2005). Each estimator is applied to linear, quadratic and cubic specifications, and each specification is estimated in a baseline form containing only the credit variable and in an extended form adding the macroeconomic controls described in Section 3. This design serves two purposes, allowing the functional form of the relationship to be selected from the data rather than imposed a priori, and allowing the static within-country association to be compared with a dynamic specification that accounts for

output persistence and for the potential endogeneity of credit. Specifications that produced no statistically significant credit coefficients under either estimator are reported in Appendix, so that the discussion below concentrates on the results that are statistically supported while the full estimation record remains available for inspection.

The functional form of each relationship was selected sequentially rather than imposed in advance, following the conventional polynomial modelling strategy. The linear specification was estimated first as a benchmark and then extended through the successive introduction of higher-order terms, with selection resting on the statistical significance of the added term together with Wald (F) tests of whether its inclusion improved model fit, both individually and jointly with the credit terms of lower order. Where the quadratic term proved significant and the joint restriction on the linear and quadratic terms was rejected, while the subsequent cubic term added no explanatory power, the quadratic specification was retained as the preferred form. Where the cubic term instead remained significant and improved fit, the same procedure was carried one degree further, and a quartic specification was estimated in order to confirm that the fourth-order term contributed nothing, which was the outcome in every case. The specification reported below for each sector and period is therefore the most parsimonious polynomial consistent with these tests, an approach that guards against the two opposing risks of imposing curvature the data do not support and of suppressing curvature the data do exhibit.

One feature of the estimation design shapes the interpretation of every coefficient reported in this section. The CIPS results presented in Section 4 indicate that the credit variables ln_ACGDP and ln_PC are integrated of order one, whereas the sectoral value added variables are stationary in levels, so the credit measures enter all equations in first differences. The estimated coefficients therefore describe the effect of the annual growth of sectoral credit relative to GDP, rather than the effect of the level of financial depth, and the ‘too much finance’ hypothesis is consequently examined in terms of the pace of credit expansion rather than its accumulated stock. This distinction is substantive rather than merely technical, since Bezemer et al. (2016) show that credit stocks and credit flows carry different growth implications,

while the credit-boom literature associates rapid expansions specifically with subsequent macroeconomic weakness (Mendoza & Terrones, 2008; Schularick & Taylor, 2012). Turning points are accordingly expressed in log-difference units and are interpreted only where they fall inside the range of credit growth actually observed in the estimation sample.

A second interpretive caution concerns the dependent variables, which measure agricultural, industrial and service value added as shares of GDP rather than in absolute terms. Movements in these shares reflect the changing sectoral composition of output, so a negative coefficient does not necessarily imply that the sector contracted in absolute terms, indicating instead that it grew more slowly than the rest of the economy. The credit measures are themselves expressed relative to GDP, so the same denominator enters both sides of every equation, and this common scaling operates in an identical fashion across the three sectoral models. Differences between the sectoral coefficients therefore cannot be attributed to the normalisation itself, which strengthens the case for reading any divergence in sign across sectors as substantive rather than as an artefact of measurement. The structural transformation literature treats such compositional shifts as a normal feature of development rather than as evidence of sectoral failure (Herrendorf et al., 2014), and the discussion below follows that convention.

5.1.1. The Agricultural Sector

5.1.1.1. The Sub-Periods (1992–2007 and 2010–2019)

Neither sub-period yields evidence that the growth of agricultural credit relative to GDP is systematically related to the agricultural share of value added. Across the linear, quadratic and cubic specifications estimated with Driscoll–Kraay standard errors, and across the corresponding two-step System GMM models, the coefficients on d_ln_ACGDP and on its higher-order terms remain statistically insignificant throughout, with the sole exception of isolated terms whose signs are not reproduced across estimators or specifications. The full estimation output for these models is reported in Appendix, so the discussion here concentrates on why the absence of a

measurable association is economically plausible rather than merely a symptom of low statistical power.

Financial development theory holds that credit supports production only when intermediaries reduce information asymmetries and channel resources towards investments capable of raising productivity, making the efficiency of allocation rather than the volume of lending the operative mechanism (Levine, 2005). Rural credit markets in much of the sample fall well short of that benchmark, since they are characterised by high transaction costs, weak collateral institutions, thin insurance provision and pronounced exposure to covariate risk, all of which limit the extent to which additional lending is translated into mechanisation, irrigation or improved inputs (Conning & Udry, 2007). Under such conditions, a rapid expansion of agricultural credit may finance short-term liquidity needs, refinancing or consumption smoothing rather than capital formation, which would leave the sectoral value added share unaffected even where credit volumes rise substantially. The insignificance of the sub-period estimates is therefore consistent with the view that agricultural finance operates through institutional channels that annual credit growth alone does not capture.

5.1.1.2. The Full Period (1992–2019)

Over the full sample, the quadratic specification estimated by fixed effects with Driscoll–Kraay standard errors is the only agricultural model in which the credit terms attain conventional significance, and it is therefore reported in the main text while the corresponding System GMM estimates are relegated to Appendix. Table 17 presents the baseline and extended versions of this specification.

Table 17. Quadratic Fixed-Effects Estimates with Driscoll–Kraay Standard Errors: Agricultural Value Added, 1992–2019

Variables	Baseline Model	Extended Model
d_ln_ACGDP	0.005 (0.009)	0.002 (0.009)
d_ln_ACGDP^2	-0.005* (0.003)	0.005** (0.002)

Table 17. (continued)

d_ln_SE	—	-0.142* (0.081)
ln_GCF	—	-0.002 (0.032)
ln_GC	—	0.130*** (0.032)
INF	—	0.0037*** (0.0009)
ln_TO	—	0.096* (0.055)
Year Fixed Effects	Yes	Yes
Constant	2.232*** (0.021)	Omitted
Model Statistics		
Observations	1,988	1,072
Countries	109	93
Within R-squared	0.2862	0.2833
F Statistic	693950.71***	161897.76***
Prob > F	0.000	0.000
Driscoll–Kraay Lag	3	3

Notes: Dependent variable: ln_AVASGDP. Fixed-effects estimates with Driscoll–Kraay robust standard errors (maximum lag = 3) in parentheses. d_ln_ACGDP² is the squared term of d_ln_ACGDP. The extended model adds d_ln_SE, ln_GCF, ln_GC, INF and ln_TO. Year fixed effects are included, and their coefficients are omitted. The constant was omitted from the extended model because of collinearity. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

In the baseline specification the linear term is positive but statistically indistinguishable from zero, while the squared term is negative and significant at the 10% level, describing a concave profile in which the marginal contribution of agricultural credit growth declines as expansion accelerates. The extended specification, which adds human capital, capital formation, government consumption, inflation and trade openness, reverses the sign of the quadratic term to a positive and more precisely estimated 0.005, implying a convex rather than concave profile once the macroeconomic environment is held constant. This reversal is the single most important feature of Table 17, since it indicates that the curvature detected in the baseline model is not a robust property of the agricultural relationship but is instead sensitive to specification and, equally, to sample composition, given that the extended model is estimated on 1,072 observations from 93 countries compared with 1,988 observations from 109 countries in the baseline.

The control variables nevertheless behave in an economically coherent manner and account for the bulk of the explanatory power in the extended model. Government consumption enters positively and highly significantly, which is consistent with the prominence of public support schemes, price interventions and subsidised input programmes in the agricultural sectors of many sample countries, while trade openness is positively signed at the 10% level, plausibly reflecting the export orientation of agriculture in lower-income economies. The negative coefficient on the growth of school enrolment is best read as a structural transformation effect rather than as evidence that education harms agriculture, since the reallocation of labour away from agriculture and towards industry and services is a well-documented feature of development (Herrendorf et al., 2014), with the accumulation of human capital plausibly accompanying that shift rather than driving the agricultural share downwards on its own. Gross capital formation is insignificant, which is unsurprising given that aggregate investment is dominated by non-agricultural activity, and the small positive inflation coefficient is most plausibly attributed to relative-price effects in a share-based dependent variable rather than to any productivity mechanism.

Because the baseline quadratic specification implies an inverted U-shaped profile, the formal test proposed by Lind and Mehlum (2010) was applied in order to establish whether the curvature is supported over the range of the data rather than being an artefact of extrapolation. Table 18 reports the outcome.

Table 18. Lind and Mehlum (2010) Inverted U-Shape Test: Agricultural Value Added, 1992–2019

Specification: $f(x) = x^2$ Estimated Turning Point: 0.502

	Lower Bound	Upper Bound
Interval	-5.662	4.376
Slope	0.065	-0.041
t-statistic	1.983	-1.442
p-value	0.029**	0.080*

Overall inverse U-shape test: t = 1.442; p = 0.080*. 80% Fieller interval for the turning point: [-0.719, 3.329].

Notes: Lind and Mehlum (2010) test for an inverted U-shaped relationship in the quadratic fixed-effects specification with Driscoll–Kraay standard errors. H_0 is a monotonic or U-shaped relationship and H_1 an inverted U-shape. The turning point is computed as $-\beta_1/(2\beta_2)$. The overall p-value of 0.080 rejects H_0 at the 10% level. The Fieller confidence interval is constructed at the 80% level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

The test rejects the null of a monotonic or U-shaped relationship at the 10% level, and the estimated turning point of 0.502 lies comfortably within the observed interval of d_ln_ACGDP, which extends from -5.662 to 4.376. The slope is positive and significant at the lower bound and negative at the upper bound, so the concavity is supported at both ends of the sampled range rather than being driven by behaviour outside it. Expressed in economically interpretable terms, the turning point corresponds to annual growth of agricultural credit relative to GDP of roughly 65%, a rate that only a small minority of country–year observations attain, which implies that the majority of the sample lies on the upward-sloping segment of the curve where additional agricultural credit growth is weakly beneficial.

Three considerations nevertheless caution against treating this inverted U as a firm finding. The overall test attains significance only at the 10% level, the 80% Fieller interval for the turning point spans [-0.719, 3.329] and therefore includes both negative values and implausibly high rates of expansion, and the curvature reverses sign once macroeconomic controls are introduced. Taken together, these features suggest that agricultural credit growth exhibits at most a weak and imprecisely located nonlinearity, and that the more defensible conclusion is the absence of a strong systematic association between the pace of agricultural credit expansion and the agricultural share of value added. This reading is consistent with the sub-period evidence in Appendix and with the argument that the effectiveness of agricultural finance depends on institutional conditions in rural credit markets rather than on lending volumes alone (Conning & Udry, 2007; Levine, 2005).

5.1.2. The Industrial Sector

The industrial sector is the only sector for which both estimators produce statistically significant credit coefficients in all three periods, which makes it the natural focus of

the periodical analysis. The discussion proceeds chronologically, examining the pre-crisis period, the post-crisis decade and the full sample in turn, and comparing the static and dynamic estimates within each period before drawing the temporal comparison in Section 5.1.4.

5.1.2.1. The Pre-Crisis Period (1992–2007)

Table 19 reports the fixed-effects estimates for 1992–2007, in which the linear specification proved to be the preferred functional form under the static estimator.

Table 19. Linear Fixed-Effects Estimates with Driscoll–Kraay Standard Errors: Industrial Value Added, 1992–2007

Variables	Baseline Model	Extended Model
d_ln_PC	0.036* (0.020)	-0.015 (0.044)
d_ln_SE	—	-0.205*** (0.063)
ln_GCF	—	0.044 (0.034)
ln_GC	—	0.101** (0.045)
INF	—	0.000062*** (0.000008)
ln_TO	—	0.176*** (0.044)
Year Fixed Effects	Yes	Yes
Constant	3.205*** (0.003)	Omitted
Observations	2,363	927
Countries	174	129
Within R-squared	0.0141	0.0924
F Statistic	36017.84***	10185.28***
Prob > F	0.000	0.000
Driscoll–Kraay Lag	3	3

Notes: Dependent variable: ln_IVASGDP. Fixed-effects estimates with Driscoll–Kraay robust standard errors (maximum lag = 3) in parentheses. The extended model adds d_ln_SE, ln_GCF, ln_GC, INF and ln_TO. Year fixed effects are included, and their coefficients are omitted. The constant was omitted from the extended model because of collinearity. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

In the baseline model the growth of private credit relative to GDP carries a positive coefficient of 0.036 that is significant at the 10% level, indicating that faster credit expansion during the pre-crisis period was weakly associated with a higher industrial share of value added. The magnitude is modest and the within R-squared of 0.0141 confirms that credit growth accounts for very little of the variation in the industrial share once country and year effects are absorbed. Once the macroeconomic controls are added the credit coefficient becomes negative and wholly insignificant, while the explanatory power of the model rises sharply to 0.0924, suggesting that the apparent credit effect in the baseline model partly proxies for the investment and trade conditions that accompany credit expansion rather than representing an independent channel.

The controls themselves are informative about the pre-crisis industrial environment. Trade openness and government consumption both enter positively and significantly, which is consistent with an industrialisation process driven by export orientation and by public demand in the decade preceding the crisis, whereas the growth of school enrolment carries a strongly negative coefficient that again reflects the reallocation of activity towards services as human capital accumulates. Gross capital formation is positively signed but imprecisely estimated, and the inflation coefficient, although highly significant, is economically negligible at 0.000062. The reduction of the estimation sample from 2,363 to 927 observations and from 174 to 129 countries when controls are added is itself consequential, since the extended sample is systematically tilted towards countries with more complete statistical reporting, so the loss of significance cannot be attributed to the controls alone. The dynamic estimates in Table 20 provide a rather different perspective, since the cubic specification was retained under System GMM and reveals nonlinearity that the static linear model cannot detect.

Table 20. Cubic Two-Step System GMM Estimates: Industrial Value Added, 1992–2007

Variables	Baseline Model	Extended Model
L.ln_IVASGDP	0.949*** (0.056)	0.957*** (0.068)
d_ln_PC	0.301 (0.244)	-0.210 (0.269)

Table 20. (continued)

d_ln_PC²	-0.075 (0.091)	-0.316 (0.312)
d_ln_PC³	-0.287** (0.130)	0.108 (0.380)
d_ln_SE	—	0.056 (0.080)
ln_GCF	—	0.075* (0.041)
ln_GC	—	-0.010 (0.012)
ln_TO	—	-0.009 (0.010)
INF	—	-0.000113*** (0.000030)
Year Fixed Effects	Yes	Yes
Constant	0.160 (0.177)	-0.000143 (0.135)
Observations	2,332	921
Countries	173	126
Instruments	29	34
AR(1) (p-value)	0.000	0.000
AR(2) (p-value)	0.623	0.802
Hansen Test (p-value)	0.731	0.833
Sargan Test (p-value)	0.570	0.710
Difference-in-Hansen (p-value)	0.352	0.670
Wald/F Statistic	153280.01***	56959.47***

Notes: Dependent variable: ln_IVASGDP. Two-step System GMM estimates with Windmeijer (2005) finite-sample-corrected robust standard errors in parentheses. d_ln_PC² and d_ln_PC³ are the squared and cubed terms of d_ln_PC. The extended model adds d_ln_SE, ln_GCF, ln_GC, INF and ln_TO. Year fixed effects are included, and their coefficients are omitted. The lagged dependent variable and the credit terms enter as collapsed GMM-style instruments at lags 2–3. The controls enter as IV-style instruments. Diagnostic entries other than Wald/F are p-values. Stata used a generalized inverse for a singular two-step moment covariance matrix, so Difference-in-Hansen results require caution. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

The diagnostics support the specification, since first-order serial correlation is present as expected while the null of second-order serial correlation cannot be rejected, and the Hansen, Sargan and Difference-in-Hansen statistics provide no evidence against the instrument set. The instrument count of 29 in the baseline model and 34 in the extended model remains far below the number of cross-sectional units, which matters because Hansen tests lose power when instruments proliferate

(Roodman, 2009a). The coefficient on the lagged dependent variable is 0.949 and highly significant, confirming that the industrial share is strongly persistent and that the static estimates in Table 19 omit an important source of dynamics.

Within this dynamic specification, the linear and quadratic credit terms are insignificant while the cubic term is negative and significant at the 5% level, producing a profile in which the marginal effect of credit growth becomes increasingly negative at high rates of expansion. The two turning points implied by the estimates are real-valued and located at -0.685 and 0.510, both lying inside the range of d_ln_PC spanned by the estimation sample, which extends from -1.781 to 1.367. The upper turning point marks a local maximum and corresponds to annual growth of private credit relative to GDP of approximately 67%, a rate situated just below the 99th percentile of the pre-crisis distribution, so the descending segment of the curve is identified from a very small number of extreme expansions. Moderate credit growth therefore appears broadly neutral or mildly favourable for the industrial share, whereas exceptionally rapid expansion is associated with a deteriorating industrial position, which is the asymmetry emphasised in the credit-boom literature (Schularick & Taylor, 2012), although Dell'Ariccia et al. (2016) find that only one boom episode in three is followed by a financial crisis, indicating that the adverse outcome is a tendency rather than a regularity. Two considerations temper this reading, since the three credit terms are only jointly significant at the 10% level and the discriminant of the derivative is imprecisely estimated, so the existence of two distinct turning points cannot be established with confidence. The extended model retains none of this curvature, which again coincides with a halving of the estimation sample, so the pre-crisis evidence for nonlinearity should be regarded as suggestive rather than decisive.

5.1.2.2. The Post-Crisis Period (2010–2019)

The post-crisis decade produces the clearest and most robust industrial results in the entire periodical analysis. Table 21 reports the cubic fixed-effects specification, in which all three credit terms are significant in both the baseline and the extended model.

Table 21. Cubic Fixed-Effects Estimates with Driscoll–Kraay Standard Errors: Industrial Value Added, 2010–2019

Variables	Baseline Model	Extended Model
d_ln_PC	-0.167*** (0.025)	-0.102*** (0.025)
d_ln_PC²	0.288*** (0.065)	0.286*** (0.081)
d_ln_PC³	0.138*** (0.028)	0.131** (0.043)
d_ln_SE	—	-0.051* (0.026)
ln_GCF	—	0.150*** (0.022)
ln_GC	—	-0.300*** (0.019)
INF	—	-0.0008 (0.0008)
ln_TO	—	-0.029 (0.029)
Year Fixed Effects	Yes	Yes
Constant	3.161*** (0.001)	Omitted
Observations	1,737	1,059
Countries	182	144
Within R-squared	0.0871	0.2396
F Statistic	496824.52***	12677.48***
Prob > F	0.000	0.000
Driscoll–Kraay Lag	3	3

Notes: Dependent variable: ln_IVASGDP. Fixed-effects estimates with Driscoll–Kraay robust standard errors (maximum lag = 3) in parentheses. d_ln_PC² and d_ln_PC³ are the squared and cubed terms of d_ln_PC. The extended model adds d_ln_SE, ln_GCF, ln_GC, INF and ln_TO. Year fixed effects are included, and their coefficients are omitted. The constant was omitted from the extended model because of collinearity. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

The coefficient pattern is remarkably stable across the two specifications, with the linear term negative, the squared term positive and the cubed term positive in both cases, and with the three credit terms jointly significant at the 1% level in the baseline model. The turning points of the cubic are estimated at -1.642 and 0.246, both real-valued and located within the observed range of d_ln_PC, and the

discriminant of the derivative is positive and significant at the 1% level, so the presence of two distinct turning points is firmly established rather than assumed. Since the median of d_ln_PC in this sample is 0.010 and its 95th percentile is 0.186, the interval over which the marginal effect is negative encompasses almost the entire mass of the data, with the upper turning point of 0.246 corresponding to annual private credit growth of roughly 28% and lying between the 95th and the 99th percentiles of the distribution. Faster credit expansion is therefore associated with a lower industrial share across the range in which countries ordinarily operate, and the relationship becomes favourable again only at rates of expansion attained by fewer than one country–year observation in twenty. The consistency of this profile across the baseline and extended models is notable, since the credit terms retain their magnitude and significance even after the addition of controls that raise the within R-squared from 0.0871 to 0.2396, indicating that the association is not merely a reduced-form reflection of investment or trade conditions.

Two of the controls deserve comment because they sharpen the interpretation. Gross capital formation enters positively and highly significantly, which confirms that physical investment supported industrial activity during the post-crisis decade and reinforces the point that the negative credit coefficient does not reflect a general weakness of capital accumulation. Government consumption, in contrast, is strongly negative, which is consistent with the composition of post-crisis fiscal expansion in many countries, where support was directed towards transfers, public services and consumption rather than towards manufacturing capacity. Trade openness and inflation are both insignificant in this period, whereas the growth of school enrolment retains its negative sign at the 10% level, again pointing to structural reallocation away from industry. Table 22 reports the corresponding dynamic estimates, for which the linear specification was retained under System GMM.

Table 22. Linear Two-Step System GMM Estimates: Industrial Value Added, 2010–2019

Variables	Baseline Model	Extended Model
L.ln_IVASGDP	0.511 (0.344)	0.815*** (0.084)
d_ln_PC	-0.298** (0.131)	-0.140 (0.250)

Table 22. (continued)

d_ln_SE	—	0.032 (0.057)
ln_GCF	—	0.080** (0.038)
ln_GC	—	-0.012 (0.024)
ln_TO	—	-0.029 (0.021)
INF	—	0.0007 (0.0013)
Year Fixed Effects	Yes	Yes
Constant	1.543 (1.078)	0.494* (0.252)
Observations	1,732	1,057
Countries	182	144
Instruments	17	22
AR(1) (p-value)	0.171	0.002
AR(2) (p-value)	0.530	0.613
Hansen Test (p-value)	0.757	0.785
Sargan Test (p-value)	0.000	0.582
Difference-in-Hansen (p-value)	0.892	0.473
Wald/F Statistic	5066.40***	45890.54***

Notes: Dependent variable: ln_IVASGDP. Two-step System GMM estimates with Windmeijer (2005) finite-sample-corrected robust standard errors in parentheses. The extended model adds d_ln_SE, ln_GCF, ln_GC, INF and ln_TO. Year fixed effects are included, and their coefficients are omitted. The lagged dependent variable and d_ln_PC enter as collapsed GMM-style instruments at lags 2–3, and the controls as IV-style instruments. Diagnostic entries other than Wald/F are p-values. Stata used a generalized inverse for a singular two-step moment covariance matrix, so Difference-in-Hansen results require caution. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

The baseline dynamic model returns a negative and significant credit coefficient of -0.298, corroborating the direction of the static estimates and indicating that faster private credit growth during the post-crisis decade was associated with a lower industrial share once persistence and endogeneity are accounted for. Two diagnostic features nevertheless require caution. The lagged dependent variable is imprecisely estimated in the baseline model, and the AR(1) statistic fails to reject the null of no first-order serial correlation in first-differenced residuals, which is the opposite of the pattern ordinarily expected and may indicate that the moment conditions are only weakly informative in this specification, although Roodman (2009b) presents the test

primarily as a check on the validity of lagged instruments rather than as a measure of identification strength. The non-robust Sargan test also rejects the overidentifying restrictions in the baseline model, although the robust Hansen test does not, and the extended model resolves both issues while rendering the credit coefficient insignificant. The dynamic evidence for the post-crisis period should therefore be read as consistent with, rather than independently confirming, the static results in Table 21.

The retention of a linear specification under System GMM, in contrast to the cubic form supported by the static estimator, is itself informative about what the dynamic model is able to identify. Successive powers of a differenced variable are noisy and strongly collinear with one another, so the deliberately parsimonious instrument set available in a ten-year panel supplies comparatively little independent variation with which to pin down the quadratic and cubic terms, both of which proved insignificant when added. The dynamic estimates for this period should accordingly be read as evidence on the direction of the credit effect rather than on its curvature, leaving the nonlinearity documented in Table 21 to rest on the static estimator alone.

The magnitude of the dynamic coefficient also merits attention, since -0.298 is close to twice the linear coefficient obtained under Driscoll–Kraay fixed effects, and the gap admits two explanations that the present design cannot separate. Conditioning on the lagged dependent variable isolates the contemporaneous response of the industrial share, whereas the static estimate blends that response with the slower adjustment embodied in persistence, so a larger short-run coefficient is to be expected on purely mechanical grounds. Alternatively, the instrumentation may be correcting attenuation arising from measurement error in annual credit growth, in which case the dynamic figure would be the less biased of the two. The instrument count of 17 for 182 countries is low enough to protect the Hansen test from the loss of power associated with instrument proliferation (Roodman, 2009a), which lends weight to its failure to reject, while the decisive rejection recorded by the Sargan statistic is most plausibly attributed to the non-robustness of that test under heteroskedasticity rather than to a genuine failure of the overidentifying restrictions.

The extended dynamic model behaves more conventionally in diagnostic terms, with a lagged dependent variable of 0.815 that is precisely estimated and an AR(1) statistic that rejects as expected, yet the credit coefficient falls to -0.140 and ceases to be significant. Gross capital formation remains positive and significant at 0.080, mirroring its role in the static specification reported in Table 21 and reinforcing the reading that physical investment rather than the pace of credit expansion is the proximate driver of industrial performance during this decade. Considered together, the two dynamic models indicate that the negative association survives the introduction of dynamics in its baseline form but not the joint imposition of controls and a substantially reduced sample, which reproduces the pattern of fragility already observed in the pre-crisis estimates.

5.1.2.3. The Full Period (1992–2019)

Estimating the industrial relationship over the complete sample combines the two regimes and yields the largest estimation samples in the analysis. Table 23 reports the cubic fixed-effects results.

Table 23. Cubic Fixed-Effects Estimates with Driscoll–Kraay Standard Errors: Industrial Value Added, 1992–2019

Variables	Baseline Model	Extended Model
d_ln_PC	0.040 (0.028)	-0.045 (0.039)
d_ln_PC²	-0.078* (0.038)	-0.086 (0.051)
d_ln_PC³	-0.045* (0.023)	-0.042* (0.023)
d_ln_SE	—	-0.257*** (0.058)
ln_GCF	—	0.196*** (0.041)
ln_GC	—	0.045 (0.066)
INF	—	0.000062*** (0.000013)
ln_TO	—	-0.001 (0.047)
Year Fixed Effects	Yes	Yes

Table 23. (continued)

Constant	3.196*** (0.004)	Omitted
Observations	4,442	2,187
Countries	186	156
Within R-squared	0.0294	0.1394
F Statistic	2153356.02***	398951.19***
Prob > F	0.000	0.000
Driscoll–Kraay Lag	3	3

Notes: Dependent variable: ln_IVASGDP. Fixed-effects estimates with Driscoll–Kraay robust standard errors (maximum lag = 3) in parentheses. d_ln_PC² and d_ln_PC³ are the squared and cubed terms of d_ln_PC. The extended model adds d_ln_SE, ln_GCF, ln_GC, INF and ln_TO. Year fixed effects are included, and their coefficients are omitted. The constant was omitted from the extended model because of collinearity. *** p < 0.01, ** p < 0.05, * p < 0.10.

The full-period profile differs instructively from the post-crisis one, since the linear term is positive though insignificant while both the squared and the cubed terms are negative, the latter remaining significant at the 10% level in the extended specification. This configuration implies a relationship that is mildly favourable at low rates of credit expansion and turns unfavourable once expansion accelerates, with a local maximum estimated at 0.216 and a lower turning point at -1.367, both lying inside the observed range of d_ln_PC. The maximum corresponds to annual private credit growth of approximately 24% and falls marginally below the 95th percentile of the full-period distribution, so the descending segment of the curve covers the upper tail of credit expansions rather than the typical experience of the sample. This location should nevertheless be treated as indicative, given that the upper turning point is imprecisely estimated with a 90% confidence interval extending from -0.076 to 0.507 and that the discriminant of the derivative is not significantly different from zero. The pattern corresponds closely to the nonlinear finance–growth relationship documented at the aggregate level by Arcand et al. (2015) and by Law and Singh (2014), with the important difference that the nonlinearity identified here operates through the pace of credit expansion rather than through the accumulated depth of the financial system.

The persistence of the negative cubic term across the baseline and extended specifications, together with a within R-squared that rises from 0.0294 to 0.1394 when controls are included, indicates that the curvature is not an artefact of omitted macroeconomic conditions. Gross capital formation carries a large positive coefficient of 0.196, confirming that investment remains the dominant proximate determinant of industrial performance, while the growth of school enrolment is strongly negative and consistent with long-run structural reallocation. Government consumption and trade openness are both insignificant over the full period, which is unsurprising given that their pre-crisis and post-crisis effects, documented in Tables 19 and 21, carry opposite signs and therefore offset one another when the regimes are pooled.

The dynamic estimates reported in Table 24 reinforce this reading.

Table 24. Cubic Two-Step System GMM Estimates: Industrial Value Added, 1992–2019

Variables	Baseline Model	Extended Model
L.ln_IVASGDP	0.911*** (0.049)	0.940*** (0.054)
d_ln_PC	0.061 (0.133)	-0.158 (0.155)
d_ln_PC²	-0.116 (0.119)	0.035 (0.397)
d_ln_PC³	-0.343** (0.157)	-0.171 (0.297)
d_ln_SE	—	0.048 (0.041)
ln_GCF	—	0.061*** (0.022)
ln_GC	—	-0.009 (0.012)
ln_TO	—	-0.014 (0.010)
INF	—	-0.000144** (0.000060)
Year Fixed Effects	Yes	Yes
Constant	0.261* (0.152)	0.094 (0.149)
Observations	4,405	2,179
Countries	186	156
Instruments	41	46
AR(1) (p-value)	0.001	0.339
AR(2) (p-value)	0.537	0.933

Table 24. (continued)

Hansen Test (p-value)	0.576	0.640
Sargan Test (p-value)	0.154	0.088
Difference-in-Hansen (p-value)	0.238	0.852
Wald/F Statistic	72066.05***	103710.34***

Notes: Dependent variable: ln_IVASGDP. Two-step System GMM estimates with Windmeijer (2005) finite-sample-corrected robust standard errors in parentheses. d_ln_PC² and d_ln_PC³ are the squared and cubed terms of d_ln_PC. The extended model adds d_ln_SE, ln_GCF, ln_GC, INF and ln_TO. Year fixed effects are included, and their coefficients are omitted. The lagged dependent variable and the credit terms enter as collapsed GMM-style instruments at lags 2–3. The controls enter as IV-style instruments. Diagnostic entries other than Wald/F are p-values. Stata used a generalized inverse for a singular two-step moment covariance matrix, so Difference-in-Hansen results require caution. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

The baseline dynamic model reproduces the negative cubic term found under the Driscoll–Kraay fixed-effects estimator, estimating it at -0.343 with significance at the 5% level, while the lagged dependent variable of 0.911 confirms the high persistence of the industrial share. The implied local maximum lies at 0.156, equivalent to annual private credit growth of about 17% and situated below the 90th percentile of the sample distribution, which places it reasonably close to the value of 0.216 obtained from the Driscoll–Kraay estimates. Neither root is individually significant in this specification, so the agreement between the two estimators should be read as convergence on the broad location of the turning region rather than as precise identification of a threshold. The diagnostics are otherwise satisfactory in the baseline specification, with second-order serial correlation absent and the Hansen, Sargan and Difference-in-Hansen tests all failing to reject, while the instrument count of 41 remains well below the 186 cross-sectional units. The agreement between the static and

The extended dynamic model retains the negative cubic coefficient but loses precision, which is consistent with the pattern observed throughout the analysis whereby the addition of controls halves the sample and weakens identification of the higher-order terms. Gross capital formation remains positive and significant at 0.061, and inflation now enters negatively, in contrast to the small positive coefficient

obtained under the Driscoll–Kraay fixed-effects estimator in Table 23. That sign reversal deserves note rather than reinterpretation, since it most likely reflects the different variation exploited by the two estimators, with the dynamic model conditioning on lagged output and thereby isolating the short-run disruptive effect of price instability. Overall, the full-period evidence supports the conclusion that the pace of private credit expansion bears a nonlinear relationship to industrial performance, in which rapid expansion is associated with a declining industrial share of value added.

5.1.3. The Service Sector

5.1.3.1. The Pre-Crisis Period and the Full Sample

The service sector yields significant credit coefficients in only one of the three periods, since neither the pre-crisis window nor the full sample produces estimates that attain conventional significance under either estimator. In the 1992–2007 models the coefficients on d_ln_PC and on its higher-order terms are uniformly small relative to their standard errors under both fixed effects and System GMM, and the dynamic estimates for that period additionally show evidence of second-order serial correlation at the 10% level, which further weakens their evidentiary value. Over the full 1992–2019 sample the static estimates are similarly insignificant, while the dynamic models encounter difficulties with the overidentifying restrictions, since the Hansen and Difference-in-Hansen tests reject in the linear and cubic specifications and the non-robust Sargan test rejects throughout. These results are reported in Appendix rather than in the main text, and the diagnostic problems recorded there provide an additional reason for restricting substantive interpretation to the post-crisis period.

The absence of a stable service-sector relationship across the longer horizons is intelligible in light of the composition of private credit. Aggregate private credit finances households, real estate and non-financial firms in proportions that vary widely across countries and over time, and the literature indicates that these components carry different implications for real activity, with enterprise credit

positively associated with growth while household credit is not (Beck et al., 2012). Bezemer et al. (2016) reach a similar conclusion by distinguishing credit to non-financial business from mortgage and other asset-market lending, showing that the latter accounts for much of the weakening of the credit–growth association since 1990. Since services encompass both highly productive business activities and sectors closely tied to property and consumption, aggregate credit growth may exert offsetting influences that leave no measurable net effect over long horizons.

5.1.3.2. The Post-Crisis Period (2010–2019)

Table 25 presents the cubic fixed-effects specification for the service sector in the post-crisis decade, which is the only service model to yield statistically supported credit coefficients. The corresponding System GMM models produced no significant credit coefficients under any of the functional forms considered and are therefore reported in Appendix, which means that the dynamic estimator can neither corroborate nor contradict the static findings discussed below.

Table 25. Cubic Fixed-Effects Estimates with Driscoll–Kraay Standard Errors: Services Value Added, 2010–2019

Variables	Baseline Model	Extended Model
d_ln_PC	0.141*** (0.025)	0.066** (0.022)
d_ln_PC²	-0.128* (0.067)	0.171* (0.088)
d_ln_PC³	-0.075** (0.028)	0.067 (0.043)
d_ln_SE	—	0.026 (0.020)
ln_GCF	—	-0.003 (0.010)
ln_GC	—	0.237*** (0.032)
INF	—	0.0007 (0.0010)
ln_TO	—	-0.099*** (0.011)
Year Fixed Effects	Yes	Yes
Constant	4.011*** (0.001)	Omitted
Observations	1,737	1,059
Countries	182	144
Within R-squared	0.1646	0.3459

Table 25. (continued)

F Statistic	532086.18***	18578.05***
Prob > F	0.000	0.000
Driscoll–Kraay Lag	3	3

Notes: Dependent variable: ln_SVASGDP. Fixed-effects estimates with Driscoll–Kraay robust standard errors (maximum lag = 3) in parentheses. d_ln_PC² and d_ln_PC³ are the squared and cubed terms of d_ln_PC. The extended model adds d_ln_SE, ln_GCF, ln_GC, INF and ln_TO. Year fixed effects are included, and their coefficients are omitted. The constant was omitted from the extended model because of collinearity. *** p < 0.01, ** p < 0.05, * p < 0.10.

The linear term is positive and highly significant in both specifications, estimated at 0.141 in the baseline model and at 0.066 once controls are added, which establishes that faster private credit growth during the post-crisis decade was associated with a higher service share of value added. This first-order result is the robust element of Table 25 and stands in direct opposition to the industrial estimates for the same period reported in Table 21, where the linear coefficient is negative and of comparable magnitude. The baseline cubic implies a local maximum at 0.405, corresponding to annual private credit growth of roughly 50%, which lies beyond the 99th percentile of d_ln_PC in this sample and therefore outside the range in which all but a handful of country–year observations are located. The service relationship is consequently positive and rising over virtually the whole of the observed distribution, and its eventual reversal is identified from too few extreme expansions to support substantive interpretation. The higher-order terms are in any case considerably less stable, since the squared term switches from negative in the baseline model to positive in the extended model and the cubed term loses significance entirely once controls are introduced, so the curvature of the service relationship cannot be characterised with confidence. Interpretation should therefore rest on the sign and significance of the first-order effect rather than on any specific turning point.

The controls clarify the mechanism to some extent. Government consumption enters strongly positively at 0.237, which is expected given that a substantial part of measured service value added consists of publicly provided education, health and administration, whereas trade openness carries a significant negative coefficient of -

0.099, consistent with greater openness shifting the composition of output towards tradable goods production. Gross capital formation is insignificant, reflecting the lower capital intensity of many service activities relative to industry, and neither inflation nor the growth of school enrolment attains significance. The rise in within R-squared from 0.1646 to 0.3459 indicates that the macroeconomic environment explains a considerably larger share of service-sector variation than of industrial variation, which is itself consistent with the view that services respond primarily to domestic demand conditions.

The contrast between the positive service coefficient and the negative industrial coefficient in the same decade is substantive rather than mechanical, since the two sectoral equations employ the same regressor and the same GDP denominator, so any influence transmitted through the scaling would displace both coefficients in a common direction rather than in opposing ones. The most plausible reading is therefore allocative, holding that post-crisis credit expansion was directed disproportionately towards real estate, consumption and service activities rather than towards manufacturing investment, thereby raising the service share while leaving industrial capacity comparatively underfunded. The evidence assembled by Beck et al. (2012) and Bezemer et al. (2016) on the differential real effects of credit according to borrower type lends support to this interpretation, although testing it directly would require disaggregated credit data that the present cross-country sample does not provide. What matters for the industrial and service outcomes documented here may therefore be the destination of lending rather than its volume, which is precisely the question that motivates the sectoral design adopted in this study.

5.1.4. Comparative Assessment of the Periodical Findings

Read together, the three sectors describe a pattern that no aggregate specification could recover. Agriculture shows no stable association between credit growth and sectoral value added in any period, with the single significant curvature reversing sign under controls and resting on a turning point whose confidence interval is extremely wide. Industry displays a significant nonlinearity in every period, although

the shape of that nonlinearity is not the same across regimes. In the pre-crisis and full-period models the marginal effect is mildly favourable at moderate rates of expansion and turns negative beyond a local maximum located at annual credit growth of roughly 24% over the full sample and roughly 67% before the crisis, whereas in the post-crisis decade the marginal effect is negative across almost the entire observed range and becomes favourable again only above approximately 28% growth. Services move in the opposite direction during the post-crisis decade, with a positive and robust first-order effect whose eventual reversal lies beyond the 99th percentile of the credit growth distribution and therefore carries little practical weight. This divergence is the central empirical contribution of the periodical analysis, since it demonstrates that a single aggregate credit coefficient would average across sectoral responses that differ in sign as well as in magnitude.

Locating the estimated turning points within the distribution of credit growth is what gives these coefficients their economic content, and the exercise yields a sharp contrast between the two non-agricultural sectors in the post-crisis decade. For industry the turning region falls between the 95th and 99th percentiles of observed credit growth, which means that the negative segment of the curve governs the experience of the overwhelming majority of country–year observations, whereas for services the corresponding point lies beyond the 99th percentile, leaving the positive segment effectively unchallenged over the sampled range. Rapid credit expansion is therefore associated with a rising service share and a falling industrial share simultaneously, and the asymmetry is not a matter of opposing curvature but of where each sector’s turning point sits relative to the rates of expansion that countries actually experience. This observation also disciplines the interpretation of the higher-order terms more generally, since a statistically significant cubic term whose turning points lie in the extreme tail describes a relationship that is, for practical purposes, monotonic over the relevant range.

The temporal dimension adds a further layer to this account. During the pre-crisis period the industrial relationship was weakly positive under the static estimator and displayed negative curvature only at high rates of expansion under the dynamic estimator, whereas in the post-crisis decade the marginal effect at ordinary rates of

credit growth turned negative and the service effect became positive and significant. The change is consistent with a shift in the composition and destination of credit, since bank lending has for several decades been directed increasingly towards real estate and other asset markets rather than towards non-financial business (Bezemer et al., 2016), a reorientation that the post-crisis policy environment may well have reinforced although the present data cannot establish as much. It is also consistent with the possibility that the relationship between finance and the real economy is regime-dependent, so that estimates obtained from a single pooled sample may conceal offsetting effects, a concern that motivates the sub-period design adopted here.

These findings bear directly on the ‘too much finance’ hypothesis, though in a form that requires care in statement. Arcand et al. (2015), Law and Singh (2014) and Cecchetti and Kharroubi (2012) locate the adverse effects of finance beyond a threshold of financial depth, whereas the nonlinearity identified here arises in the growth of credit, since the credit variables are integrated of order one and enter the estimated equations in first differences. The industrial results are therefore better described as evidence that the pace of credit expansion matters for industrial performance, with faster expansion associated with a weakening industrial position over most of the observed range, than as a direct confirmation of a depth threshold, and in that respect they align with the credit-boom literature (Mendoza & Terrones, 2008; Schularick & Taylor, 2012). Establishing whether the two forms of nonlinearity share a common mechanism would require joint modelling of stocks and flows, and this remains an avenue for further work rather than a claim supported by the present estimates.

The policy implications follow the same sectoral logic that motivates the thesis as a whole. If credit growth raises the service share while depressing the industrial share, then blanket measures to expand or restrain aggregate credit will have consequences for the composition of output that a one-size-fits-all financial policy is unlikely to anticipate, still less to control. What matters for productive transformation is the destination of lending, since the classic argument that financial development supports growth by relaxing external finance constraints applies most forcefully to

manufacturing activities with high dependence on external funds (Rajan & Zingales, 1998), and since the efficiency of allocation rather than the volume of intermediation determines the real return to financial deepening (Levine, 2005). The results reported here accordingly support a more granular and sector-aware approach to credit policy, in which the pace of expansion and the sectoral allocation of lending are monitored jointly rather than subsumed within a single aggregate indicator.

Several limitations qualify these conclusions and should be borne in mind when the sectoral estimates are compared with the country-level evidence presented in the analysis of Türkiye that follows in the next chapter. The dependent variables are shares rather than levels, so the results describe compositional change and cannot by themselves establish whether credit growth raises or lowers absolute sectoral output. The extended specifications rest on samples roughly half the size of the baseline ones, which limits the strength of any inference drawn from changes in significance between the two. Private credit is measured in aggregate, so the allocative interpretation advanced above remains a conjecture consistent with the evidence rather than a tested proposition. Finally, the higher-order polynomial terms are identified from a comparatively small number of extreme observations, which is why the analysis interprets turning points only where they lie within the supported range of the data and treats the direction of curvature as more informative than its precise location.

5.2. Income-Level Analysis

The periodical analysis reported in the preceding section established that the association between credit growth and sectoral value added is not stable over time, shifting appreciably between the pre-crisis and post-crisis windows. A complementary question concerns whether that association also varies across stages of development, since countries at different income levels differ systematically in financial depth, institutional capacity, productive structure and the efficiency with which savings are channelled into investment. To address this question, the empirical models were re-estimated separately by income group over the full 1992–2019 period, using the World Bank's official income classification to allocate countries to

three groups (World Bank, 2026b). The low-income group comprises economies classified as low income, the middle-income group combines the lower-middle and upper-middle categories, and the high-income group consists of economies in the high-income category. Merging the two middle categories preserves a sufficient number of country groups for the dynamic estimator while retaining the substantive distinction between economies still building financial and productive capacity and those that have largely completed that transition.

The rationale for this partition rests on a well-established strand of the finance–growth literature. Rioja and Valev (2004a) argue that the effect of financial development is not uniform but varies across three regions, being uncertain at very low levels of development, large and positive at intermediate levels and positive but smaller at high levels. Demetriades and Law (2006) reach a compatible conclusion through an institutional lens, reporting that financial development exerts its strongest influence in middle-income countries and its weakest influence in low-income ones, where finance unaccompanied by sound institutions may deliver few durable benefits. Aizenman et al. (2015) extend the argument to the sectoral level, documenting substantial sectoral heterogeneity and marked regional differences in the effect of financial depth on sectoral output growth, including negative effects in several sectors. These contributions together imply that pooling economies of very different income levels may average out offsetting effects, motivating the disaggregation adopted here.

The estimation strategy replicates that of the periodical analysis in every respect other than the sample split. Because the CIPS results reported in the methodology chapter indicate that the credit variables are integrated of order one, credit enters all specifications in first differences, so the estimates below describe the association between the pace of credit expansion and sectoral value added rather than between the level of financial depth and that value added. The distinction matters for interpretation, since Sahay et al. (2015) caution that the speed of financial deepening carries risks of its own once it outpaces supervisory and regulatory capacity. For each sector and income group, linear, quadratic and cubic forms were estimated both with fixed effects and Driscoll–Kraay standard errors and with the two-step System GMM

estimator, and the specification reported in the main text is the one in which the highest-order credit term attains conventional significance. Where a quadratic form is selected, the Lind and Mehlum (2010) test is reported immediately beneath the corresponding table, and where a cubic form is selected the stationary points of the fitted curve are computed by the delta method and evaluated against the empirical support of the regressor. Specifications in which no credit term is significant, together with those whose diagnostics fail, are collected in Appendix and referred to rather than reproduced here. The results for the agricultural, industrial and service sectors are presented in sequence below, and within each sector the low-income group is examined first, followed by the middle-income and high-income groups.

5.2.1. The Agricultural Sector

Among the three income groups, only the low-income sample produces a statistically significant credit term in the agricultural equations, and only under the fixed-effects estimator. Table 26 presents the quadratic fixed-effects estimates with Driscoll–Kraay standard errors for low-income countries over 1992–2019.

Table 26. Quadratic Fixed-Effects Estimates with Driscoll–Kraay Standard Errors: Agricultural Value Added, Low-Income Countries, 1992–2019

Variables	Model (1) 1992–2019 Baseline	Model (2) 1992–2019 With Controls
d_ln_ACGDP	0.007 (0.013)	0.009 (0.014)
d_ln_ACGDP²	-0.009** (0.004)	0.003 (0.005)
d_ln_SE	—	-0.160 (0.142)
ln_GCF	—	0.072 (0.097)
ln_GC	—	0.098* (0.056)
INF	—	0.001 (0.003)
ln_TO	—	-0.227 (0.143)
Constant	3.466*** (0.025)	Omitted
Year Fixed Effects	Yes	Yes

Table 26. (continued)

Country Fixed Effects	Yes	Yes
Observations	276	112
Number of Countries	14	9
Within R^2	0.1603	0.3786
F Statistic	789.75***	11435.43***
Driscoll–Kraay Lags	3	3

Notes: Dependent variable: ln_AVASGDP. Fixed-effects estimates with Driscoll–Kraay robust standard errors (maximum lag = 3) in parentheses. d_ln_ACGDP is the first difference of ln_ACGDP and d_ln_ACGDP² its squared term. Model (2) adds d_ln_SE, ln_GCF, ln_GC, INF and ln_TO. Country and year fixed effects are included, and year coefficients are omitted. Model (2) uses a smaller sample (112 observations and 9 countries versus 276 and 14), so coefficient differences may also reflect sample composition. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

In the baseline specification the linear term on agricultural credit growth is small and statistically indistinguishable from zero, whereas the squared term is negative and significant at the 5% level, indicating that the relationship between agricultural credit growth and agricultural value added is concave rather than proportional. The within R-squared of 0.16 is modest but not unusual for a first-differenced regressor in a sector whose output responds to weather, prices and policy as much as to finance.

Once the control set is added in Model (2), the squared term changes sign and loses significance entirely, and government consumption is the only control to attain even marginal significance, suggesting that the estimated curvature is fragile. The estimation sample also contracts from 276 observations across 14 countries to 112 observations across 9 countries, so the difference between the two columns may reflect the composition of the surviving panel as much as the influence of the controls.

Table 27 reports the Lind and Mehlum (2010) test applied to the baseline specification.

Table 27. Lind and Mehlum (2010) Inverted U-Shape Test: Agricultural Value Added, Low-Income Countries, 1992–2019

Specification: $f(x) = x^2$ Estimated Turning Point: 0.414

	Lower Bound	**Upper Bound**
Interval	-5.662	3.646
Slope	0.109	-0.058
t-statistic	3.084	-1.507
p-value	0.002***	0.072*

Overall inverse U-shape test: t = 1.507; p = 0.072*. 80% Fieller interval for the turning point: [-0.400, 2.739].

Notes: Lind and Mehlum (2010) test for an inverted U-shaped relationship in the quadratic fixed-effects specification with Driscoll–Kraay standard errors for the low-income sample, 1992–2019. H_0 is a monotonic or U-shaped relationship and H_1 an inverted U-shape. The turning point is computed as $-\beta_1/(2\beta_2)$. The overall p-value of 0.072 rejects H_0 at the 10% level but not at the 5% level. The Fieller confidence interval is constructed at the 80% level. *** p < 0.01, ** p < 0.05, * p < 0.10.

The test locates the turning point at 0.414, well inside the observed interval, with a positive and highly significant slope at the lower bound and a negative slope at the upper bound that reaches significance only at the 10% level. The overall statistic therefore rejects monotonicity at the 10% level but not at the 5% level, and the 80% Fieller interval of -0.400 to 2.739 is wide enough to accommodate almost any location in the upper part of the distribution. Because the estimated maximum lies far above the median rate of agricultural credit growth in this group, the great majority of low-income observations sit on the upward-sloping segment of the fitted curve, implying that faster agricultural credit growth tends to be associated with higher agricultural value added across the range in which these economies actually operate. This reading accords with the assessment that financial constraints in smallholder agriculture originate in the absence of collateralisable assets and in the reluctance to pledge assets that are vital to livelihoods, leaving substantial gaps in rural financial services that institutional innovation has not yet closed (World Bank, 2007). The weakness of the upper-bound slope and the collapse of the quadratic term under controls nonetheless counsel against treating the concavity as an established property of the data rather than as a tentative indication.

Neither the middle-income nor the high-income agricultural equations yield a significant credit coefficient in any functional form or under either estimator, and the corresponding tables are reported in Appendix. The System GMM results for the low-income sample are relegated there as well, since the Hansen statistic equals 1.000 in all three specifications while the instrument count of 34 to 40 exceeds the 14 available country groups, a configuration that Roodman (2009) identifies as instrument proliferation weakening the test of joint instrument validity rather than confirming it. Fixed-effects estimation consequently supplies the only interpretable evidence for agriculture. The absence of any measurable association in richer economies is unsurprising, given that agriculture accounts for a small and declining share of output there, so that variation in the pace of agricultural credit growth may simply be too narrow a margin to register in aggregate sectoral value added. One qualification applies to the whole of this subsection. The agricultural credit series is available for considerably fewer countries than the private credit series used in the industrial and service equations, with 14 low-income and 59 middle-income groups against 23 and 99 respectively, so the agricultural results rest on a narrower empirical base than those reported below.

5.2.2. The Industrial Sector

The industrial equations produce significant credit terms in all three income groups under fixed-effects estimation, although the selected functional form differs across them. Table 28 reports the cubic fixed-effects estimates with Driscoll–Kraay standard errors for low-income countries.

Table 28. Cubic Fixed-Effects Estimates with Driscoll–Kraay Standard Errors: Industrial Value Added, Low-Income Countries, 1992–2019

Variables	Model (1) Cubic Baseline	Model (2) Cubic with Controls
d_ln_PC	-0.095 (0.077)	-0.118 (0.157)
d_ln_PC²	0.108 (0.077)	0.764 (0.769)
d_ln_PC³	0.287** (0.123)	-0.855 (1.208)
d_ln_SE	—	-0.247 (0.245)

Table 28. (continued)

ln_GCF	—	-0.202*** (0.062)
ln_GC	—	0.353*** (0.102)
INF	—	-0.004 (0.003)
ln_TO	—	0.225 (0.139)
Year Fixed Effects	Yes	Yes
Constant	2.821*** (0.013)	1.574*** (0.403)
Observations	521	191
Countries	23	16
Within R-squared	0.1312	0.5085
F Statistic	4097.65***	1417.11***
Prob > F	0.000	0.000
Driscoll–Kraay Lag	3	3

Notes: Dependent variable: ln_IVASGDP. Fixed-effects estimates with Driscoll–Kraay robust standard errors (maximum lag = 3) in parentheses. d_ln_PC is the first difference of ln_PC, and d_ln_PC² and d_ln_PC³ its squared and cubed terms. Model (2) adds d_ln_SE, ln_GCF, ln_GC, INF and ln_TO. Country and year fixed effects are included, and year coefficients are omitted. Model (2) uses a smaller sample (191 observations and 16 countries versus 521 and 23), so coefficient differences may also reflect sample composition. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

The cubic term is positive and significant at the 5% level while the linear and quadratic terms are individually insignificant, a pattern characteristic of polynomial specifications in which the curvature rather than the average slope carries the information. Joint tests support this reading. The hypothesis that the cubic term is zero is rejected at the 3% level, the quadratic and cubic terms jointly at the 1.4% level and all three credit terms jointly at the 2.9% level, indicating that private credit growth enters the industrial equation in a non-monotonic form even though no single coefficient dominates. Adding controls in Model (2) removes all significance from the credit terms and reverses the sign of the cubic coefficient, while gross capital

formation enters negatively and government consumption positively, both at the 1% level. The sample falls from 521 observations across 23 countries to 191 observations across 16 countries, which again makes the two columns imperfectly comparable.

Because a cubic specification was selected, the stationary points of the fitted curve were computed by the delta method and evaluated against the empirical support of the regressor. The two roots lie at -0.480 and 0.230, with 90% confidence intervals of -0.719 to -0.241 and 0.012 to 0.449 respectively, and both fall comfortably inside the estimation-sample range of private credit growth, which extends from -1.089 to 1.018. With a positive cubic coefficient the fitted relationship rises up to the first stationary point, declines between the two and rises again beyond the second. The interquartile range of private credit growth in this group runs from -0.067 to 0.133 around a median of 0.041, so the overwhelming majority of low-income observations lie on the declining middle segment, implying that faster private credit growth tends to accompany a lower industrial value-added share across the range in which these economies typically operate. The upward segment beyond 0.230 lies above the 90th percentile of the distribution and therefore rests on comparatively few observations. Notably, the discriminant of the first-order condition is 0.094 with a 90% confidence interval running from -0.059 to 0.247, so the hypothesis that the derivative has no real roots cannot be rejected, and the two stationary points are best read as the most likely location of curvature rather than as firmly identified thresholds.

The System GMM results for low-income industry are reported in Appendix and carry no interpretive weight. The credit terms were dropped altogether in the quadratic and cubic specifications, the Hansen statistic equals 1.000 in all three models, and the instrument count of 35 to 41 again exceeds the 23 country groups, leaving the overidentification tests uninformative in the sense described by Roodman (2009).

Table 29 turns to the middle-income group, which constitutes the largest of the three samples with 99 countries in the baseline specification.

Table 29. Cubic Fixed-Effects Estimates with Driscoll–Kraay Standard Errors: Industrial Value Added, Middle-Income Countries, 1992–2019

Variables	Model (1) Cubic Baseline	Model (2) Cubic with Controls
d_ln_PC	0.051* (0.027)	-0.051** (0.023)
d_ln_PC²	-0.060 (0.046)	-0.053 (0.065)
d_ln_PC³	-0.049* (0.026)	-0.024 (0.031)
d_ln_SE	—	-0.128** (0.052)
ln_GCF	—	0.221*** (0.048)
ln_GC	—	-0.162** (0.070)
INF	—	0.000062*** (0.000017)
ln_TO	—	0.058 (0.050)
Year Fixed Effects	Yes	Yes
Constant	3.221*** (0.005)	Omitted
Observations	2,417	1,141
Countries	99	82
Within R-squared	0.0294	0.1820
F Statistic	1026638.50***	81751.35***
Prob > F	0.000	0.000
Driscoll–Kraay Lag	3	3

Notes: Dependent variable: ln_IVASGDP. Fixed-effects estimates with Driscoll–Kraay robust standard errors (maximum lag = 3) in parentheses. d_ln_PC is the first difference of ln_PC, and d_ln_PC² and d_ln_PC³ its squared and cubed terms. Model (2) adds d_ln_SE, ln_GCF, ln_GC, INF and ln_TO. Country and year fixed effects are included, and year coefficients are omitted. The constant in Model (2) was omitted because of collinearity. Model (2) uses a smaller sample (1,141 observations and 82 countries versus 2,417 and 99), so coefficient differences may also reflect sample composition. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Here the linear term is positive and the cubic term negative, both significant at the 10% level, while the quadratic term is insignificant. The joint test of all three credit terms is rejected at the 4.5% level, whereas the quadratic and cubic terms jointly are not, indicating that the evidence for curvature is weaker than in the low-income

sample. The addition of controls produces a substantive change rather than a mere loss of precision. The linear credit term reverses sign and becomes negative at the 5% level, the cubic term loses significance, and gross capital formation, government consumption, secondary enrolment growth and inflation all enter significantly. This reversal deserves emphasis, since it suggests that a substantial part of the positive baseline association may operate through channels captured by the control set rather than through credit growth itself.

The stationary points implied by the baseline cubic lie at 0.309 and -1.130, with 90% confidence intervals of -0.098 to 0.715 and -1.501 to -0.760, and both fall within the observed range of private credit growth, which spans -2.249 to 1.367 in this group. The negative cubic coefficient reverses the ordering found in the low-income sample, so the fitted curve declines below the lower stationary point, rises between the two and declines again beyond the upper one. Since the lower stationary point falls below the first percentile of the distribution, the effective working range of the data lies almost entirely on the rising segment, with the descent beginning only above 0.309 – a value close to the 95th percentile of 0.286. The fitted relationship in middle-income countries is therefore approximately concave over the region in which the data are dense, with the marginal contribution of credit growth turning negative in the upper tail. As in the low-income case, the discriminant is not statistically distinguishable from zero, its 90% interval running from -0.003 to 0.026, so the exact location of the maximum should be treated with caution. The middle-income sample is also the only one in which the dynamic estimator yields a significant and diagnostically acceptable credit term, allowing the static and dynamic results to be compared directly. Table 30 reports the two-step System GMM estimates for this group.

Table 30. Quadratic Two-Step System GMM Estimates: Industrial Value Added, Middle-Income Countries, 1992–2019

Variables	Quadratic Baseline	Quadratic + Controls
L.ln_IVASGDP	0.960*** (0.058)	0.942*** (0.270)
d_ln_PC	0.014 (0.122)	-0.431 (0.335)
d_ln_PC²	-0.268** (0.102)	-0.108 (0.251)

Table 30. (continued)

d_ln_SE	—	0.148 (0.140)
ln_GCF	—	0.085 (0.068)
ln_GC	—	-0.015 (0.075)
INF	—	-0.000191* (0.000103)
ln_TO	—	0.010 (0.032)
Year Fixed Effects	Yes	Yes
Constant	0.148 (0.189)	-0.064 (0.812)
Observations	2,401	1,138
Countries	99	82
Instruments	38	43
AR(1) (p-value)	0.000	0.039
AR(2) (p-value)	0.478	0.719
Hansen Test (p-value)	0.235	0.529
Sargan Test (p-value)	0.000	0.810
Difference-in-Hansen (p-value)	0.136	0.277
Wald/F Statistic	77101.43***	8456.13***

Notes: Dependent variable: ln_IVASGDP. Two-step System GMM estimates with Windmeijer (2005) finite-sample-corrected robust standard errors in parentheses. d_ln_PC² is the squared term of d_ln_PC. Model (2) adds d_ln_SE, ln_GCF, ln_GC, INF and ln_TO. Year fixed effects are included, and their coefficients are omitted. The lagged dependent variable and the credit terms enter as collapsed GMM-style instruments at lags 2–3. The controls enter as IV-style instruments. Diagnostic entries other than Wald/F are p-values. Stata used a generalized inverse for a singular two-step moment covariance matrix, so Difference-in-Hansen results require caution. Samples differ across columns (2,401 observations and 99 countries versus 1,138 and 82), so coefficient differences may also reflect sample composition. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

The lagged dependent variable enters with a coefficient of 0.960, confirming the strong persistence of the industrial value-added share, and the squared credit term is negative and significant at the 5% level while the linear term is not. The diagnostics

of the baseline model are satisfactory in the respects that matter most for the validity of the estimator. AR(2) does not reject the absence of second-order serial correlation, neither the Hansen nor the Difference-in-Hansen test rejects the overidentifying restrictions, and the 38 instruments remain well below the 99 country groups, so the instrument proliferation problem that undermines the low-income models does not arise here. The Sargan test rejects, but that statistic is not robust to heteroskedasticity and is reported for completeness alongside its robust counterpart. As in the fixed-effects results, the credit terms lose significance once the full control set is introduced and the sample contracts to 1,138 observations across 82 countries. Table 31 reports the Lind and Mehlum (2010) test applied to the baseline quadratic specification.

Table 31. Lind and Mehlum (2010) Inverted U-Shape Test: Industrial Value Added, Middle-Income Countries, 1992–2019

Specification: $f(x) = x^2$ Estimated Turning Point: 0.026

	Lower Bound	**Upper Bound**
Interval	-2.249	1.367
Slope	1.220	-0.719
t-statistic	2.569	-2.373
p-value	0.006***	0.010***

Overall inverse U-shape test: t = 2.373; p = 0.010***. 80% Fieller interval for the turning point: [-0.304, 0.370].

Notes: Lind and Mehlum (2010) test for an inverted U-shaped relationship in the quadratic System GMM specification for the industrial sector in middle-income countries, 1992–2019. H_0 is a monotonic or U-shaped relationship and H_1 an inverted U-shape. The turning point, computed as $-\beta_1/(2\beta_2)$, lies within the observed interval, where the positive slope at the lower bound and the negative slope at the upper bound are both significant. The overall p-value of 0.00981 rejects H_0 at the 1% level. The Fieller confidence interval is constructed at the 80% level. Stars are based on unrounded p-values: *** p < 0.01, ** p < 0.05, * p < 0.10.

The test rejects the null of a monotonic or U-shaped relationship at the 1% level, with a positive and significant slope at the lower bound, a negative and significant slope at the upper bound and an estimated maximum at 0.026. That value sits almost exactly at the median rate of private credit growth in middle-income countries,

implying that roughly half of the observations in this group lie beyond the point at which the marginal association turns negative. The two estimators consequently agree on the shape of the relationship in the upper range while disagreeing on where the maximum falls, the fixed-effects cubic placing it at 0.309 and the dynamic quadratic at 0.026. The 80% Fieller interval of -0.304 to 0.370 nonetheless encompasses the fixed-effects estimate, so the disagreement concerns the precision with which the threshold can be located rather than the existence of the concavity itself.

These middle-income results provide the clearest support in this study for the ‘too much finance’ proposition, and they do so in the sector for which that proposition has the strongest theoretical rationale. Arcand et al. (2015) report that financial depth begins to exert a negative influence on output growth once credit to the private sector approaches the size of the economy, while Law and Singh (2014) identify a comparable threshold with a dynamic panel threshold estimator and conclude that more finance is not invariably beneficial. Cecchetti and Kharroubi (2015) supply a mechanism that speaks directly to industry, arguing that rapid financial expansion disproportionately benefits projects that are easy to pledge as collateral but less productive, and reporting that financial growth harms industries with less tangible assets or higher research intensity most severely. This mechanism may be particularly relevant to middle-income economies, in which manufacturing remains a substantial share of output while supervisory frameworks are still maturing and where supervisory capacity may not expand as rapidly as credit, consistent with the broader risk identified by Sahay et al. (2015). Table 32 presents the corresponding estimates for high-income countries, where a quadratic form is the only specification in which the credit terms attain significance.

Table 32. Quadratic Fixed-Effects Estimates with Driscoll–Kraay Standard Errors: Industrial Value Added, High-Income Countries, 1992–2019

Variables	Model (1) Quadratic Baseline	Model (2) Quadratic with Controls
d_ln_PC	-0.049 (0.039)	-0.222*** (0.069)
d_ln_PC²	-0.110*** (0.032)	-0.150 (0.129)

Table 32. (continued)

d_ln_SE	—	-0.166 (0.122)
ln_GCF	—	0.261*** (0.052)
ln_GC	—	0.092 (0.136)
INF	—	0.000423 (0.001535)
ln_TO	—	-0.019 (0.084)
Year Fixed Effects	Yes	Yes
Constant	3.341*** (0.004)	Omitted
Observations	1,467	852
Countries	62	57
Within R-squared	0.1896	0.3583
F Statistic	1831281.19***	297731.32***
Prob > F	0.000	0.000
Driscoll–Kraay Lag	3	3
Lind–Mehlum U-test (p-value)	0.005***	—

Notes: Dependent variable: ln_IVASGDP. Fixed-effects estimates with Driscoll–Kraay robust standard errors (maximum lag = 3) in parentheses. d_ln_PC is the first difference of ln_PC and d_ln_PC² its squared term. Model (2) adds d_ln_SE, ln_GCF, ln_GC, INF and ln_TO. Country and year fixed effects are included, and year coefficients are omitted. The constant in Model (2) was omitted because of collinearity. For Model (1), the Lind–Mehlum test rejects the null of a monotonic or U-shaped relationship ($t = 2.732$, $p = 0.00548$) and supports an inverted U-shape with a turning point of -0.224 within the observed interval, whereas the squared term is insignificant in Model (2), so the nonlinearity is not robust to the controls. Model (2) also uses a smaller sample (852 observations and 57 countries versus 1,467 and 62), so coefficient differences may also reflect sample composition. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

The squared credit term is negative and significant at the 1% level in the baseline model, while the linear term is negative but insignificant. In Model (2) the pattern

changes materially, as the linear term becomes negative and significant at the 1% level, the squared term loses significance, and gross capital formation is the only control to enter significantly. Read together, the two columns suggest that the relationship in high-income countries is better described as broadly negative than as genuinely hump-shaped, since the curvature does not survive the inclusion of controls whereas the negative level effect strengthens. Table 33 reports the Lind and Mehlum (2010) test for the baseline specification.

Table 33. Lind and Mehlum (2010) Inverted U-Shape Test: Industrial Value Added, High-Income Countries, 1992–2019

Specification: $f(x) = x^2$ Estimated Turning Point: -0.224

	Lower Bound	**Upper Bound**
Interval	-1.781	0.835
Slope	0.342	-0.233
t-statistic	3.701	-2.732
p-value	<0.001***	0.005***

Overall inverse U-shape test: t = 2.732; p = 0.005***. 80% Fieller interval for the turning point: [-0.391, 0.009].

Notes: Lind and Mehlum (2010) test for an inverted U-shaped relationship in the quadratic fixed-effects specification with Driscoll–Kraay standard errors for the industrial sector in high-income countries, 1992–2019. H_0 is a monotonic or U-shaped relationship and H_1 an inverted U-shape. The turning point, computed as $-\beta_1/(2\beta_2)$, lies within the observed interval [-1.781, 0.835], where the slope is positive and significant at the lower bound and negative and significant at the upper bound. The overall p-value of 0.00548 rejects H_0 at the 1% level. The Fieller confidence interval is constructed at the 80% level. and remains within the observed range. Although it includes zero, that concerns uncertainty about the location of the turning point rather than the slope-based test. Stars are based on unrounded p-values: *** p < 0.01, ** p < 0.05, * p < 0.10.

The test rejects the null at the 1% level, with a positive and strongly significant slope at the lower bound and a negative and significant slope at the upper bound, placing the turning point at -0.224 within the observed interval of -1.781 to 0.835. The location of that maximum is the substantively important feature of the result. Because it lies below zero, essentially the whole range of positive private credit growth falls on the declining branch of the fitted curve, implying that in high-income

economies faster credit growth tends to be associated with a lower industrial value-added share across nearly every value the data actually take. The 80% Fieller interval of -0.391 to 0.009 includes zero, yet this reflects uncertainty about the precise position of the maximum rather than doubt about the slope-based test, which the interval does not overturn.

Several strands of the literature help to rationalise this pattern. Rodrik (2016) documents that the hump-shaped relationship between industrialisation and income has shifted downwards and closer to the origin, so that countries run out of industrialisation opportunities sooner and at lower income levels, a structural trend against which any positive credit effect would have to work. The composition of credit is likely to matter as well. Beck et al. (2012) show that enterprise credit is positively associated with economic growth whereas household credit is not, and that the household share of credit is higher in more urbanised economies and in those with smaller manufacturing sectors, while Mian et al. (2017) find that increases in the household-debt ratio predict lower subsequent output growth in a panel of thirty mostly advanced economies. Since the household-credit share tends to increase with income and is higher in economies with smaller manufacturing sectors, aggregate private credit growth in high-income countries need not primarily reflect lending to industrial firms, which may account for the negative association estimated over the empirically relevant range. Rousseau and Wachtel (2011) offer a complementary reading in documenting the weakening or apparent disappearance of the effect of financial deepening on growth in more recent samples. The System GMM estimates for high-income industry appear in Appendix and are not treated as benchmark evidence, since the null of no second-order serial correlation is rejected at the 10% level in all three specifications, with AR(2) p-values of 0.047, 0.072 and 0.068.

5.2.3. The Service Sector

The service equations are the least responsive of the three sectors to variation in credit growth. Significance is confined to the middle-income group and to the linear fixed-effects specification, with neither quadratic nor cubic terms reaching conventional levels in any income group. Table 34 reports these estimates.

Table 34. Linear Fixed-Effects Estimates with Driscoll–Kraay Standard Errors: Services Value Added, Middle-Income Countries, 1992–2019

Variables	Model (1) Linear Baseline	Model (2) Linear with Controls
d_ln_PC	0.034* (0.019)	0.025 (0.019)
d_ln_SE	—	0.024 (0.064)
ln_GCF	—	0.056*** (0.011)
ln_GC	—	0.051** (0.019)
INF	—	-0.00005*** (0.00001)
ln_TO	—	-0.082*** (0.021)
Year Fixed Effects	Yes	Yes
Constant	3.843*** (0.003)	3.856*** (0.112)
Observations	2,361	1,127
Countries	99	81
Within R-squared	0.2012	0.3746
F Statistic	485974.12***	30391.10***
Prob > F	0.000	0.000
Driscoll–Kraay Lag	3	3

Notes: Dependent variable: ln_SVASGDP. Fixed-effects estimates with Driscoll–Kraay robust standard errors (maximum lag = 3) in parentheses. d_ln_PC is the first difference of ln_PC. Model (2) adds d_ln_SE, ln_GCF, ln_GC, INF and ln_TO. Country and year fixed effects are included, and year coefficients are omitted. Model (2) uses a smaller sample (1,127 observations and 81 countries versus 2,361 and 99), so coefficient differences may also reflect sample composition. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

The credit coefficient of 0.034 is positive and significant at the 10% level in the baseline model, indicating a modest association between faster private credit growth and a higher service value-added share. Once controls are added, the coefficient falls to 0.025 and loses significance, while gross capital formation and government consumption enter positively, inflation and trade openness negatively, lifting the within R-squared from 0.20 to 0.37. The evidence for a credit effect in services is therefore weak and not robust to conditioning, and nothing in these estimates suggests a threshold beyond which further credit growth becomes detrimental.

For low- and high-income countries no credit term is significant in any functional form under either estimator, and the relevant tables are reported in Appendix. The

System GMM results add little in these groups. In the low-income sample every credit term was dropped during estimation, the Hansen statistic equals 1.000 and the instruments outnumber the country groups, while in the high-income sample the Sargan test rejects in all three specifications and the AR(2) statistic is borderline in the cubic model. The general absence of nonlinearity in services may reflect the heterogeneity of the aggregate itself, which combines trade, transport, public administration and financial intermediation within a single value-added measure. Financial intermediation is a component of that aggregate, so an expansion of the financial sector raises measured service output directly even where it draws skilled resources away from other activities. This interpretation extends the resource-allocation mechanism developed by Cecchetti and Kharroubi (2015), although the resulting composition effect within aggregate service value added is an inference of the present analysis rather than a result directly tested in their study. Such a composition effect may offset any diminishing returns operating within non-financial services, producing the flat and largely insignificant profile observed here.

5.2.4. Comparative Assessment of the Income-Level Findings

Reading the three sectors together, a coherent developmental pattern emerges from the income-level estimates. Agricultural credit growth registers only in low-income countries, where agriculture remains a substantial share of output and where credit constraints on producers are most binding, whereas private credit growth registers in industry across all three income groups. Services show little response at any income level. Within industry the estimated shape shifts as income rises, from an N-shaped profile whose dominant middle segment slopes downward in low-income countries, through an approximately concave profile whose maximum sits in the upper tail in middle-income countries, to a relationship that is negative over almost the entire observed range in high-income countries. The comparison is not a simple ordering, since the low-income profile is N-shaped rather than concave and its two stationary points do not correspond to the single maximum estimated elsewhere. What holds consistently across the three groups is that the interval over which faster credit growth accompanies industrial expansion is narrowest in high-income countries, where the estimated maximum of -0.224 lies below zero and the whole positive range of credit growth falls on the declining branch.

This graded pattern is broadly consistent with the three-region view of Rioja and Valev (2004b) and with the finding of Demetriades and Law (2006) that financial development is most potent at middle income levels and weakest at low ones, although the present results locate the strongest evidence of harm rather than of benefit in the middle-income group. The divergence is attributable in part to the choice of regressor, since the analysis here concerns the growth rate of credit rather than its level, a distinction that Sahay et al. (2015) treat as consequential for both growth and stability. Aizenman et al. (2015) provide the closest sectoral parallel, reporting substantial sectoral heterogeneity alongside marked regional differences in the effect of financial depth on output growth and confirming a negative influence in several sectors. The sectoral contrast documented here – concavity concentrated in industry, absent in services and only weakly present in agriculture – is what an aggregate specification would obscure, and it reinforces the argument advanced throughout this study that the composition of credit matters at least as much as its quantity.

Two features of these results carry implications for how nonlinear finance–growth models should be specified. First, the selected functional form varies not only across sectors but also across income groups within the same sector, so imposing a single quadratic specification on a pooled panel would misrepresent at least two of the three groups. Second, the highest-order credit term loses significance under the full control set in every specification reported above, and in three of them a credit coefficient also reverses sign, indicating that the estimated curvature is sensitive to conditioning information and to the accompanying reduction in sample size. Both features caution against reading any single aggregate threshold as a general property of the finance–growth relationship, and they support the disaggregated design adopted throughout this study.

A final qualification concerns the interaction between the two dimensions of heterogeneity examined in this chapter. The income-level models pool the entire 1992–2019 period, so they cannot separate developmental differences from the temporal shifts documented in the periodical analysis, and the middle-income concavity in particular may partly reflect the credit contractions that followed the

global financial crisis in emerging markets. Estimating the two dimensions jointly would require a panel considerably longer or wider than the one available here, and that question is noted as an avenue for further work rather than resolved in this section.

CHAPTER 6

SECTORAL DATA FOR THE TÜRKİYE ANALYSIS

6.1. Data Sources and Sample Period

The sectoral credit data employed in this chapter are drawn from the Banking Regulation and Supervision Agency (BRSA, 2026), covering total cash loans extended by deposit, participation, development and investment banks operating in the Turkish banking sector, including their foreign branches. This coverage is what makes a genuinely sector-specific analysis of credit volumes possible, since it records the credit allocated to each branch of economic activity rather than an economy-wide aggregate. The emphasis on cash loans is deliberate, as these balances represent credit actually extended to the real sector, thereby permitting comparisons of financing conditions across activities that differ markedly in their reliance on bank intermediation.

Sectoral output is measured using gross domestic product data published by the Turkish Statistical Institute (TurkStat, 2026). TurkStat's GDP figures, compiled through the production approach, express value added as the difference between gross output and intermediate consumption for resident producer units, thereby isolating the contribution each activity makes to aggregate production. Combining the sectoral credit series with the sectoral value-added series allows the link between banking-system financing and productive performance to be examined at a level of granularity that aggregate credit-to-GDP measures cannot reach.

The analysis covers the period from the first quarter of 2005 to the first quarter of 2026, using quarterly observations that track both the allocation of credit and the evolution of production across activities. Quarterly frequency is well suited to the

research question, since it is fine enough to capture the short-run adjustment of credit and output while remaining consistent with the publication calendar of the national accounts. This span also encompasses several distinct macroeconomic episodes in Türkiye, allowing sectoral credit and production to be observed under widely varying financing conditions.

6.2. Sectoral Credit Data: Definition, Coverage and Frequency Conversion

The total cash loans variable used to proxy financial development comprises both performing loans and non-performing receivables. Cash loans, in the narrow sense, denote funds extended by banks to natural and legal persons that remain classified as performing receivables, recorded as credit balances in accounts 100 to 161 of the Uniform Chart of Accounts. These balances therefore constitute the core of the credit stock channelled through the banking system to firms and households, providing the most direct available measure of bank financing at the sectoral level.

The second component of the variable, non-performing receivables, consists of exposures reclassified because of a deterioration in repayment performance. Such receivables become non-performing under the Regulation on the Procedures and Principles for the Classification of Loans and Provisions to Be Set Aside for Such Loans, and are recorded in accounts 170 to 177 of the Uniform Chart of Accounts. Including them means that the credit measure captures not only currently performing exposures but also balances transferred to follow-up accounts, giving a fuller picture of the total credit the banking sector has committed to each activity. This broader definition is preferable for the present purpose, since excluding impaired exposures would understate credit precisely in those sectors and periods where financial stress is most acute.

The BRSA publishes credit data at monthly frequency for the period from January 2005 to April 2026, with each release referring to the preceding month. These figures are stock rather than flow measures, so each observation records the total outstanding loan balance on a given reporting date instead of the volume of new lending extended during the month. Recognising this distinction is essential, because the

stock character of the series determines how it should be converted to the quarterly frequency at which output is observed. Since the national accounts are published quarterly, the monthly credit series had to be aligned with the production data before estimation. Given the stock nature of the credit variable, the observation for the final month of each quarter was taken as that quarter's credit stock, rather than averaging the three monthly values. March observations therefore represent the first quarter, June the second, September the third and December the fourth, yielding a quarterly series that records the loan balance outstanding in each branch of activity at the close of the period. This conversion rule is consistent with the accounting logic of the underlying data, since averaging a stock over three months would produce a quantity with no clear balance-sheet interpretation. End-of-quarter values instead preserve the level of the credit stock at the point that corresponds to the close of the production period, making the two series conceptually comparable. The procedure also removes the frequency mismatch between the monthly credit data and the quarterly output data, allowing all variables to be observed within a single time frame.

The sectoral breakdown follows the BRSA's classification of credit by branch of economic activity, organised according to the functional groups of banks operating in the Turkish banking sector and based on the NACE Rev. 2.1 activity codes (BRSA, 2026; Eurostat, 2025). NACE Rev. 2.1, the first update of NACE Rev. 2, was established in 2023 and has been applied progressively across statistical domains since 2025, so the supervisory credit breakdown follows the updated classification while the national accounts aggregates matched to it in Table 36 are reported on the NACE Rev. 2 sections set out in Table 35. Credit amounts are expressed in thousand Turkish liras throughout. The BRSA series used here are not seasonally adjusted, so the credit variables entering the analysis are the raw stock values published for each reporting period.

6.3. Sectoral Value-Added Data

The dependent variable is constructed from the gross domestic product statistics compiled by TurkStat at current prices, meaning that output is valued at the market prices prevailing in the period of measurement. The series therefore expresses the

nominal size of economic activity and the current monetary value of the value added generated in each branch, and is likewise reported in thousand Turkish liras. Measuring credit and output in the same nominal units keeps the two series on a comparable basis, which matters in an economy where price dynamics have varied substantially over the sample.

TurkStat's national accounts are released on a quarterly basis, with the first quarter covering January to March, the second April to June, the third July to September and the fourth October to December. This structure permits economic activity to be followed at three-month intervals within the year, and it is the reason the monthly credit series had to be aggregated to the same periodicity, as described above. Output is measured through the production approach, which quantifies the value added created by resident producer units within a given period by deducting intermediate consumption from the gross value of goods and services produced. Because this method assigns value added directly to the branch in which production takes place, it provides the appropriate basis for assessing each sector's contribution to total output and for matching that contribution against the credit extended to the same branch.

TurkStat's production-based series are compiled in accordance with the United Nations System of National Accounts (United Nations et al., 2009) and the European System of Accounts (European Parliament & Council of the European Union, 2013), which secures international comparability of the resulting aggregates. Within this framework, economic activities are partitioned according to NACE Rev. 2, the statistical classification of economic activities in the European Community (Eurostat, 2008), enabling value added to be tracked consistently across agriculture, industry and services as well as their sub-branches. Table 35 reports the NACE Rev. 2 sections used in TurkStat's reporting.

Table 35. NACE Rev. 2 Economic Activity Sections Used by TurkStat

Code	Economic activity section
A	Agriculture, Forestry and Fishing
B	Mining and Quarrying

Table 35. (continued)

C	Manufacturing
D	Electricity, Gas, Steam and Air Conditioning Supply
E	Water Supply; Sewerage, Waste Management and Remediation Activities
F	Construction
G	Wholesale and Retail Trade; Repair of Motor Vehicles and Motorcycles
H	Transportation and Storage
I	Accommodation and Food Service Activities
J	Information and Communication
K	Financial and Insurance Activities
L	Real Estate Activities
M	Professional, Scientific and Technical Activities
N	Administrative and Support Service Activities
O	Public Administration and Defence; Compulsory Social Security
P	Education
Q	Human Health and Social Work Activities
R	Arts, Entertainment and Recreation
S	Other Service Activities
T	Activities of Households as Employers

Notes: NACE Rev. 2 is the statistical classification of economic activities used for sectoral reporting in the European Statistical System (Eurostat, 2008). The table reproduces the TurkStat economic activity sections used in constructing the data.

No seasonal adjustment has been applied to the output data used in this study, so the analysis relies on the raw series as published by TurkStat. This choice mirrors the treatment of the credit variable, which likewise enters as an unadjusted stock value, and it therefore keeps the two sides of the sectoral matching on identical processing

terms. Both credit and production variables are consequently analysed in their originally published form, with no adjustment applied to one series that has not been applied to the other.

6.4. Concordance between Sectoral Credit and Value-Added Data

Although the output and credit data both follow the NACE Rev. 2 framework, their sectoral detail is not identical. TurkStat's national accounts aggregate activities into groupings dictated by the production approach and by the level at which value added can be reliably estimated, whereas the BRSA credit statistics disaggregate lending into a considerably finer set of borrower categories. Sharing a classification system therefore does not guarantee that sector definitions coincide, and a matching procedure is required before the two sources can be combined.

The divergence is most pronounced within industrial activities. In the national accounts, several branches are reported jointly within a broad industry aggregate, while manufacturing can also be followed separately, so the same activity may appear at two different levels of aggregation. To ensure that each sector used in the analysis represents an identical scope of activity on both the credit and the output side, an additional adjustment was applied to the TurkStat data: the C 'Manufacturing' series was subtracted from the B C D E 'Manufacturing, Mining and Quarrying and Other Industries' series, producing a B D E 'Industry excluding Manufacturing' series. This constructed series covers mining and quarrying, electricity, gas, steam and air conditioning supply, together with water supply, sewerage, waste management and remediation activities, and it can therefore be analysed as a distinct industrial segment.

This step is essential for the internal consistency of the sectoral matching. Had the B C D E aggregate been used directly, manufacturing would have remained embedded within it, so that the same activity would have been counted twice once the C series was also treated as a separate sector. Subtracting C from the aggregate removes the overlap and confines credit and production variables to mutually exclusive activity boundaries, thereby allowing manufacturing and the remaining industrial activities to enter the empirical model as genuinely distinct sectors.

Matching for the remaining sectors proceeded from the economic content of each BRSA credit category, with every credit item or sub-item assigned to the NACE Rev. 2 activity group under which the corresponding value added is reported. In some cases several credit sub-items mapped onto a single output sector, while in others sub-sectors monitored separately by the supervisor were absorbed into broader national-accounts groupings, reflecting the different reporting priorities of a prudential authority and a statistical office. Proceeding item by item in this way minimises the inconsistencies that would otherwise arise from the unequal granularity of the two sources.

The procedure materially improves the comparability of the credit and production variables in the estimations that follow. If a credit variable referred to one branch of activity while the corresponding output variable referred to a broader or partially overlapping sector, any estimated coefficient would confound the financing of one activity with the production of another, making interpretation unreliable. The matching stage is therefore not a presentational convenience but a precondition for identifying the relationship between financial development and sectoral production within coherent activity boundaries.

Table 36 sets out the correspondence between the sectoral total cash loan categories published by the BRSA and the production-based output classifications of TurkStat, indicating which activity group each credit category and sub-category was assigned to. This mapping is what permits the credit and output series to be integrated into a single sectoral framework, and it yields the eleven matched sectors that constitute the cross-sectional dimension of the panel used throughout the analysis.

Table 36. Concordance between BRSA Sectoral Credit Categories and TurkStat Sectoral Classifications

BRSA sectoral credit category	Matched TurkStat sectoral classification
Agriculture, Hunting and Forestry	—
Agriculture	A. Agriculture, Forestry and Fishing
Hunting	A. Agriculture, Forestry and Fishing

Table 36. (continued)

Timber and Forest Products	A. Agriculture, Forestry and Fishing
Fishing	A. Agriculture, Forestry and Fishing
Mining and Quarrying	—
Extraction of Energy-Producing Minerals	B D E. Industry excluding Manufacturing
Extraction of Non-Energy-Producing Minerals	B D E. Industry excluding Manufacturing
Manufacturing	—
Food, Beverages and Tobacco Manufacturing	C. Manufacturing
Textiles and Textile Products Manufacturing	C. Manufacturing
Leather and Leather Products Manufacturing	C. Manufacturing
Wood and Wood Products Manufacturing	C. Manufacturing
Pulp, Paper and Paper Products Manufacturing	C. Manufacturing
Coke, Refined Petroleum Products and Nuclear Fuel Manufacturing	C. Manufacturing
Chemicals and Chemical Products Manufacturing	C. Manufacturing
Rubber and Plastic Products Manufacturing	C. Manufacturing
Other Non-Metallic Mineral Products Manufacturing	C. Manufacturing
Basic Metals and Fabricated Metal Products Manufacturing	C. Manufacturing
Machinery and Equipment Manufacturing	C. Manufacturing

Table 36. (continued)

Electrical and Optical Equipment Manufacturing	C. Manufacturing
Transport Equipment Manufacturing	C. Manufacturing
Ship and Boat Building	C. Manufacturing
Manufacture of Motor Vehicles and Other Transport Equipment	C. Manufacturing
Manufacturing Not Elsewhere Classified	C. Manufacturing
Electricity, Gas and Water Production and Distribution	B D E. Industry excluding Manufacturing
Construction	F. Construction
Wholesale and Retail Trade; Repair of Motor Vehicles; Personal and Household Goods	—
Retail Sale of Motor Vehicles and Automotive Fuels	G H I. Wholesale and Retail Trade, Transportation and Storage, Accommodation and Food Service Activities
Wholesale Trade and Commission Trade	G H I. Wholesale and Retail Trade, Transportation and Storage, Accommodation and Food Service Activities
Retail Trade and Personal Goods	G H I. Wholesale and Retail Trade, Transportation and Storage, Accommodation and Food Service Activities
Hotels and Restaurants / Tourism	—
Hotels	G H I. Wholesale and Retail Trade, Transportation and Storage, Accommodation and Food Service Activities
Restaurants	G H I. Wholesale and Retail Trade, Transportation and Storage, Accommodation and Food Service Activities

Table 36. (continued)

Other Tourism Activities	G H I. Wholesale and Retail Trade, Transportation and Storage, Accommodation and Food Service Activities
Transportation, Storage and Communication	—
Rail Transport	G H I. Wholesale and Retail Trade, Transportation and Storage, Accommodation and Food Service Activities
Road Passenger Transport	G H I. Wholesale and Retail Trade, Transportation and Storage, Accommodation and Food Service Activities
Road Freight Transport	G H I. Wholesale and Retail Trade, Transportation and Storage, Accommodation and Food Service Activities
Water Transport	G H I. Wholesale and Retail Trade, Transportation and Storage, Accommodation and Food Service Activities
Air Transport	G H I. Wholesale and Retail Trade, Transportation and Storage, Accommodation and Food Service Activities
Other Transport Activities and Storage	G H I. Wholesale and Retail Trade, Transportation and Storage, Accommodation and Food Service Activities
Communication	J. Information and Communication
Financial Intermediation	—
Monetary Institutions	K. Financial and Insurance Activities
Other Financial Intermediaries	K. Financial and Insurance Activities

Table 36. (continued)

Voluntary Insurance and Pension Services	K. Financial and Insurance Activities
Other Financial Intermediation	K. Financial and Insurance Activities
Real Estate Brokerage, Renting and Business Activities	—
Real Estate Brokerage	L. Real Estate Activities
Renting of Transport Equipment, Machinery and Equipment	L. Real Estate Activities
Computer and Related Activities	J. Information and Communication
Research, Consultancy, Advertising and Other Business Activities	M N. Professional, Scientific, Technical, Administrative and Support Service Activities
Public Administration, Defence and Compulsory Social Security	O P Q. Public Administration and Defence, Education, Human Health and Social Work Activities
Education	O P Q. Public Administration and Defence, Education, Human Health and Social Work Activities
Human Health and Social Work Activities	O P Q. Public Administration and Defence, Education, Human Health and Social Work Activities
Other Services	—
Sewerage and Waste Management	B D E. Industry excluding Manufacturing
Activities of Membership Organizations	R S T U. Other Service Activities
Arts, Entertainment and Sports Activities	R S T U. Other Service Activities
Other Personal Service Activities	R S T U. Other Service Activities
Activities of Households as Employers	R S T U. Other Service Activities
Activities of Extraterritorial Organisations and Bodies	R S T U. Other Service Activities

Notes: BRSA denotes the Banking Regulation and Supervision Agency of Türkiye and TurkStat the Turkish Statistical Institute. A dash indicates a higher-level BRSA category heading whose TurkStat counterpart is assigned at the sub-category level. The classification 'B D E. Industry excluding Manufacturing' applies to BRSA categories that fall within industry but are not matched to 'C. Manufacturing'.

CHAPTER 7

EMPIRICAL METHODOLOGY FOR THE TÜRKİYE ANALYSIS

7.1. Definition of Variables and Functional Form

This chapter examines how financial development shapes sectoral economic performance within a nonlinear framework applied at the level of individual branches of activity. Financial development, proxied here by total cash credit (TCC) extended by private-sector deposit banks, participation banks and development and investment banks, is treated as a sector-specific rather than an economy-wide magnitude. The central question is whether sectoral credit expansion influences sectoral value added in a linear and symmetric manner, or whether its contribution is conditional on the level of credit already in place. Sectoral value added enters as ln_GDP in levels and as D.ln_GDP in growth form, allowing the long-run association between financing and production to be separated from the short-run dynamics of sectoral growth.

The modelling strategy accommodates sectoral heterogeneity and potential nonlinearity by supplementing the linear credit term with its quadratic (TCC^2) and cubic (TCC^3) transformations. The quadratic term identifies whether the relationship is U-shaped or inverted-U-shaped, indicating a single point at which the marginal contribution of credit changes sign. The cubic term admits richer configurations such as N-shaped or inverted-N-shaped paths, implying that the effect of financial development may reverse more than once across the credit distribution. Rajan and Zingales (1998) show that sectors differ systematically in their dependence on external finance, providing a theoretical reason to expect that the curvature of the credit–output relationship will itself vary across branches of activity.

The approach thus moves beyond estimating a single average slope for the sectoral effects of financial development. It permits the empirical identification of threshold levels and structural turning points at which the direction of the relationship may change, making it possible to test how far credit expansion supports production before its contribution weakens or reverses. This design also aligns with the theoretical position that the influence of finance on the real economy is context-dependent, varying both across sectors and with the volume of credit already intermediated.

Rather than confining attention to linear coefficients, the specification therefore treats asymmetry as an object of estimation in its own right. Rising sectoral credit may exert only a limited influence on value added up to a point, beyond which a stronger, weaker or opposite effect emerges. Rioja and Valev (2004a) demonstrate that the growth contribution of financial development differs across stages of economic development and across the channels through which growth is generated, suggesting that its effects are conditional rather than uniform. Sahay et al. (2015) reach a comparable conclusion for emerging markets, arguing that the gains from financial deepening remain substantial but are bounded by limits on both the size and the speed of that deepening. Adopting a nonlinear sectoral specification accordingly relaxes the assumption that financial development affects every sector with the same sign and magnitude irrespective of credit levels.

All variables enter the estimations in natural logarithms. The logarithmic transformation compresses the pronounced scale differences and high variability characteristic of nominal macroeconomic series, which is particularly relevant here because both credit and value added rise steeply over the sample period. Taking logarithms yields more balanced series, stabilises their distributions and limits the leverage exerted by extreme observations on the estimated coefficients. It does not, however, resolve heteroskedasticity, which must still be assessed through formal diagnostic testing and, where necessary, addressed through robust inference (Wooldridge, 2010).

A further advantage of the logarithmic form concerns economic interpretation. When level variables are expressed in logarithms, the estimated coefficients read as

elasticities, so that a one percent change in sectoral credit can be related directly to the proportional change in sectoral value added. Interpretation requires more care in the growth specification, where the dependent variable is D.ln_GDP and therefore measures the logarithmic change – the approximate growth rate – of sectoral value added. In that case the coefficients express how proportional movements in credit relate to sectoral growth rather than to the level of production, so that the logarithmic transformation both standardises the scale of the variables and sharpens the economic content of the estimated relationship.

7.2. Cross-Sectional Dependence and Unit-Root Tests

7.2.1. Cross-Sectional Dependence Test

In panels built from macroeconomic and financial variables, the assumption that cross-sectional units evolve independently is rarely tenable. Where economic linkages are strong, a shock originating in one unit may propagate to others, or a common factor may move all units simultaneously, generating what the literature terms cross-sectional dependence. Such dependence typically originates in global crises, financial cycles, monetary policy shifts, exchange rate shocks, co-movement in demand and supply conditions or synchronised policy changes. In sectoral applications the problem is likely to be more acute still, since branches of the same economy share a single financing environment and are tied together through input–output relations.

Testing for cross-sectional dependence is therefore a necessary preliminary to unit root, cointegration and long-run coefficient analysis rather than an optional diagnostic. Ignoring it amounts to assuming away common shocks and spillovers, an assumption whose failure distorts first-generation panel unit root tests and standard estimators alike. Pesaran (2007) shows that conventional panel unit root tests lose accuracy once cross-sectional dependence is present, and proposes an alternative built on cross-sectional averages. The dependence test consequently guides the choice of econometric method at every subsequent stage of the analysis.

The present study applies the CD test of Pesaran (2004) to establish whether cross-sectional dependence characterises the eleven principal sectors of the Turkish economy. The test evaluates the null of cross-sectional independence by averaging the pairwise correlation coefficients of the residuals across units, so that a significant statistic indicates systematic co-movement. Under the null the sectors behave independently, while under the alternative they respond to shared movements or shocks. De Hoyos and Sarafidis (2006) show that dependence of this kind is common in multi-unit panels and provide the diagnostic tools for detecting it, so that both model selection and subsequent inference can be informed by the outcome. Pesaran (2015) further distinguishes weak from strong dependence, the consequences for any given estimator depending on the degree of dependence and on whether the estimator models the common-factor structure; strong dependence is the more consequential case for estimators that do not.

The Pesaran CD statistic takes the following general form:

$$CD = \sqrt{[\ 2T / (N(N-1))\ ]} \cdot \Sigma(i=1 \text{ to } N-1)\ \Sigma(j=i+1 \text{ to } N)\ \hat{\rho}_{ij} \quad (18)$$

Here N denotes the number of cross-sectional units, T the length of the time dimension and $\hat{\rho}_{ij}$ the estimated pairwise correlation coefficient between units i and j. In this application N corresponds to the eleven sectors representing the Turkish economy, while T records the number of quarterly observations available over the sample period. Because the statistic accumulates the co-movement of every sector pair, it measures the average degree of dependence across the panel as a whole. The test is particularly pertinent here, given that sectoral production and credit are determined within a single national macroeconomic environment in which changes in credit conditions, interest rates, inflationary pressure, exchange rates, domestic demand and policy decisions may affect all branches at once.

Table 37 reports the cross-sectional dependence test results for the variables used in the study, in both their level and logarithmic forms. The null of cross-sectional independence is rejected at the 1% level for every series examined, indicating statistically significant dependence across sectors in GDP, TCC, ln_GDP and

ln_TCC alike. Sectoral value added and sectoral credit in Türkiye therefore do not evolve independently, but co-move to a substantial degree as a consequence of shared macroeconomic conditions, common financial developments and the productive linkages that bind sectors together.

Table 37. Cross-Sectional Dependence (CD) Test Results: Türkiye Sectoral Panel

Variables	**Pesaran CD-test Statistic**	**p-value**	**Average Correlation**	**Absolute Correlation**
GDP	65.38	0.000***	0.951	0.951
TCC	67.62	0.000***	0.989	0.989
ln_GDP	67.07	0.000***	0.975	0.975
ln_TCC	67.06	0.000***	0.981	0.981

Notes: The null hypothesis is cross-sectional independence across panel units. Tests are conducted across 11 economic sectors (N = 11) over a strongly balanced panel of 86 periods (T = 86). *** $p < 0.01$.

The evidence extends beyond the formal significance of the test statistics. The average pairwise correlation coefficients range from 0.951 to 0.989 across the four series, revealing co-movement among the eleven sectors that is not merely detectable but close to complete. Correlations of this magnitude imply that sectors within the Turkish economy respond jointly to national shocks affecting both production and credit utilisation, so that credit expansion, output growth or contraction in one branch is unlikely to remain confined to it, propagating instead through intersectoral linkages and shared financing conditions.

Three methodological implications follow for the estimations reported below. First, conventional panel techniques predicated on sectoral independence are inappropriate for these data, since their standard errors and, in dynamic settings, their point estimates would be distorted. Second, stationarity must be assessed using second-generation panel unit root tests that accommodate dependence explicitly, for which the CADF/CIPS framework of Pesaran (2007) is well suited. Third, the estimator

itself must incorporate common factors, intersectoral linkages and simultaneous national shocks if the coefficients are to be interpreted as sector-specific.

The CD test results thus establish pronounced cross-sectional dependence among the sectoral value added and credit series for Türkiye, confirming that these sectors are embedded in a single macroeconomic and financial environment. Far from constituting a preliminary formality, the test outcome forms a load-bearing element of the study's methodological framework, determining the choice of unit root test, the specification of the model and the estimator ultimately adopted.

7.2.2. Panel Unit-Root Test

Because the dependence tests reveal strong co-movement among the eleven sectors, first-generation panel unit root tests – which presuppose independence across units – were judged unsuitable for assessing the stationarity of the variables. Tests of this kind may deliver misleading conclusions about the order of integration precisely because they abstract from the common shocks and sectoral interlinkages documented above. The integration properties of the series are therefore established using the second-generation Cross-Sectionally Augmented IPS (CIPS) test of Pesaran (2007).

The CIPS procedure generalises the standard IPS test by augmenting each unit-specific regression with information that absorbs the common-factor structure. Pesaran (2007) shows that, in the presence of dependence, the ADF equation for each unit should include the cross-sectional mean of the variable together with its lagged differences, yielding the Cross-Sectionally Augmented Dickey–Fuller (CADF) regressions. Averaging the resulting CADF statistics across units produces the CIPS statistic, providing a testing framework appropriate to panels exhibiting both cross-sectional dependence and heterogeneous dynamics.

The null hypothesis of the CIPS test is that the panel series contain a unit root and are therefore non-stationary, while the alternative allows stationarity in at least some cross-sectional units. Rejection of the null leads the variable to be treated as

stationary. Stationarity is examined here in levels and, where the level test fails to reject, in first differences. Every variable is tested under two deterministic specifications – one including a constant only, and one including both a constant and a deterministic trend – so that stationarity around a mean and stationarity around a deterministic trend are each assessed rather than assumed.

Lag length selection materially affects the reliability of unit root inference. Specifying too few lags leaves autocorrelation in the residuals, whereas specifying too many consumes degrees of freedom and erodes the power of the test. The CIPS equations are consequently estimated with variable-specific lag orders reflecting the dynamic structure of each series, selected using the Akaike Information Criterion and the Schwarz Bayesian Information Criterion. The AIC of Akaike (1974) balances goodness of fit against the number of estimated parameters, while the SBIC of Schwarz (1978) imposes a heavier penalty on complexity and therefore favours more parsimonious specifications. A common lag order per variable was obtained from the mode of the AIC and SBIC selections across cross-sections, giving four lags for GDP and ln_GDP, the indicators of economic activity, and one lag for TCC and ln_TCC, the private-sector cash credit variables.

Table 38 reports the CIPS results for the variables in level form. GDP, ln_GDP and TCC emerge as stationary in levels, with the unit root null rejected at the 1% and 5% levels under the relevant specifications. Sectoral value added, its logarithm and total cash credit in levels are accordingly classified as I(0), indicating that they may enter the econometric model in their original form without further transformation.

Table 38. Pesaran (2007) CIPS Panel Unit Root Test Results: Türkiye Sectoral Panel

Variables	Optimal Lags	Constant (Intercept)		Constant & Trend	
		t-bar stat.	p-value	t-bar stat.	p-value
GDP	4	−3.412***	0.000	−3.820***	0.000
ln_GDP	4	−2.710***	0.000	−2.992***	0.006
TCC	1	−2.249**	0.044	−3.398***	0.000
ln_TCC	1	−2.155*	0.086	−2.246	0.660
Δln_TCC	1	−5.909***	0.000	−6.210***	0.000

Notes: The null hypothesis is that the series contains a unit root. Tests are conducted across 11 economic sectors (N = 11). To accommodate heterogeneous dynamics, variable-specific lag lengths were selected from the mode of the Akaike and Schwarz Bayesian information criteria across cross-sections. Every variable was tested under both the constant and the constant-and-trend specification. Δ is the first difference operator. *, ** and *** denote significance at the 10%, 5% and 1% levels.

The natural logarithm of total cash credit behaves differently in level form. Although ln_TCC displays marginal evidence of stationarity at the 10% level when only a constant is included, the unit root null cannot be rejected once a deterministic trend is added, and the associated p-value of 0.660 leaves no basis for treating the series as stationary. Since a deterministic trend plausibly captures the systematic upward drift of nominal credit over two decades, the trend specification is the more informative of the two, and ln_TCC is therefore classified as non-stationary in levels.

To establish the order of integration precisely, the CIPS test was repeated on the first difference of the series. The unit root null is rejected at the 1% level for Δln_TCC under both deterministic specifications, with p-values of 0.000 in each case. The credit variable is consequently non-stationary in levels but stationary after first differencing, confirming its classification as I(1) and requiring that it enter the estimated equations in differenced form.

The panel therefore exhibits a mixed integration structure, with GDP, ln_GDP and TCC stationary in levels and ln_TCC integrated of order one, while no variable is integrated of order two. This configuration is decisive for model selection, since ARDL-based procedures are valid for combinations of I(0) and I(1) regressors but not for I(2) series. The bounds testing framework of Pesaran et al. (2001) relies on critical values derived exclusively for I(0) and I(1) variables, so the absence of second-order integration is a precondition for its application. With that condition satisfied, the analysis may proceed to the CS-ARDL specification developed below.

Three conclusions for the empirical strategy follow from the unit root results. First, the use of a second-generation test is warranted by the cross-sectional dependence documented in the previous subsection, so the stationarity classification itself rests on an appropriate testing framework. Second, the mixture of I(0) and I(1) variables

justifies an ARDL-based approach capable of estimating short-run and long-run relationships jointly. Third, the strength of the dependence across sectors requires a specification augmented with cross-sectional averages rather than a conventional panel ARDL, for which the dynamic common correlated effects framework of Chudik and Pesaran (2015) provides the appropriate estimator.

Taken together, the CIPS results supply a firm basis for the modelling decisions that follow. Classifying GDP, ln_GDP and TCC as I(0) and ln_TCC as I(1) establishes the mixed integration structure of the data, while the absence of I(2) variables confirms the admissibility of the CS-ARDL framework. These outcomes thus determine not only the form in which each variable enters the equations, but also the class of estimator capable of separating short-run adjustment from long-run equilibrium.

7.3. The Endogeneity Problem and the Choice of Estimator

7.3.1. Panel Granger Causality Analysis

Having established the integration properties of the variables, the analysis turns to the direction of the relationship between financial development and sectoral economic performance. The purpose at this stage is to determine whether sectoral credit expansion carries information useful for predicting sectoral value added and, conversely, whether sectoral value added helps predict credit expansion. The Dumitrescu and Hurlin (2012) panel causality test is used for this purpose, extending the framework of Granger (1969) to panel settings in which causal patterns may differ across cross-sectional units.

Granger causality does not establish structural or theoretical causation, but assesses whether past values of one variable improve the prediction of another. The results reported here are accordingly interpreted as evidence of predictive causality in the Granger sense. If credit expansion Granger-causes sectoral value added, past credit movements contribute significantly to explaining the current level of value added; if the causality runs in the opposite direction, changes in sectoral size carry information about subsequent credit developments.

The Dumitrescu and Hurlin (2012) test was selected because it does not impose a common causal structure on every unit in the panel. In sectoral analysis the relationship between finance and performance need not display the same direction or strength everywhere, since the influence of credit expansion on production is plausibly stronger in capital-intensive branches than in activities that depend little on bank intermediation. This expectation follows directly from Rajan and Zingales (1998), who document systematic differences across industries in their reliance on external finance. By allowing causality to be present in some sectors and absent in others, the test accommodates precisely the heterogeneity that a study of eleven distinct branches of the Turkish economy is designed to expose.

Under the null hypothesis there is no causal relationship from one variable to the other in any cross-sectional unit, while the alternative permits causality in at least some units. Panel-level inference is based on the average of the individual Wald statistics computed sector by sector, together with its standardised counterpart. The procedure thus evaluates whether a common causal tendency emerges within a heterogeneous panel, without requiring that the relationship hold uniformly.

Stationarity is a precondition for VAR-based causality testing, since running the procedure on non-stationary series may generate spurious evidence of predictive links. The variables entering the causality analysis were therefore selected in accordance with the CIPS results reported above. Sectoral value added enters directly as ln_GDP, being stationary in levels, whereas total cash credit enters as Δln_TCC, since ln_TCC is stationary only after first differencing. The differenced credit variable measures the proportional change in credit volume, allowing credit dynamics to be modelled in stationary form.

Table 39 presents the panel causality results for Δln_TCC and ln_GDP. Bidirectional Granger causality is detected between private-sector credit growth and the level of sectoral value added. The null that credit expansion does not Granger-cause sectoral value added is rejected at the 1% level, with a p-value of 0.000, indicating that past movements in credit contribute significantly to explaining current sectoral value added. This pattern is consistent with the supply-leading view associated with Patrick

(1966), under which the development of financial intermediation precedes and facilitates real activity, although it demonstrates predictive rather than structural causality.

Table 39. Dumitrescu–Hurlin (2012) Panel Granger Causality Test Results: Türkiye Sectoral Panel

Null Hypothesis (H_0)	Lags	W-bar	Z-bar tilde	p-value	Decision
Δln_TCC does not Granger-cause ln_GDP	1	2.839	4.067	0.000***	Reject H_0
ln_GDP does not Granger-cause Δln_TCC	1	2.701	3.759	0.000***	Reject H_0

Notes: Tests follow the Dumitrescu and Hurlin (2012) procedure, which accommodates cross-sectional heterogeneity in the causal structure. The sample is fully balanced from 2005q2 across 11 economic sectors. Δ is the first difference operator. *** $p < 0.01$.

The second null, that sectoral value added does not Granger-cause credit expansion, is likewise rejected at the 1% level, indicating that past sectoral output carries predictive information for credit growth. Sectors generating larger volumes of production and value added plausibly face greater financing requirements, translating into stronger credit demand from the banking system. This direction corresponds to the demand-following hypothesis, under which the expansion of real activity elicits a supply response from financial intermediaries rather than the reverse (Patrick, 1966).

Detecting causality in both directions points to a reciprocal and dynamic interaction rather than a one-way link. Credit expansion supplies a financing channel capable of supporting production and investment across sectors, while growth in sectoral value added raises the demand for bank credit, so the two processes may operate concurrently. The relationship between financial development and sectoral activity is therefore unlikely to be adequately described by either the supply-leading or the demand-following account taken alone, both channels appearing to contribute to a feedback loop between the financial and the real economy.

These findings carry direct consequences for the modelling stages that follow. Bidirectional predictive causality implies that credit may help explain sectoral performance while sectoral performance simultaneously shapes credit dynamics, so any specification that treats credit as strictly exogenous risks misattributing the direction of the association. A framework capable of representing both short-run and long-run interactions is therefore preferable to a static, unidirectional model, and the causality results provide the empirical justification for examining the finance–output relationship within the CS-ARDL framework developed below.

In summary, the Dumitrescu–Hurlin results establish a strong two-way predictive relationship between private-sector credit growth and sectoral value added across the eleven principal sectors of the Turkish economy. The direction running from credit to output is consistent with the supply-leading hypothesis, while the reverse direction supports the demand-following hypothesis, together indicating a feedback mechanism sustained by the continuing interaction between financial development and sectoral production.

7.3.2. Baseline Fixed Effects Model and Diagnostic Tests

Before proceeding to dynamic estimation, a static fixed effects model is estimated to establish a preliminary reading of the relationship between the variables and to evaluate the diagnostic properties of the specification. Fixed effects provide a natural point of departure in panel analysis because they absorb unobserved, sector-specific and time-invariant characteristics – production technology, structural capital intensity, institutional arrangements, barriers to entry or the prevailing financing environment – that would otherwise contaminate the estimates (Wooldridge, 2010). The model therefore serves as a benchmark against which the influence of unobserved fixed sectoral differences is held constant.

The static specification is nonetheless a diagnostic device rather than the estimator on which the substantive conclusions rest. The preceding sections have documented strong cross-sectional dependence and bidirectional Granger causality, indicating that the link between sectoral activity and credit dynamics is shaped by common

macroeconomic shocks, intersectoral linkages and lagged responses rather than by a static, unidirectional mechanism. The fixed effects model can accordingly identify a first-order pattern in the data, but it is not adequate for definitive econometric inference.

Model construction also reflects the CIPS results. Because ln_TCC is non-stationary in levels, including it directly would expose the regression to spurious inference, in which high explanatory power and apparently significant coefficients arise between series that share no genuine economic relationship – a problem documented systematically by Granger and Newbold (1974) and formalised by Phillips (1986). The credit variable therefore enters in first-differenced form as Δln_TCC, which is stationary and measures the logarithmic change in credit volume, capturing the pace of credit expansion. Sectoral value added, being stationary in levels, enters directly as ln_GDP.

The baseline fixed effects specification is written as follows:

$$ln_GDP_{it} = \alpha_i + \beta\, \Delta ln_TCC_{it} + \varepsilon_{it} \quad (19)$$

Here *i* indexes sectors, *t* the time dimension, α_i the sector-specific fixed effects and ε_{it} the error term. The α_i term controls for time-invariant sectoral characteristics, while β measures the association between credit expansion and the level of sectoral value added. Because the specification is static, this coefficient records only the contemporaneous relationship, capturing neither the lagged effects of credit expansion, nor the persistence generated by past values of sectoral value added, nor the dependence across sectors induced by common shocks. It should therefore be read as a restricted benchmark preceding the dynamic estimation strategy adopted in the main analysis.

Table 40 reports the baseline fixed effects estimates together with the Modified Wald test for groupwise heteroskedasticity, applied to determine whether the variance of the error term is constant across cross-sectional units. Under the null the error variances are equal across sectors, supporting the assumption of groupwise

homoskedasticity, whereas the alternative allows them to differ (Baum, 2001). The test statistic is $\chi^2 = 3.110$ with a p-value of 0.989, so the null cannot be rejected and there is no material evidence of groupwise heteroskedasticity in the residuals of the baseline model.

Table 40. Baseline Fixed-Effects Estimates and Diagnostic Test Results: Türkiye Sectoral Panel

Dependent Variable: ln_GDP	**Coefficient**	**Std. Error**	**t-Statistic**	**p-value**
Δln_TCC	0.711	0.466	1.520	0.128
Constant	17.698***	0.054	327.230	0.000
Model Diagnostics	Obs. = 924	N = 11	Overall $R^2 = 0.0039$	
Heteroskedasticity Test	Modified Wald	$\chi^2 = 3.110$	p-value = 0.989	

Notes: The estimation is a static fixed-effects (within) regression. The Modified Wald test evaluates the null of groupwise homoskedasticity across cross-sectional units. Δ is the first difference operator. *** $p < 0.01$.

The high p-value returned by the Modified Wald test indicates that groupwise heteroskedasticity is not a material concern in the residuals of the baseline model, but it does not address other econometric hazards, spurious regression foremost among them. That risk is mitigated separately by excluding ln_TCC in level form and entering the credit variable in the stationary first difference consistent with its order of integration. The heteroskedasticity test and the stationarity adjustment therefore function as two distinct safeguards, neither substituting for the other.

Turning to the estimates themselves, the coefficient on Δln_TCC is statistically insignificant, its p-value of 0.128 offering no support for a strong contemporaneous link between credit growth and the level of sectoral value added within a static framework. The explanatory power of the model is correspondingly minimal, with an overall R^2 of 0.0039 indicating that almost none of the variation in sectoral value added is accounted for. The relationship between credit expansion and

sectoral production is evidently not something a static fixed effects specification can capture.

This outcome should not, however, be read as evidence that financial development and sectoral value added are unrelated. The earlier results on cross-sectional dependence and panel Granger causality reveal pronounced common movements across sectors together with bidirectional predictive links between finance and real performance. The insignificance of the Δln_TCC coefficient therefore indicates the limitations of the estimator rather than the absence of a relationship, pointing towards lagged effects, common factors and intersectoral dependence that the static model cannot represent.

Three specific limitations become apparent at this point. First, excluding lagged values of the dependent variable prevents the model from capturing the persistence of sectoral value added, even though macroeconomic and sectoral variables typically adjust with delay – the effect of credit expansion on investment, productive capacity and value added unfolding progressively rather than within the quarter in which borrowing occurs. Second, the standard fixed effects estimator does not model the cross-sectional dependence generated by common macroeconomic factors, so that the influence of national financial conditions, monetary policy, exchange rate movements and demand shocks documented by the CD test remains outside the specification, leaving it incompletely specified. Third, the bidirectional causality identified by the Dumitrescu–Hurlin test suggests that restricting the relationship to a contemporaneous flow from credit to value added is too confining, since feedback mechanisms appear to operate in both directions.

These considerations motivate the move to the Cross-Sectionally Augmented Autoregressive Distributed Lag framework, which addresses cross-sectional dependence and dynamic adjustment simultaneously. The Common Correlated Effects approach of Pesaran (2006) introduces cross-sectional averages into the estimation to control for the unobserved common factors generating dependence, while Chudik and Pesaran (2015) extend the method to heterogeneous dynamic panels containing lagged dependent variables and weakly exogenous regressors.

Because the CS-ARDL specification represents both the common-factor structure and the lag structure of the data, it constitutes a more appropriate framework than static fixed effects given the diagnostic evidence assembled above.

It bears emphasising that the CS-ARDL framework does not eliminate endogeneity altogether. What it offers is a more defensible estimation environment than a static model, accommodating common factors, cross-sectional dependence and dynamic adjustment while distinguishing short-run from long-run effects. The improvement is therefore substantial but partial, and the coefficients reported below should be interpreted as conditional associations within a dynamic heterogeneous panel rather than as fully identified structural parameters.

The baseline model consequently yields two useful conclusions. The Modified Wald test indicates that groupwise heteroskedasticity does not compromise inference in this specification, while the insignificance of the credit coefficient and the negligible explanatory power demonstrate that the finance–output relationship is not recoverable within a static fixed effects framework. Read alongside the cross-sectional dependence and causality results, these outcomes confirm that the static model is a preliminary and restricted device, and that an estimator combining common factors, dynamic lags and a separation of short- and long-run effects is required for the substantive analysis.

7.3.3. Limitations of System GMM and the Case for CS-ARDL

Introducing a lagged dependent variable into a panel regression has well-known methodological consequences. The lagged dependent variable becomes correlated with the composite error term once fixed effects are present, so that conventional within estimation delivers biased coefficients. Nickell (1981) identifies this dynamic panel bias formally, showing that it is a function of the length of the time dimension. Since sectoral value added is expected to depend on its own past values, static panel methods are inadequate for the present application and a dynamic estimator is required. Beyond the bias arising from dynamics, the study has produced empirical evidence of a potential two-way relationship between financial development and

sectoral performance. The Dumitrescu and Hurlin (2012) test reported above establishes bidirectional predictive causality between private-sector credit growth and sectoral value added, implying that credit expansion helps forecast sectoral output while sectoral output helps forecast credit expansion. The finance–output link should therefore be treated as reciprocal and dynamic rather than as a one-way influence from an external regressor. Because Granger causality does not establish structural causation, these results are best read as strong grounds for taking reverse causality and feedback seriously in specification, rather than as proof of a particular causal architecture.

The literature has developed Generalised Method of Moments estimators precisely to address dynamic panels containing lagged dependent variables, fixed effects and regressors that fail the assumption of strict exogeneity. Arellano and Bover (1995) propose a more efficient instrumental variable framework for error-components models, while Blundell and Bond (1998) develop the System GMM estimator by adding moment conditions in levels, recognising that estimators relying solely on differenced equations perform poorly when the series are highly persistent. System GMM has consequently become a standard response to the joint problem of endogeneity, fixed effects and dynamic specification.

The performance of System GMM depends heavily, however, on the dimensional structure of the panel. Roodman (2009a, 2009b) emphasises that difference and system GMM estimators are designed for panels with a small time dimension and a large cross-section, in which the number of instruments remains modest relative to the number of units. The present dataset departs sharply from that configuration, comprising eleven sectors observed over eighty-six quarters, so that the time dimension exceeds the cross-sectional dimension by a substantial margin. The data are therefore incompatible with the design for which System GMM performs best.

A closely related practical difficulty concerns the proliferation of instruments. Roodman (2009a) shows that the instrument count expands rapidly in GMM estimation, particularly in the system variant, and that an excessive instrument set may overfit the endogenous regressors while weakening the Hansen test of

overidentifying restrictions. In such circumstances very high or near-unity Hansen p-values do not confirm instrument validity but rather signal a loss of test power, producing a diagnostic impression that is reassuring and uninformative in equal measure. Considerable caution is therefore warranted whenever the instrument count is large relative to the number of cross-sectional units, as it would necessarily be here.

A further limitation is that System GMM does not model strong cross-sectional dependence arising from common macroeconomic shocks or unobserved common factors, unlike estimators built on cross-sectional averages such as CCE and CS-ARDL. Time dummies can absorb common shocks only under the restrictive assumption that all units respond to them identically, which is difficult to sustain when sectors differ in credit sensitivity, capital intensity and exposure to external demand. Given the strength of the dependence documented in Section 8.2 and the T $\gg$ N structure of the panel, an estimator that models the common-factor structure explicitly is the more defensible choice.

The unit root results add a further consideration, though not the decisive one. The variables display mixed orders of integration, with some series I(0) and others I(1) but none I(2), a configuration that suits ARDL-based methods capable of estimating short- and long-run relationships jointly. This does not, however, disqualify System GMM as such, since the estimator can be applied to variables of mixed integration order. The binding constraints in this application are the dimensional structure of the panel, the strength of the cross-sectional dependence and the risk of instrument proliferation, with the mixed integration structure serving to confirm the suitability of CS-ARDL rather than to rule out the alternative on its own.

The study therefore adopts the CS-ARDL approach, which belongs to the Common Correlated Effects tradition and addresses dependence arising from unobserved common factors through the inclusion of cross-sectional averages. Pesaran (2006) develops the CCE framework for panels with a multifactor error structure, while Chudik and Pesaran (2015) extend it to heterogeneous dynamic models containing lagged dependent variables and weakly exogenous regressors. Chudik et al. (2016)

subsequently examine cross-sectionally augmented dynamic estimators for long-run coefficients, demonstrating their effectiveness in recovering long-run relationships in the presence of correlated errors.

The principal attraction of CS-ARDL in this setting lies in its capacity to address several econometric difficulties simultaneously. Cross-sectional averages absorb the dependence generated by common shocks, the autoregressive distributed lag structure permits lagged values of both dependent and explanatory variables to enter, and the error correction representation separates transitory dynamics from equilibrium relationships. The last of these is essential to the research question, since whether credit influences sectoral value added differently over different horizons is precisely what the chapter sets out to establish.

The framework nevertheless has limits that should be stated plainly. Although CS-ARDL improves upon static panel models by representing dynamic adjustment, cross-sectional dependence and common factors, it does not resolve endogeneity in full. Including multiple lags captures feedback and adjustment mechanisms between the variables, yet some sources of simultaneity may persist structurally. The estimates should accordingly be interpreted within a robust empirical framework that acknowledges the dynamic panel structure and the common-factor environment, rather than as evidence that all identification problems have been settled.

In conclusion, the decision to estimate CS-ARDL rather than System GMM rests on the dimensional structure of the data and the diagnostic evidence assembled in the preceding sections. System GMM is designed for panels with a short time dimension and a large cross-section, whereas the present panel comprises eighty-six quarters and eleven sectors, so that instrument proliferation would compromise the Hansen test and the reliability of the estimates. The CS-ARDL framework instead matches the requirements of the analysis, accommodating cross-sectional dependence, dynamic lag structure and the distinction between short- and long-run effects, and it is therefore the more appropriate method for examining the relationship between financial development and sectoral value added in Türkiye.

7.4. Dynamic Common Correlated Effects and the CS-ECM Model

To analyse short- and long-run relationships between sectoral value added and financial development, the study employs dynamic panel methods that accommodate cross-sectional dependence and sectoral heterogeneity jointly. The diagnostic evidence reported above reveals pronounced dependence among the eleven sectors of the Turkish economy, plausibly reflecting shared macroeconomic conditions, a common financing environment and dense production linkages between branches. Estimators that treat sectors as independent would therefore misstate both the precision and, in dynamic settings, the magnitude of the estimated coefficients. The analysis accordingly adopts the Dynamic Common Correlated Effects framework, which represents unobserved common factors and the synchronised movements they generate across sectors.

Estimation is carried out using the xtdcce2 package for Stata developed by Ditzen, which implements the dynamic common correlated effects approach of Chudik and Pesaran (2015). Ditzen (2018) documents that the command estimates heterogeneous coefficients, augments the specification with cross-sectional averages and their lags to absorb dependence, and reports post-estimation dependence diagnostics. This combination is well matched to the present application, since it allows sectors to respond differently to financial development instead of imposing a single coefficient across the panel.

7.4.1. CS-ARDL Model Specification and Baseline Structure

The empirical strategy centres on the CS-ARDL specification, which admits lagged values of both the dependent and the explanatory variables and thereby permits their dynamic interaction to be examined over successive quarters. Cross-sectional averages enter alongside these lags to control for unobserved common shocks – interest rates, exchange rates, inflation, aggregate credit conditions and shifts in domestic demand – that may move several sectors at once. The general nonlinear CS-ARDL model, expressed in the level and mixed forms of the variables and encompassing both short- and long-run relationships, may be written as:

$$ln_GDP_{it} = \alpha_i + \Sigma(j=1 \text{ to } p)\ \lambda_{ij}\ ln_GDP_{i,t-j} + \Sigma(j=0 \text{ to } q)\ \beta'_{ij}\ X_{i,t-j} + \Sigma(l=0 \text{ to } p_z)\ \gamma'_{il}\ \bar{Z}_{t-l} + \varepsilon_{it} \quad (20)$$

The parameters of this equation are defined as follows. The term $\ln_GDP_{it}$ denotes the natural logarithm of value added in sector *i* at time *t* and constitutes the dependent variable, with the subscript *i* indexing the cross-sectional dimension of sectors and *t* the quarterly time dimension. The intercept α_i captures sector-specific unobserved fixed effects and therefore takes the form of a heterogeneous constant. The parameters *p* and *q* denote the optimal lag orders for the dependent and the independent variables respectively, while λ_{ij} collects the autoregressive coefficients attached to the lagged dependent variable. The vector X_{it} contains the independent variables through which nonlinear financialisation effects are identified, taking the form $[X_{it}, X^2_{it}]$ in the quadratic model and $[X_{it}, X^2_{it}, X^3_{it}]$ in the cubic model, with β_{ij} the corresponding coefficient vector for their contemporaneous and lagged values. The vector $\bar{Z}_t$ contains the cross-sectional averages included in line with Pesaran (2006) to control for cross-sectional dependence and common macroeconomic shocks, p_z denotes the lag order selected for those averages – implemented empirically through the cr_lags(2) option – and γ_{il} the associated parameter vector. The disturbance ε_{it} is assumed independent across sectors and Gaussian.

7.4.2. CS-ECM Specification and the Separation of Short- and Long-Run Dynamics

The CS-ARDL model is estimated in Cross-Sectionally Augmented Error Correction form, which sharpens the distinction between short-run and long-run relationships among variables of differing integration order. Ditzen (2021) shows that the CS-DL, CS-ARDL and CS-ECM representations can all be implemented within the xtdcce2 framework to recover long-run effects in models exhibiting cross-sectional dependence. The error correction representation is preferred here because it estimates the speed at which deviations from equilibrium are corrected while simultaneously reporting the transitory response to credit shocks.

The error correction specification estimated in Stata under the long-run restriction, which separates short-run shocks from long-run structural adjustment, is given by:

$$\Delta ln_GDP_{it} = \alpha_i + \varphi_i \left(ln_GDP_{i,t-1} - \theta'_i X_{i,t} \right) + \Sigma(j=1 \text{ to } p-1)\, \lambda^*_{ij}\, \Delta ln_GDP_{i,t-j} + \Sigma(j=0 \text{ to } q-1)\, \beta^{*\prime}_{ij}\, \Delta X_{i,t-j} + \Sigma(l=0 \text{ to } p_z)\, \gamma^{*\prime}_{il}\, \Delta \bar{Z}_{t-l} + u_{it} \quad (21)$$

The operator Δ denotes the first difference, so that $\Delta X_t = X_t - X_{t-1}$. The term Δln_GDP_{it} measures the growth rate of value added in sector *i* at time *t* and serves as the short-run dependent variable. The parameter φ_i is the error correction coefficient, or speed of adjustment, recording how rapidly the system returns to its long-run equilibrium following a shock; cointegration requires this coefficient to be negative and statistically significant, and it corresponds empirically to the coefficient on L.ln_GDP. The vector θ_i collects the long-run coefficients, from which the long-run effects of credit and of its quadratic and cubic transformations, together with the associated turning points, are recovered. The expression $(ln_GDP_{i,t-1} - \theta'_i X_{i,t})$ constitutes the error correction term $ECT_{i,t-1}$ and measures deviations from long-run equilibrium. The parameters λ^*_{ij} are the lag coefficients of the differenced dependent variable and capture short-run stationary dynamics, while $\Delta X_{i,t-j}$ contains the first differences of the independent variables, including the polynomial credit terms, and therefore represents transitory credit shocks. The coefficients β^*_{ij} measure the short-run response of growth to those shocks, $\Delta \bar{Z}_{t-l}$ denotes the lagged cross-sectional averages included to absorb dependence, and u_{it} is the white-noise error vector.

This structure allows the model to represent dynamic adjustment driven by the past values of each sector alongside the cross-sectional dependence generated by common macroeconomic and financial shocks. The first difference of the dependent variable records short-run movements in sectoral production, the differences of the independent variables capture period-to-period variation in credit expansion, and the lagged levels identify the long-run equilibrium relationship. Short-run and long-run parameters are consequently estimated within a single specification rather than obtained from separate models, so that the two horizons are directly comparable.

7.4.3. Controlling for Cross-Sectional Dependence and Post-Estimation Diagnostics

Cross-sectional dependence is addressed by including the cross-sectional averages of the model variables, following the Common Correlated Effects methodology of

Pesaran (2006). These averages proxy for the unobserved common factors driving joint movement across units, so that their inclusion removes from the error term the influence of macroeconomic and financial conditions affecting several sectors at once. In the estimations reported below this is implemented through the cr() option of xtdcce2.

Contemporaneous averages alone may be insufficient in dynamic settings. Chudik and Pesaran (2015) demonstrate that, when the specification contains a lagged dependent variable, a sufficient number of lags of the cross-sectional averages must also be included for the estimates to be consistent, the required number increasing with the length of the time dimension. The cr_lags() option is used accordingly, so that common factors evolving over time are absorbed rather than left in the residuals.

The appropriate lag order was determined by post-estimation diagnostics rather than imposed ex ante. Specifications with one, two and three lags of the cross-sectional averages were estimated, and the Pesaran CD test was applied to the residuals of each. Two lags proved sufficient to remove the evidence of residual dependence, with CD test p-values of 0.5443, 0.6794 and 0.4114 for the linear, quadratic and cubic specifications respectively. Since these values lie well above conventional thresholds, the null of cross-sectional independence in the residuals cannot be rejected, and cr_lags(2) is therefore adopted as the baseline specification throughout the analysis.

The CS-ECM/DCCE strategy employed in this study consequently rests on three methodological commitments. The estimator is CCE-based, incorporating common factors in response to the pronounced dependence documented among sectors. The coefficients are heterogeneous, acknowledging that sectors may respond differently to financial development and to macroeconomic shocks. And the model is estimated in error correction form, so that short-run dynamics and long-run equilibrium relationships are identified separately. Implemented through xtdcce2, this configuration matches the structure of the data, the outcome of the diagnostic testing and the objectives of the research question.

7.5. Testing for Heterogeneity: The MG and PMG Approaches

A central specification question in dynamic panel modelling concerns whether long- and short-run parameters should be treated as homogeneous or heterogeneous across units. Imposing identical behaviour on all units can raise estimation efficiency where the restriction is justified, but becomes untenable when structural differences between sectors, countries or regions are pronounced. The choice between the Pooled Mean Group and the Mean Group estimators therefore carries consequences for the economic assumptions embedded in the model as much as for its statistical properties.

The Mean Group estimator of Pesaran and Smith (1995) estimates a separate dynamic regression for each cross-sectional unit and averages the resulting coefficients across units. Long-run coefficients, short-run coefficients, error correction terms and error variances are all permitted to differ, so that the estimator imposes no common structure beyond the specification itself. This flexibility makes it well suited to panels whose units differ structurally, since each sector may display coefficients reflecting its own production technology, financing requirements, investment intensity and sensitivity to credit.

The Pooled Mean Group estimator of Pesaran et al. (1999) occupies an intermediate position between the Mean Group estimator and fully pooled models. It constrains the long-run coefficients to be identical across units while allowing short-run coefficients, error correction terms and error variances to vary, thereby positing that sectors share a common long-run equilibrium relationship but adjust towards it at differing speeds and through differing channels. Where the assumption of a common long-run relationship is theoretically defensible, PMG is the more parsimonious and efficient of the two estimators.

Given the research question and the structure of the data, however, the restriction that long-run coefficients be identical across sectors is difficult to defend. Sectors in Türkiye differ substantially in production technology, capital utilisation, access to finance, credit requirements, investment horizons and borrowing capacity.

Manufacturing, construction, trade, financial services and agriculture are unlikely to share a common elasticity of value added with respect to credit, whether in the short run or in equilibrium. Representing the long-run effect of financial development through a single coefficient would therefore conceal precisely the differences the study is designed to expose.

The point carries additional force because the analysis centres on the nonlinear character of financialisation and on sector-specific credit saturation. In nonlinear specifications the effect of financial development depends not only on the sign and magnitude of the coefficients but also on the turning points implied by the quadratic and cubic terms, which is where the substantive interpretation resides. Constraining long-run coefficients to equality across sectors would impose a common turning point as well, making sector-specific saturation levels unidentifiable by construction. Under the PMG restriction the long-run relationship is forced into a uniform shape, creating a risk of over-pooling that would obscure the structural differences of interest.

The Mean Group estimator is therefore adopted in preference to PMG. Because it does not require equality of long-run coefficients, the effects of financial development on sectoral value added can be estimated according to the dynamics of each branch of activity. The linear, quadratic and cubic influences of the credit variable are consequently identified sector by sector, allowing sector-specific financialisation effects and turning points to be examined within a flexible framework consistent with the study's emphasis on heterogeneity and nonlinearity.

This choice reflects theoretical reasoning as much as technical convenience. If the influence of financial development on the real economy varies with production structure, capital composition, investment intensity and the capacity to absorb credit, then a specification imposing common long-run coefficients is misspecified rather than merely restrictive. The capacity of the Mean Group estimator to recover sector-specific long- and short-run dynamics without imposing homogeneity makes it the coherent choice for the empirical strategy pursued here.

The Mean Group estimator is accordingly used to assess sector-specific effects of financial development and sector-specific credit saturation within the structural context of each branch. The influence of financialisation on sectoral value added is not compressed into a single common long-run coefficient, but is examined instead through the linear and nonlinear relationships that emerge from the internal dynamics of each sector.

7.6. Analysis of Threshold Values (Turning Points)

The study proceeds from the premise that the relationship between financial development and sectoral value added need not be linear, and therefore estimates quadratic and cubic specifications alongside the linear benchmark. The purpose of the nonlinear models is to determine whether the effect of an expanding credit volume on sectoral performance changes beyond identifiable thresholds. This relaxes the assumption that financial development exerts an effect of constant sign and magnitude at all credit levels, permitting instead U-shaped, inverted-U-shaped, N-shaped or inverted-N-shaped configurations to emerge from the data.

Computing turning points accurately is essential to interpreting these specifications. A turning point marks the value of the independent variable at which its marginal effect on the dependent variable changes sign, identifying the credit level at which the influence of financial development on sectoral value added shifts from positive to negative, from negative to positive, or through a more complex sequence of reversals. Turning points are therefore not merely technical quantities but policy-relevant magnitudes, indicating the thresholds beyond which financial development may exert a stronger, weaker or adverse influence on sectoral production. The quadratic specification represents the relationship between financial development and sectoral value added as:

$$y = \beta_1 x + \beta_2 x^2 \quad (22)$$

Here y denotes the sectoral value added or growth indicator and x the financial development variable, while β_1 and β_2 are the coefficients on the linear and quadratic terms. The marginal effect is obtained as the first derivative with respect to x:

$$\partial y/\partial x = \beta_1 + 2\beta_2 x \quad (23)$$

The turning point occurs where the marginal effect equals zero, so that the condition to be solved is:

$$\beta_1 + 2\beta_2 x = 0 \quad (24)$$

The turning point of the quadratic model is therefore given by:

$$TP = -\beta_1 / (2\beta_2) \quad (25)$$

When β_2 is negative the relationship is inverted-U-shaped and the turning point identifies a maximum, beyond which the positive contribution of financial development to sectoral value added begins to decline. Credit volume raises value added up to that threshold, after which the marginal effect diminishes and may turn negative. When β_2 is positive the relationship is U-shaped and the turning point identifies a minimum, so that the effect of financial development may be negative or negligible at low credit levels before becoming positive once a sufficient scale is reached. Lind and Mehlum (2010) caution that the sign of the quadratic coefficient alone does not establish the presence of a genuine U-shaped or inverted-U-shaped relationship, since the turning point must also lie within the observed data range – a requirement Haans et al. (2016) likewise stress when interpreting curvilinear specifications.

The cubic specification permits a more elaborate nonlinear relationship and takes the form:

$$y = \beta_1 x + \beta_2 x^2 + \beta_3 x^3 \quad (26)$$

Differentiating with respect to x yields the marginal effect:

$$\partial y/\partial x = \beta_1 + 2\beta_2 x + 3\beta_3 x^2 \quad (27)$$

Turning points are found by setting this derivative equal to zero:

$$\beta_1 + 2\beta_2 x + 3\beta_3 x^2 = 0 \quad (28)$$

Because the resulting expression is quadratic, the cubic function may possess two distinct turning points, obtained from:

$$TP_{1,2} = [\, -2\beta_2 \pm \sqrt{(2\beta_2)^2 - 12\beta_3\beta_1}\,] \,/\, (6\beta_3) \quad (29)$$

The expression beneath the square root is the discriminant, and its sign determines the shape of the estimated relationship. A positive discriminant implies two real turning points, consistent with an N-shaped or inverted-N-shaped path; a zero discriminant yields a single tangential root; and a negative discriminant means that no real turning point exists, so that the marginal effect never changes sign. A cubic function always possesses exactly one inflection point, located at $x = -\beta_2/(3\beta_3)$, where the second derivative vanishes and the curvature of the relationship reverses. Where the discriminant is negative that inflection point is the only structural feature the function exhibits, since the marginal effect never changes sign and the relationship remains monotone across the whole credit range. Interpreting a cubic specification therefore requires attention not only to the signs of the coefficients but also to whether the turning points are real and whether they fall within the range of the observed data.

The sign of β_3 shapes the overall configuration of the relationship, a positive value being associated with N-shaped paths and a negative value with inverted-N-shaped paths under particular coefficient combinations. Reading the shape from that sign alone would nonetheless be misleading. The combined marginal effect of the linear, quadratic and cubic terms must be evaluated jointly, the sign and reality of the roots established, and their position relative to the observed credit distribution verified before any curve description is adopted. Where the discriminant is negative, the relationship should be characterised through its inflection point rather than through a pair of turning points, the curvature reversing while the slope retains its sign, so that the appropriate description is one of changing momentum instead of regime reversal.

Turning points are computed from the estimated coefficients of the quadratic and cubic specifications, the former using the linear and quadratic terms and the latter using all three. The procedure identifies algebraically the credit level at which the marginal effect of financial development on sectoral value added changes direction, and it is applied separately to the short-run and long-run components of the CS-ECM estimates, since the two horizons carry distinct economic interpretations.

One further consideration governs interpretation. Because the financial development variable enters in natural logarithms, the turning points recovered from the quadratic and cubic specifications are expressed on a logarithmic scale and must be transformed before they can be read as credit volumes. The required transformation is exponential:

$$TCC^* = e^{(TP)} \quad (30)$$

Here TP denotes the turning point on the logarithmic scale and TCC* the corresponding value of the financial development variable on its original monetary scale. Since the credit data are measured in thousand Turkish liras, the exponentiated value must be interpreted in the same units and is reported below in billion Turkish liras for readability. This transformation allows thresholds derived from a logarithmic specification to be expressed directly as credit volumes, rendering them interpretable for policy purposes rather than confined to the estimation scale.

In summary, the turning points identified in the nonlinear specifications locate the credit levels at which the influence of financial development on sectoral value added changes direction. The quadratic model establishes whether the relationship possesses a single maximum or minimum, while the cubic model permits a more elaborate structure with two distinct turning points or, where the discriminant is negative, a monotone path whose inflection point marks only a change in curvature. Computing these thresholds therefore allows the sectoral effects of financial development to be assessed through the marginal effects prevailing at different credit levels, rather than through coefficient signs alone.

CHAPTER 8

EMPIRICAL RESULTS FOR THE TÜRKİYE ANALYSIS

This section examines the empirical relationship between sectoral credit and economic activity in Türkiye, asking whether sectoral credit supports output growth uniformly and linearly or whether its effects differ across sectors, time horizons and credit levels. The dependent variable throughout is the quarterly change in the logarithm of sectoral GDP (ΔlnGDP), which measures sectoral output growth. Financial development is proxied by the total cash credit allocated to each sector, a variable whose role in the empirical model differs between the short-run and long-run components of the CS-ECM specification.

That distinction is fundamental to reading the results correctly. In the short-run component, credit enters through ΔlnTCC, with the nonlinear specifications adding $\Delta[(\ln TCC)^2]$ and $\Delta[(\ln TCC)^3]$. These terms are the first differences of the squared and cubed transformations of the logarithmic credit level, not the square and cube of ΔlnTCC itself. Short-run estimates therefore do not measure the effect of the credit stock, but record how changes in the logarithm of credit relate to quarterly changes in sectoral GDP. A positive or negative short-run coefficient consequently describes the contemporaneous association between shifting sectoral credit conditions and output growth rather than the equilibrium influence of accumulated credit.

The long-run component, by contrast, is expressed in levels. Financial development enters through lnTCC in the linear specification, extended by the polynomial terms $(\ln TCC)^2$ and $(\ln TCC)^3$ in the quadratic and cubic models. Long-run estimates therefore describe the equilibrium relationship between sectoral credit levels and sectoral economic activity within an error correction framework in which deviations from that equilibrium feed into subsequent quarterly output growth, as set out in

Equation 21. This structure is particularly valuable for the present question, since it isolates the dynamics of credit growth from the equilibrium consequences of credit accumulation.

The CS-ARDL framework in error correction form is appropriate for two reasons. It separates short-run from long-run parameters, which matters when credit influences sectoral output both through transitory liquidity effects and through more durable mechanisms of capital accumulation. It also accommodates the cross-sectional dependence generated by shared macroeconomic shocks, which in the Turkish case are substantial: sectoral output and credit are plausibly driven jointly by monetary tightening, inflation, exchange rate movements, credit policy and shifts in aggregate demand. Neglecting these common influences would risk attributing macroeconomic disturbances to sector-specific credit variation.

The models therefore include cross-sectional averages and their lags, following the correlated effects approach of Pesaran (2006) and its dynamic heterogeneous extension by Chudik and Pesaran (2015). This augmentation matters especially in sectoral panels with dense inter-sectoral linkages, since sectors do not evolve in isolation and a shock to credit, domestic demand, inflation or the exchange rate may reach manufacturing, construction, trade, real estate and services simultaneously. Filtering out such unobserved common factors is what allows the remaining variation to be interpreted as sector-specific.

The estimation sample comprises eleven sectoral groups and 902 observations. In each of the linear, quadratic and cubic specifications the estimated error correction term is negative and statistically significant, confirming that deviations from long-run equilibrium are corrected progressively and validating the CS-ECM representation in which the dependent variable appears in first differences while the long-run relationship is expressed in levels. Residual diagnostics for cross-sectional dependence further indicate that including cross-sectional averages with two lags is sufficient to absorb the common-factor structure, so the reported coefficients can be read as sector-specific effects net of economy-wide shocks.

The findings are presented through three successive specifications – a linear benchmark, a quadratic model and a cubic model – a progression that reflects the underlying theoretical debate rather than a purely technical sequence. The established literature emphasises the growth-promoting role of financial development, with financial systems understood to raise growth by mobilising savings, improving the allocation of resources, facilitating risk diversification, lowering information and transaction costs and easing firms' access to external funding (Beck et al., 2000; King & Levine, 1993; Levine, 2005). On this view, rising sectoral credit should raise sectoral output by relaxing financing constraints and thereby supporting investment, working capital and capacity expansion.

A substantial body of more recent work qualifies that expectation. Financial deepening may stimulate activity initially, yet excessive credit has been associated with debt overhang, inefficient investment, speculative behaviour, resource misallocation and financial instability. This reasoning underpins the 'too much finance' hypothesis of Arcand et al. (2015) and is consistent with the threshold evidence of Law and Singh (2014) and the nonlinear patterns documented by Samargandi et al. (2015). Cecchetti and Kharroubi (2012) reach a similar conclusion, reporting that the growth contribution of financial sector size becomes negative beyond a point, while their subsequent work traces the mechanism to the disproportionate financing of high-collateral, low-productivity projects and to the reallocation of skilled labour towards finance (Cecchetti & Kharroubi, 2015). Sahay et al. (2015) argue in the same direction that the benefits of deepening in emerging markets are bounded by limits on both size and pace. A strictly linear specification, by imposing a constant marginal effect of credit across the entire distribution, is therefore likely to be too restrictive.

The analysis accordingly moves from the linear model to the quadratic and cubic specifications. The linear model establishes what would be concluded if the marginal effect of credit were constant. The quadratic model introduces a single turning point, permitting a test for inverted-U or U-shaped relationships between sectoral credit and sectoral activity. The cubic model admits two potential turning points and therefore captures N-shaped or inverted-N-shaped dynamics. This graduated approach is

essential given that the chapter asks not only whether credit affects output but whether its influence varies across sectors, across time horizons and across stages of financial development.

Interpretation follows the structure of the model throughout. Short-run results describe the association between changes in the logarithmic credit terms and quarterly sectoral output growth, while long-run results describe the equilibrium relationship between sectoral credit levels and sectoral activity. The turning points recovered from the quadratic and cubic specifications identify the credit levels at which the marginal effect changes direction. Framing the results in this way moves the analysis beyond asking whether credit is beneficial or harmful, towards identifying the conditions under which sectoral credit supports growth, ceases to do so, or begins to impede it.

8.1. Linear CS-ECM Results

The linear CS-ECM specification provides the initial benchmark for assessing whether the marginal effect of sectoral credit is constant across sectors and credit levels. It includes ΔlnTCC as the short-run component and lnTCC in levels as the long-run component. While this framework offers a useful empirical foundation, it assumes a uniform relationship between credit and sectoral output growth irrespective of the volume of credit, the character of the sector – capital-intensive or service-oriented – and the horizon over which effects are measured. Table 41 reports the short-run and long-run estimates obtained from this specification.

Table 41. Linear CS-ARDL (CS-ECM) Short-Run and Long-Run Estimates: Türkiye Sectoral Panel

	Short-Run Estimates	**Long-Run Estimates**
Sector Code	**(1) ΔlnTCC**	**(2) lnTCC**
A	−1.030** (0.424)	0.848*** (0.271)
BDE	−0.095 (0.181)	1.320*** (0.415)
C	−0.483 (0.319)	0.119 (1.532)

Table 41. (continued)

F	0.056 (0.425)	−1.707 (1.399)
GHI	−0.291* (0.174)	0.814** (0.316)
J	−0.128 (0.088)	0.037 (0.081)
K	0.010 (0.129)	0.008 (0.197)
L	0.084 (0.052)	0.068 (0.181)
MN	0.012 (0.093)	0.030 (0.039)
OPQ	−0.178* (0.095)	0.016 (0.079)
RSTU	−0.295 (0.194)	0.265* (0.149)
Mean Group	**−0.213** (0.097)**	**0.165 (0.230)**
Obs. = 902	Groups = 11	ECT = −0.620*** (0.120)
R^2(MG) = 0.850	F = 11.88	CD = 0.61 p(CD) = 0.544

Notes: Column (1) reports short-run and Column (2) long-run estimates, with standard errors in parentheses. ECT is the error correction term. Sector codes are: A = Agriculture, Forestry and Fisheries; BDE = Mining and Quarrying and Other Industries; C = Manufacturing; F = Construction; GHI = Wholesale and Retail Trade, Transportation and Storage, Accommodation and Food Service Activities; J = Information and Communication; K = Financial and Insurance Activities; L = Real Estate Activities; MN = Professional, Scientific, Technical, Administrative and Support Service Activities; OPQ = Public Administration and Defence, Education, Human Health and Social Services Activities; RSTU = Other Services. ***, ** and * denote significance at the 1%, 5% and 10% levels.

At the Mean Group level the short-run coefficient on ΔlnTCC is negative and significant at the 5% level, estimated at −0.213. Because the dependent variable is ΔlnGDP, this coefficient records the average association between quarterly sectoral output growth and the quarterly logarithmic change in sectoral credit, and should not be read as the effect of the credit level itself. Within this linear short-run specification, a larger quarterly increase in sectoral credit is on average associated with weaker quarterly sectoral GDP growth once common factors have been absorbed by the cross-sectionally augmented structure.

This negative average coefficient calls for careful interpretation, since it does not imply that credit systematically damages sectoral activity. It records an average short-run association estimated under a linear restriction. In the short run, credit growth may coincide with adjustment costs, rising debt-service obligations, liquidity

pressure or financing needs arising precisely when demand is weak. Firms may borrow not only to invest but to cover working-capital gaps, roll over existing liabilities or absorb cost pressures, so that ΔlnTCC can rise while sectoral output growth is flat or falling. The negative coefficient therefore describes a contemporaneous short-run association rather than supporting any general conclusion that credit accumulation is harmful.

The sectoral estimates reveal that this average is unevenly distributed across the economy. Statistically significant negative short-run coefficients appear in Agriculture, Forestry and Fisheries (A), in Wholesale and Retail Trade, Transportation and Storage, Accommodation and Food Service Activities (GHI) and in Public Administration and Defence, Education, Human Health and Social Services Activities (OPQ). These three sectors account for most of the negative average, while the remaining sectors return coefficients that are statistically indistinguishable from zero, indicating that the credit–output association varies considerably across branches even within a linear framework.

The strongest short-run negative association is found in sector A, where the coefficient on ΔlnTCC is −1.030 and significant at the 5% level. In agriculture, a rise in the quarterly logarithmic change of credit is associated with lower sectoral output growth in the same quarter. This pattern is plausibly connected to the production structure of the sector, which is shaped by seasonal cycles, biological lags, weather conditions and delayed revenue realisation. Credit drawn in one quarter may therefore finance inputs, inventories, machinery or liquidity requirements without generating measurable value added within the same period, so that the negative coefficient reflects the temporal gap between borrowing and output realisation rather than unproductive use of funds.

The short-run coefficient is likewise negative and significant at the 10% level in GHI, though smaller in magnitude than in agriculture. Since this group covers trade, transportation, storage, accommodation and food services, its credit use is closely tied to working capital, inventories, logistics and prevailing demand conditions. A negative short-run coefficient may indicate that credit growth in these

activities partly reflects liquidity pressure or weak demand rather than immediate productivity gains, as firms borrow to sustain operations, finance inventories or maintain cash flow when costs rise or demand becomes volatile. Borrowing of this kind supports continuity without necessarily raising value added within the same quarter.

A comparable reading applies to OPQ, where the short-run coefficient is negative and significant at the 10% level. This group comprises public administration, education, health and social services, whose output is shaped by institutional arrangements, public expenditure patterns, budgetary rules and constraints on service capacity as much as by market financing conditions. Where borrowing serves liquidity management or supports service provision that materialises with delay, a short-run increase in the logarithmic change of credit need not translate into a contemporaneous rise in value added.

In the long-run component of the linear model the Mean Group coefficient on lnTCC is positive at 0.165 but statistically insignificant. This coefficient does not imply that the dependent variable changes from ΔlnGDP to a level, since quarterly sectoral output growth remains the dependent variable in every specification. It represents instead the equilibrium component of the error correction mechanism, indicating how sectoral credit levels relate to lagged sectoral output within the adjustment process that governs subsequent growth. Its insignificance means that the linear model does not identify a stable long-run relationship between credit and output across the sectoral panel as a whole.

Behind that insignificant average, however, the sector-specific long-run estimates differ markedly. Several sectors return positive and significant long-run coefficients, the strongest being Mining and Quarrying and Other Industries (BDE) at 1.320, significant at the 1% level. In this group the long-run level of sectoral credit is positively associated with equilibrium output, an outcome that is economically coherent given the capital intensity of the underlying activities. Credit plausibly finances fixed investment, capacity expansion and infrastructure, whose returns accrue over extended horizons rather than within a single quarter.

Agriculture, Forestry and Fisheries (A) also returns a positive and significant long-run coefficient of 0.848 at the 1% level, in direct contrast to its negative short-run estimate. The divergence suggests that agricultural credit may carry short-run adjustment costs while still contributing positively to long-run equilibrium, the sector requiring time to convert borrowed funds into productive output. This is consistent with the view that credit-financed inputs, machinery, technology adoption and productivity-enhancing investment in agriculture raise output only with a lag, becoming more consequential as the investment horizon lengthens.

GHI likewise displays a positive long-run coefficient, estimated at 0.814 and significant at the 5% level. Short-run credit growth in this group is negatively associated with quarterly output growth, yet long-run credit levels are positively associated with equilibrium output. The contrast is economically meaningful, since trade, transportation, accommodation and food services depend heavily on working capital, logistics, inventories and demand conditions. Credit may therefore support long-run expansion by financing scale, operational continuity and service capacity even where short-run borrowing signals liquidity pressure.

Other Services (RSTU) returns a positive long-run coefficient of 0.265, significant at the 10% level. Although weaker in both magnitude and significance than the coefficients for BDE, A and GHI, the estimate indicates that the long-run contribution of credit varies even among service activities. The long-run coefficients for C, F, J, K, L, MN and OPQ, by contrast, are statistically insignificant, confirming that the equilibrium relationship between credit and sectoral output differs substantially across sectors and cannot be represented by a single average effect.

The negative and statistically significant error correction term of −0.620 indicates that deviations from the long-run relationship are corrected over time, confirming that estimation in CS-ECM form is appropriate. The residual cross-sectional dependence test, with a p-value of 0.544, further indicates that the inclusion of cross-sectional averages and their lags absorbs the shared shocks affecting the panel. The short-run and long-run coefficients may therefore be interpreted as sector-specific relationships purged of the common factors captured by the cross-sectionally augmented structure.

The contrast between a negative and significant short-run Mean Group coefficient and a positive but insignificant long-run coefficient suggests that the relationship between sectoral credit and output growth may not be linear at all. If the effect of credit varies across the distribution, a linear specification will average positive and negative regimes into a single slope that describes neither. Credit may raise growth by relaxing financing constraints while becoming less effective, or actively harmful, once it generates excessive leverage, debt overhang or inefficient allocation. Turning points of this kind lie beyond the reach of a linear model, which can only return a weak, insignificant or misleading average.

This reading is consistent with the wider nonlinear finance–growth literature. Arcand et al. (2015) find that financial depth ceases to contribute positively to output growth beyond a threshold, while Law and Singh (2014) show that the growth effect of finance depends on the level of financial development, with an optimal range beyond which additional finance yields diminishing returns. Samargandi et al. (2015) identify an inverted-U relationship between financial development and growth in middle-income countries. Read against this literature, the linear estimates reported here are best treated as a benchmark rather than as a description of the finance–growth relationship in the Turkish sectoral economy.

Two conclusions follow from the linear specification. Short-run increases in the logarithmic change of sectoral credit are on average negatively associated with quarterly sectoral output growth, an average driven principally by sectors A, GHI and OPQ. The long-run Mean Group coefficient remains statistically insignificant even though several sectors return positive and significant long-run estimates. Taken together, these results indicate that the effect of sectoral credit is conditional on both horizon and sector, providing the empirical justification for the quadratic and cubic specifications that allow the marginal effect of credit to vary with its level.

8.2. Quadratic CS-ECM Results: Evidence of Threshold Effects

The quadratic CS-ECM specification demonstrates that the relationship between sectoral credit and output growth in Türkiye is not adequately described by a single

slope. The linear model provides a reference point but confines the effect of credit to a constant marginal contribution, and it therefore cannot detect changes in that contribution once credit reaches particular levels. Adding the squared logarithmic credit term allows the marginal effect to depend on the prevailing level of sectoral credit, which is precisely what the 'too much finance' hypothesis requires, since that hypothesis holds that financial development supports growth at moderate levels while becoming less effective, or damaging, once credit exceeds an economy's capacity to absorb it.

The quadratic model is estimated in CS-ECM form, with short-run and long-run components identified separately. In the short-run component the relevant regressors are $\Delta \ln TCC$ and $\Delta[(\ln TCC)^2]$, so that the estimated coefficients describe the relationship between quarterly sectoral GDP growth and changes in the logarithmic credit polynomial rather than the direct contribution of the credit stock. The distinction matters, because $\Delta[(\ln TCC)^2]$ is not the square of $\Delta \ln TCC$ but the first difference of the squared log-credit level. The short-run quadratic estimates therefore reflect movements in the credit polynomial rather than powers of the quarterly credit growth rate. Table 42 reports the results of the quadratic specification.

Table 42. Quadratic CS-ARDL (CS-ECM) Short-Run and Long-Run Estimates: Türkiye Sectoral Panel

	Short-Run Estimates		**Long-Run Estimates**	
Sector Code	**(1) ΔlnTCC**	**(2) ΔlnTCC²**	**(3) lnTCC**	**(4) lnTCC²**
A	8.221* (4.628)	−0.267* (0.139)	−2.932* (1.725)	0.105** (0.048)
BDE	−0.546 (1.071)	0.013 (0.031)	0.196 (0.580)	0.035** (0.016)
C	2.570 (2.560)	−0.081 (0.066)	8.886 (6.675)	−0.217 (0.161)
F	1.609 (2.792)	−0.045 (0.075)	4.566 (4.097)	−0.143* (0.086)
GHI	4.385*** (1.206)	−0.125*** (0.032)	1.451 (1.102)	−0.016 (0.024)
J	2.530** (1.174)	−0.086** (0.037)	−1.295** (0.514)	0.044*** (0.017)

Table 42. (continued)

K	−0.302 (1.924)	0.009 (0.056)	7.403* (3.960)	−0.217* (0.116)
L	−0.223 (0.318)	0.011 (0.011)	0.137 (0.357)	−0.004 (0.011)
MN	−0.483 (0.928)	0.015 (0.029)	0.343 (0.334)	−0.009 (0.010)
OPQ	4.121*** (0.974)	−0.129*** (0.029)	0.111 (0.391)	0.000 (0.011)
RSTU	0.577 (2.960)	−0.027 (0.088)	2.502 (1.696)	−0.067 (0.051)
Mean Group	**2.042** (0.823)**	**−0.065** (0.026)**	**1.943* (1.095)**	**−0.044 (0.032)**
Obs. = 902	Groups = 11	ECT = −0.634*** (0.122)	R^2(MG) = 0.860	F = 11.24
CD = 0.41	p(CD) = 0.679			

Notes: Columns (1)–(2) report short-run and Columns (3)–(4) long-run estimates, with standard errors in parentheses. ECT is the error correction term. Sector codes are: A = Agriculture, Forestry and Fisheries; BDE = Mining and Quarrying and Other Industries; C = Manufacturing; F = Construction; GHI = Wholesale and Retail Trade, Transportation and Storage, Accommodation and Food Service Activities; J = Information and Communication; K = Financial and Insurance Activities; L = Real Estate Activities; MN = Professional, Scientific, Technical, Administrative and Support Service Activities; OPQ = Public Administration and Defence, Education, Human Health and Social Services Activities; RSTU = Other Services. ***, ** and * denote significance at the 1%, 5% and 10% levels.

At the Mean Group level the short-run coefficient on ΔlnTCC is positive and significant at the 5% level, estimated at 2.042, while the coefficient on Δ[(lnTCC)²] is negative and also significant at the 5% level, estimated at −0.065. A positive linear term combined with a negative quadratic term describes an inverted-U relationship in the short run. In economic terms, when the underlying log-credit level is comparatively low, changes in sectoral credit are positively associated with quarterly sectoral output growth, whereas the incremental contribution of further credit expansion declines as credit levels rise and eventually turns negative.

This finding materially alters the interpretation carried over from the linear model. There, the short-run Mean Group coefficient was negative, inviting the conclusion

that credit growth is generally detrimental in the short run. The quadratic specification shows that conclusion to be incomplete. The effect of credit is not uniformly negative but conditional on the position of each sector relative to its own threshold. Below the turning point, credit may raise output growth by easing liquidity constraints, financing working capital, supporting inventories and sustaining short-term production or service provision. Above it, further credit growth may be associated with heavier repayment burdens, less efficient allocation of funds, speculative or non-productive borrowing and greater financial vulnerability.

The estimated short-run Mean Group turning point is 15.79 in logarithmic terms, corresponding to approximately 7.23 billion TL. This value is obtained from the quadratic relationship in the log-credit polynomial and converted to its monetary equivalent by exponentiating the logarithmic threshold, as set out in Equation 30. Because the turning point lies within the observed range of lnTCC, it is empirically meaningful rather than an extrapolation beyond the support of the data (Haans et al., 2016; Lind & Mehlum, 2010). The result indicates that the short-run capacity of the Turkish sectoral economy to absorb credit is bounded, with further credit growth beyond a certain stock level associated with weaker sectoral output growth.

The Mean Group result aligns with the broader nonlinear finance–growth literature. Arcand et al. (2015) argue that financial depth ceases to contribute positively to output growth beyond a threshold, while Law and Singh (2014) identify an optimal range of financial development past which additional finance no longer supports growth. Samargandi et al. (2015) document an inverted-U pattern in middle-income countries. The present results are consistent with this evidence while extending it in a specifically sectoral direction, indicating that the threshold effect is not solely a macroeconomic regularity but also operates in the relationship between sector-specific credit and sector-specific output growth.

The diagnostics support the quadratic specification. The error correction term is negative and statistically significant at −0.634, confirming that deviations from the long-run relationship are corrected progressively. Residual cross-sectional dependence becomes statistically insignificant once cross-sectional averages and two

lags are included, with a CD p-value of 0.679, indicating that this structure effectively absorbs the shocks shared across sectors. The short-run inverted-U relationship should therefore be read as a sector-level phenomenon identified net of the unobserved common factors driving the Turkish economy as a whole.

At the sectoral level, the short-run quadratic estimates reveal considerable variation behind the Mean Group average. Statistically significant inverted-U patterns appear in four sectors: Agriculture, Forestry and Fisheries (A); Wholesale and Retail Trade, Transportation and Storage, Accommodation and Food Service Activities (GHI); Information and Communication (J); and Public Administration and Defence, Education, Human Health and Social Services Activities (OPQ). In each, the coefficient on $\Delta \ln TCC$ is positive while that on $\Delta[(\ln TCC)^2]$ is negative, and in each the implied turning point lies within the sample range, indicating that the short-run nonlinearity reflects genuine sectoral patterns rather than an artefact of averaging.

In sector A the short-run coefficients follow the expected inverted-U configuration and are significant at the 10% level, with $\Delta \ln TCC$ estimated at 8.221 and $\Delta[(\ln TCC)^2]$ at -0.267. The implied turning point is 15.41 in logarithmic terms, approximately 4.94 billion TL. This comparatively low threshold indicates limited short-run credit absorption capacity in agriculture, forestry and fisheries, which is consistent with the structural characteristics of the sector: agricultural output is governed by seasonal cycles, biological lags, weather uncertainty and delayed revenue realisation. Credit-financed expenditure may therefore fund inputs, inventories, machinery or liquidity needs without generating value added within the same quarter, while beyond the threshold debt servicing and financing costs may outweigh whatever immediate output gains are realised.

The low short-run threshold in sector A should not be read as evidence that agricultural credit is unproductive in general. It indicates rather that the timing and scale of credit are decisive. Moderate increases in credit may relax financing constraints, whereas rapid or excessive short-run credit growth may generate financial strain before the productive benefits materialise. This reading is consistent with the divergence between short- and long-run outcomes in agriculture, since the

long-run quadratic estimates for the same sector describe a U-shaped relationship. Agriculture therefore appears fragile over short horizons while retaining the capacity for adjustment and recovery over longer ones.

Information and Communication (J) records the lowest short-run turning point among the sectors displaying significant inverted-U patterns. The coefficients are significant at the 5% level, with ΔlnTCC estimated at 2.530 and Δ[(lnTCC)²] at −0.086, implying a turning point of 14.72 in logarithmic terms, approximately 2.46 billion TL. This low threshold suggests that the sector does not require large volumes of conventional bank credit to generate short-run output gains. Given its knowledge-intensive character and its reliance on intangible assets, human capital, digital infrastructure and innovation, additional bank credit beyond a modest level may not translate proportionally into immediate value added.

The result for sector J also speaks to the collateral basis of bank lending. Banks extend credit more readily to activities backed by tangible collateral, whereas information and communication activities rest disproportionately on intangible assets whose value is not always recognised in lending decisions. Beyond a relatively low credit threshold, therefore, further short-run credit growth appears less effective in raising current output. The short-run inverted-U pattern observed here should be interpreted as evidence of a low credit-volume threshold rather than as an argument that finance is unimportant to technology- and knowledge-intensive activities.

In OPQ, which covers public administration, defence, education, human health and social services, the short-run coefficients are significant at the 1% level, with ΔlnTCC estimated at 4.121 and Δ[(lnTCC)²] at −0.129. The implied turning point is 15.98 in logarithmic terms, approximately 8.68 billion TL, placing this sector above J and A but well below GHI. An inverted-U relationship is economically plausible here, since output in these activities is not determined by market-based credit expansion alone: public expenditure, institutional rules, service capacity, staffing constraints and regulatory frameworks may all shape value added independently of financing conditions.

Credit growth therefore supports OPQ activities only up to a point. At lower credit levels, additional financing may sustain education, health, administrative and social service provision, whereas beyond the threshold further short-run credit growth becomes less effective. This may reflect financing needs arising from delayed payments, institutional liquidity management or the expansion of activities whose output does not respond immediately to bank credit. The negative quadratic term indicates that the marginal benefit of short-run credit declines once the threshold is passed, consistent with the non-financial constraints that bind output in these activities.

GHI displays the highest short-run quadratic threshold in the panel. Both short-run coefficients are significant at the 1% level, with ΔlnTCC estimated at 4.385 and Δ[(lnTCC)²] at −0.125, implying a turning point of 17.51 in logarithmic terms, approximately 40.23 billion TL. This threshold is an order of magnitude above those for A, J and OPQ, indicating that GHI possesses a considerably greater short-run capacity to absorb credit than the other sectors exhibiting inverted-U dynamics.

The finding is economically coherent. GHI encompasses wholesale and retail trade, transportation, storage, accommodation and food services, activities tightly bound to working capital, inventories, logistics, distribution networks, tourism demand and domestic consumption. Credit can therefore support the regulation of operations, the financing of inventories, the maintenance of trade flows and the sustaining of service capacity across a wide range of volumes. Relative to agriculture or information and communication, GHI appears able to convert credit growth into value added over a broader interval, which is why its turning point occurs at a substantially higher level.

The negative quadratic term in GHI is nonetheless significant, indicating that even the sector with the greatest short-run absorption capacity faces limits. Credit expansion is associated with rising output only up to the threshold, beyond which the relationship reverses. Activities heavily dependent on working capital and liquidity are therefore not immune to the risks of excessive credit growth: once credit rises disproportionately relative to the sector's capacity to generate value added, the benefits of liquidity support may be offset by higher debt-servicing costs,

overtrading, inventory accumulation or heightened exposure to fluctuations in demand.

Taken together, the short-run quadratic results indicate that the Turkish sectoral economy is not governed by a single credit threshold. The Mean Group value of approximately 7.23 billion TL describes an overall tendency, yet the sectoral thresholds around it differ by more than an order of magnitude. J and A reach their critical points at comparatively low credit levels, OPQ occupies an intermediate position and GHI exhibits substantially greater absorption capacity. This dispersion is central to interpreting the results, indicating that a sector's response to credit expansion depends on its production structure, financing requirements, output cycle, collateral quality and its capacity to convert borrowed funds into value added within a short horizon.

The long-run quadratic results are more heterogeneous and less uniform than their short-run counterparts. At the Mean Group level the coefficient on lnTCC is positive and weakly significant at the 10% level, estimated at 1.943, while the coefficient on $(\ln TCC)^2$ is negative but statistically insignificant. The long-run estimates therefore provide no strong support for an inverted-U relationship holding across the sectoral panel as a whole. This contrasts with the short-run quadratic model, which identifies a clear average inverted-U pattern, and suggests that long-run nonlinearity is a sector-specific rather than an economy-wide phenomenon.

The long-run coefficients must be read within the CS-ECM framework. They do not imply that the dependent variable becomes the level of sectoral GDP, which remains $\Delta \ln GDP$ throughout. They represent instead the equilibrium component of the error correction mechanism, describing how sectoral credit levels and their squared transformations relate to lagged sectoral output within the adjustment process that generates subsequent quarterly growth. The long-run results therefore characterise the equilibrium role of credit levels, while the short-run results describe the influence of changes in the logarithm of credit.

In Financial and Insurance Activities (K), the long-run coefficients describe a statistically significant inverted-U relationship at the 10% level, with lnTCC

estimated at 7.403 and $(\ln TCC)^2$ at −0.217. The implied turning point is 17.08 in logarithmic terms, approximately 26.07 billion TL. Credit levels are therefore positively associated with the long-run equilibrium of financial and insurance activities up to that threshold, after which the marginal relationship turns negative.

The long-run inverted-U pattern in sector K is of particular interest for the 'too much finance' debate. Because K comprises financial and insurance activities themselves, the result indicates that moderate expansion of the financial sector may support economic activity while further expansion beyond a threshold diminishes its growth contribution. Over-expansion of financial activity may divert resources towards intermediation, speculation or balance-sheet management rather than productive investment. Cecchetti and Kharroubi (2015) advance precisely this argument, showing that a rapidly expanding financial sector depresses total factor productivity growth by favouring projects that are easily collateralised but comparatively unproductive, and by drawing skilled labour away from the real economy – an earlier study by the same authors having established the underlying threshold relationship (Cecchetti & Kharroubi, 2012). This is consistent with the pattern identified here.

Sectors A and J, by contrast, display statistically significant U-shaped long-run relationships. In sector A the coefficient on lnTCC is negative and significant at the 10% level while that on $(\ln TCC)^2$ is positive and significant at the 5% level, implying a turning point of 13.97 in logarithmic terms, approximately 1.16 billion TL. A U-shaped configuration means that the long-run relationship between credit levels and sectoral output begins negative and becomes positive beyond the turning point. This differs from the short-run inverted-U pattern in the same sector, underlining the importance of distinguishing short-run credit-growth dynamics from long-run credit-level effects.

The U-shaped long-run pattern in agriculture plausibly reflects the sector's adjustment process. At low levels, credit may generate costs without corresponding productivity gains, since small or fragmented borrowing is often directed towards liquidity rather than transformative investment, yielding a weak or negative association with long-run equilibrium. Once credit reaches sufficient scale, it can

finance machinery, irrigation, storage, technology adoption and modernisation, whose benefits accrue over extended horizons rather than in quarterly output. The U-shape for sector A therefore suggests that agricultural credit requires both scale and time before its long-run contribution becomes visible.

Sector J likewise exhibits a statistically significant U-shaped long-run relationship, with the coefficient on lnTCC negative and significant at the 5% level and that on $(\ln TCC)^2$ positive and significant at the 1% level. The implied turning point is 14.76 in logarithmic terms, approximately 2.56 billion TL. At lower credit levels the long-run association between credit and the sector's equilibrium output is negative, turning positive once credit passes a comparatively low threshold.

This pattern is consistent with the fixed costs and scalability characteristic of information and communication activities. Investment in digital infrastructure, software, data systems, telecommunications and related technologies typically involves substantial initial outlays and uncertain returns. At low credit levels, financing may be insufficient to generate meaningful productivity gains while still adding to debt burdens, whereas beyond a threshold credit may become adequate to fund scalable investment and realise network-related productivity effects. The long-run U-shape therefore indicates that the sector benefits from credit only where the scale of financing permits structural expansion, rather than through fragmented or short-term borrowing.

These U-shaped long-run results warrant careful interpretation. The 'too much finance' literature is predominantly concerned with inverted-U relationships, in which finance supports growth initially and impedes it beyond a threshold. A U-shape describes a different process, in which credit is costly or ineffective at low levels and becomes beneficial once it reaches sufficient scale. The long-run U-shapes identified for A and J should therefore not be generalised to the economy as a whole, but read as sector-specific adjustment processes reflecting production lags, fixed investment requirements, modernisation needs and the possibility that small-scale credit is simply insufficient to generate long-run productivity gains.

Some long-run curvature is also detected in BDE and F, though the evidence falls short of establishing complete quadratic threshold relationships. In BDE the coefficient on $(\ln TCC)^2$ is positive and significant at the 5% level while the linear term is insignificant, indicating curvature in the long-run association without a turning point identified by both polynomial terms jointly. The BDE result should accordingly be read as partial evidence of long-run nonlinearity rather than as a well-determined threshold.

A similar qualification applies to Construction (F), where the coefficient on $(\ln TCC)^2$ is negative and significant at the 10% level while the linear term is insignificant. This indicates negative curvature without constituting a statistically robust inverted-U relationship. Such curvature is economically plausible given the nature of construction activity, since credit-financed construction may support growth over certain ranges while losing effectiveness where leverage, overbuilding or speculative demand predominate. Because the two quadratic coefficients do not jointly delineate an interpretable threshold structure, however, the evidence remains partial rather than conclusive.

In the remaining sectors the long-run quadratic coefficients provide no strong statistical evidence of nonlinear threshold effects. This does not imply that credit is irrelevant in those branches, but that the quadratic specification does not recover a stable long-run U-shaped or inverted-U-shaped relationship. The absence of significant quadratic effects may reflect weaker reliance on bank credit, more complex adjustment processes, the predominance of non-financial determinants of output or the need for a more flexible functional form to characterise the long-run relationship adequately.

Overall, the quadratic results deliver a more granular account than the linear benchmark. The central finding is a clearly inverted-U short-run Mean Group relationship, in which changes in sectoral credit initially support quarterly output growth before their contribution turns negative beyond a threshold of approximately 7.23 billion TL. At the sectoral level this short-run threshold ranges widely, from roughly 2.46 billion TL in J to 40.23 billion TL in GHI. The long-run patterns are

more varied still: K displays an inverted-U configuration, A and J display U-shaped relationships, and BDE and F exhibit curvature without a fully identified threshold structure.

Two implications follow for the discussion that closes this chapter. First, the negative short-run result obtained from the linear model does not tell the whole story, since credit appears beneficial below sector-specific thresholds and harmful only beyond them. Second, the dispersion of turning points across sectors indicates that the capacity to absorb credit varies substantially within the Turkish economy. The sectoral effect of credit therefore depends not only on how much credit is extended but on the character of the sector, the horizon over which the effect is measured and the sector's position relative to its own threshold.

8.3. Cubic CS-ECM Results: Sectoral Asymmetry and Complex Long-Run Dynamics

The cubic CS-ECM specification extends the analysis by admitting a more flexible functional form for the finance–growth relationship. The quadratic model identifies a single turning point and is therefore restrictive wherever the marginal effect of credit changes direction more than once. The cubic specification accommodates such trajectories: a positive–negative–positive coefficient sequence may generate an N-shaped curve, while a negative–positive–negative sequence may generate an inverted-N shape. Configurations of this kind imply that the effect of credit is not merely nonlinear but may differ across successive stages of financial development.

The underlying economic logic is straightforward. In some sectors credit may initially raise output growth by relaxing financing constraints, supporting working capital or enabling investment, while beyond a certain point additional credit becomes less effective or harmful as debt costs, leverage, misallocation and speculative behaviour accumulate. At still higher levels the relationship may turn positive again if large investment projects begin to yield returns, if economies of scale are realised or if credit-financed expansion becomes embedded in the sector's productive structure, producing an N-shaped path. An inverted-N configuration

describes the converse: credit is associated with weak or negative performance at low levels, improves over an intermediate range and deteriorates again once debt burdens or inefficient allocation come to dominate.

Cubic results require particular caution in interpretation. Coefficient sign sequences such as positive–negative–positive or negative–positive–negative do not by themselves establish an N-shaped or inverted-N-shaped relationship in the strict sense of two turning points. A cubic function possesses two genuine turning points only where its derivative has two real roots, and where the discriminant is negative the function exhibits changing curvature or shifting momentum without local maxima and minima. In that case the correct description is an inflection point, located where the second derivative vanishes and the curvature of the relationship reverses while the slope retains its sign. This distinction proves decisive for sectors A, BDE and F, whose estimated coefficients yield inflection points rather than two real turning points.

As in the preceding specifications, the cubic model is estimated in CS-ECM form and the dependent variable remains ΔlnGDP throughout. The short-run cubic coefficients therefore describe the relationship between quarterly sectoral GDP growth and changes in the logarithmic credit polynomial – ΔlnTCC, Δ[(lnTCC)²] and Δ[(lnTCC)³] – rather than the effects of the credit level itself. It bears repeating that Δ[(lnTCC)²] and Δ[(lnTCC)³] are not the square and cube of ΔlnTCC, but the first differences of the squared and cubed transformations of the underlying log-credit level. The estimated short-run dynamics accordingly concern movements in the polynomial transformation of credit rather than powers of the quarterly credit growth rate. Table 43 reports the short-run and long-run estimates of the cubic specification.

At the Mean Group level the short-run cubic model provides only limited evidence of a complete cubic relationship across the sectoral panel. The coefficient on ΔlnTCC is positive at 10.235 and weakly significant at the 10% level, while that on Δ[(lnTCC)²] is negative at −0.522 and likewise weakly significant. The coefficient on Δ[(lnTCC)³] is positive at 0.008 but statistically insignificant. Short-run dynamics

at the panel level are therefore closer to a quadratic than to a fully cubic form, the first two terms reproducing the familiar inverted-U configuration while the third-order term receives no statistical support.

Table 43. Cubic CS-ARDL (CS-ECM) Short-Run and Long-Run Estimates: Türkiye Sectoral Panel

	Short-Run Estimates			**Long-Run Estimates**		
Sector	**(1) ΔlnTCC**	**(2) ΔlnTCC²**	**(3) ΔlnTCC³**	**(4) lnTCC**	**(5) lnTCC²**	**(6) lnTCC³**
A	44.041 (86.364)	−2.367 (4.952)	0.041 (0.094)	12.888 (8.520)	−0.769* (0.466)	0.016* (0.008)
BDE	−13.246 (14.647)	0.757 (0.848)	−0.015 (0.016)	24.360*** (6.713)	−1.388*** (0.390)	0.028*** (0.008)
C	9.919 (40.156)	−0.442 (2.072)	0.006 (0.035)	−55.004*** (12.522)	2.829*** (0.624)	−0.048*** (0.010)
F	12.733 (45.406)	−0.715 (2.530)	0.013 (0.047)	−40.582* (24.683)	2.453* (1.392)	−0.050* (0.027)
GHI	45.167** (20.426)	−2.320** (1.083)	0.039** (0.019)	−12.771*** (4.831)	0.734*** (0.250)	−0.014*** (0.004)
J	−7.745 (15.609)	0.543 (0.969)	−0.013 (0.020)	−6.865 (6.247)	0.395 (0.396)	−0.007 (0.008)
K	0.430 (21.686)	−0.026 (1.234)	0.001 (0.023)	−20.854 (27.020)	1.347 (1.480)	−0.029 (0.027)
L	0.585 (2.463)	−0.039 (0.160)	0.001 (0.003)	9.944*** (2.790)	−0.655*** (0.185)	0.014*** (0.004)
MN	10.264 (13.253)	−0.628 (0.793)	0.013 (0.016)	−3.506 (3.033)	0.227 (0.183)	−0.005 (0.004)
OPQ	−5.555 (16.268)	0.413 (0.937)	−0.010 (0.018)	−6.001* (3.197)	0.349* (0.185)	−0.007* (0.004)
RSTU	15.988 (48.069)	−0.921 (2.792)	0.017 (0.054)	4.215 (11.770)	−0.157 (0.667)	0.002 (0.013)
Mean Group	**10.235* (5.812)**	**−0.522* (0.316)**	**0.008 (0.006)**	**−8.561 (7.030)**	**0.488 (0.394)**	**−0.009 (0.008)**
Obs. = 902	Groups = 11	ECT = −0.709*** (0.103)	R^2(MG) = 0.860	F = 10.14	CD = −0.82	p(CD) = 0.411

Notes: Columns (1)–(3) report short-run and Columns (4)–(6) long-run estimates, with standard errors in parentheses. ECT is the error correction term. Sector codes are: A = Agriculture, Forestry and Fisheries; BDE = Mining and Quarrying and Other

Industries; C = Manufacturing; F = Construction; GHI = Wholesale and Retail Trade, Transportation and Storage, Accommodation and Food Service Activities; J = Information and Communication; K = Financial and Insurance Activities; L = Real Estate Activities; MN = Professional, Scientific, Technical, Administrative and Support Service Activities; OPQ = Public Administration and Defence, Education, Human Health and Social Services Activities; RSTU = Other Services. ***, ** and * denote significance at the 1%, 5% and 10% levels.

The Mean Group short-run results thus reinforce rather than contradict the quadratic findings. For the sectoral panel as a whole, short-run credit-growth dynamics are better described by a single-threshold adjustment process than by multiple reversals, the average short-run effect of credit rising and then falling rather than oscillating. This is consistent with the earlier evidence of a statistically significant inverted-U relationship in the short-run quadratic model, and introducing a third-order term does not materially alter that conclusion at the level of the panel average.

The diagnostics again support the specification. The error correction term is negative and statistically significant at −0.709, indicating that deviations from the long-run relationship are corrected progressively. The residual cross-sectional dependence test returns an insignificant statistic with a p-value of 0.411, confirming that the inclusion of cross-sectional averages and their lags absorbs the unobserved common factors. The cubic estimates should therefore be interpreted within a framework that already accounts for economy-wide shocks and sectoral interdependence.

The principal departure from the weak Mean Group short-run result occurs in GHI, where all three short-run cubic coefficients are significant at the 5% level and follow a positive–negative–positive sequence, with ΔlnTCC estimated at 45.167, $\Delta[(\ln TCC)^2]$ at −2.320 and $\Delta[(\ln TCC)^3]$ at 0.039. This configuration implies an N-shaped short-run relationship in the credit-growth dynamics of the sector, with two turning points located at 17.28 and 22.29 in logarithmic terms, corresponding to approximately 32.02 billion TL and 4,795.78 billion TL respectively.

The first turning point in GHI carries the greater economic content, marking the level at which the positive association between short-run credit changes and sectoral output growth begins to weaken. Since GHI comprises wholesale and retail trade,

transportation, storage, accommodation and food services, the finding is intuitive: these activities depend heavily on liquidity, working capital, inventories, logistics, tourism and domestic demand, so credit can support value added over a wide interval by financing inventories, trade flows, transport capacity and service provision. The first turning point nonetheless indicates that even in credit-dependent activities the marginal benefit of additional credit ultimately diminishes.

The second turning point for GHI lies substantially higher, at approximately 4,795.78 billion TL. Although this value falls within the observed range of lnTCC, it sits close to the upper tail of the credit distribution and therefore warrants cautious interpretation. It should not be read as evidence that very large credit volumes are welfare-enhancing, but as an indication that the shape of the fitted cubic function changes again at extreme credit levels. In practical terms, GHI possesses a strong short-run capacity to convert credit growth into value added, yet that capacity is bounded and becomes markedly nonlinear at the highest observed levels of credit.

The long-run cubic estimates display considerably greater sectoral heterogeneity. At the Mean Group level none of the long-run cubic coefficients is statistically significant, the estimates for lnTCC, $(\ln TCC)^2$ and $(\ln TCC)^3$ being −8.561, 0.488 and −0.009 respectively, all statistically indistinguishable from zero. There is therefore no uniform long-run cubic relationship across the sectoral panel: the Turkish sectoral economy does not follow a single N-shaped or inverted-N-shaped long-run finance–growth path, and long-run cubic effects operate at the level of individual sectors instead.

As before, these coefficients must be read within the CS-ECM framework rather than as a levels regression. The dependent variable remains quarterly sectoral output growth, $\Delta \ln GDP$, and the long-run coefficients represent the equilibrium component of the error correction mechanism, describing how sectoral credit and its polynomial transformations relate to lagged sectoral output within the adjustment process generating subsequent growth. The long-run cubic estimates therefore characterise

the shape of the credit–output relationship operating through the error correction term, not a static association between output levels and credit levels.

Among individual sectors, BDE returns a positive–negative–positive long-run sequence, with coefficients of 24.360 for lnTCC, −1.388 for $(\ln TCC)^2$ and 0.028 for $(\ln TCC)^3$, all significant at the 1% level. Although the sign pattern initially suggests an N-shaped relationship, the derivative of the fitted cubic function does not possess two real roots. The estimated coefficients instead imply an inflection point at 16.50 in logarithmic terms, approximately 14.62 billion TL, so BDE should not be characterised as displaying two distinct regime switches.

The appropriate reading of BDE is that the long-run association between credit and sectoral output changes in slope or momentum around the inflection point rather than passing through a local maximum or minimum. This is economically coherent for mining and quarrying; electricity, gas, steam and air conditioning supply; and water supply, sewerage, waste management and remediation activities. Such industries are characterised by high fixed costs, indivisible investments, infrastructure requirements, long planning horizons and delayed returns, so credit may support long-run capacity expansion and infrastructure financing while its influence evolves unevenly rather than smoothly. The inflection point therefore signals a change in the intensity of the credit–output relationship rather than a reversal in its direction.

Real Estate Activities (L) exhibits the clearest long-run N-shaped configuration in the cubic model, with coefficients of 9.944 for lnTCC, −0.655 for $(\ln TCC)^2$ and 0.014 for $(\ln TCC)^3$, all significant at the 1% level. The two turning points occur at 14.12 and 16.44 in logarithmic terms, corresponding to approximately 1.36 billion TL and 13.84 billion TL. Unlike BDE, the derivative here possesses two real roots, so the N-shaped description is warranted rather than approximate.

The long-run N-shaped pattern in real estate carries substantial economic meaning. At relatively low credit levels, rising sectoral credit may support real estate activity by financing purchases, construction, property development and collateral-based expansion. Beyond the first turning point the association weakens and turns negative, plausibly reflecting the accumulation of leverage, speculative pressure or the

declining productivity of additional credit. Beyond the second turning point it becomes positive once more, an outcome that should be interpreted cautiously since it may reflect the systemic weight of real estate, the role of collateral values in sustaining credit cycles or the delayed realisation of large investments. A positive association at high credit levels is not evidence of a benign outcome, given that credit-driven real estate booms may coincide with asset-price inflation, mounting financial vulnerability and boom–bust dynamics.

The real estate result connects directly to the credit-cycle literature, in which collateral values and borrowing capacity reinforce one another. Kiyotaki and Moore (1997) model an economy in which durable assets serve both as factors of production and as collateral for borrowing, showing that the dynamic interaction between credit limits and asset prices constitutes a transmission mechanism through which shocks persist, amplify and spill over to other sectors, so that rising property values relax collateral constraints and permit further borrowing and activity. Bezemer et al. (2016) provide complementary empirical evidence, reporting that credit directed towards real estate and other asset markets carries a negative growth coefficient while credit flows to non-financial business remain growth-enhancing. The N-shaped pattern in sector L should therefore be read as evidence of a complex and highly credit-sensitive long-run relationship rather than as an indication that additional credit reliably benefits real estate.

Agriculture, Forestry and Fisheries (A) also returns a positive–negative–positive long-run sequence, though on weaker evidence than BDE or L. The linear term is positive but statistically insignificant, while the squared and cubic terms are significant only at the 10% level. The derivative of the fitted function does not yield two real turning points, the estimated inflection point lying at 16.26 in logarithmic terms, approximately 11.52 billion TL. Agriculture should therefore not be treated as displaying a fully established N-shaped long-run relationship.

The results for sector A remain informative nonetheless. They indicate that the marginal effect of the credit level varies across the credit distribution even though two conventional turning points cannot be identified, which is consistent with the

quadratic findings for the same sector. In the short run agriculture displays a comparatively low inverted-U threshold, pointing to limited short-run absorption capacity, while the long-run quadratic model suggests a U-shaped relationship in which agricultural credit becomes more productive once it reaches sufficient scale. The cubic inflection point adds a further dimension, indicating that the long-run credit–output relationship in agriculture involves shifts in momentum rather than a simple monotonic or single-threshold structure.

Manufacturing (C) displays an inverted-N configuration, with long-run coefficients of −55.004 for lnTCC, 2.829 for $(\ln TCC)^2$ and −0.048 for $(\ln TCC)^3$, all significant at the 1% level. The two turning points are located at 18.35 and 20.67 in logarithmic terms, corresponding to approximately 92.80 billion TL and 951.93 billion TL. The long-run relationship between credit levels and manufacturing output therefore declines initially, rises between the two thresholds and declines again beyond the second.

This result is economically consequential, since manufacturing typically depends on external financing for machinery, technology, working capital and capacity expansion. The inverted-N shape indicates that credit is not unconditionally beneficial in this sector. At low credit levels, borrowing may be too limited to finance meaningful productive transformation while still imposing financial costs. Over the intermediate range, credit appears to support growth, plausibly because it reaches the scale required to fund investment and expand production. At very high levels the relationship turns negative again, potentially reflecting excessive leverage, inefficient investment, debt overhang or the financing of low-productivity firms.

The terminal negative segment in manufacturing is consistent with the broader concern that excessive or poorly allocated credit may impair productivity. Caballero et al. (2008) demonstrate that continued bank lending to weak or insolvent firms delays restructuring and distorts the allocation of resources across an industry. Although the present analysis does not identify such firms directly, the inverted-N pattern in manufacturing is compatible with the possibility that very high credit levels sustain less efficient structures and impede sectoral adjustment. Manufacturing

therefore exemplifies a long-run relationship in which credit is productive only within a bounded range.

GHI displays a statistically significant long-run inverted-N relationship, with coefficients of −12.771 for lnTCC, 0.734 for $(\ln TCC)^2$ and −0.014 for $(\ln TCC)^3$, all significant at the 1% level. The estimated turning points are 14.48 and 21.76 in logarithmic terms, corresponding to approximately 1.94 billion TL and 2,828.45 billion TL. The sector therefore exhibits an N-shaped relationship in the short run and an inverted-N-shaped relationship in the long run, a reversal that is among the most striking results in the panel.

The contrast between the short-run and long-run cubic patterns in GHI underlines the importance of separating credit-growth dynamics from credit-level effects. In the short run, changes in credit may support trade, transportation, accommodation and food services through liquidity and working-capital channels, since these activities require continuous financing for inventories, logistics, operating costs and service provision. Over longer horizons, accumulated credit levels interact with demand cycles, cost pressures and the build-up of debt in more complex ways. The long-run inverted-N pattern indicates that GHI benefits from credit within a bounded interval, with both very low and very high credit levels associated with weaker equilibrium outcomes.

The second long-run turning point in GHI, at approximately 2,828.45 billion TL, is notably high and requires careful interpretation given its proximity to the upper end of the observed credit range. It nonetheless falls within the sample and should not be dismissed as an out-of-sample artefact. Read together, the short-run and long-run estimates establish GHI as among the most credit-sensitive sectors of the Turkish economy, its relationship with credit differing substantially according to the horizon over which it is measured.

OPQ likewise displays an inverted-N configuration in the long run, with coefficients of −6.001 for lnTCC, 0.349 for $(\ln TCC)^2$ and −0.007 for $(\ln TCC)^3$, all significant at the 10% level. The two turning points occur at 14.48 and 21.21 in logarithmic terms, corresponding to approximately 1.95 billion TL and 1,631.88 billion TL. Since OPQ

covers public administration, defence, education, health and social services, the relationship between sectoral credit and output is likely to be mediated by institutional and public-sector factors, output in these activities depending on public budgets, regulatory rules, staffing, service capacity and administrative frameworks as much as on market demand or private credit conditions.

The inverted-N pattern in OPQ should therefore not be interpreted in the same terms as in manufacturing or trade. At low credit levels, additional credit may not raise output appreciably, since productivity and service delivery depend on institutional capacity and public funding. Over the intermediate range, credit may support the maintenance or expansion of education, health and social services. At very high levels its effectiveness may decline again, plausibly because sectoral output is constrained by non-financial factors. The OPQ result thus reinforces the general finding that credit effects are sector-specific and cannot be transferred across activities with different institutional foundations.

Construction (F) returns a long-run cubic sequence alternating between negative and positive signs, with coefficients of −40.582 for lnTCC, 2.453 for $(\ln TCC)^2$ and −0.050 for $(\ln TCC)^3$, all significant at the 10% level and describing an inverted-N configuration. As in sectors A and BDE, however, the derivative analysis yields no two real turning points, the estimated inflection point lying at 16.43 in logarithmic terms, approximately 13.69 billion TL. Construction therefore does not display a conventional inverted-N relationship with two distinct thresholds.

The inflection-point analysis for construction indicates a change in the strength of the long-run credit–output relationship rather than a complete reversal across regimes. This is economically plausible given the character of the activity, since construction is highly credit-sensitive and closely tied to collateral values, real estate cycles, leverage and long-duration projects (Kiyotaki & Moore, 1997). Credit-financed construction may support activity over the short and medium term while losing effectiveness where overbuilding, speculative demand or excessive leverage predominate. The negative–positive–negative sequence, combined with the inflection

point, therefore indicates nonlinear credit-driven dynamics without supporting a two-threshold description.

The remaining sectors do not display statistically robust cubic dynamics. In sector J the long-run cubic coefficients are insignificant, whereas the quadratic model identifies a long-run U-shaped relationship, indicating that a second-order specification captures the nonlinearity in Information and Communication adequately and that adding a cubic term yields no statistical gain. The same holds for sector K, where the long-run cubic coefficients are insignificant while the quadratic model identifies a long-run inverted-U pattern, so that financial and insurance activities are better represented by the quadratic than by the cubic specification.

MN and RSTU likewise display no statistically significant cubic nonlinearity. For MN, covering professional, scientific, technical, administrative and support services, this suggests that the link between sectoral credit and output is less directly reflected in aggregate sectoral bank credit volumes, these activities depending more on human capital, knowledge intensity, contractual service relationships and organisational capacity than on large-scale bank financing. For RSTU, the linear model returns a weakly positive long-run coefficient while neither the quadratic nor the cubic specification identifies a stable nonlinear threshold structure, indicating a long-run relationship with credit that is weakly linear or limited rather than distinctly nonlinear.

Overall, the cubic results refine rather than overturn the findings of the linear and quadratic specifications. At the Mean Group level the cubic model corroborates the short-run inverted-U pattern identified by the quadratic model, the average short-run cubic term being statistically insignificant and therefore providing no evidence of a full cubic relationship across the panel. Sector-specific results nonetheless reveal considerable long-run complexity: Real Estate Activities follows a clear N-shaped path; Manufacturing, GHI and OPQ display inverted-N configurations; and BDE, A and F exhibit inflection points rather than conventional two-turning-point curves.

These findings bear directly on the discussion of sectoral credit policy. The relationship between credit and output is heterogeneous across sectors and

conditional on the horizon of analysis, with some sectors adequately described by quadratic thresholds while others require cubic specifications to capture long-run shifts in the credit–output relationship. The cubic evidence therefore indicates that sectoral credit effects cannot be reduced to a single national threshold or to a uniformly positive or negative characterisation, credit instead supporting, impeding or reshaping sectoral output according to the sector, the credit level and the time horizon under consideration.

8.4. Overall Interpretation and Discussion

Taken as a whole, the empirical results indicate that the relationship between sectoral credit and sectoral GDP growth in Türkiye is heterogeneous, horizon-dependent and nonlinear. The linear, quadratic and cubic CS-ECM estimates should be read as complementary rather than competing evidence. The linear model establishes what would be concluded under a constant marginal effect of credit; the quadratic model demonstrates that this effect changes beyond a threshold in the short run; and the cubic model shows that, over longer horizons, several sectors follow trajectories that a single turning point cannot represent.

The linear specification returns a negative and statistically significant short-run association between changes in log sectoral credit and sectoral GDP growth, so that a larger quarterly logarithmic increase in sectoral credit coincides with weaker quarterly output growth. This should not be construed as evidence that credit is harmful in general. It illustrates instead what happens when a relationship that is potentially nonlinear and demonstrably heterogeneous is compressed into a single average slope. The linear model is therefore valuable precisely because it exposes the limits of a strictly linear reading.

The long-run Mean Group coefficient in the linear model is positive but statistically insignificant, indicating that no stable long-run relationship between sectoral credit levels and the equilibrium component of sectoral output can be identified across the panel as a whole. That insignificance conceals substantial sectoral variation, since BDE, A, GHI and RSTU all return positive and statistically significant long-run coefficients. The absence of a significant average effect therefore reflects the

dispersion of sector-specific effects rather than the irrelevance of credit, and it cannot be summarised by a single economy-wide coefficient.

Introducing the quadratic terms alters the interpretation appreciably. At the Mean Group level the short-run quadratic model identifies a statistically significant inverted-U relationship, the coefficient on ΔlnTCC being positive and significant and that on $\Delta[(\ln TCC)^2]$ negative and significant. Initial increases in the logarithmic credit terms therefore support quarterly sectoral GDP growth, while their contribution turns negative beyond a threshold estimated at 15.79 in logarithmic terms, approximately 7.23 billion TL. Because this turning point falls within the sample range, it constitutes concrete evidence of a bounded short-run capacity to absorb credit in the Turkish sectoral economy.

This result is central to interpreting the findings as a whole, since it establishes that the negative short-run coefficient of the linear model is an artefact of aggregation rather than a description of the underlying relationship. Credit does not uniformly depress growth in the short run; its influence varies with its level. Below the estimated threshold, increases in credit tend to support sectoral output growth, plausibly by relaxing liquidity constraints, financing inventories, supporting working capital and sustaining continuity of production or service provision. Above it, further expansion is associated with slower growth, potentially reflecting rising debt-service burdens, inefficient allocation of funds, balance-sheet vulnerability, speculative borrowing or the simple inability of a sector to convert additional debt into immediate value added.

The pattern is consistent with the 'too much finance' hypothesis. Arcand et al. (2015) show that financial depth ceases to promote growth beyond a threshold, Law and Singh (2014) demonstrate that the finance–growth link exhibits an optimal level rather than a monotonically positive relationship, and Samargandi et al. (2015) identify an inverted-U pattern in middle-income countries. Sahay et al. (2015) reach a comparable conclusion for emerging markets, arguing that deepening remains beneficial but is bounded in both size and speed. The present results support this literature while extending it in an important respect: the threshold effect is not exclusively macroeconomic but operates at the sectoral level, with the critical

threshold differing markedly from one branch of activity to another. The sector-specific quadratic estimates make that heterogeneity explicit. Significant short-run inverted-U relationships appear in A, GHI, J and OPQ, yet their turning points differ by more than an order of magnitude. Information and Communication (J) reaches its threshold at approximately 2.46 billion TL and Agriculture, Forestry and Fisheries (A) at roughly 4.94 billion TL, while OPQ turns at around 8.68 billion TL and GHI at approximately 40.23 billion TL. Such dispersion indicates that the capacity to absorb credit varies substantially across sectors within a single national economy.

The comparatively low thresholds in A and J indicate greater sensitivity to short-run credit growth. In agriculture, production cycles, biological lags, seasonality and delayed output realisation may prevent credit from being converted into value added within the quarter in which it is drawn. In information and communication, the low threshold plausibly reflects the knowledge-intensive and intangible-asset-driven character of the sector, in which large-scale bank credit is not the principal determinant of short-run output. The considerably higher threshold in GHI, by contrast, matches the operational dynamics of trade, transportation, storage, accommodation and food services, activities dependent on working capital, inventories, logistics and demand conditions, in which credit can support output across a wider interval before its marginal contribution declines.

The long-run quadratic estimates add a further layer to the analysis. At the Mean Group level the long-run quadratic relationship is not strongly supported, the coefficient on lnTCC being positive and only weakly significant while the squared term remains insignificant. Sector-specific results nonetheless reveal distinct long-run patterns. Financial and Insurance Activities (K) displays an inverted-U configuration with a turning point at approximately 26.07 billion TL, indicating that even within the financial sector there exists a credit level beyond which further expansion ceases to raise the sector's contribution to output. This finding is consistent with the argument that an oversized financial sector may divert resources from more productive uses and depress aggregate efficiency (Cecchetti & Kharroubi, 2012, 2015).

Sectors A and J, by contrast, exhibit U-shaped long-run relationships within the quadratic model, which is particularly notable given that both also display relatively low short-run inverted-U thresholds. The combination implies that short-run credit fluctuations and long-run credit levels operate through different mechanisms in these sectors. Rapid credit growth may initially generate adjustment costs or financial strain before any output response materialises, whereas over longer horizons credit of sufficient scale may finance modernisation, fixed investment, technological upgrading or capacity expansion. A sector that appears fragile in the short run may therefore still derive substantial benefit from credit over a longer horizon.

The cubic results reveal greater long-run complexity still. No statistically significant common long-run cubic relationship is detected at the Mean Group level, yet several individual sectors display significant cubic dynamics. The complexity of the finance–growth relationship is therefore not an economy-wide average phenomenon but the product of sector-specific adjustment processes. Some sectors are adequately represented by a quadratic specification, while others require a cubic form to characterise the long-run credit–output relationship accurately.

Real Estate Activities (L) follows a clear N-shaped long-run path, in which credit initially raises sectoral value added, subsequently sees its marginal contribution weaken and turn negative, and then becomes positively associated again beyond the second turning point. The configuration accords with economic intuition, given the close ties between real estate activity and collateral values, credit cycles, asset prices and leverage (Kiyotaki & Moore, 1997). The terminal positive segment should not be read as evidence that very high credit levels are beneficial, since in real estate a positive association at high credit levels may coincide with asset-price inflation, speculative behaviour and mounting financial vulnerability. Bezemer et al. (2016) reinforce this caution, finding that credit directed to real estate and other asset markets is associated with weaker growth than credit extended to non-financial business.

Manufacturing (C), GHI and OPQ display inverted-N-shaped long-run relationships. In manufacturing, the pattern indicates that credit is associated with weaker

outcomes at low levels, supports growth over an intermediate range and becomes negatively associated again at very high levels. This matters because manufacturing relies substantially on external finance for machinery, investment, working capital and capacity expansion. The inverted-N configuration implies that credit becomes productive only once it reaches a sufficient scale, while excessive credit may ultimately impair performance through debt overhang, inefficient investment or the financing of low-productivity activity.

The inverted-N pattern in GHI is especially notable because the same sector displays an N-shaped relationship in the short-run cubic model. The contrast underscores the importance of distinguishing the short-run and long-run components of the CS-ECM framework. In the short run, credit fluctuations may support GHI through liquidity and working-capital channels, whereas over longer horizons accumulated credit levels interact with demand cycles, logistics capacity, cost pressures and debt accumulation in more intricate ways. GHI therefore emerges as a sector in which the credit–output relationship is both economically significant and strongly conditional on the horizon of analysis.

OPQ likewise exhibits an inverted-N long-run pattern, though its interpretation differs from that of manufacturing or trade, since the group covers public administration, education, health and social services. Output in these activities responds to institutional arrangements, public expenditure, service capacity and regulation, so the long-run influence of credit is unlikely to operate primarily through market-based investment. The inverted-N configuration indicates that credit is not irrelevant in this sector, but that its influence is mediated by non-financial constraints and institutional frameworks that bound the response of output to financing conditions.

BDE, A and F require the most cautious interpretation. Although their cubic coefficients display sign sequences resembling N-shaped or inverted-N-shaped curves, the derivatives of the fitted functions do not yield two real turning points, the calculations instead identifying inflection points. These sectors should therefore not be classified as displaying conventional N-shaped or inverted-N-shaped

relationships, the cubic functions implying changes in momentum or intensity around the inflection point rather than regime reversals. For BDE this plausibly reflects large fixed investment requirements and delayed returns in mining and quarrying; electricity, gas, steam and air conditioning supply; and water supply, sewerage, waste management and remediation activities. For A it is consistent with the dependence of agricultural credit effects on scale, timing and production lags. For F, the inflection-point structure matches the credit-sensitive character of construction and the tendency for leverage-driven expansion to become less productive over time.

The evidence assembled here therefore admits no one-size-fits-all account of the finance–growth relationship in Türkiye. The influence of credit varies with the sector, the time horizon, the functional form and the sector's position relative to its own threshold. A single sector may display a negative linear coefficient in the short run, a quadratic threshold effect at an intermediate level of aggregation and a distinct nonlinear adjustment pattern in the long run. The linear, quadratic and cubic results must consequently be interpreted jointly rather than treated as competing candidates for a single correct specification.

The findings also fix the sense in which 'credit expansion' is used throughout this chapter. In the short-run results the regressors are changes in the logarithmic credit terms – $\Delta \ln TCC$, $\Delta[(\ln TCC)^2]$ and $\Delta[(\ln TCC)^3]$ – which describe how credit-growth dynamics relate to quarterly sectoral output growth. Because the differenced polynomial terms are products of credit growth and the credit level, however, the short-run marginal effect is itself a function of the prevailing level of log credit, which is what allows short-run thresholds to be expressed as credit volumes at all. The long-run results instead concern the level of sectoral credit and its polynomial transformations as they enter the error correction component. Interpretation therefore differs between the two horizons even though $\Delta \ln GDP$ remains the dependent variable in every specification.

Economically, the results suggest that credit supports sectoral growth where it is matched to a sector's capacity to convert financing into value added. That capacity depends on production technology, investment horizon, working-capital

requirements, collateral structures, demand conditions and the weight of public or institutional determinants of output. Sectors with short production cycles or substantial working-capital needs use credit differently from those with long gestation periods or predominantly intangible assets. Sectors whose activity is tied to collateral and asset prices may in turn display more pronounced nonlinearities than those driven principally by human capital or public expenditure (Beck et al., 2012; Bezemer et al., 2016).

Beyond sector-specific thresholds, additional credit may weaken output growth through several channels. It may raise debt-service costs, constrain firms' investment capacity, finance low-return activities, fuel speculative dynamics or postpone necessary restructuring. In manufacturing, excessive credit may sustain inefficient firms and depress measured productivity, an interpretation consistent with the mechanism documented by Caballero et al. (2008). In real estate and construction it may amplify collateral cycles and asset-price fluctuations. In services and public activities, additional credit is likely to have limited effect where output is constrained by demand, institutional capacity or other non-financial factors.

The policy implication is not that credit should be expanded or restricted in the aggregate. The results argue instead for a sector-differentiated approach to financial development, since policies anchored solely in aggregate credit indicators risk overlooking substantial variation in sectoral absorption capacity. A uniform credit ceiling would be excessively restrictive for sectors with high productive capacity and insufficiently binding for those that reach their limits at comparatively low credit levels. By the same logic, broad-based credit expansion may benefit some sectors while generating instability in others.

Macroprudential policy should therefore attend to the composition of credit across sectors rather than to aggregate credit volume alone. Sector-specific credit limits, the maturity structure of lending, the intended use of credit and each sector's demonstrated capacity to convert credit into value added all constitute relevant design parameters for more targeted financial regulation. This orientation is consistent with the evidence of Beck et al. (2012) and Bezemer et al. (2016), both of

which show that the growth consequences of credit depend on where it is allocated rather than on how much of it is extended. The results reported here indicate specifically that the growth contribution of credit depends on whether it finances productive investment, working capital, modernisation and capacity building, or instead supports debt rollover, speculative activity and low-efficiency expansion.

In summary, the empirical evidence supports a nonlinear and sector-specific characterisation of the finance–growth relationship in Türkiye. A simple linear model would suggest a negative short-run effect together with a statistically insignificant long-run average. The quadratic specification shows the short-run relationship to be inverted-U-shaped, with a Mean Group threshold of approximately 7.23 billion TL. The cubic specification reveals substantial variation in long-run dynamics across sectors, with individual branches displaying N-shaped, inverted-N-shaped or inflection-point structures. The influence of sectoral credit on output growth therefore depends on sector-specific conditions rather than operating uniformly, credit promoting growth only where it is aligned with a sector's absorption capacity, its productive use and the timing of its deployment.

CHAPTER 9

DISCUSSION AND CONCLUSION

This dissertation set out to re-examine the relationship between the quantitative dimension of financial development and economic performance without assuming that the effects of credit are linear, uniform across sectors, stable over time or invariant to the stage of development. Two complementary investigations were conducted to that end. The first estimated the association between the pace of credit expansion and the sectoral composition of value added in a cross-country panel covering 1992 to 2019, comparing the pre-crisis expansion with the post-crisis decade and the full sample, and comparing economies grouped by World Bank income classification (World Bank, 2026b), with coverage reaching 186 countries in the industrial and service models and 109 countries in the agricultural models. The second constructed a matched quarterly panel for Türkiye, pairing the total cash credit extended to each of eleven activity groups with the value added that activity generates, which permits short-run adjustment and long-run equilibrium to be estimated separately while common macroeconomic shocks and sector-specific slopes are accommodated within a single framework.

The central conclusion is that no single coefficient, and no universal credit threshold, can characterise the relationship between credit and real activity. The evidence supports neither the proposition that additional credit raises output proportionally in every sector nor the contrary proposition that credit expansion is generally harmful, indicating instead that the sign, the magnitude and the functional form of the relationship depend jointly on the sector receiving the funds, the pace at which credit expands relative to that sector's capacity to absorb it, the stage of development, the prevailing financial regime and the horizon over which the effect is measured. This position extends rather than contradicts the movement of the literature away from

monotonic accounts of finance and growth towards diminishing returns, credit composition and context-dependent effects (Arcand et al., 2015; Botev et al., 2019; Law & Singh, 2014; Rioja & Valev, 2004a). What the present study adds is the demonstration that these complications are not peripheral qualifications to an otherwise stable relationship but constitute the structure of the relationship itself.

Two features of the empirical design govern the interpretation of every result discussed below. The first concerns the margin at which finance is measured. Because private credit by deposit money banks and credit to agriculture, forestry and fishing are integrated of order one while the sectoral value-added shares are stationary in levels, the credit variables enter the cross-country equations in first differences, so the estimated parameters describe the pace of credit expansion rather than the accumulated depth of the financial system. The turning points recovered from these specifications are therefore located in credit-growth space and are not comparable with the depth thresholds of eighty to one hundred per cent of GDP reported in the aggregate ‘too much finance’ literature (Arcand et al., 2015; Cecchetti & Kharroubi, 2012; Law & Singh, 2014). The Türkiye chapter is the only part of the analysis in which both margins are recovered simultaneously, since the error-correction representation retains sectoral credit levels in the long-run component while modelling changes in the logarithmic credit terms in the short-run component. Even there the separation is not absolute, because the differenced polynomial terms are products of credit growth and the prevailing credit level, which is precisely what allows the short-run turning points to be expressed as credit volumes.

The second feature concerns what the dependent variables measure. In the international models the outcome is sectoral value added expressed as a share of GDP, so a negative coefficient indicates that a sector grew more slowly than the rest of the economy rather than that it contracted in absolute terms, an interpretation the structural transformation literature treats as a normal feature of development rather than as evidence of sectoral failure (Herrendorf et al., 2014). In the Türkiye models both credit and output are observed at sectoral resolution, yet the series are nominal, unadjusted for seasonality and expressed in current prices, and the credit measure records the stock of total cash loans including non-performing receivables. The study

accordingly measures the volume and the pace of credit intermediation rather than its quality, its maturity structure, its price or its allocation at the borrower level. Mechanisms invoked throughout the discussion – collateral quality, allocative efficiency, debt rollover, technological capacity and institutional strength – should therefore be read as explanations proposed in the literature and consistent with the pattern of estimates, not as channels identified directly within the models.

9.1. Answers to the Research Questions

The first question asked whether the pace of credit expansion bears a stable relationship to sectoral value added or whether its marginal contribution changes as expansion accelerates. The answer is unambiguous for industry and qualified elsewhere. Industry is the only sector in which both estimators return statistically significant credit coefficients in all three periods, and the fitted relationship is nonlinear in every one of them, requiring a cubic form under the static estimator in the post-crisis decade and over the full sample. Services in the post-crisis decade display a positive first-order effect whose implied reversal lies beyond the ninety-ninth percentile of observed credit growth, which makes the relationship monotonic for practical purposes despite significant higher-order terms. Agriculture yields curvature in only one specification, and that curvature reverses sign once macroeconomic controls are introduced. The marginal contribution of credit growth is therefore variable rather than constant, though the evidence for variability is strong in industry, weak in agriculture and confined to a single decade in services.

The second question concerned whether these patterns are common across sectors or specific to each, and the periodical results answer it decisively. In the post-crisis decade the same regressor, the same denominator and the same estimation window produce a negative linear coefficient of -0.167 for industry and a positive coefficient of 0.141 for services, an asymmetry that cannot be attributed to the common GDP scaling, since a scaling artefact would displace both coefficients in the same direction. Sectors differ not merely in the location of their turning points but in the order of the polynomial their data support and in the sign of the association itself. This divergence is the central empirical result of the cross-country analysis,

indicating that an aggregate credit coefficient averages over sectoral responses that differ qualitatively and not only in magnitude.

The third question asked whether the relationship changed after the global financial crisis. It did, and the change is sectorally specific rather than a uniform shift in level. Before the crisis the static estimator returned a weakly positive industrial coefficient of 0.036 that did not survive the addition of controls, while the dynamic estimator identified negative curvature confined to very rapid expansions, with a local maximum at 0.510 lying just below the ninety-ninth percentile of the pre-crisis distribution. In the post-crisis decade the marginal association turned negative across almost the entire observed range, the upper turning point of 0.246 sitting between the ninety-fifth and ninety-ninth percentiles, and services simultaneously acquired the positive and robust first-order effect they display in no other period. Faster credit expansion accompanied a rising service share and a falling industrial share within the same decade and the same sample, a configuration consistent with a reorientation of bank lending towards property, consumption and service activities (Bezemer et al., 2016), although the aggregate credit measure available here cannot establish that reorientation directly.

The fourth question concerned heterogeneity across stages of development, and the income-level estimates confirm that no single finance–sector curve applies across low-, middle- and high-income economies. Agricultural credit growth registers only in the low-income group, where a weakly supported inverted U places the maximum at 0.414 and leaves the great majority of observations on the upward-sloping segment. Industry registers in all three groups, but with different functional forms and with the estimated maximum shifting steadily leftward as income rises, from a cubic profile whose dominant middle segment slopes downward in low-income economies, through an approximately concave profile in middle-income economies with the maximum at 0.309 under the static estimator and at 0.026 under the dynamic one, to a maximum of -0.224 in high-income economies that lies below zero and therefore places the whole positive range of credit growth on the declining branch. Services respond only in the middle-income group, and only weakly. The middle-income industrial evidence provides the clearest support in this study for the ‘too

much finance' proposition, and it arises in the sector for which that proposition has the strongest theoretical rationale, since manufacturing remains a substantial share of output in these economies while supervisory capacity is still maturing (Cecchetti & Kharroubi, 2015; Sahay et al., 2015).

The fifth question asked whether the estimated relationships survive the introduction of dynamic persistence and the instrumentation of potentially endogenous credit, and the answer is partial in a way that is itself informative. In industry the dynamic estimator corroborates the static one in every period, reproducing the negative cubic term over the full sample at -0.343 with an implied local maximum of 0.156 close to the static value of 0.216, and returning a negative linear coefficient of -0.298 in the post-crisis decade. In agriculture and services the System GMM specifications yield no significant credit coefficients, and in several income-level models the instrument count relative to the number of country groups leaves the overidentification tests uninformative in the sense described by Roodman (2009a, 2009b). The lagged dependent variable is estimated between roughly 0.91 and 0.96 in every dynamic specification whose diagnostics are satisfactory, confirming that sectoral shares are highly persistent and that credit growth operates alongside rather than independently of that persistence. Fixed-effects estimation with Driscoll–Kraay standard errors remains reliable under heteroskedasticity, serial correlation and cross-sectional dependence (Driscoll & Kraay, 1998), so the static results retain their standing as conditional associations, but the divergence between estimators restricts how far any of these associations can be given a causal reading.

Turning to the national evidence, the sixth question asked how changes in sectoral credit relate to short-run output growth, what long-run relationship links sectoral credit levels to sectoral activity and whether sectors differ in their capacity to absorb credit productively. The linear specification returns a negative and significant Mean Group short-run coefficient of -0.213 alongside a positive but insignificant long-run coefficient of 0.165, a combination that would suggest a puzzle if either estimate described the panel adequately. Neither does. The quadratic specification shows the negative short-run average to be an artefact of imposing a constant slope, since the short-run relationship is inverted U-shaped, with credit changes supporting quarterly

output growth up to approximately 7.23 billion Turkish liras and their contribution turning negative beyond it. The insignificant long-run average likewise conceals rather than summarises the sectoral evidence, given that mining, quarrying and other industries, agriculture, the trade-related group and other services all return positive and significant long-run coefficients, the largest being 1.320 for mining, quarrying and other industries. Absorptive capacity differs by more than an order of magnitude across activities, the short-run turning points running from roughly 2.46 billion liras in information and communication to 40.23 billion liras in trade, transport, accommodation and food services.

The seventh question runs across both settings and follows from the integration properties of the data, asking whether the pace of credit expansion and the level of financial depth are separable margins carrying different implications for real activity. The Turkish estimates answer it directly, since the same sector may display opposite curvature at the two margins. Agriculture and information and communication combine a short-run inverted U, in which changes in credit support output only up to a comparatively low threshold, with a U-shaped long-run profile in which credit levels are associated with weaker equilibrium output until they reach sufficient scale to finance modernisation, fixed investment or technological upgrading. These are not contradictory findings but descriptions of different objects, and their divergence corroborates the encompassing result of Loayza and Rancière (2006), who reconcile a positive long-run relationship between intermediation and growth with a largely negative short-run relationship linked to financial fragility. Treating the stock of credit and its rate of expansion as interchangeable would therefore misstate the evidence at both horizons.

9.2. The Cross-Country Evidence in Perspective

9.2.1. Agriculture: Weak Nonlinearity and the Primacy of Institutional Conditions

The agricultural results are the weakest in the cross-country analysis, and the most useful thing that can be said about them is why weakness is the expected outcome

rather than a symptom of insufficient statistical power. Neither sub-period yields evidence that the growth of agricultural credit relative to GDP is systematically related to the agricultural share of value added, and over the full sample only the quadratic fixed-effects specification attains conventional significance. The Lind and Mehlum (2010) test rejects monotonicity at the ten per cent level and locates the maximum at 0.502, equivalent to annual agricultural credit growth of roughly sixty-five per cent, with a positive and significant slope at the lower bound of the data and a negative slope at the upper bound. That configuration would ordinarily support an inverted U, yet the eighty per cent Fieller (1954) interval for the turning point spans -0.719 to 3.329 and thus accommodates both negative values and implausibly rapid expansions, while the curvature reverses sign once human capital, capital formation, government consumption, inflation and trade openness are held constant.

The defensible reading is that agricultural credit growth exhibits at most a weak and imprecisely located nonlinearity. Financial development theory holds that credit raises production only where intermediaries reduce information asymmetries and direct resources towards investments capable of raising productivity, which makes the efficiency of allocation rather than the volume of lending the operative mechanism (Levine, 2005). Rural credit markets in much of the sample depart substantially from that benchmark, being characterised by high transaction costs, weak collateral institutions, thin insurance provision and pronounced exposure to covariate risk, all of which limit the extent to which additional lending is converted into mechanisation, irrigation or improved inputs (Conning & Udry, 2007). Under such conditions a rapid expansion of agricultural credit may finance liquidity needs, refinancing or consumption smoothing rather than capital formation, leaving the sectoral share unaffected even where credit volumes rise substantially. Behaviour of this kind is not captured by an annual credit-growth variable, which is one reason why the institutional character of rural finance should be regarded as the more informative object of study than the aggregate lending volume.

The control variables in the extended specification support this reading indirectly, since they rather than credit account for most of the explanatory power. Government consumption enters positively and highly significantly, consistent with the

prominence of public support schemes, price interventions and subsidised input programmes in the agricultural sectors of many sample countries, while trade openness is positively signed, plausibly reflecting the export orientation of agriculture in lower-income economies. The negative coefficient on the growth of school enrolment is best read as a structural transformation effect rather than as evidence that education damages agriculture, since the reallocation of labour away from agriculture towards industry and services is a well-documented feature of development (Herrendorf et al., 2014). Read as a whole, the agricultural evidence indicates that the pace of lending is not the margin along which agricultural finance operates, and that the relevant constraints lie in the institutional architecture of rural credit markets.

9.2.2. Industry: Regime-Dependent Nonlinearity and the Location of Turning Points

Industry is the sector in which the pace of credit expansion matters most, and the shape of the relationship changes across financial regimes in a way that carries substantive content. Three features distinguish the industrial results. The nonlinearity is present in every period and under both estimators, the higher-order terms retain their magnitude and significance when controls are added, and the location of the turning points relative to the empirical distribution of credit growth differs sharply between the pre-crisis and post-crisis windows. The last of these features is what gives the coefficients their economic content, since a significant cubic term whose turning points lie in the extreme tail describes a relationship that is, over the range countries actually experience, effectively monotonic.

In the pre-crisis window the static estimator returns a weakly positive linear coefficient that does not survive the addition of controls, while the dynamic estimator identifies a significant negative cubic term implying a local maximum at 0.510, or annual private credit growth of about sixty-seven per cent. Because that rate sits just below the ninety-ninth percentile of the pre-crisis distribution, the descending segment of the fitted curve is identified from a very small number of extreme expansions, so moderate credit growth appears broadly neutral or mildly favourable

while exceptionally rapid expansion accompanies a deteriorating industrial position. The asymmetry corresponds to the pattern emphasised in the credit-boom literature, in which rapid expansions of credit and leverage raise the probability of subsequent macroeconomic weakness (Mendoza & Terrones, 2008; Schularick & Taylor, 2012), though Dell'Ariccia et al. (2016) report that only about one boom episode in three is followed by a financial crisis, which places the adverse outcome as a tendency rather than a regularity.

The post-crisis decade produces the clearest industrial results in the analysis and inverts that picture. The coefficient pattern is stable across the baseline and extended models, the discriminant of the derivative is positive and significant at the one per cent level, so two distinct turning points are established rather than assumed, and the upper turning point of 0.246 lies between the ninety-fifth and ninety-ninth percentiles of a distribution whose median is 0.010. The interval over which the marginal association is negative therefore encompasses almost the entire mass of the data, and the relationship becomes favourable again only at rates of expansion attained by fewer than one country–year observation in twenty. That this profile survives the addition of controls which raise the within R-squared from 0.087 to 0.240 indicates that the association is not simply a reduced-form reflection of investment or trade conditions, a reading reinforced by the strongly positive coefficient on gross capital formation, which confirms that physical investment supported industrial activity throughout the same decade.

Over the full sample both estimators agree on the broad shape while differing on the precise location of the maximum, the static model placing it at 0.216 and the dynamic model at 0.156, corresponding to annual credit growth of approximately twenty-four and seventeen per cent respectively. Neither root is individually significant in the dynamic specification and the discriminant is not distinguishable from zero in the static one, so the agreement is best described as convergence on the broad location of a turning region rather than as identification of a threshold. This pattern corresponds closely to the nonlinear finance–growth relationship documented at the aggregate level by Arcand et al. (2015) and by Law and Singh (2014), with the substantial difference that the nonlinearity identified here operates through the pace

of credit expansion rather than through accumulated financial depth. Whether the two forms of nonlinearity share a common mechanism cannot be settled by these estimates and would require the joint modelling of stocks and flows.

The composition of lending offers the most plausible account of why the industrial association weakened after 2008. Cecchetti and Kharroubi (2015) argue that a rapidly expanding financial sector draws skilled labour away from other activities and favours projects that are easy to collateralise over those that are more productive, with research-intensive industries and firms holding few tangible assets suffering the largest losses, a mechanism that bears directly on manufacturing. Beck et al. (2012) show that the cross-country association between credit and growth is driven mainly by enterprise lending rather than household lending, and Bezemer et al. (2016) report that credit flows to non-financial business carry positive growth effects while mortgage and other asset-market lending do not. Borio et al. (2016) add the productivity dimension, finding that credit booms tend to reallocate labour towards sectors with lower productivity growth. Each of these mechanisms would produce the pattern observed here, and none can be tested with an aggregate credit measure, which is why the industrial results should be read as establishing the association and narrowing the field of candidate explanations rather than as identifying the channel.

9.2.3. Services: A Positive First-Order Effect and Concealed Internal Heterogeneity

The service sector yields significant credit coefficients in only one of the three periods, and that concentration is itself informative. Over the pre-crisis window and the full sample the credit terms are uniformly small relative to their standard errors under both estimators, while the dynamic models encounter difficulties with the overidentifying restrictions that further limit their evidentiary value. The absence of a stable long-horizon relationship is intelligible in light of what aggregate private credit finances, since households, real estate and non-financial firms are funded in proportions that vary widely across countries and over time, and these components carry different implications for real activity (Beck et al., 2012; Bezemer et al., 2016). Because services encompass both highly productive business activities and branches

closely tied to property and consumption, aggregate credit growth may exert offsetting influences within the sector that leave no measurable net effect.

In the post-crisis decade the linear term is positive and highly significant in both specifications, estimated at 0.141 in the baseline model and 0.066 once controls are added, which establishes that faster private credit growth accompanied a higher service share of value added. The implied local maximum at 0.405 lies beyond the ninety-ninth percentile of the distribution, so the relationship is positive and rising over virtually the whole of the observed range, and the higher-order terms are in any case unstable, the squared term switching sign between specifications and the cubed term losing significance once controls enter. Interpretation should accordingly rest on the sign and the significance of the first-order effect rather than on any specific turning point, which is an instance of the more general discipline this study applies to polynomial specifications.

The composition of the service aggregate deserves particular emphasis, because financial intermediation is itself a component of measured service value added. An expansion of the financial sector therefore raises the service share directly, even where it draws skilled resources away from other activities, which means that the positive service coefficient may combine a genuine demand-side association with a mechanical composition effect. This reading extends the resource-allocation mechanism developed by Cecchetti and Kharroubi (2015), although the composition effect within aggregate service value added is an inference drawn here rather than a proposition tested in their study. The Turkish evidence discussed in Section 9.4 supports the concern, since financial and insurance activities there display an inverted U-shaped long-run profile while information and communication display a U-shaped one, so the branches aggregated within the international service variable behave in ways a common coefficient cannot represent.

9.3. Development Stage, Structural Transformation and the Limits of a Common Curve

The income-level estimates establish that pooling economies at very different stages of development under a single functional form misrepresents at least two of the three

groups. Greenwood and Jovanovic (1990) model the joint evolution of financial structure and income distribution, Rioja and Valev (2004a) argue that the effect of financial development is uncertain at very low levels, large at intermediate levels and positive but smaller at high levels, and Rioja and Valev (2004b) show that the sources of the effect shift from capital accumulation towards productivity as development proceeds. Demetriades and Law (2006) reach a compatible conclusion through an institutional lens, reporting that finance exerts its strongest influence in middle-income economies and its weakest in low-income ones. The results reported here are broadly consistent with this graded view while diverging from it in one respect that follows from the choice of regressor, since the strongest evidence of harm rather than of benefit is located in the middle-income group, and the object of measurement is the growth rate of credit rather than its level – a distinction Sahay et al. (2015) treat as consequential for both growth and stability.

In low-income economies credit expansion begins from a position of pronounced scarcity, yet the industrial estimates place the working range of the data on the declining middle segment of a cubic profile whose stationary points lie at -0.480 and 0.230. Since the interquartile range of private credit growth in this group runs from -0.067 to 0.133 around a median of 0.041, faster credit expansion tends to accompany a lower industrial share across the range in which these economies typically operate, while the ascending segment beyond 0.230 rests on comparatively few observations. Agriculture behaves differently, the estimated maximum of 0.414 lying far above the median rate of agricultural credit growth so that most observations sit on the upward-sloping segment, which accords with the assessment that constraints in smallholder agriculture originate in the absence of collateralisable assets and in the reluctance to pledge assets vital to livelihoods (World Bank, 2007). Neither result should be read as evidence that low-income economies suffer from too much finance in the aggregate. They indicate rather that even modest rates of credit expansion may encounter diminishing returns where infrastructure, market access, technology and institutional capacity do not expand alongside lending.

The middle-income group produces the most interpretable evidence in the income-level analysis and the strongest support for the ‘too much finance’ proposition.

The static cubic specification places the descent above 0.309, close to the ninety-fifth percentile of 0.286, while the dynamic quadratic specification returns a Lind and Mehlum (2010) test that rejects monotonicity at the one per cent level with a maximum at 0.026 – almost exactly the median rate of private credit growth in this group – implying that roughly half of the observations lie beyond the point at which the marginal association turns negative. The two estimators thus agree on the shape of the relationship in the upper range while disagreeing on where the maximum falls, and the eighty per cent Fieller (1954) interval of -0.304 to 0.370 encompasses the static estimate, so the disagreement concerns the precision with which the threshold can be located rather than the existence of the concavity. Middle-income economies are precisely where this configuration would be expected, combining substantial manufacturing sectors with supervisory frameworks that may not expand as rapidly as credit (Cecchetti & Kharroubi, 2015; Sahay et al., 2015).

High-income economies display the most restrictive interval of all, the estimated maximum of -0.224 lying below zero so that essentially the whole range of positive credit growth falls on the declining branch, while the curvature does not survive the addition of controls and the negative level effect strengthens. Several strands of the literature rationalise this pattern without recourse to the claim that finance has ceased to matter in advanced economies. Rodrik (2016) documents that the hump-shaped relationship between industrialisation and income has shifted downwards and closer to the origin, so that economies exhaust their industrialisation opportunities sooner and at lower income levels, a structural trend against which any positive credit effect would have to work. The composition of credit is likely to matter as well, since the household share of lending tends to rise with income and is higher in economies with smaller manufacturing sectors (Beck et al., 2012), while increases in the household-debt ratio predict lower subsequent output growth in a panel of mostly advanced economies (Mian et al., 2017). Aggregate private credit growth in high-income countries need not, on this account, primarily reflect lending to industrial firms, and Rousseau and Wachtel (2011) offer the complementary observation that the estimated effect of financial deepening on growth weakens appreciably in more recent samples.

The structural transformation reading is essential rather than decorative, because the international dependent variables are shares. An increase in the service share necessarily alters the relative position of agriculture and industry even where the absolute output of all three sectors is rising, so a negative coefficient may record reallocation rather than contraction. McMillan and Rodrik (2011) demonstrate that structural change may either raise or lower aggregate productivity depending on which sectors attract resources, which is why the compositional shifts documented here cannot by themselves be classified as beneficial or damaging. The contribution of this study at that intersection is to show that the pace of credit expansion is associated with the direction of compositional change, faster expansion accompanying a rising service share and a falling industrial share in the post-crisis decade, while establishing whether that reallocation raised or lowered productivity would require output, employment and productivity data that the present design does not incorporate.

9.4. The Türkiye Evidence in Perspective

9.4.1. Common Factors, Feedback and the Choice of Framework

The Türkiye analysis was designed to address two constraints that the international investigation cannot overcome, namely the absence of internationally comparable sectoral credit data for industry and services and the difficulty of separating short-run adjustment from long-run equilibrium in annual sectoral shares. Matching the total cash loans reported by the Banking Regulation and Supervision Agency to the production-based value added published by the Turkish Statistical Institute across eleven activity groups aligns the financing side with the producing side, so that the credit each activity receives is paired with the value added that activity generates. The quarterly panel runs from the first quarter of 2005 to the first quarter of 2026 and yields 902 observations, covering a period that includes rapid financial deepening, repeated credit cycles, the expansion of publicly supported guarantee schemes, episodes of exchange-rate and inflation instability, a pronounced credit response to the pandemic and a subsequent reliance on selective lending and macroprudential instruments, which makes Türkiye an unusually informative setting for examining whether the destination and the scale of credit matter as much as its volume.

The diagnostic evidence establishes that Turkish sectors cannot be treated as independent cross-sectional units, since they share monetary conditions, exchange-rate movements, domestic demand, inflation, regulation and dense input–output linkages. Strong cross-sectional dependence is accordingly an economic property of the system rather than a statistical inconvenience, and the cross-sectionally augmented framework addresses it by using cross-sectional averages and their lags as proxies for unobserved common factors while retaining sector-specific slopes (Chudik & Pesaran, 2015; Pesaran, 2006). The residual dependence tests return insignificant statistics in all three specifications, with p-values of 0.544, 0.679 and 0.411, indicating that two lags of the cross-sectional averages absorb the common-factor structure adequately, while the error-correction term is negative and significant throughout, ranging from -0.620 to -0.709, so deviations from the long-run relationship are progressively corrected.

The panel causality evidence reinforces the choice of framework and simultaneously limits the interpretation of the estimates. Bidirectional predictive causality between sectoral credit growth and sectoral value added is detected at the one per cent level in both directions, which corresponds to the supply-leading and demand-following mechanisms of Patrick (1966) operating jointly rather than to either acting alone. Since the test establishes predictive precedence within a heterogeneous panel rather than structural causation (Dumitrescu & Hurlin, 2012), the subsequent estimates should be read as dynamic conditional relationships rather than as identified effects of exogenous credit supply. The same consideration, together with a time dimension of eighty-six quarters against eleven sectors, is what rules out System GMM in this application, since instrument proliferation would weaken the overidentification tests (Roodman, 2009a, 2009b) and the common factors driving the dependence would remain unmodelled.

9.4.2. Short-Run Adjustment, Long-Run Equilibrium and the Divergence Between Horizons

The linear specification appears at first to deliver a simple message and in fact demonstrates why a single average slope is uninformative in this panel. The Mean

Group short-run coefficient of -0.213 indicates that larger quarterly increases in sectoral credit are on average associated with weaker contemporaneous output growth, an average driven principally by agriculture, by the trade, transport, accommodation and food services group and by public and social services. Credit-financed inputs, inventories, machinery, construction and organisational change may impose costs before generating measurable value added, and firms may borrow precisely when cash flow is weak, so a negative contemporaneous coefficient is consistent with both an investment lag and a reverse demand effect. Neither reading supports the conclusion that credit accumulation damages sectoral activity, and the quadratic results show the linear average to be a restriction rather than a finding.

The long-run component of the linear model illustrates the same point from the opposite direction, since the Mean Group coefficient of 0.165 is positive but insignificant while four sectors return positive and significant long-run coefficients. Mining, quarrying and other industries record the largest at 1.320, an outcome that is economically coherent given the capital intensity of mining, energy and water-related activities, in which credit plausibly finances fixed investment, capacity expansion and infrastructure whose returns accrue over extended horizons. Agriculture returns 0.848 in direct contrast to its negative short-run estimate, the trade-related group 0.814 and other services 0.265. The insignificance of the average therefore reflects the dispersion of sector-specific effects rather than the irrelevance of credit, which is precisely the kind of result that an economy-wide summary would suppress.

Agriculture is the clearest illustration of horizon dependence in the panel and merits separate comment. Its short-run coefficient is negative and its long-run coefficient positive and significant, a divergence that the production structure of the sector makes intelligible, since biological lags, seasonal revenue and delayed output realisation prevent credit drawn in one quarter from generating value added within the same quarter. The cross-country and Turkish agricultural results should not be compared numerically, because the international model relates agricultural credit relative to GDP to the agricultural share of GDP while the Turkish model relates nominal sectoral credit to sectoral value added within an error-correction framework. Read qualitatively rather than numerically, however, the two settings converge on the

same proposition, namely that the effectiveness of agricultural finance depends on scale and timing rather than on lending volume alone.

9.4.3. Sector-Specific Thresholds and the Capacity to Absorb Credit

The quadratic specification changes the interpretation of the linear results and supplies the central Turkish finding. At the Mean Group level the coefficient on the change in log credit is positive at 2.042 and that on the differenced squared term negative at -0.065, both significant at the five per cent level, describing an inverted U in which changes in credit support quarterly output growth up to a turning point of approximately 7.23 billion Turkish liras and detract from it beyond. Because the turning point lies within the observed range of the credit variable it is empirically meaningful rather than an extrapolation, a requirement that the methodological literature on U-shaped relationships treats as necessary for interpretation (Haans et al., 2016; Lind & Mehlum, 2010). Below the threshold credit may ease liquidity constraints and finance inventories, payrolls, inputs and continuity of production, whereas beyond it the association may reflect debt-service burdens, distress borrowing, weaker allocation or the simple inability of a sector to convert additional funds into output within the quarter.

The dispersion of the sectoral thresholds around that average is more consequential than the average itself. Information and communication reaches its turning point at approximately 2.46 billion liras, agriculture at 4.94 billion, public and social services at 8.68 billion and the trade, transport, accommodation and food services group at 40.23 billion, a range of more than an order of magnitude within a single national economy. The low threshold in information and communication is consistent with the intangible character of the sector’s assets, since banks lend more readily against tangible collateral while software, data, human capital and network infrastructure are less readily pledged, and Hsu et al. (2014) report that credit-market development does not support innovation in externally dependent and technology-intensive industries in the way that equity-market development does. The high threshold in the trade-related group matches its dependence on working capital, inventories, logistics, distribution networks and tourism demand, activities in which credit can support

value added across a wide interval. Even that sector displays a significant negative quadratic term, indicating that the greatest observed absorptive capacity in the panel is nonetheless bounded.

The long-run quadratic estimates identify nonlinearity as a sector-specific rather than an economy-wide property, since the Mean Group squared term is insignificant while several sectors return well-determined profiles. Financial and insurance activities display an inverted U with a maximum near 26.07 billion liras, a result that speaks directly to the 'too much finance' debate because the sector in question is the financial sector itself, and because over-expansion of financial activity may divert resources towards intermediation, balance-sheet management or speculation rather than productive investment (Cecchetti & Kharroubi, 2012, 2015). Agriculture and information and communication instead display U-shaped long-run profiles, with turning points near 1.16 and 2.56 billion liras, in which credit appears costly or ineffective at low levels and productive only once it reaches the scale required to finance modernisation, fixed investment or technological upgrading. A U-shape describes a different process from the one the 'too much finance' literature examines and should not be generalised beyond the sectors in which it is estimated.

The cubic specification refines this account rather than overturning it. At the Mean Group level the third-order term is insignificant, so the panel average remains best described by the single-threshold adjustment identified in the quadratic model, yet several sectors display long-run trajectories that one turning point cannot represent. Real estate follows a clear N-shaped path with turning points near 1.36 and 13.84 billion liras, a configuration consistent with the mechanism formalised by Kiyotaki and Moore (1997), in which durable assets serving simultaneously as factors of production and as collateral transmit and amplify shocks through the interaction of credit limits and asset prices. The terminal positive segment should not be read as benign, given the association between credit-driven property booms and asset-price inflation and the evidence that credit directed towards real estate and other asset markets carries weaker growth effects than credit extended to non-financial business (Bezemer et al., 2016). Manufacturing, the trade-related group and public and social

services display inverted-N configurations in which credit is associated with weaker outcomes at low levels, supports activity over an intermediate range and becomes negatively associated again at very high levels, a pattern compatible in manufacturing with the possibility that very large exposures sustain less efficient structures and delay adjustment (Caballero et al., 2008).

Three sectors require the interpretive discipline that this study applies throughout. Agriculture, industry excluding manufacturing and construction return coefficient sign sequences resembling N-shaped or inverted-N-shaped curves, yet the discriminants of their derivatives are negative, so the fitted functions possess inflection points rather than two genuine turning points. Every cubic function has exactly one inflection point, and a negative discriminant means that this inflection is the only structural feature of the curve, marking a change in the momentum of the relationship rather than a reversal in its direction. Reading the sign sequence directly off the coefficients would have produced three spurious regime descriptions, which is the clearest demonstration in the dissertation of why functional form must be evaluated through the roots and the discriminant of the derivative rather than inferred from significance patterns.

9.5. Integrating the International and Turkish Evidence

The two investigations differ in data structure, dependent variable, financial indicator, frequency and estimator, so their numerical turning points are not commensurable and should not be compared as though they measured the same quantity. What they can be compared on is structure, and on four propositions they agree. Sectoral disaggregation alters the substantive conclusions rather than refining them at the margin. Nonlinearity is common but not universal, appearing strongly in industry and in several Turkish branches while remaining weak or absent elsewhere. The relationship depends on horizon and on developmental context. And average coefficients conceal economically consequential heterogeneity in both settings, the insignificant Turkish long-run average and the within-country explanatory power of between 0.059 and 0.122 recorded by the aggregate preliminary analysis being two expressions of the same phenomenon.

Agriculture converges across the two settings once the difference in measurement is respected. The international full-sample estimates describe a weakly supported concave profile whose maximum lies far above the typical rate of agricultural credit growth, so most observations sit where faster expansion is mildly favourable, while the Turkish estimates combine a negative short-run coefficient, a positive and significant long-run coefficient, a comparatively low short-run threshold and a U-shaped long-run profile. These findings are not in tension. Together they indicate that agricultural credit may relax binding constraints while its benefits are delayed, nonlinear and conditional on the scale and the timing of funding, and that neither a negative contemporaneous coefficient nor a diminishing marginal effect at high credit levels can be read as evidence that agricultural finance is generally unproductive.

The industrial comparison is where the Turkish evidence contributes most to the interpretation of the cross-country results. The international industry variable combines mining, utilities, construction and manufacturing, activities that differ in capital intensity, investment horizon, collateral structure and working-capital requirement, and the Turkish disaggregation shows how differently those components behave. Industry excluding manufacturing returns a strong positive long-run coefficient and a cubic function with an inflection point rather than turning points, manufacturing displays an inverted-N profile with a negative terminal segment, and construction exhibits an inflection point consistent with its exposure to collateral values and leverage cycles. Higher-order international relationships in industry are therefore likely to arise, at least in part, from the aggregation of sector-specific processes that differ in shape as well as in sign, which is an empirical contribution of the pairing rather than a conjecture.

Services demonstrate the same point more sharply still. The international estimates return a positive and robust first-order effect in the post-crisis decade, whereas the Turkish service branches move in different directions and with different functional forms, the trade-related group combining a negative short-run coefficient with a positive long-run one and the highest short-run threshold in the panel, information and communication combining the lowest short-run threshold with a U-shaped long-

run profile, financial and insurance activities displaying an inverted U, real estate an N-shape, public and social services an inverted N and professional and administrative services no significant linear relationship at all. A positive aggregate service coefficient is therefore an average across structurally dissimilar activities and should not be taken to indicate that service branches benefit uniformly from faster credit growth.

The combined evidence also fixes the sense in which the phrase ‘too much finance’ is used in this dissertation. The results do not identify a level of financial development beyond which finance becomes damaging in general, and they were not designed to, since the credit variables enter the international models in first differences. What they identify is that the relationship between credit and real activity may weaken, disappear or reverse where the pace or the scale of lending exceeds the capacity of a production structure to absorb it within a given horizon, and that this capacity varies with investment opportunity, production lags, collateral availability, technology, demand conditions and institutional arrangements. This interpretation is compatible with the mechanism proposed by Cecchetti and Kharroubi (2015), with the allocation evidence of Beck et al. (2012), Mian et al. (2017) and Müller and Verner (2024), and with the finding of Ductor and Grechyna (2015) that the growth effect of financial development turns negative where credit expansion outpaces real output growth. Because the present analysis measures credit volume rather than loan-level allocation, these remain interpretive mechanisms consistent with the estimates rather than channels the models identify.

9.6. Contributions of the Dissertation

The first contribution is conceptual and consists in joining the finance–growth and structural transformation literatures, treating financial development not merely as a determinant of aggregate growth but as a force capable of reshaping the sectoral composition of production. This reframing shifts the question from whether credit is beneficial in general towards how its benefits and costs are distributed across productive activities, and the empirical results give the reframing content rather than leaving it programmatic, since faster credit expansion accompanied a rising

service share and a falling industrial share within the same decade and the same sample.

The second contribution concerns measurement and follows from taking the integration properties of the data seriously rather than treating them as a technicality. Because the credit variables are integrated of order one while the sectoral value-added shares are stationary, credit enters the international equations in first differences and the estimated nonlinearities are located in credit-growth space rather than in credit-depth space. The distinction between the stock of credit and its rate of expansion is well established in the crisis and business-cycle literature (Jordà et al., 2013; Loayza & Rancière, 2006; Schularick & Taylor, 2012), yet it has not been carried systematically into sectoral analyses of value added, where the two margins are frequently treated as interchangeable. The Türkiye chapter goes further by recovering both margins within one specification, retaining sectoral credit levels in the long-run component while modelling credit growth in the short-run component, which allows their divergence to be observed directly rather than inferred.

The third contribution is empirical and rests on the pairing of breadth with depth. Agriculture, industry and services are examined within a common framework that permits relationships to differ across pre-crisis, post-crisis and full-sample periods and across income groups, while the Türkiye analysis adds resolution by matching eleven credit–output pairs at quarterly frequency over a period encompassing the principal shifts in the country's financial and policy environment. The national component also extends a Turkish literature that has typically related sectoral credit to aggregate output, worked with shorter periods or imposed linearity (Kandemir et al., 2018; Koç, 2015; Sarı & Konukman, 2023; Sever & Han, 2015), and it complements firm-level evidence indicating that guarantee schemes and financing constraints affect firms and sectors differently (Akçiğit et al., 2021; Bağır & Seven, 2021).

The fourth contribution is methodological. Functional form is selected sequentially rather than assumed, turning points are recovered from marginal effects rather than read off coefficient signs, quadratic profiles are subjected to formal slope tests and

interpreted only where the maximum lies within the supported range of the data (Lind & Mehlum, 2010), and cubic profiles are evaluated through the roots and the discriminant of the derivative so that an inflection point is not mistaken for a pair of turning points. Static estimates are compared with dynamic ones throughout, and the Turkish analysis addresses common factors, sectoral heterogeneity, bidirectional feedback and the separation of horizons explicitly. Locating each estimated turning point within the empirical distribution of the regressor is a further element of the same discipline, since a statistically significant higher-order term whose turning points lie in the extreme tail describes a relationship that is, for practical purposes, monotonic over the range countries actually experience.

The fifth contribution concerns the geometry of the nonlinearity rather than its existence. That the relationship between credit and real activity departs from linearity is by now a familiar conclusion, but the departure documented here does not take the form the ‘too much finance’ literature has made canonical, in which the contribution of finance rises, peaks at an interior threshold and declines thereafter (Arcand et al., 2015; Cecchetti & Kharroubi, 2012; Law & Singh, 2014; Samargandi et al., 2015). The aggregate estimates reported in Chapter 3 are concave in every period examined, yet their maxima lie at or below zero, so the ascending branch falls entirely within the range of credit contraction and the configuration cannot be read as one in which expansion supports growth up to a threshold. Disaggregation indicates why an aggregate specification may return a shape of this kind, since sectors differ not only in the location of their turning points but in the order of polynomial their data support and in the sign of the association, and the same sector may reverse its curvature between horizons. An aggregate turning point is therefore a compositional average of sectoral relationships differing in shape as well as in sign, and it need not correspond to the behaviour of any individual sector.

The sixth contribution is directed at policy and follows from the preceding five. Aggregate credit targets and uniform lending rules presuppose that sectors convert finance into value added at broadly similar rates, and the evidence assembled here indicates that they do not. A rate of expansion that is beneficial in one activity may exceed the absorptive capacity of another, and a negligible economy-wide

association may coexist with strong positive, negative or nonlinear relationships at the sectoral level, which makes the destination of lending at least as consequential as its volume.

9.7. Policy Implications

The principal implication is that credit policy cannot be evaluated by aggregate volume alone. Volume continues to matter, particularly where financing constraints bind, but the same expansion may produce different outcomes depending on the receiving sector, the prevailing level of credit and the horizon over which the outcome is assessed. Uniform growth targets and undifferentiated lending incentives implicitly assume comparable absorptive capacity across recipients, and the dispersion of Turkish thresholds by more than an order of magnitude indicates how strong that assumption is.

In low-income and agriculture-dependent economies the expansion of formal financial services remains valuable, since most observations in that group lie on the segment where faster agricultural credit growth accompanies higher agricultural value added. Credit policy in such settings should nonetheless be coordinated with irrigation, transport, storage, extension services, insurance, technology and access to output markets, because the estimated concavity indicates that lending cannot indefinitely substitute for absent complementary conditions (World Bank, 2007). The estimated turning point marks the range in which the marginal contribution changes and in which the use and the structure of additional lending warrant closer examination, not a limit beyond which credit should be withheld.

In middle-income economies the evidence supports differentiation by stage and by sector rather than the pursuit of a single economy-wide optimum. Credit may finance industrial upgrading and the growth of modern services while simultaneously reinforcing construction, household borrowing or low-productivity non-tradable activity, and Müller and Verner (2024) document that booms disproportionately favour non-tradable sectors and that such expansions are systematically associated with subsequent slowdowns. Because the middle-income industrial estimates place

the point at which the marginal association turns negative close to the median rate of private credit growth in that group, the case for monitoring the composition of lending alongside its aggregate rate is stronger in these economies than in either of the others.

In high-income economies the weak independent association between credit growth and industrial value added suggests that bank credit expansion is an insufficient instrument at advanced stages, where innovation, intangible investment, specialised finance, equity markets and institutional efficiency plausibly carry greater weight. Hsu et al. (2014) show that equity-market development supports innovation in externally dependent and technology-intensive industries while credit-market development does not, which implies that a predominantly bank-based system may finance conventional capital accumulation effectively while remaining a weaker instrument for innovation-intensive activity. The positive service association observed in the post-crisis decade should be read with the Turkish heterogeneity in view, since measures directed at real estate or financial services are unlikely to produce the effects of measures directed at information, professional, health or trade services.

The Turkish results bear directly on selective credit programmes, guarantee schemes, state-bank lending and macroprudential design. Sector-specific thresholds are diagnostic instruments rather than regulatory limits, being estimated from nominal historical data, conditional on model and sample, and unable to distinguish performing from impaired exposure, new lending from inherited stock or productive investment from refinancing. Their value lies in identifying activities where rapid credit growth indicates that the purpose, maturity, collateral structure and borrower balance sheets behind that growth merit closer scrutiny. Policy evaluation should also be horizon-sensitive, since programmes assessed solely on contemporaneous output may appear ineffective in agriculture, infrastructure and other activities with long production or investment lags, while a short-run expansion should not be judged successful merely because it raises current activity if it leaves higher leverage without durable capacity. Evidence that household leverage and rapid credit growth tend to precede weaker medium-term activity underlines the point (Mian et al., 2017; Schularick & Taylor, 2012).

9.8. Limitations

Several limitations define the scope of the conclusions. The first concerns the international dependent variables, which are sectoral value-added shares of GDP. This choice is deliberate and follows from the structural transformation framing, providing comparable sectoral measures across a wide set of countries, but shares are not direct indicators of absolute output or productivity, and a change in a share may reflect developments within a sector, differing growth rates across sectors or shifts in relative prices. The international results accordingly describe how the pace of credit expansion is associated with the relative weight of agriculture, industry and services.

Second, internationally comparable sector-specific credit data are available only for agriculture. The industrial and service models therefore use private credit by deposit money banks, so their coefficients describe how overall banking depth relates to sectoral composition rather than the effect of credit extended to the sector in question. The allocative interpretation advanced in Sections 9.2 and 9.5 is consistent with the estimates and with the credit-composition literature, yet it remains an interpretation rather than a tested proposition within this design.

Third, the cross-country panel is unbalanced and the agricultural sample is narrower, since credit and value-added data overlap for fewer countries, so data availability shapes the composition of the estimated samples. The extended specifications also rest on samples roughly half the size of the baseline ones, which limits the inference that can be drawn from changes in significance between the two columns. The three broad sectors additionally conceal substantial internal heterogeneity, as the Turkish evidence demonstrates.

Fourth, the estimators address the principal econometric difficulties without delivering full causal identification. Driscoll–Kraay standard errors support inference under heteroskedasticity, serial correlation and cross-sectional dependence (Driscoll & Kraay, 1998), yet fixed-effects estimates may still be affected by reverse causality and omitted time-varying factors. System GMM addresses persistence and potentially endogenous regressors through internal moment conditions (Blundell &

Bond, 1998), while the standard errors of the two-step estimates are corrected for finite-sample downward bias following Windmeijer (2005), though several specifications return insignificant credit coefficients or diagnostics that limit their weight, and large instrument sets reduce the power of overidentification tests (Roodman, 2009a, 2009b). In the Türkiye chapter the causality tests establish predictive precedence rather than structural causation, and the error-correction estimates describe conditional short-run and long-run relationships rather than exogenous credit-supply shocks.

Fifth, polynomial specifications are flexible approximations rather than structural functional forms, and higher-order terms may be influenced by sample range, outliers, collinearity and specification. The analysis mitigates this by testing turning points formally, by interpreting them only within the supported range and by distinguishing inflection points from genuine reversals, but no estimated threshold should be treated as a fixed parameter, and the divergence between static and dynamic estimates indicates that the location of a turning point depends on how persistence and controls are handled.

Sixth, income groups are broad categories that reduce the risk of imposing a single coefficient on economies at very different stages while leaving within-group heterogeneity unaddressed, since countries of similar income may differ in legal systems, financial inclusion, banking concentration, policy environment, export composition and institutional quality. The classification is also applied as of a single World Bank vintage, so movements between categories during the sample period are not reflected. The income-level models additionally pool the whole of 1992 to 2019, so developmental differences and the temporal shifts documented in the periodical analysis cannot be separated within that design.

Seventh, the Turkish series are measured at current prices without seasonal adjustment. Unadjusted output series were used deliberately, since the sectoral credit data are published only in unadjusted form and combining officially adjusted output with unadjusted credit would introduce inconsistent seasonal treatment across the two sides of the relationship. Consistency is thereby preserved at the cost of leaving

seasonal variation embedded in the estimates, and the turning points expressed in Turkish liras reflect the price level and the scale of the sample period, so they are interpretable within the sample without transferring to other periods as fixed real magnitudes. Using quarter-end credit stocks is appropriate for a balance-sheet variable but does not capture intra-quarter lending activity, and because total cash credit combines performing loans with non-performing receivables the estimates do not separate active lending from the accumulation of impaired exposure.

Eighth, the concordance between supervisory credit categories and national accounts activity groups improves consistency without achieving an exact correspondence, since some categories must be combined and the series for industry excluding manufacturing is constructed by subtraction. The panel also contains eleven sectors, a dimension well suited to the cross-sectionally augmented framework given the length of the time series, though the standard asymptotics underlying the dependence, unit-root and common-correlated-effects procedures are derived for larger cross-sections, and the design cannot resolve subsectoral or firm-level heterogeneity.

Finally, the study concentrates by design on the quantitative dimension of financial development and does not measure interest rates, maturity, currency denomination, collateral, borrower size, loan purpose, bank ownership, guarantee status, inclusion, efficiency or stability. These dimensions may well explain why similar credit volumes produce different outcomes, and the conclusions reached here concern the scale, the pace and the nonlinear structure of credit intermediation rather than the quality of finance.

9.9. Directions for Further Research

The most valuable extension of the international analysis would be the construction of internationally comparable sectoral credit data covering manufacturing, construction and the major service branches. At present only agricultural credit can be matched directly to agricultural value added, while the industrial and service models rely on aggregate private credit, which prevents the allocative interpretation

from being tested. Data identifying the sectoral destination of bank lending would establish whether the nonlinear patterns documented in industry and services arise from credit extended to those sectors or from broader shifts in financial depth, and would allow the differing behaviour of enterprise, household and asset-market lending to be examined at sectoral resolution (Beck et al., 2012; Bezemer et al., 2016).

A second extension would combine sectoral credit with a wider set of sectoral outcomes. Value-added shares are informative about structural change and compositional shifts but silent about absolute output and productivity, so models estimating real sectoral value added, employment, labour productivity, total factor productivity, investment and export performance jointly would separate financial effects on reallocation from effects on production levels and efficiency. Because a sector may raise its absolute output while losing share to faster-growing activities, examining these outcomes together would clarify whether the compositional shifts documented here correspond to productivity-enhancing or productivity-reducing reallocation in the sense analysed by McMillan and Rodrik (2011).

Disaggregation beyond the three broad sectors is the natural continuation of the Turkish evidence, which shows how much variation the standard categories conceal. Manufacturing, construction, real estate, information and communication, financial services and public and social services differ substantially in financing structure, collateral base, adjustment speed and productivity potential, and comparable international data would allow those differences to be examined systematically. Incorporating input–output linkages would strengthen such an analysis further, since credit extended to one sector may influence production in others through suppliers, customers, investment demand and working-capital chains, and spatial or network specifications could distinguish direct from indirect transmission.

For Türkiye the treatment of nominal and seasonal variation is the most immediate priority. Rather than substituting officially adjusted output series for the unadjusted ones, a consistent approach would apply the same seasonal adjustment, calendar specification and outlier treatment to the raw credit and output series, and would

convert both to real terms using general or sector-specific deflators where suitable price indices exist. That would allow turning points to be expressed in constant prices and would establish whether the nonlinear structures documented here are stable once inflation and seasonality are removed.

The measurement of sectoral credit admits further refinement. Because total cash credit combines performing loans with non-performing receivables, estimating the two components separately would establish whether the observed nonlinearities are driven mainly by changes in performing credit, by the accumulation of impaired receivables or by shifts in the composition of total exposure, while the ratio of non-performing receivables to total sectoral credit could serve as an indicator of deteriorating credit quality or sectoral financial stress. Distinguishing newly issued loans from inherited stock, and decomposing credit by currency, maturity, interest-rate structure, collateral type, borrower size, bank ownership and guarantee status, would help explain why comparable rates of credit growth generate different sectoral outcomes.

Matched bank–firm or loan-level administrative data would represent the most substantial advance. Linking credit records to firm productivity, investment, employment, sales, exports, balance-sheet structure and survival would identify which borrowers convert additional finance into productive capacity and which use it primarily for liquidity management, refinancing or the maintenance of existing liabilities, and would help separate movements in credit demand from shifts in credit supply. Policy measures including the expansion of the Credit Guarantee Fund, macroprudential adjustments, state-bank lending programmes and sector-specific incentives generate variation that, under appropriate assumptions, could support difference-in-differences, instrumental-variable, event-study or local-projection designs capable of estimating the dynamic effects of such interventions on allocation and on real outcomes.

Regime dependence deserves separate treatment in both settings. The Turkish sample spans financial reform, global crises, exchange-rate instability, large-scale credit campaigns, the pandemic shock and subsequent monetary and macroprudential

adjustment, so specifications allowing for structural breaks, regime switching or time-varying parameters could establish whether sectoral absorptive capacity changes during periods of financial stress, rapid inflation, monetary tightening or targeted credit expansion. Rolling-window and recursive estimation would indicate whether the turning points identified here describe durable sectoral characteristics or particular episodes. The international analysis raises the parallel question of estimating developmental and temporal heterogeneity jointly, which would require a panel longer or wider than the one available here.

Methodologically, the polynomial specifications adopted in this dissertation could be compared with less restrictive alternatives, including panel threshold models, spline regressions, smooth-transition specifications, semiparametric functions, quantile regressions and time-varying parameter models. Such comparisons would establish whether the turning points recovered here survive when curvature is not constrained to a quadratic or cubic form, and quantile methods in particular would indicate whether credit operates differently in weakly performing and rapidly expanding sectors. Out-of-sample forecasting, rolling validation and sensitivity analysis across alternative windows would further establish whether the documented nonlinearities carry predictive content beyond the estimation sample.

The institutional and structural determinants of the relationship warrant more direct examination as well. Income groups proxy for a set of underlying mechanisms without identifying any of them, so future cross-country work could model the interaction between credit and regulatory quality, contract enforcement, financial inclusion, banking competition, trade exposure, human capital, infrastructure and technological capacity, while allowing countries to move between categories over time. Finally, extending the analysis beyond the quantitative dimension of financial development – incorporating borrowing costs, maturity, collateral requirements, borrower access, bank soundness and loan purpose alongside credit volume – would address the mechanisms that the present design can invoke but not identify, clarifying when a larger stock of credit raises sectoral value added and when additional lending produces diminishing, negligible or negative returns.

9.10. Conclusion

This dissertation began from a familiar and, as it turns out, insufficiently discriminating proposition, namely that financial development measured by credit intermediation exerts a stable and positive influence on economic performance. The evidence assembled in the two empirical investigations indicates why that proposition requires qualification rather than rejection. Credit matters, and in several settings it matters considerably, but its relationship with real activity is neither uniform nor monotonic, varying across agriculture, industry and services, across low-, middle- and high-income economies, between the pre-crisis and post-crisis regimes, between short and long horizons and, within a single national economy, across eleven distinct branches of production.

The cross-country evidence establishes that the sectors respond to the pace of credit expansion in ways no aggregate specification could recover. Agriculture displays no stable association in any period, with its single significant curvature reversing sign under controls and resting on an imprecisely located turning point. Industry displays significant nonlinearity in every period and under both estimators, with the interval over which faster expansion accompanies industrial growth narrowing as income rises until, in high-income economies, the estimated maximum falls below zero. Services move in the opposite direction during the post-crisis decade, faster credit growth accompanying a rising service share across virtually the whole observed range. Dynamic estimation confirms the persistence of sectoral shares and corroborates the industrial results while leaving the agricultural and service estimates unsupported, which indicates that credit growth operates within a broader process of structural adjustment rather than as an independent force acting upon it.

The Turkish evidence supplies the resolution that the international panel cannot. A negative average short-run coefficient conceals an inverted U-shaped short-run relationship with a Mean Group threshold near 7.23 billion Turkish liras, and an insignificant average long-run coefficient conceals four sectors with positive and significant long-run estimates. Sectoral thresholds differ by more than an order of magnitude, long-run profiles range from inverted U through U to N-shaped and

inverted-N configurations, and three sectors possess inflection points rather than genuine turning points. Absorptive capacity is therefore not a national characteristic but a sectoral one, shaped by production lags, asset structure, working-capital requirements, collateral quality, regulation, demand conditions and the credit level already in place.

The conclusion this supports is conditional rather than categorical. The volume and the pace of credit may support sectoral activity where financing constraints bind and where borrowed resources are deployed within productive capacity, while the marginal contribution may diminish, disappear or reverse where expansion outruns the capacity of a sector to absorb it, where borrowing reflects leverage or distress, or where complementary institutional conditions are absent. Since these mechanisms are inferred from the pattern of estimates rather than measured directly, the turning points reported throughout should be treated as points of departure for closer investigation rather than as lending limits.

Effective financial policy accordingly requires that volume be assessed alongside destination, timing, sectoral structure and macroeconomic context. Aggregate financial depth remains an informative summary indicator, yet it cannot by itself explain how finance supports the real economy, since an economy-wide expansion may conceal underfunding in some activities and excessive leverage in others. By estimating the relationship separately by sector, by income level, by period and by horizon, and by insisting that the shape of each relationship be established rather than assumed, this dissertation offers a framework for moving past the question of whether more finance is beneficial or harmful. The more informative question is therefore not whether an economy has too much finance, but ‘too much finance for whom’ – for which sectors, at which stage of development and over which horizon additional credit ceases to be converted into durable value creation.

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APPENDICES

A. SUPPLEMENTARY RESULTS: MODELS WITH GOVERNMENT CONSUMPTION INTERACTION

Table A.1. Linear Fixed-Effects Estimates with Driscoll–Kraay Standard Errors and GOV Interaction: Agricultural, Industrial and Services Value Added, 2005–2019

Variables	Model (1) 2005–2019 Agriculture	Model (2) 2005–2019 Industry	Model (3) 2005–2019 Services
d_ln_ACGDP	0.412 (1.247)	—	—
d_ln_PC	—	-1.689 (4.067)	-1.313 (4.134)
GOV	0.474 (0.677)	-0.551* (0.287)	0.505* (0.254)
ACGDP_GOV	-0.039 (0.472)	—	—
PC_GOV	—	-0.002 (1.619)	0.845 (1.501)
Year Fixed Effects	Yes	Yes	Yes
Country Fixed Effects	Yes	Yes	Yes
Observations	557	1,015	1,011
Number of Countries	47	82	81
Within R^2	0.218	0.035	0.087
Driscoll–Kraay Lags	3	3	3

Notes: Dependent variables are AVASGDP in Model (1), IVASGDP in Model (2) and SVASGDP in Model (3). Fixed-effects estimates with Driscoll–Kraay robust standard errors (lag = 3) in parentheses. d_ln_ACGDP and d_ln_PC are the first differences of ln_ACGDP and ln_PC, and ACGDP_GOV and PC_GOV the corresponding interaction terms. Country and year fixed effects are included, and year coefficients are omitted. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A.2. Linear Two-Step System GMM Estimates with GOV Interaction: Agricultural, Industrial and Services Value Added, 2005–2019

Variables	Model (1) 2005–2019 Agriculture	Model (2) 2005–2019 Industry	Model (3) 2005–2019 Services
L.Dep. variable	0.914*** (0.152)	0.903*** (0.093)	1.056*** (0.125)
Credit variable	-4.198 (4.302)	-2.766 (21.673)	24.284 (44.179)
GOV	0.710 (0.953)	-0.818 (1.211)	0.824 (1.396)
Credit × GOV	1.457 (1.535)	1.305 (5.916)	-11.254 (15.520)
Year Fixed Effects	Yes	Yes	Yes
Constant	-1.232 (5.036)	4.628 (5.513)	-4.299 (6.503)
Observations	555	1009	1005
Countries	47	82	81
Instruments	27	27	27
AR(1) (p-value)	0.001	0.040	0.015
AR(2) (p-value)	0.693	0.625	0.869
Hansen Test (p-value)	0.668	0.249	0.357
Sargan Test (p-value)	0.085	0.000	0.000
Difference-in-Hansen (p-value)	0.391	0.300	0.330
Wald/F Statistic	718.34***	4179.23***	20565.05***

Notes: Dependent variables are AVASGDP, IVASGDP and SVASGDP in Models (1)–(3). Two-step System GMM estimates with Windmeijer (2005) finite-sample-corrected robust standard errors in parentheses. The credit variable is d_ln_ACGDP in Model (1) and d_ln_PC in Models (2)–(3), and Credit × GOV denotes ACGDP_GOV and PC_GOV. Year fixed effects are included, and their coefficients are omitted. The lagged dependent variable, credit, GOV and the interaction enter as collapsed GMM-style instruments at lags 2–3, and year indicators as standard instruments. The 2013 indicator in Models (1)–(2) and the 2012 indicator in Model (3) were omitted because of collinearity. Diagnostic entries are p-values. Stata used a generalized inverse for a singular two-step moment covariance matrix, so Difference-in-Hansen results require caution. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A.3. Quadratic Fixed-Effects Estimates with Driscoll–Kraay Standard Errors and GOV Interaction: Agricultural, Industrial and Services Value Added, 2005–2019

Variables	Model (1) 2005–2019 Agriculture	Model (2) 2005–2019 Industry	Model (3) 2005–2019 Services
d_ln_ACGDP	-0.149 (0.984)	—	—
d_ln_ACGDP²	-0.578* (0.316)	—	—
d_ln_PC	—	-2.867 (4.066)	-2.305 (3.869)
d_ln_PC²	—	12.625 (13.548)	2.220 (8.536)
GOV	0.388 (0.733)	-0.436 (0.378)	0.546* (0.262)
ACGDP_GOV	0.155 (0.398)	—	—
ACGDP²_GOV	0.165 (0.098)	—	—
PC_GOV	—	0.551 (1.500)	1.019 (1.315)
PC²_GOV	—	-4.695 (5.385)	-1.422 (3.418)
Year Fixed Effects	Yes	Yes	Yes
Country Fixed Effects	Yes	Yes	Yes
Observations	557	1,015	1,011
Number of Countries	47	82	81
Within R²	0.220	0.039	0.091
Driscoll–Kraay Lags	3	3	3

Notes: Dependent variables are AVASGDP in Model (1), IVASGDP in Model (2) and SVASGDP in Model (3). Fixed-effects estimates with Driscoll–Kraay robust standard errors (lag = 3) in parentheses. d_ln_ACGDP and d_ln_PC are the first differences of ln_ACGDP and ln_PC, and d_ln_ACGDP² and d_ln_PC² their squared terms. ACGDP_GOV and PC_GOV denote the interactions of GOV with the linear credit terms, and ACGDP²_GOV and PC²_GOV those with the squared terms. Country and year fixed effects are included, and year coefficients are omitted. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A.4. Quadratic Two-Step System GMM Estimates with GOV Interaction: Agricultural, Industrial and Services Value Added, 2005–2019

Variables	Model (1) 2005–2019 Agriculture	Model (2) 2005–2019 Industry	Model (3) 2005–2019 Services
L.Dep. variable	0.895*** (0.102)	0.870*** (0.322)	0.934*** (0.106)
Credit variable	-3.012 (3.058)	-34.469 (81.703)	43.466 (51.340)
Credit variable²	0.785 (1.005)	14.773 (127.282)	-15.564 (116.239)
GOV	1.221 (1.202)	-1.751 (3.891)	0.440 (2.273)
Credit × GOV	1.164 (1.166)	9.462 (27.856)	-14.249 (17.483)
Credit² × GOV	-0.371 (0.424)	-4.343 (50.789)	7.374 (53.329)
Year Fixed Effects	Yes	Yes	Yes
Constant	-2.344 (4.721)	8.117 (18.642)	2.349 (8.652)
Observations	555	1,009	1,005
Countries	47	82	81
Instruments	33	33	33
AR(1) (p-value)	0.001	0.122	0.031
AR(2) (p-value)	0.647	0.689	0.784
Hansen Test (p-value)	0.646	0.855	0.533
Sargan Test (p-value)	0.195	0.001	0.000
Difference-in-Hansen (p-value)	0.541	0.721	0.680
Wald/F Statistic	477.98***	337.99***	2997.47***

Notes: Dependent variables are AVASGDP, IVASGDP and SVASGDP in Models (1)–(3). Two-step System GMM estimates with Windmeijer (2005) finite-sample-corrected robust standard errors in parentheses. The credit variable is d_ln_ACGDP in Model (1) and d_ln_PC in Models (2)–(3), with squared terms d_ln_ACGDP2 and d_ln_PC2. Credit × GOV denotes ACGDP_GOV or PC_GOV, and Credit² × GOV denotes ACGDP2_GOV or PC2_GOV. Year fixed effects are included, and their coefficients are omitted. The lagged dependent variable, the credit terms, GOV and the interactions enter as collapsed GMM-style instruments at lags 2–3, and year indicators as standard instruments. Stata omitted the 2013 indicator in Model (1) and the 2012 indicator in Models (2)–(3) because of collinearity. Diagnostic entries are p-values. Stata used a generalized inverse for a singular two-step moment covariance matrix, so Difference-in-Hansen results require caution. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

B. SUPPLEMENTARY RESULTS: AGGREGATE GDP PER CAPITA GROWTH MODELS

Table B.1. Linear, Quadratic and Cubic Two-Step System GMM Estimates: GDP per Capita Growth, 1992–2007

Variables	Model (1) 1992-2007 Linear	Model (2) 1992-2007 Quadratic	Model (3) 1992-2007 Cubic
L.GDPPG	0.470 (0.305)	0.539 (0.385)	0.153 (0.275)
d_ln_PC	3.281 (6.137)	2.609 (8.703)	16.035* (8.413)
d_ln_PC2	—	-0.535 (9.932)	2.714 (6.666)
d_ln_PC3	—	—	-10.032 (6.657)
Year Effects	Yes	Yes	Yes
Observations	2,600	2,600	2,600
Number of Countries	179	179	179
Number of Instruments	23	26	29
GMM Lag Range	2-3	2-3	2-3
Collapsed Instruments	Yes	Yes	Yes
AR(1) Test p-value	0.014	0.025	0.013
AR(2) Test p-value	0.322	0.345	0.713
Sargan Test p-value	0.000	0.000	0.000
Hansen Test p-value	0.088	0.087	0.136
Difference-in-Hansen p-value	0.020	0.028	0.071

Notes: Dependent variable: GDPPG. Two-step System GMM estimates with Windmeijer (2005) finite-sample-corrected robust standard errors in parentheses. All specifications are estimated for 1992–2007. L.GDPPG and the included d_ln_PC polynomial terms enter as collapsed GMM-style instruments at lags 2–3, and year indicators as standard instruments. Year coefficients and constants are omitted.

Diagnostic entries are p-values, and Difference-in-Hansen refers to the exogeneity test for the level-equation instruments. Stata used a generalized inverse for a singular two-step moment covariance matrix, so Difference-in-Hansen results require caution. *** p < 0.01, ** p < 0.05, * p < 0.10.

Table B.2. Linear, Quadratic and Cubic Two-Step System GMM Estimates: GDP per Capita Growth, 2010–2019

Variables	Model (1) 2010-2019 Linear	Model (2) 2010-2019 Quadratic	Model (3) 2010-2019 Cubic
L.GDPPG	-0.409** (0.201)	-0.374 (0.310)	-0.251 (0.225)
d_ln_PC	4.756 (6.291)	11.062 (8.809)	-0.410 (9.369)
d_ln_PC2	—	-10.539 (28.509)	10.576 (19.650)
d_ln_PC3	—	—	-21.565 (56.258)
Year Effects	Yes	Yes	Yes
Observations	1,748	1,748	1,748
Number of Countries	183	183	183
Number of Instruments	17	20	23
GMM Lag Range	2-3	2-3	2-3
Collapsed Instruments	Yes	Yes	Yes
AR(1) Test p-value	0.294	0.265	0.609
AR(2) Test p-value	0.227	0.449	0.364
Sargan Test p-value	0.091	0.463	0.876
Hansen Test p-value	0.308	0.181	0.094
Difference-in-Hansen p-value	0.551	0.719	0.052

Notes: Dependent variable: GDPPG. Two-step System GMM estimates with Windmeijer (2005) finite-sample-corrected robust standard errors in parentheses. All specifications are estimated for 2010–2019. L.GDPPG and the included d_ln_PC polynomial terms enter as collapsed GMM-style instruments at lags 2–3, and year indicators as standard instruments. Year coefficients and constants are omitted.

Diagnostic entries are p-values, and Difference-in-Hansen refers to the exogeneity test for the level-equation instruments. Stata used a generalized inverse for a singular two-step moment covariance matrix, so Difference-in-Hansen results require caution. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

C. SUPPLEMENTARY RESULTS: PERIODICAL ANALYSIS

Table C.1. Linear, Quadratic and Cubic Fixed-Effects Estimates with Driscoll–Kraay Standard Errors: Agricultural Value Added, 1992–2007

Variables	Model (1) 1992–2007 Baseline	Model (2) 1992–2007 Quadratic	Model (3) 1992–2007 Cubic
d_ln_ACGDP	0.004 (0.012)	0.006 (0.011)	0.016 (0.024)
d_ln_ACGDP²	—	-0.007 (0.012)	-0.005 (0.010)
d_ln_ACGDP³	—	—	-0.002 (0.003)
Year Fixed Effects	Yes	Yes	Yes
Country Fixed Effects	Yes	Yes	Yes
Observations	793	793	793
Number of Countries	86	86	86
Within R^2	0.315	0.315	0.316
Driscoll–Kraay Lags	3	3	3

Notes: Dependent variable: ln_AVASGDP. Fixed-effects estimates with Driscoll–Kraay robust standard errors (maximum lag = 3) in parentheses. d_ln_ACGDP is the first difference of ln_ACGDP, and d_ln_ACGDP² and d_ln_ACGDP³ its squared and cubed terms. Country and year fixed effects are included, and year coefficients are omitted. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table C.2. Linear, Quadratic and Cubic Two-Step System GMM Estimates: Agricultural Value Added, 1992–2007

Variables	Model (1) 1992–2007 Baseline	Model (2) 1992–2007 Quadratic	Model (3) 1992–2007 Cubic
L.ln_AVASGDP	1.010*** (0.016)	1.009*** (0.016)	1.011*** (0.015)
d_ln_ACGDP	-0.081 (0.124)	-0.042** (0.020)	-0.175 (0.107)
d_ln_ACGDP²	—	0.002 (0.013)	-0.003 (0.021)
d_ln_ACGDP³	—	—	0.010 (0.011)
Year Fixed Effects	Yes	Yes	Yes
Constant	-0.075** (0.034)	-0.075** (0.032)	-0.072** (0.032)
Observations	785	785	785
Countries	86	86	86
Instruments	22	25	28
AR(1) (p-value)	0.046	0.059	0.046
AR(2) (p-value)	0.201	0.269	0.175
Hansen Test (p-value)	0.266	0.358	0.639
Sargan Test (p-value)	0.576	0.738	0.894
Difference-in-Hansen (p-value)	0.224	0.178	0.584
Wald/F Statistic	40044.73***	37517.66***	26342.05***

Notes: Dependent variable: ln_AVASGDP. Two-step System GMM estimates with Windmeijer (2005) finite-sample-corrected robust standard errors in parentheses. d_ln_ACGDP is the first difference of ln_ACGDP, and d_ln_ACGDP² and d_ln_ACGDP³ its squared and cubed terms. Year fixed effects are included, and their coefficients are omitted. The lagged dependent variable and the credit terms enter as collapsed GMM-style instruments at lags 2–3. AR(1), AR(2), Hansen, Sargan and Difference-in-Hansen entries are p-values. Stata used a generalized inverse for a singular two-step moment covariance matrix, so Difference-in-Hansen results require caution. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table C.3. Linear, Quadratic and Cubic Fixed-Effects Estimates with Driscoll–Kraay Standard Errors: Agricultural Value Added, 2010–2019

Variables	Model (1) 2010–2019 Baseline	Model (2) 2010–2019 Quadratic	Model (3) 2010–2019 Cubic
d_ln_ACGDP	0.007 (0.011)	0.008 (0.010)	0.016 (0.018)
d_ln_ACGDP²	—	0.008 (0.005)	0.006 (0.008)
d_ln_ACGDP³	—	—	-0.003 (0.004)
Year Fixed Effects	Yes	Yes	Yes
Country Fixed Effects	Yes	Yes	Yes
Observations	1,018	1,018	1,018
Number of Countries	108	108	108
Within R^2	0.066	0.067	0.067
Driscoll–Kraay Lags	3	3	3

Notes: Dependent variable: ln_AVASGDP. Fixed-effects estimates with Driscoll–Kraay robust standard errors (maximum lag = 3) in parentheses. d_ln_ACGDP is the first difference of ln_ACGDP, and d_ln_ACGDP² and d_ln_ACGDP³ its squared and cubed terms. Country and year fixed effects are included, and year coefficients are omitted. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table C.4 Linear, Quadratic and Cubic Two-Step System GMM Estimates: Agricultural Value Added, 2010–2019

Variables	Model (1) 2010–2019 Baseline	Model (2) 2010–2019 Quadratic	Model (3) 2010–2019 Cubic
L.ln_AVASGDP	1.023*** (0.167)	0.942*** (0.088)	0.941*** (0.072)
d_ln_ACGDP	0.054 (0.061)	0.125 (0.090)	0.138 (0.135)
d_ln_ACGDP²	—	-0.098 (0.114)	-0.181* (0.107)
d_ln_ACGDP³	—	—	0.007 (0.052)
Year Fixed Effects	Yes	Yes	Yes
Constant	-0.077 (0.306)	0.082 (0.161)	0.093 (0.138)

Observations	1,017	1,017	1,017
Countries	108	108	108
Instruments	17	20	23
AR(1) (p-value)	0.000	0.037	0.059
AR(2) (p-value)	0.103	0.474	0.733
Hansen Test (p-value)	0.076	0.285	0.293
Sargan Test (p-value)	0.001	0.341	0.418
Difference-in-Hansen (p-value)	0.115	0.798	0.569
Wald/F Statistic	29315.89***	6421.29***	8594.60***

Notes: Dependent variable: ln_AVASGDP. Two-step System GMM estimates with Windmeijer (2005) finite-sample-corrected robust standard errors in parentheses. d_ln_ACGDP is the first difference of ln_ACGDP, and d_ln_ACGDP² and d_ln_ACGDP³ its squared and cubed terms. Year fixed effects are included, and their coefficients are omitted. The lagged dependent variable and the credit terms enter as collapsed GMM-style instruments at lags 2–3. AR(1), AR(2), Hansen, Sargan and Difference-in-Hansen entries are p-values. Stata used a generalized inverse for a singular two-step moment covariance matrix, so Difference-in-Hansen results require caution. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table C.5. Linear, Quadratic and Cubic Two-Step System GMM Estimates: Agricultural Value Added, 1992–2019

Variables	Model (1) 1992–2019 Baseline	Model (2) 1992–2019 Quadratic	Model (3) 1992–2019 Cubic
L.ln_AVASGDP	0.972*** (0.021)	0.977*** (0.020)	0.968*** (0.023)
d_ln_ACGDP	0.005 (0.039)	0.069** (0.029)	0.069 (0.064)
d_ln_ACGDP²	—	0.007 (0.014)	0.007 (0.018)
d_ln_ACGDP³	—	—	-0.000004 (0.006848)

Year Fixed Effects	Yes	Yes	Yes
Constant	0.049 (0.043)	0.035 (0.040)	0.053 (0.044)
Observations	1,979	1,979	1,979
Countries	109	109	109
Instruments	34	37	40
AR(1) (p-value)	0.002	0.002	0.002
AR(2) (p-value)	0.101	0.070	0.090
Hansen Test (p-value)	0.865	0.770	0.849
Sargan Test (p-value)	0.928	0.716	0.688
Difference-in-Hansen (p-value)	0.653	0.537	0.714
Wald/F Statistic	13260.14***	11658.34***	11269.25***

Notes: Dependent variable: ln_AVASGDP. Two-step System GMM estimates with Windmeijer (2005) finite-sample-corrected robust standard errors in parentheses. d_ln_ACGDP is the first difference of ln_ACGDP, and d_ln_ACGDP² and d_ln_ACGDP³ its squared and cubed terms. Year fixed effects are included, and their coefficients are omitted. The lagged dependent variable and the credit terms enter as collapsed GMM-style instruments at lags 2–3. AR(1), AR(2), Hansen, Sargan and Difference-in-Hansen entries are p-values. Stata used a generalized inverse for a singular two-step moment covariance matrix, so Difference-in-Hansen results require caution. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table C.6. Linear, Quadratic and Cubic Fixed-Effects Estimates with Driscoll–Kraay Standard Errors: Services Value Added, 1992–2007

Variables	Model (1) 1992–2007 Baseline	Model (2) 1992–2007 Quadratic	Model (3) 1992–2007 Cubic
d_ln_PC	0.008 (0.022)	0.008 (0.022)	0.009 (0.022)
d_ln_PC²	—	0.001 (0.016)	0.001 (0.017)

d_ln_PC³	—	—	-0.002 (0.008)
Year Fixed Effects	Yes	Yes	Yes
Country Fixed Effects	Yes	Yes	Yes
Observations	2,298	2,298	2,298
Number of Countries	172	172	172
Within R^2	0.039	0.039	0.039
Driscoll–Kraay Lags	3	3	3

Notes: Dependent variable: ln_SVASGDP. Fixed-effects estimates with Driscoll–Kraay robust standard errors (maximum lag = 3) in parentheses. d_ln_PC is the first difference of ln_PC, and d_ln_PC² and d_ln_PC³ its squared and cubed terms. Country and year fixed effects are included, and year coefficients are omitted. *** p < 0.01, ** p < 0.05, * p < 0.10.

Table C.7. Linear, Quadratic and Cubic Two-Step System GMM Estimates: Services Value Added, 1992–2007

Variables	Model (1) 1992–2007 Baseline	Model (2) 1992–2007 Quadratic	Model (3) 1992–2007 Cubic
L.ln_SVASGDP	0.905*** (0.066)	0.921*** (0.067)	0.912*** (0.078)
d_ln_PC	0.010 (0.097)	0.043 (0.069)	-0.074 (0.156)
d_ln_PC²	—	0.027 (0.055)	0.002 (0.049)
d_ln_PC³	—	—	0.051 (0.073)
Year Fixed Effects	Yes	Yes	Yes
Constant	0.370 (0.264)	0.303 (0.268)	0.349 (0.313)
Observations	2,261	2,261	2,261
Countries	172	172	172
Instruments	23	26	29

AR(1) (p-value)	0.000	0.000	0.000
AR(2) (p-value)	0.038	0.086	0.034
Hansen Test (p-value)	0.373	0.480	0.400
Sargan Test (p-value)	0.515	0.667	0.666
Difference-in-Hansen (p-value)	0.314	0.915	0.424
Wald/F Statistic	417656.57***	262988.99***	471349.16***

Notes: Dependent variable: ln_SVASGDP. Two-step System GMM estimates with Windmeijer (2005) finite-sample-corrected robust standard errors in parentheses. d_ln_PC is the first difference of ln_PC, and d_ln_PC² and d_ln_PC³ its squared and cubed terms. Year fixed effects are included, and their coefficients are omitted. The lagged dependent variable and the credit terms enter as collapsed GMM-style instruments at lags 2–3. AR(1), AR(2), Hansen, Sargan and Difference-in-Hansen entries are p-values. Stata used a generalized inverse for a singular two-step moment covariance matrix, so Difference-in-Hansen results require caution. At the 10% level, the null of no second-order serial correlation is rejected in all three specifications, so the estimates require caution. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table C.8. Linear, Quadratic and Cubic Two-Step System GMM Estimates: Services Value Added, 2010–2019

Variables	Model (1) 2010–2019 Baseline	Model (2) 2010–2019 Quadratic	Model (3) 2010–2019 Cubic
L.ln_SVASGDP	0.689*** (0.150)	0.609 (0.371)	0.794*** (0.299)
d_ln_PC	-0.082 (0.115)	-0.013 (0.084)	-0.003 (0.122)
d_ln_PC²	—	0.330 (0.459)	0.232 (0.723)
d_ln_PC³	—	—	-0.706 (0.576)
Year Fixed Effects	Yes	Yes	Yes
Constant	1.270** (0.601)	1.539 (1.475)	0.809 (1.190)
Observations	1,732	1,732	1,732
Countries	182	182	182
Instruments	17	20	23

AR(1) (p-value)	0.041	0.280	0.234
AR(2) (p-value)	0.746	0.731	0.445
Hansen Test (p-value)	0.795	0.879	0.366
Sargan Test (p-value)	0.025	0.858	0.257
Difference-in-Hansen (p-value)	0.662	0.533	0.568
Wald/F Statistic	136448.74***	53326.42***	41768.68***

Notes: Dependent variable: ln_SVASGDP. Two-step System GMM estimates with Windmeijer (2005) finite-sample-corrected robust standard errors in parentheses. d_ln_PC² and d_ln_PC³ are the squared and cubed terms of d_ln_PC. Year fixed effects are included, and their coefficients are omitted. The lagged dependent variable and the credit terms enter as collapsed GMM-style instruments at lags 2–3. Diagnostic entries are p-values. Stata used a generalized inverse for a singular two-step moment covariance matrix, so Difference-in-Hansen results require caution. In Model (1) the non-robust Sargan test rejects the overidentifying restrictions whereas the robust Hansen test does not, and AR(1) is not rejected in Models (2) and (3), unlike the pattern ordinarily expected for first-differenced residuals. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table C.9. Linear, Quadratic and Cubic Fixed-Effects Estimates with Driscoll–Kraay Standard Errors: Services Value Added, 1992–2019

Variables	Model (1) 1992–2019 Baseline	Model (2) 1992–2019 Quadratic	Model (3) 1992–2019 Cubic
d_ln_PC	0.014 (0.018)	0.013 (0.017)	0.011 (0.019)
d_ln_PC²	—	-0.007 (0.017)	-0.004 (0.024)
d_ln_PC³	—	—	0.003 (0.010)
Year Fixed Effects	Yes	Yes	Yes
Country Fixed Effects	Yes	Yes	Yes
Observations	4,375	4,375	4,375
Number of Countries	185	185	185

Within R^2	0.1451	0.1452	0.1452
Driscoll–Kraay Lags	3	3	3

Notes: Dependent variable: ln_SVASGDP. Fixed-effects estimates with Driscoll–Kraay robust standard errors (maximum lag = 3) in parentheses. d_ln_PC is the first difference of ln_PC, and d_ln_PC² and d_ln_PC³ its squared and cubed terms. Country and year fixed effects are included, and year coefficients are omitted. *** p < 0.01, ** p < 0.05, * p < 0.10.

Table C.10. Linear, Quadratic and Cubic Two-Step System GMM Estimates: Services Value Added, 1992–2019

Variables	Model (1) 1992–2019 Baseline	Model (2) 1992–2019 Quadratic	Model (3) 1992–2019 Cubic
L.ln_SVASGDP	0.800*** (0.126)	0.756*** (0.139)	0.774*** (0.160)
d_ln_PC	-0.037 (0.083)	-0.071 (0.082)	-0.055 (0.117)
d_ln_PC²	—	0.104 (0.095)	-0.000 (0.070)
d_ln_PC³	—	—	0.030 (0.068)
Year Fixed Effects	Yes	Yes	Yes
Constant	0.797 (0.506)	0.971* (0.561)	0.889 (0.635)
Observations	4,331	4,331	4,331
Countries	185	185	185
Instruments	35	38	41
AR(1) (p-value)	0.001	0.000	0.002
AR(2) (p-value)	0.356	0.307	0.371
Hansen Test (p-value)	0.037	0.126	0.069
Sargan Test (p-value)	0.001	0.003	0.015
Difference-in-Hansen (p-value)	0.066	0.380	0.097
Wald/F Statistic	160507.55***	61456.34***	96668.49***

Notes: Dependent variable: ln_SVASGDP. Two-step System GMM estimates with Windmeijer (2005) finite-sample-corrected robust standard errors in parentheses. d_ln_PC is the first difference of ln_PC, and d_ln_PC² and d_ln_PC³ its squared and

cubed terms. Year fixed effects are included, and their coefficients are omitted. The lagged dependent variable and the credit terms enter as collapsed GMM-style instruments at lags 2–3. Year indicators enter as standard instruments. AR(1), AR(2), Hansen, Sargan and Difference-in-Hansen entries are p-values. Stata used a generalized inverse for a singular two-step moment covariance matrix, so Difference-in-Hansen results require caution. At the 10% level AR(2) is satisfactory in all models, but Hansen and Difference-in-Hansen reject the relevant nulls in Models (1) and (3), and the non-robust Sargan test rejects the overidentifying restrictions in all models. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

D. SUPPLEMENTARY RESULTS: INCOME-LEVEL ANALYSIS

Table D.1. Linear, Quadratic and Cubic Two-Step System GMM Estimates: Agricultural Value Added, Low-Income Countries, 1992–2019

Variables	Model (1) Linear	Model (2) Quadratic	Model (3) Cubic
L.ln_AVASGDP	1.004*** (0.005)	1.003*** (0.008)	0.999*** (0.009)
d_ln_ACGDP	-0.100 (0.114)	0.060 (0.078)	0.145* (0.081)
d_ln_ACGDP²	—	-0.013 (0.034)	-0.032 (0.037)
d_ln_ACGDP³	—	—	-0.021 (0.019)
Year Fixed Effects	Yes	Yes	Yes
Constant	Omitted	Omitted	Omitted
Observations	276	276	276
Countries	14	14	14
Instruments	34	37	40
AR(1) (p-value)	0.026	0.016	0.017
AR(2) (p-value)	0.916	0.607	0.567
Hansen Test (p-value)	1.000	1.000	1.000
Sargan Test (p-value)	0.211	0.667	0.826
F Statistic	10782.90***	13321.76***	29338.53***

Notes: Dependent variable: ln_AVASGDP. Two-step System GMM estimates with Windmeijer (2005) finite-sample-corrected robust standard errors in parentheses. d_ln_ACGDP is the first difference of ln_ACGDP, and d_ln_ACGDP² and d_ln_ACGDP³ its squared and cubed terms. Year fixed effects are included, and their coefficients are omitted. The lagged dependent variable and the credit terms enter as collapsed GMM-style instruments at lags 2–3. AR(1), AR(2), Hansen and Sargan entries are p-values. Stata reported both a large instrument count relative to the sample and a singular two-step moment covariance matrix, for which a generalized inverse was used. The number of instruments (34–40) also exceeds the 14 country groups and every Hansen p-value equals 1.000, so the Hansen tests and the estimates require substantial caution. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table D.2. Linear, Quadratic and Cubic Fixed-Effects Estimates with Driscoll–Kraay Standard Errors: Agricultural Value Added, Middle-Income Countries, 1992–2019

Variables	Model (1) 1992–2019 Baseline	Model (2) 1992–2019 Quadratic	Model (3) 1992–2019 Cubic
d_ln_ACGDP	-0.006 (0.014)	-0.005 (0.013)	-0.016 (0.024)
d_ln_ACGDP²	—	0.004 (0.005)	0.008 (0.009)
d_ln_ACGDP³	—	—	0.005 (0.005)
Constant	2.479*** (0.025)	2.479*** (0.025)	2.480*** (0.025)
Year Fixed Effects	Yes	Yes	Yes
Country Fixed Effects	Yes	Yes	Yes
Observations	1,003	1,003	1,003
Number of Countries	59	59	59
Within R^2	0.2398	0.2399	0.2404
F Statistic	248609.87***	422692.68***	1.64e+07***
Driscoll–Kraay Lags	3	3	3

Notes: Dependent variable: ln_AVASGDP. Fixed-effects estimates with Driscoll–Kraay robust standard errors (maximum lag = 3) in parentheses. d_ln_ACGDP is the first difference of ln_ACGDP, and d_ln_ACGDP² and d_ln_ACGDP³ its squared and cubed terms. Country and year fixed effects are included, and year coefficients are omitted. The sample is restricted to middle-income countries (income_class = 2). *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table D.3. Linear, Quadratic and Cubic Two-Step System GMM Estimates: Agricultural Value Added, Middle-Income Countries, 1992–2019

Variables	Model (1) Linear	Model (2) Quadratic	Model (3) Cubic
L.ln_AVASGDP	0.924*** (0.054)	0.904*** (0.064)	0.959*** (0.073)
d_ln_ACGDP	0.044 (0.103)	0.061 (0.104)	0.181 (0.187)
d_ln_ACGDP²	—	-0.024 (0.066)	0.006 (0.093)
d_ln_ACGDP³	—	—	-0.038 (0.043)
Year Fixed Effects	Yes	Yes	Yes
Constant	0.171 (0.123)	0.219 (0.146)	0.085 (0.167)
Observations	999	999	999
Countries	59	59	59
Instruments	34	37	40
AR(1) (p-value)	0.000	0.001	0.001
AR(2) (p-value)	0.328	0.476	0.678
Hansen Test (p-value)	0.617	0.588	0.363
Sargan Test (p-value)	0.543	0.414	0.580
Difference-in-Hansen (p-value)	0.270	0.799	0.200
F Statistic	12848.59***	20060.80***	20673.35***

Notes: Dependent variable: ln_AVASGDP. Two-step System GMM estimates with Windmeijer (2005) finite-sample-corrected robust standard errors in parentheses. d_ln_ACGDP is the first difference of ln_ACGDP, and d_ln_ACGDP² and d_ln_ACGDP³ its squared and cubed terms. Year fixed effects are included, and their coefficients are omitted. The lagged dependent variable and the credit terms enter as collapsed GMM-style instruments at lags 2–3. AR(1), AR(2), Hansen, Sargan and Difference-in-Hansen entries are p-values. Stata used a generalized inverse for a singular two-step moment covariance matrix, so Difference-in-Hansen results require caution. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table D.4. Linear, Quadratic and Cubic Fixed-Effects Estimates with Driscoll–Kraay Standard Errors: Agricultural Value Added, High-Income Countries, 1992–2019

Variables	Model (1) 1992–2019 Baseline	Model (2) 1992–2019 Quadratic	Model (3) 1992–2019 Cubic
d_ln_ACGDP	0.009 (0.032)	0.014 (0.038)	0.024 (0.048)
d_ln_ACGDP²	—	-0.008 (0.012)	-0.005 (0.012)
d_ln_ACGDP³	—	—	-0.001 (0.003)
Constant	1.173*** (0.026)	1.173*** (0.026)	1.173*** (0.026)
Year Fixed Effects	Yes	Yes	Yes
Country Fixed Effects	Yes	Yes	Yes
Observations	694	694	694
Number of Countries	35	35	35
Within R^2	0.4715	0.4720	0.4721
F Statistic	196776.37***	199877.81***	144083.13***
Driscoll–Kraay Lags	3	3	3

Notes: Dependent variable: ln_AVASGDP. Fixed-effects estimates with Driscoll–Kraay robust standard errors (maximum lag = 3) in parentheses. d_ln_ACGDP is the first difference of ln_ACGDP, and d_ln_ACGDP² and d_ln_ACGDP³ its squared and cubed terms. Country and year fixed effects are included, and year coefficients are omitted. The sample is restricted to high-income countries (income_class = 3). *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table D.5. Linear, Quadratic and Cubic Two-Step System GMM Estimates: Agricultural Value Added, High-Income Countries, 1992–2019

Variables	Model (1) Linear	Model (2) Quadratic	Model (3) Cubic
L.ln_AVASGDP	0.975*** (0.026)	0.975*** (0.034)	1.298*** (0.257)
d_ln_ACGDP	-0.095 (0.071)	-0.131 (0.098)	0.046 (0.047)
d_ln_ACGDP²	—	0.026 (0.028)	-0.297*** (0.098)
d_ln_ACGDP³	—	—	0.062*** (0.019)
Year Fixed Effects	Yes	Yes	Yes
Constant	0.028 (0.020)	-0.051** (0.021)	-0.086 (0.118)
Observations	689	689	689
Countries	35	35	35
Instruments	34	37	40
AR(1) (p-value)	0.042	0.053	0.072
AR(2) (p-value)	0.189	0.178	0.149
Hansen Test (p-value)	0.131	0.411	0.952
Sargan Test (p-value)	0.314	0.524	0.854
Difference-in-Hansen (p-value)	0.063	0.251	1.000
F Statistic	3948.77***	18048.34***	8.57***

Notes: Dependent variable: ln_AVASGDP. Two-step System GMM estimates with Windmeijer (2005) finite-sample-corrected robust standard errors in parentheses. d_ln_ACGDP is the first difference of ln_ACGDP, and d_ln_ACGDP² and d_ln_ACGDP³ its squared and cubed terms. Year fixed effects are included, and their coefficients are omitted. The lagged dependent variable and the credit terms enter as collapsed GMM-style instruments at lags 2–3. AR(1), AR(2), Hansen, Sargan and Difference-in-Hansen entries are p-values. Stata used a generalized inverse for a singular two-step moment covariance matrix, so Difference-in-Hansen results require caution. Stata also warned that the instrument count may be large relative to the sample in the quadratic and cubic models, so the diagnostics, particularly the Difference-in-Hansen statistic of the cubic model, require caution. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table D.6. Linear, Quadratic and Cubic Two-Step System GMM Estimates: Industrial Value Added, Low-Income Countries, 1992–2019

Variables	Model (1) Linear	Model (2) Quadratic	Model (3) Cubic
L.ln_IVASGDP	0.988*** (0.014)	0.986*** (0.052)	0.955*** (0.065)
d_ln_PC	0.294 (0.362)	Omitted	Omitted
d_ln_PC²	—	Omitted	Omitted
d_ln_PC³	—	—	Omitted
Year Fixed Effects	Yes	Yes	Yes
Constant	Omitted	Omitted	Omitted
Observations	519	519	519
Countries	23	23	23
Instruments	35	38	41
AR(1) (p-value)	0.004	0.209	0.005
AR(2) (p-value)	0.289	0.313	0.373
Hansen Test (p-value)	1.000	1.000	1.000
Sargan Test (p-value)	0.001	0.010	0.003
F Statistic	14668.14***	48901.15***	2236.74***
Prob > F	0.000	0.000	0.000

Notes: Dependent variable: ln_IVASGDP. Two-step System GMM estimates with Windmeijer (2005) finite-sample-corrected robust standard errors in parentheses. d_ln_PC is the first difference of ln_PC, and d_ln_PC² and d_ln_PC³ its squared and cubed terms. A dash indicates that a term was not included and 'Omitted' that Stata removed the regressor during estimation, so both credit terms were omitted from the quadratic model and all three from the cubic model. Year fixed effects are included, and their coefficients are omitted. The lagged dependent variable and the credit terms enter as collapsed GMM-style instruments at lags 2–3. AR(1), AR(2), Hansen and Sargan entries are p-values. Stata reported a large instrument count relative to the sample and a singular two-step moment covariance matrix, for which a generalized inverse was used. The number of instruments (35–41) also exceeds the 23 country groups and every Hansen p-value equals 1.000, so the diagnostics and the estimates require substantial caution. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table D.7. Linear, Quadratic and Cubic Two-Step System GMM Estimates: Industrial Value Added, High-Income Countries, 1992–2019

Variables	Model (1) Linear	Model (2) Quadratic	Model (3) Cubic
L.ln_IVASGDP	0.934*** (0.042)	0.940*** (0.036)	0.932*** (0.027)
d_ln_PC	0.127 (0.198)	0.175 (0.359)	-0.022 (0.172)
d_ln_PC²	—	0.057 (0.283)	0.308 (0.378)
d_ln_PC³	—	—	0.233 (0.282)
Year Fixed Effects	Yes	Yes	Yes
Constant	0.186 (0.127)	0.162 (0.100)	0.200** (0.083)
Observations	1,448	1,448	1,448
Countries	62	62	62
Instruments	35	38	41
AR(1) (p-value)	0.000	0.000	0.000
AR(2) (p-value)	0.047†	0.072†	0.068†
Hansen (p-value)	0.780	0.189	0.739
Sargan (p-value)	0.757	0.818	0.885
Diff.-in-Hansen (p-value)	0.756	0.469	0.767
F Statistic	82800.96***	63991.16***	152156.57***

Notes: Dependent variable: ln_IVASGDP. Two-step System GMM estimates with Windmeijer (2005) finite-sample-corrected robust standard errors in parentheses. d_ln_PC is the first difference of ln_PC, and d_ln_PC² and d_ln_PC³ its squared and cubed terms. Year fixed effects are included, and their coefficients are omitted. The lagged dependent variable and the credit terms enter as collapsed GMM-style instruments at lags 2–3. AR(1), AR(2), Hansen, Sargan and Difference-in-Hansen entries are p-values. † The null of no second-order serial correlation is rejected at the 10% level in every model (AR(2) p = 0.047, 0.072 and 0.068), so none of the specifications fully satisfies this System GMM validity condition and the estimates should not be treated as benchmark evidence. Stata used a generalized inverse for a singular two-step moment covariance matrix, so Difference-in-Hansen results require caution. *** p < 0.01, ** p < 0.05, * p < 0.10.

Table D.8. Linear, Quadratic and Cubic Fixed-Effects Estimates with Driscoll–Kraay Standard Errors: Services Value Added, Low-Income Countries, 1992–2019

Variables	Model (1) Linear	Model (2) Quadratic	Model (3) Cubic
d_ln_PC	-0.042 (0.049)	-0.041 (0.054)	-0.028 (0.072)
d_ln_PC²	—	0.007 (0.083)	-0.002 (0.085)
d_ln_PC³	—	—	-0.042 (0.084)
Year Fixed Effects	Yes	Yes	Yes
Constant	3.719*** (0.006)	3.719*** (0.007)	3.720*** (0.008)
Observations	519	519	519
Countries	22	22	22
Within R-squared	0.0615	0.0616	0.0617
F Statistic	1908.67***	2760.01***	1738.25***
Prob > F	0.000	0.000	0.000
Driscoll–Kraay Lag	3	3	3

Notes: Dependent variable: ln_SVASGDP. Fixed-effects estimates with Driscoll–Kraay robust standard errors (maximum lag = 3) in parentheses. d_ln_PC is the first difference of ln_PC, and d_ln_PC² and d_ln_PC³ its squared and cubed terms. Country and year fixed effects are included, and year coefficients are omitted. All three models use the same sample of 519 observations from 22 low-income countries. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table D.9. Linear, Quadratic and Cubic Two-Step System GMM Estimates: Services Value Added, Low-Income Countries, 1992–2019

Variables	Model (1) Linear	Model (2) Quadratic	Model (3) Cubic
L.ln_SVASGDP	1.005*** (0.005)	1.006*** (0.005)	1.003*** (0.008)
d_ln_PC	Omitted	Omitted	Omitted
d_ln_PC²	—	Omitted	Omitted
d_ln_PC³	—	—	Omitted
Year Fixed Effects	Yes	Yes	Yes
Constant	Omitted	Omitted	Omitted
Diagnostic Statistics			
Observations	517	517	517
Countries	22	22	22
Instruments	35	38	41
AR(1) (p-value)	0.003	0.006	0.009
AR(2) (p-value)	0.199	0.204	0.289
Hansen Test (p-value)	1.000	1.000	1.000
Sargan Test (p-value)	0.000	0.000	0.038
F Statistic	87136.29***	23267.08***	95336.19***
Prob > F	0.000	0.000	0.000

Notes: Dependent variable: ln_SVASGDP. Two-step System GMM estimates with Windmeijer (2005) finite-sample-corrected robust standard errors in parentheses. d_ln_PC is the first difference of ln_PC, and d_ln_PC² and d_ln_PC³ its squared and cubed terms. A dash indicates that a term was not included and 'Omitted' that Stata removed the regressor, and all credit terms included in the three specifications were omitted, so their effects cannot be identified. Year fixed effects are included, and their coefficients are omitted. The lagged dependent variable and the credit terms enter as collapsed GMM-style instruments at lags 2–3. AR(1), AR(2), Hansen and Sargan entries are p-values. Although AR(2) does not reject the absence of second-order serial correlation, the instruments (35–41) exceed the 22 country groups, all Hansen p-values equal 1.000, all Sargan tests reject at the 10% level, and a generalized inverse was used for a singular two-step moment covariance matrix, so these specifications should not be treated as reliable evidence. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table D.10. Linear, Quadratic and Cubic Two-Step System GMM Estimates: Services Value Added, Middle-Income Countries, 1992–2019

Variables	Model (1) Linear	Model (2) Quadratic	Model (3) Cubic
L.ln_SVASGDP	0.828*** (0.121)	0.788*** (0.092)	0.861*** (0.089)
d_ln_PC	0.011 (0.058)	-0.014 (0.076)	0.025 (0.104)
d_ln_PC²	—	0.060 (0.071)	-0.048 (0.078)
d_ln_PC³	—	—	0.038 (0.088)
Year Fixed Effects	Yes	Yes	Yes
Constant	0.690 (0.480)	0.835** (0.366)	0.548 (0.355)
Observations	2,340	2,340	2,340
Countries	99	99	99
Instruments	35	38	41
AR(1) (p-value)	0.006	0.002	0.002
AR(2) (p-value)	0.819	0.808	0.761
Hansen Test (p-value)	0.392	0.593	0.344
Sargan Test (p-value)	0.105	0.328	0.140
F Statistic	245822.57***	145250.28***	305591.68***
Prob > F	0.000	0.000	0.000

Notes: Dependent variable: ln_SVASGDP. Two-step System GMM estimates with Windmeijer (2005) finite-sample-corrected robust standard errors in parentheses. d_ln_PC is the first difference of ln_PC, and d_ln_PC² and d_ln_PC³ its squared and cubed terms. A dash indicates that a term was not included in the corresponding specification. Year fixed effects are included, and their coefficients are omitted. The lagged dependent variable and the credit terms enter as collapsed GMM-style instruments at lags 2–3. AR(1), AR(2), Hansen and Sargan entries are p-values. AR(1) is significant in all three specifications whereas AR(2) does not reject the absence of second-order serial correlation, neither Hansen nor Sargan rejects the overidentifying restrictions at the 10% level, and the instrument count remains below the number of country groups. Stata used a generalized inverse for a singular two-step moment covariance matrix, so Difference-in-Hansen results require caution. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table D.11. Linear, Quadratic and Cubic Fixed-Effects Estimates with Driscoll–Kraay Standard Errors: Services Value Added, High-Income Countries, 1992–2019

Variables	Model (1) Linear	Model (2) Quadratic	Model (3) Cubic
d_ln_PC	0.021 (0.027)	0.022 (0.030)	0.018 (0.037)
d_ln_PC²	—	0.003 (0.024)	0.022 (0.054)
d_ln_PC³	—	—	0.014 (0.035)
Year Fixed Effects	Yes	Yes	Yes
Constant	4.017*** (0.002)	4.017*** (0.002)	4.017*** (0.002)
Observations	1,458	1,458	1,458
Countries	62	62	62
Within R-squared	0.2251	0.2251	0.2252
F Statistic	2143471.56***	1485466.55***	1253322.71***
Prob > F	0.000	0.000	0.000
Driscoll–Kraay Lag	3	3	3

Notes: Dependent variable: ln_SVASGDP. Fixed-effects estimates with Driscoll–Kraay robust standard errors (maximum lag = 3) in parentheses. d_ln_PC is the first difference of ln_PC, and d_ln_PC² and d_ln_PC³ its squared and cubed terms. Country and year fixed effects are included, and year coefficients are omitted. All three models use the same sample of 1,458 observations from 62 high-income countries, and none of the credit terms is significant at the 10% level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table D.12. Linear, Quadratic and Cubic Two-Step System GMM Estimates: Services Value Added, High-Income Countries, 1992–2019

Variables	Model (1) Linear	Model (2) Quadratic	Model (3) Cubic
L.ln_SVASGDP	0.585** (0.232)	0.857*** (0.128)	0.910*** (0.085)
d_ln_PC	-0.086 (0.163)	0.026 (0.155)	0.090 (0.103)
d_ln_PC²	—	-0.035 (0.081)	-0.397 (0.342)
d_ln_PC³	—	—	-0.230 (0.222)
Year Fixed Effects	Yes	Yes	Yes
Constant	1.708* (0.954)	0.583 (0.527)	0.369 (0.348)
Observations	1,437	1,437	1,437
Countries	62	62	62
Instruments	35	38	41
AR(1) (p-value)	0.010	0.001	0.000
AR(2) (p-value)	0.136	0.138	0.099
Hansen Test (p-value)	0.568	0.366	0.431
Sargan Test (p-value)	0.021	0.000	0.013
F Statistic	43162.65***	1270000.00***	916535.04***
Prob > F	0.000	0.000	0.000

Notes: Dependent variable: ln_SVASGDP. Two-step System GMM estimates with Windmeijer (2005) finite-sample-corrected robust standard errors in parentheses. d_ln_PC is the first difference of ln_PC, and d_ln_PC² and d_ln_PC³ its squared and cubed terms. A dash indicates that a term was not included in the corresponding specification. Year fixed effects are included, and their coefficients are omitted. The lagged dependent variable and the credit terms enter as collapsed GMM-style instruments at lags 2–3. AR(1), AR(2), Hansen and Sargan entries are p-values. AR(1) rejects the absence of first-order serial correlation in all three specifications, and AR(2) does not reject the absence of second-order serial correlation in the linear and quadratic models, whereas the cubic model rejects at the 10% level ($p = 0.099$). Hansen does not reject the overidentifying restrictions and the instrument count remains below the number of country groups, but Sargan rejects in all three models. Stata used a generalized inverse for a singular two-step moment covariance matrix, so Difference-in-Hansen results require caution. The cubic specification in particular requires caution. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

E. CURRICULUM VITAE

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F. TURKISH SUMMARY / TÜRKÇE ÖZET

Kimin İçin Çok Fazla Finans? Kredi-Çıktı Bağlantısını Yeniden Düşünmek

Finansal gelişme ile ekonomik performans arasındaki ilişki bir yüzyıldan uzun süredir iktisatçıların gündeminde kalmıştır ve güncelliğini korumasının nedeni, verilen yanıtın toplumların sermaye tahsisini nasıl örgütlediğine doğrudan dokunmasıdır. Schumpeter (1911/1934), kalkınma sürecinin merkezine bankacıyı yerleştirmiş ve kredinin girişimciye henüz kazanılmamış kaynaklar üzerinde tasarrufta bulunma imkânı verdiğini savunmuştur. Modern literatür bu sezgiyi, finansal sistemlerin yerine getirdiği bir işlevler kümesine dönüştürerek tasarrufların harekete geçirilmesi, bilgi ve işlem maliyetlerinin azaltılması, yatırımın izlenmesi, riskin çeşitlendirilmesi ve fonların en yüksek getirili kullanımlara yönlendirilmesi biçiminde tanımlamıştır (Levine, 2005). Üç on yıl boyunca biriken ülkeler arası kanıtlar bu anlatıyı geniş ölçüde desteklemiştir. King ve Levine (1993) finansal derinlik ölçütlerinin sonraki dönem büyümesini, sermaye birikimini ve verimlilik artışını öngördüğünü, Levine ve Zervos (1998) bankacılık ile hisse senedi piyasası gelişiminin birbirinden bağımsız açıklayıcı güç taşıdığını, Rajan ve Zingales (1998) ise dış finansmana bağımlı sanayilerin finansal açıdan gelişmiş ekonomilerde orantısız biçimde daha hızlı büyüdüğünü göstermiştir. Beck ve diğerleri (2000) finansal aracılığın reel faaliyete ulaştığı daha sağlam kanalın fiziksel sermaye birikimi değil toplam faktör verimliliği olduğunu belirlerken, Levine ve diğerleri (2000) hukuk kökeni farklılıkları yoluyla ayrıştırılan dışsal bileşenin uzun dönem büyümeyle olumlu ilişkili olduğunu ortaya koymuştur.

Bu çalışmalar bütünüyle okunduğunda finansal gelişme, uzun dönem ekonomik performansın edilgen bir yansıması değil belirleyicisi olarak konumlanmıştır. Ne var ki söz konusu birikim, çoğunlukla tekdüze bir ilişki tarif ettiği biçiminde yorumlanmış, yani daha derin finansal sistemlerin her gelişme düzeyinde daha iyi reel sonuçlar ürettiği varsayılmıştır. 2008 küresel finans krizi ise bu ön kabulün

kırılganlığını açığa çıkarmıştır. Krizi önceleyen yıllardaki hızlı kredi genişlemesi, kaynakların yanlış tahsisi, kaldıraç birikimi, spekülatif pozisyonların büyümesi ve nihayetinde reel faaliyette sert daralmalarla birlikte gerçekleşmiş, bu da araştırma gündemini finansın büyümeyi teşvik edip etmediği sorusundan hangi koşullarda, hangi büyüklüklerde ve hangi ufuklarda teşvik ettiği sorusuna doğru kaydırmıştır.

Bu yeniden yönelimden doğan literatür genellikle "çok fazla finans" hipotezi çevresinde düzenlenmekte ve finansal gelişmenin marjinal katkısının derinleşme düzeyleri boyunca sabit olmadığını ileri sürmektedir. Cecchetti ve Kharroubi (2012) finansal sektör büyüklüğünün büyümeye katkısının belirli bir noktadan sonra negatife döndüğünü bildirmiş, Arcand ve diğerleri (2015) eşiği özel kredinin ekonominin büyüklüğüne yaklaştığı düzeye yerleştirmiş, Law ve Singh (2014) dinamik panel eşik tahmincisiyle benzer bir sonuca ulaşmış, Samargandi ve diğerleri (2015) ise orta gelirli ekonomilerden oluşan bir örneklemde ters U biçimli bir örüntü belgelemiştir. Tersine dönüşü açıklamak üzere çeşitli mekanizmalar önerilmiştir. Finansman kısıtları gevşedikten sonra ilave borç verme, üretken yatırım yerine tüketime, gayrimenkule ya da mevcut yükümlülüklerin yeniden finansmanına yönelebilmekte; bu da kredinin bileşimini kendi başına sonuç doğuran bir unsur hâline getirmektedir, çünkü işletme kredileri büyümeyle ilişkiliyken hanehalkı kredileri aynı ilişkiyi taşımamaktadır (Beck vd., 2012). Cecchetti ve Kharroubi (2015) bir kaynak kanalı tanımlayarak hızla genişleyen bir finansal sektörün nitelikli işgücünü diğer faaliyetlerden çektiğini ve teminatlandırılması kolay projeleri daha üretken olanlara tercih ettiğini, en büyük kaybı ise araştırma yoğun sanayilerin ve maddi varlığı az olan firmaların üstlendiğini savunmaktadır.

Zayıf borçlulara kredi verilmeye devam edilmesi yeniden yapılanmayı geciktirebilmekte ve sektör düzeyinde verimliliği baskılayabilmektedir; nitekim Caballero ve diğerleri (2008) bu örüntüyü Japonya için belgelemiştir. Makroekonomik düzeyde ise hızlı kredi ve kaldıraç artışı sistemik kriz olasılığını yükseltmektedir (Schularick ve Taylor, 2012). Bununla birlikte eşiklerin kendisi örnekleme, döneme ve tahminciye karşı duyarlı çıkmakta, Botev ve diğerleri (2019) finansal gelişmenin, ötesinde her koşulda zararlı hâle geldiği evrensel bir devrilme noktası bulunmadığını raporlamaktadır. Bu da ilişkinin tek bir küresel parametreyle

yönetilen değil, doğrusal olmayan ve bağlama duyarlı bir yapı taşıdığını düşündürmektedir.

Literatürde çok daha az ilgi görmüş ikinci bir güçlük, bu tezin analizini belirleyici biçimde şekillendirmiştir. Doğrusal olmayan literatürün neredeyse tamamı finansal gelişmeyi bir stok olarak, yani ödenmemiş kredinin GSYİH'ye oranı biçiminde ölçmektedir; oysa kredi oranları, düzey üzerinden düzey regresyonlarının yorumlanabildiği koşulları nadiren sağlayan yüksek düzeyde ısrarcı serilerdir. Trend içeren iki serinin birbiri üzerine regresyonunda istatistiksel olarak anlamlı katsayılar iktisadi ilişkileri değil ortak trendleri yansıtabilir; bu tehlike sahte regresyonun klasik ele alınışında ortaya konmuştur (Granger ve Newbold, 1974). Yatay kesit bağımlılığına izin veren ikinci nesil panel birim kök yöntemleri, özellikle Pesaran (2007) tarafından geliştirilen yatay kesitle genişletilmiş test, ortak şokların yaygın olduğu makroekonomik panellerde uygun tanılayıcıyı sunmaktadır. Bu testin tez için derlenen veriye uygulanması açık bir sonuç vermiştir: tarım, sanayi ve hizmetlerin sektörel katma değer payları düzeylerinde durağanken, mevduat bankalarınca sağlanan özel kredi ile tarım, ormancılık ve balıkçılığa açılan kredi birinci dereceden bütünleşiktir.

Bu bulgunun sonucu olarak kredi değişkenleri tahmin edilen her denkleme birinci farklarıyla girmekte ve ilgi duyulan parametre finansal derinliği değil kredi genişlemesinin hızını tanımlamaktadır. Söz konusu ikame, yalnızca ölçü biriminin değil analiz nesnesinin de değişmesi anlamına gelir ve geniş bir literatür kredinin stok ile akımının farklı davrandığını göstermektedir. Loayza ve Rancière (2006), finansal derinliğin olumlu etkilerini bildiren büyüme literatürü ile kredi büyüklüklerinin bankacılık krizlerinin en iyi öngörücüleri arasında yer aldığını bildiren kriz literatürü arasındaki görünürdeki çelişkiyi ele almış ve uzun ile kısa dönem etkilerini kapsayan bir modelden hareketle, aracılık ile üretim artışı arasındaki olumlu uzun dönem ilişkisinin finansal kırılganlıkla bağlantılı ve büyük ölçüde olumsuz bir kısa dönem ilişkisiyle bir arada bulunduğu sonucuna varmıştır. Beck ve diğerleri (2014) finansal sistemlerin aracılık işlevini toplam büyüklüğünden ayırarak uyumlu bir sonuca ulaşmakta, aracılık faaliyetlerinin uzun ufukta büyümeyi yükselttiğini ve oynaklığı azalttığını, diğer boyutlardaki genişlemenin ise uzun dönemde reel bir etkisinin bulunmadığını bildirmektedir.

En doğrudan kanıt Ductor ve Grechyna (2015) ile Bezemer ve diğerlerinden (2016) gelmektedir. İlk çalışma, finansal gelişmenin büyüme etkisinin özel kredinin reel üretim artışına göre genişleme hızına bağlı olduğunu ve kredi reel ekonomiyi geride bıraktığında etkinin negatife döndüğünü göstermekte; ikincisi kredi stoklarını kredi akımlarından ayırarak finansal olmayan işletmelere yönelen akımlar için pozitif büyüme etkileri raporlarken konut ve diğer varlık piyasası kredileri için aynı sonucu bulmamaktadır. Uzun dönemli makro-finansal tarihten gelen kanıtlar da kredi genişlemesinin hızını kendi başına bir inceleme nesnesi saymayı desteklemektedir. Schularick ve Taylor (2012) kredi artışını finansal krizlerin öncü göstergesi olarak belirlemiş, Mendoza ve Terrones (2008) kredi patlamalarına eşlik eden makroekonomik ve firma düzeyindeki düzenlilikleri belgelemiş, Jordà ve diğerleri (2013) kredi yoğun genişlemelerin daha derin resesyonlar ve daha yavaş toparlanmalarla izlendiğini göstermiştir. Borio ve diğerleri (2016) tartışmayı verimliliğe taşıyarak kredi patlamalarının işgücünü verimlilik artışı daha düşük sektörlere kaydırma eğiliminde olduğunu, Müller ve Verner (2024) ise patlamaların ticarete konu olmayan faaliyetleri orantısız biçimde beslediğini ve bu genişlemelerin sistematik olarak sonraki yavaşlamalarla ilişkilendiğini ortaya koymaktadır.

Buradan doğrudan çıkan bir yorum sonucu tezin tamamını yönetmektedir. Birinci farklarla kurulmuş bir modelden elde edilen dönüm noktası, bir değişim oranı olarak ifade edildiği için kredi büyümesi uzayında yer alır ve dolayısıyla Arcand ve diğerlerinın (2015), Law ve Singh'in (2014) ya da Cecchetti ve Kharroubi'nin (2012) bildirdiği GSYİH'nin yüzde sekseni ile yüzde yüzü arasındaki derinlik eşikleriyle karşılaştırılamaz. Kredinin marjinal katkısının belirli bir genişleme oranının ötesinde negatife döndüğü yönündeki bir bulgu, sektörlerin yeni borçlanmayı hangi hızda soğurabildiğine ilişkindir; bir ekonominin hangi finansal derinlik düzeyinde aşırı bankacılığa sahip olduğuna değil. İki büyüklüğün karıştırılması hem sonucu hem de onun politika içeriğini yanlış ifade eder.

Doğrusal olmayış ve ölçüm limitinin seçimi güçlüğün yalnızca bir bölümüdür. Geleneksel finans–büyüme çalışmaları finansal tarafta toplulaştırma yapmakta, ekonomi geneli bir kredi ölçütü ise teknoloji, yatırım ufku, teminat yapısı, işletme sermayesi gereksinimi ve katma değer üretme kapasitesi bakımından birbirinden

ayrışan faaliyetlere açılan borcu bir araya toplamaktadır. Tarımsal modernizasyon, sınai makine, bilgi teknolojisi, inşaat, finansal aracılık ya da hanehalkı tüketimi için kullandırılan kredinin eşdeğer reel etkiler üretmesi makul biçimde beklenemez ve bu sorun düzeylerden büyüme oranlarına geçilerek çözülmez, çünkü toplulaştırılmış bir kredi genişleme oranı da tıpkı toplulaştırılmış bir stok gibi kullanım yerleri üzerinden ortalama almaktadır. Tahsis literatürü bu noktayı birkaç yönden somutlaştırmıştır: Beck ve diğerleri (2012) kredi ile büyüme arasındaki ülkeler arası ilişkinin esas olarak işletme kredilerinden kaynaklandığını, Mian ve diğerleri (2017) hanehalkı borcundaki genişlemelerin orta vadeli yavaşlamaları öngördüğünü, Bahadir ve Gumus (2016) daha önce toplam krediye atfedilen döngüsel izin hanehalkı borçlanmasında işletme borçlanmasına kıyasla belirgin biçimde güçlü olduğunu, Bai ve diğerleri (2018) ise banka kredisinin bileşiminin daha üretken borçlulara kayması hâlinde sektör verimliliğinin yükseldiğini göstermiştir.

Hsu ve diğerleri (2014) yapısal bir boyut eklemekte, hisse senedi piyasası gelişiminin dışa bağımlı ve teknoloji yoğun sanayilerde yeniliği desteklerken kredi piyasası gelişiminin aynı işlevi görmediğini bildirmektedir; bu da ağırlıklı olarak banka temelli bir sistemin geleneksel sermaye birikimini etkin biçimde finanse ederken yenilik yoğun faaliyet için daha zayıf bir araç olarak kalabileceğini ima etmektedir. Çıktı tarafındaki toplulaştırma ise simetrik bir sorun doğurmaktadır. Toplam üretim, ortak bir finansal şoka farklı tepki veren sektörleri bir araya getirdiğinden bir faaliyetteki genişleme bir diğerindeki daralmayı dengeleyebilir ve toplamı büyük ölçüde değişmemiş bırakabilir; üretimin bileşimindeki kaymalar ise tek bir büyüme oranının içinde görünmez kalır. Bu nedenle toplulaştırılmış bir model, hangi sektörlerin daha hızlı kredi genişlemesinden yararlandığını, hangilerinin azalan getirilerle karşılaştığını ve hangilerinin ilave borcu üretken biçimde soğurma kapasitesini tüketmiş olabileceğini belirleyemez.

Tez bu sorunu, bağımlı değişken olarak sektörel katma değeri alarak ele almakta ve başlangıçta bir ölçüm açıklaması yapmaktadır; çünkü ülkeler arası modellerde sektörel katma değer GSYİH içindeki pay olarak ifade edilmektedir. Böyle bir paydaki değişim, faaliyetin mutlak üretimindeki değil göreli ağırlığındaki değişimi gösterir, dolayısıyla negatif bir katsayı sektörün daraldığını değil ekonominin geri

kalanına göre daha yavaş büyüdüğünü ima eder. Yapısal dönüşüm literatürü bu durumu sektörel başarısızlık kanıtı olarak değil kalkınmanın olağan bir özelliği olarak ele almaktadır. Aynı literatür ayrımın neden analitik olarak önemli olduğunu da açıklamaktadır. Büyüme yalnızca toplam üretimdeki bir artışı değil, işgücünün, sermayenin ve üretimin tarımdan sanayiye ve ardından hizmetlere yeniden tahsisini içerir (Herrendorf vd., 2014; McMillan ve Rodrik, 2011) ve finansal gelişme, her sektörün karşılaştığı göreli finansman koşullarını değiştirerek bu süreci biçimlendirebilir.

Sermaye yoğun sanayi tipik olarak uzun ufuklu dış finansmana bağımlıdır, tarım mevsimsel döngülere, biyolojik gecikmelere ve ortak riske uygun araçlar gerektirir, modern hizmetlerin çoğu ise teminatlandırılabilir tesisten çok maddi olmayan varlıklara, beşerî sermayeye ve yeniliğe dayanır. Yapısal değişimin bir aşamasına iyi uyarlanmış bir finansal sistemin bir sonraki aşamaya aynı ölçüde hizmet etmesi gerekmez ve bu aşamanın kendisi de zorlaşmıştır, zira Rodrik (2016) sanayileşme ile gelir arasındaki hörgüç biçimli ilişkinin aşağıya ve başlangıç noktasına doğru kaydığını, böylece ekonomilerin sanayileşme fırsatlarını daha erken ve daha düşük gelir düzeylerinde tükettiğini belgelemektedir. Ülkeler arası heterojenlik bu değerlendirmeleri güçlendirmektedir. Kredinin üretken katma değere dönüştürülmesi kurumsal kalite, finansal kapsayıcılık, teknolojik kapasite, beşerî sermaye, altyapı ve denetim etkinliğine bağlıdır ve bağlayıcı kısıt kalkınma aşamasına göre farklılaşır.

Greenwood ve Jovanovic (1990) finansal yapı ile gelir dağılımının birlikte evrimini modellemiş, Rioja ve Valev (2004a) finansal gelişmenin etkisinin çok düşük düzeylerde belirsiz, orta düzeylerde büyük, yüksek düzeylerde ise pozitif fakat daha küçük olduğunu savunmuş, Demetriades ve Law (2006) kurumsal bir mercekle uyumlu bir sonuca ulaşarak finansın en güçlü etkisini orta gelirli, en zayıf etkisini ise düşük gelirli ekonomilerde gösterdiğini bildirmiştir. Aizenman ve diğerleri (2015) tartışmayı sektörel düzeye taşıyarak finansal derinliğin üretim artışı üzerindeki etkisinde kayda değer sektörel ve bölgesel heterojenlik belgelemekte, birkaç sektörde negatif etkiler de raporlamaktadır. Sahay ve diğerleri (2015) farklandırılmış bir modele en doğrudan dokunan hususu eklemekte ve finansal derinleşmenin hızının denetim ve düzenleme kapasitesini aştığında kendine özgü riskler taşıdığı konusunda

uyarmaktadır. Zamansal bağlam da aynı nedenle önem taşımaktadır; nitekim Rousseau ve Wachtel (2011) finansal derinleşmenin büyüme üzerindeki tahmin edilen etkisinin daha yakın tarihli örneklemlerde zayıfladığını, bazı modellerde ise görünürde ortadan kalktığını belgelemektedir.

Bu nedenle dört önermenin ayrı ayrı değil birlikte incelenmesi gerekmektedir. Kredi ile reel faaliyet arasındaki ilişki doğrusal olmayabilir ve genişleme hızlandıkça yönü ile şiddeti değişebilir. Finansın ölçüldüğü limit sonuç doğurucudur, çünkü birikmiş kredi stoku ile onun genişleme hızı farklı ampirik özelliklere sahip ayrı nesnelerdir. Kredinin ve çıktının bileşimi belirleyicidir, zira toplulaştırılmış göstergeler büyüklük kadar işaret bakımından da ayrışabilen sektöre özgü dinamikleri gizlemektedir. Son olarak finansın etkisi bağlama bağlıdır ve kalkınma aşamaları, tarihsel dönemler ile kurumsal ortamlar arasında değişmektedir. Tezde ileri sürülen merkezi iddia bu dördünün birlikte ciddiye alınmasından doğmaktadır: finansal gelişme ile ekonomik performans arasındaki ilişkiyi ne tek bir katsayı ne de evrensel bir kredi eşiği nitelendirebilir; bu ilişkinin sektöre, kalkınma aşamasına, döneme ve zaman ufkuna göre ayrı ayrı değerlendirilmesi gerekir.

Ampirik çalışma birbirine bağlı bir dizi soru etrafında düzenlenmiştir. Kredi genişlemesinin hızı sektörel katma değerle istikrarlı bir ilişki taşımakta mıdır, yoksa marjinal katkısı genişleme hızlandıkça değişmekte midir? Tarımda, sanayide ve hizmetlerde gözlenen doğrusal ve doğrusal olmayan örüntüler sektörler arasında ortak mıdır yoksa her birine mi özgüdür? İlişki küresel kriz sonrasında değişmiş midir ve değiştiyse hangi yönde ve hangi faaliyetler için? Farklı gelir düzeylerindeki ekonomiler farklı fonksiyonel biçimler ve farklı dönüm noktaları sergilemekte midir? Tahmin edilen ilişkiler dinamik sürekliliğin modele dâhil edilmesine ve içsel olabilecek kredinin araçlanmasına dayanmakta mıdır? Ulusal kanıta geçildiğinde, sektörel kredideki değişimler tek bir ekonomi içinde kısa dönem üretim artışıyla nasıl ilişkilidir, sektörel kredi düzeylerini sektörel faaliyete bağlayan uzun dönem denge ilişkisi nedir ve sektörler ilave krediyi üretken biçimde soğurma kapasitesi bakımından sistematik olarak ayrışmakta mıdır? Son soru iki çerçeveyi birden kesmekte ve verinin bütünleşme özelliklerinden doğarak kredi genişlemesinin hızı ile finansal derinliğin düzeyinin, reel faaliyet için farklı sonuçlar taşıyan ayrıştırılabilir limitler olup olmadığını sormaktadır.

Tez bu soruları birbirine bağlı iki ampirik incelemeyle ele almaktadır. Birincisi geniş uluslararası kapsam, ikincisi ulusal derinlik sağlamaktadır. Bu eşleştirme, ülkeler arası bir panelin karşılaştırmalı erişimini tek bir ekonomi içinde elde edilebilen kurumsal ve sektörel çözünürlükle birleştirmekte; uluslararası analizin başlıca sektörler, gelir grupları ve finansal ortamlar arasında sistematik farklılıklar bulunup bulunmadığını saptamasına imkân verirken, Türkiye analizinin sektörel heterojenliği daha detaylı çözünürlükte incelemesine ve kısa dönem uyarlanmasını uzun dönem dengesinden ayırmasına olanak tanımaktadır. Her iki ortamda birden görünen sonuçlar daha geniş bir anlam taşımakta, ayrışmalar ise ulusal kurumların, politika tasarımının ve sektörel bileşimin etkisine işaret etmektedir.

Uluslararası analiz, 1992 ile 2019 arasını kapsayan dengesiz bir yıllık panel kullanmakta ve Dünya Bankası'nın her gelir kategorisinden ekonomileri içermektedir (Dünya Bankası, 2026); sanayi ve hizmet modellerinde 186 ülkeye, sektöre özgü kredi verisinin sınırladığı tarım modellerinde ise 109 ülkeye ulaşılmaktadır. Örneklemin 2019'da sonlandırılmasının nedeni, COVID-19 salgınının olağanüstü ve eşitsiz dağılmış bozulmalarının tahmin penceresinin dışında bırakılmasıdır; zira bu yıllar üretime, kredi arzına, kamu harcamalarına ve çıktının bileşimine aynı anda gelen şokları birleştirmektedir. Tüm dönem, 1992–2007 ve 2010–2019 için ayrı tahminlerle desteklenmekte, böylece kriz öncesi genişleme onu izleyen on yılla karşılaştırılabilmektedir. 2008 ve 2009 yılları krizin en akut evresine ve doğrudan politika tepkisine karşılık geldiğinden ve iki rejimden hiçbirine atanamayacağından alt dönemlerin dışında bırakılmıştır.

Ekonomik performans, Uluslararası Standart Sanayi Sınıflaması izlenerek üç sektöre ayrıştırılmakta (Birleşmiş Milletler, 2008) ve bağımlı değişkenler ormancılık ile balıkçılık dâhil tarımda, inşaat dâhil sanayide ve hizmetlerde yaratılan ve her biri GSYİH payı olarak ifade edilen katma değerin doğal logaritması biçiminde tanımlanmaktadır. Finansal gelişme, veri erişilebilirliğine ve sektörel uygunluğa göre ölçülmektedir. Tarım modelleri, finansmanı finanse edilen faaliyete doğrudan bağlayan ve GSYİH ile normalleştirilen tarım, ormancılık ve balıkçılık kredisini kullanırken; sanayi ve hizmet modelleri, geniş bir uluslararası örneklem için sektöre özgü kredi verisi bulunmadığından mevduat bankalarınca sağlanan özel kredinin

GSYİH'ye oranını kullanmaktadır. Bütün kredi ölçütleri yukarıda açıklanan nedenlerle logaritmalarının birinci farkıyla modele girmektedir. Modeller ayrıca beşerî sermayeyi, fiziksel yatırımı, kamu faaliyetinin ölçeğini, makroekonomik istikrarı ve dışa açıklığı temsil etmek üzere ortaöğretim okullaşma oranını, gayrisafi sermaye oluşumunu, kamu tüketimini, enflasyonu ve ticari açıklığı içermekte; okullaşma oranı da logaritmik düzey serisi birim kök taşıdığından birinci farkıyla yer almaktadır.

Tahmin stratejisi, yatay kesit bağımlılığı, eğim heterojenliği, değişen varyans, ardışık bağımlılık ve karma bütünleşme dereceleri sergileyen panelin tanılayıcı özelliklerine göre belirlenmiştir. Temel modeller bu nedenle ülke ve yıl sabit etkileriyle ve değişen varyans, ardışık bağımlılık ve ülkeler arası genel bağımlılık biçimleri altında güvenilirliğini koruyan Driscoll–Kraay standart hatalarıyla tahmin edilmektedir (Driscoll ve Kraay, 1998). Blundell ve Bond'u (1998) izleyen ve Windmeijer (2005) sonlu örneklem düzeltmesini içeren iki aşamalı Sistem GMM tahminleri, üretimdeki sürekliliği, ters nedenselliği ve ihmal edilen değişken içselliğini ele almakta; geçerlilikleri Arellano–Bond ardışık bağımlılık testleri (Arellano ve Bond, 1991) ile Hansen, Sargan ve Fark-Hansen istatistikleri (Hansen, 1982; Sargan, 1958) üzerinden, araç sayısının artmasının yol açtığı test gücü kaybına açıkça dikkat edilerek değerlendirilmektedir (Roodman, 2009).

Fonksiyonel biçim dayatılmak yerine veriden seçilmekte, doğrusal, kuadratik, kübik ve dördüncü dereceden modeller sırayla tahmin edilerek yalnızca ilave terimin istatistiksel olarak anlamlı olduğu ve uyumu iyileştirdiği durumlarda korunmaktadır. Anlamlı bir kare terimi tek başına gerçek bir dönüm noktası kurmadığından kuadratik modeller Lind ve Mehlum'un (2010) önerdiği teste tabi tutulmakta; bu test eğimin verinin alt sınırında pozitif, üst sınırında negatif olmasını ve dönüm noktasının gözlenen aralığın içinde bulunmasını gerektirmektedir. Kübik modeller ise birinci türevlerinin kökleri ve diskriminantı üzerinden değerlendirilmekte, böylece bir eğri ancak iki gerçek dönüm noktası belirlenebildiğinde N biçimli ya da ters N biçimli olarak tanımlanmaktadır. Her dönüm noktasının açıklayıcı değişkenin ampirik dağılımı içindeki yerinin belirlenmesi bu disiplinin bir başka unsurudur, zira dönüm noktaları uç kuyrukta kalan istatistiksel olarak anlamlı bir yüksek dereceli

terim, ülkelerin fiilen deneyimlediği aralıkta pratik olarak tekdüze bir ilişkiyi tarif etmektedir.

Toplulaştırılmış bir ön analiz, kendi başına bir bulgu olmaktan çok sektörel tasarımı gerekçelendirmektedir. Kişi başına GSYİH büyümesinin logaritmik özel kredinin birinci farkı üzerine tahmini her dönemde tutarlı biçimde iç bükey bir ilişki üretmekte, kare terimi her dönemde negatif ve yüksek düzeyde anlamlı çıkmaktadır. Dönüm noktası hesaplamaları büyümeyi maksimize eden kredi genişleme oranını 1992–2007 için logaritmik olarak −0,023'te, tüm dönem için −0,276'da konumlandırmakta; bunlar kredi-GSYİH oranında kabaca yüzde iki ve yüzde yirmi dört düzeyinde daralmalara karşılık gelmektedir. Ülke içi açıklayıcı güç ise 0,059 ile 0,122 arasında kalarak mütevazı seyretmektedir ki bu durum kendi başına bilgi vericidir; tek bir ekonomi geneli kredi ölçütünün finansın reel faaliyete ulaşabileceği mekanizmaların yalnızca bir kısmını yakaladığını ve bunların belirlenmesi için ayrıştırma gerektiğini göstermektedir.

Sektörel sonuçlar bu beklentiyi doğrulamakta, üç sektör hiçbir toplulaştırılmış modelin değerlendiremeyeceği bir örüntü tarif etmektedir. Tarım, her iki alt dönemde de kredi genişlemesinin hızı ile tarımın katma değer payı arasında istikrarlı bir ilişki sergilememekte; kredi terimlerinin anlamlılığa ulaştığı tek model olan Driscoll–Kraay standart hatalarıyla tahmin edilmiş tüm dönem kuadratik modeli ise incelendiğinde kırılgan çıkmaktadır. Lind ve Mehlum (2010) testi monotonluğu yüzde on düzeyinde reddetmekte ve dönüm noktası hesaplamaları ise maksimumu yıllık yaklaşık yüzde altmış beşlik tarımsal kredi artışına denk gelen 0,502'ye yerleştirmektedir; ancak makroekonomik kontroller eklendiğinde eğrilik işaret değiştirmekte ve dönüm noktası için hesaplanan Fieller (1954) aralığı −0,719 ile 3,329 arasında uzanarak hem negatif değerleri hem de makul olmayan ölçüde hızlı genişlemeleri kapsamaktadır. Savunulabilir okuma, tarımsal kredi artışının en fazla zayıf ve konumu belirsiz bir doğrusal olmayışa sahip olduğudur ve bu, tarımsal finansmanın etkinliğinin kredi hacminden çok kırsal kredi piyasalarının kurumsal koşullarına bağlı olduğu savıyla uyumludur (Conning ve Udry, 2007).

Sanayi, her iki tahmincinin de üç dönemin tamamında istatistiksel olarak anlamlı kredi katsayıları ürettiği tek sektördür ve ilişkinin biçimi finansal rejimler arasında

içerik taşıyan bir biçimde değişmektedir. Kriz öncesi pencerede statik tahminci, kontrollerin eklenmesine dayanmayan zayıf pozitif bir doğrusal katsayı verirken, dinamik tahminci 0,510'da yerel bir maksimum ima eden anlamlı ve negatif bir kübik terim belirlemektedir; bu, yıllık yaklaşık yüzde altmış yedilik özel kredi artışına karşılık gelmekte ve kriz öncesi dağılımın doksan dokuzuncu yüzdeliğinin hemen altında yer almaktadır. Dolayısıyla ılımlı kredi artışı büyük ölçüde nötr ya da hafif olumlu görünürken, olağanüstü hızlı genişleme sanayi konumunun gerilemesine eşlik etmektedir. Bu asimetri, kredi ve kaldıraçtaki hızlı genişlemelerin sonraki makroekonomik zayıflık olasılığını yükselttiği yönündeki kredi patlaması literatürüyle örtüşmektedir (Mendoza ve Terrones, 2008; Schularick ve Taylor, 2012); ancak Dell'Ariccia ve diğerleri (2016) patlama epizotlarının yaklaşık üçte birinin finansal krizle izlendiğini bildirerek olumsuz sonucu bir düzenlilikten çok bir eğilim olarak konumlandırmaktadır.

Kriz sonrası on yıl, dönemsel analizin en açık sonuçlarını üretmekte ve bu tabloyu tersine çevirmektedir. Katsayı örüntüsü temel ve genişletilmiş modeller arasında istikrarlıdır, türevin diskriminantı yüzde bir düzeyinde pozitif ve anlamlıdır, dolayısıyla iki ayrı dönüm noktası varsayılmak yerine kurulmaktadır ve 0,246'daki üst dönüm noktası, medyanı 0,010 olan bir dağılımın doksan beşinci ile doksan dokuzuncu yüzdelikleri arasında yer almaktadır. Marjinal ilişkinin negatif olduğu aralık böylece verinin neredeyse tamamını kapsamakta; ilişki ancak yirmi ülke-yıl gözleminden azının ulaştığı genişleme oranlarında yeniden olumlu hâle gelmektedir. Bu profilin, model içi belirlilik katsayısını 0,087'den 0,240'a yükselten kontrollerin eklenmesine dayanması, ilişkinin yatırım ya da ticaret koşullarının indirgenmiş biçimdeki bir yansımasından ibaret olmadığını göstermektedir; gayrisafi sermaye oluşumunun güçlü pozitif katsayısı da aynı on yıl boyunca fiziksel yatırımın sınai faaliyeti desteklediğini doğrulayarak bu okumayı pekiştirmektedir.

Tüm örneklemde her iki tahminci genel biçim üzerinde uzlaşırken maksimumun tam yeri konusunda ayrışmakta, statik model onu 0,216'ya, dinamik model 0,156'ya yerleştirmektedir; bunlar sırasıyla yaklaşık yüzde yirmi dört ve yüzde on yedilik yıllık kredi artışına karşılık gelmektedir. Dinamik modelde köklerin hiçbiri tek başına anlamlı değildir ve statik modelde diskriminant sıfırdan ayırt

edilememektedir, dolayısıyla uzlaşma bir eşiğin belirlenmesinden çok bir dönüş bölgesinin genel konumunda buluşma olarak nitelendirilmelidir. Kredi bileşimi, sanayideki ilişkinin 2008 sonrasında neden zayıfladığına dair en makul açıklamayı sunmaktadır. Cecchetti ve Kharroubi'nin (2015) kaynak kanalı, Beck ve diğerlerinin (2012) işletme kredisi bulgusu, Bezemer ve diğerlerinin (2016) akım ayrıştırması ve Borio ve diğerlerinin (2016) verimlilik bulgusu burada gözlenen örüntüyü üretebilecek niteliktedir; ancak hiçbiri toplulaştırılmış bir kredi ölçütüyle sınanamaz, bu nedenle sınai sonuçlar kanalı belirlemekten çok ilişkiyi kurmakta ve aday açıklamaların alanını daraltmaktadır.

Hizmetler ters yönde hareket etmekte ve bu karşıtlık dönemsel analizin merkezî ampirik bulgusunu oluşturmaktadır. Ne kriz öncesi pencere ne de tüm örneklem her iki tahminci altında anlamlı kredi katsayıları vermekte, buna karşılık kriz sonrası on yılda doğrusal terim hem temel hem de genişletilmiş modelde pozitif ve yüksek düzeyde anlamlı çıkmakta, temel modelde 0,141, kontroller eklendiğinde 0,066 olarak tahmin edilmektedir. Uygulanan kübik modelin ima ettiği tersine dönüş 0,405'te, yani dağılımın doksan dokuzuncu yüzdeliğinin ötesinde yer aldığından ilişki gözlenen aralığın neredeyse tamamında pozitif ve artan seyretmekte, eğriliği ise güvenle nitelendirilememektedir. Asıl önemli olan, aynı açıklayıcı değişkenin, aynı paydanın ve aynı on yılın sanayide negatif, hizmetlerde pozitif bir katsayı üretmesidir; ölçeklendirmeye atfedilemeyecek bu asimetri tahsise işaret etmekte ve kriz sonrası borç verilmesinin sınai kapasiteden çok gayrimenkule, tüketime ve hizmet faaliyetlerine yönelmiş olabileceğini düşündürmektedir.

Örneklemin gelir düzeyine göre bölünmesi, kalkınma aşamaları boyunca tek bir finans–büyüme eğrisinin geçerli olmadığını doğrulamaktadır. Tarımsal kredi artışı yalnızca düşük gelirli ekonomilerde kayda geçmekte, burada zayıf desteklenen bir ters U maksimumu 0,414'e yerleştirmekte ve gözlemlerin büyük çoğunluğu eğrinin yukarı eğimli bölümünde kalmaktadır; bu da daha hızlı tarımsal kredi artışının bu ekonomilerin fiilen faaliyet gösterdiği aralıkta daha yüksek tarımsal katma değere eşlik etme eğiliminde olduğunu ima etmekte ve küçük ölçekli tarım finansmanındaki kısıtların sürdüğü değerlendirmesiyle uyuşmaktadır (Dünya Bankası, 2007). Sanayi üç grupta da kayda geçmekte, ancak farklı fonksiyonel biçimlerle ve daha önemlisi

tahmin edilen maksimumun gelir yükseldikçe istikrarlı biçimde sola kaymasıyla belirmektedir.

Düşük gelirli ülkelerde durağan noktaları −0,480 ve 0,230'da bulunan kübik bir profil, verinin çalışma aralığını azalan orta bölüme yerleştirmektedir; bu grupta özel kredi artışının çeyrekler arası açıklığı 0,041 medyanı çevresinde −0,067 ile 0,133 arasında seyrettiğinden, daha hızlı kredi genişlemesi bu ekonomilerin tipik olarak faaliyet gösterdiği aralıkta daha düşük bir sanayi payına eşlik etme eğilimindedir. Orta gelirli ülkelerde uydurulan eğri, gözlemlerin yoğunlaştığı bölgede yaklaşık olarak iç bükeydir; statik tahminci maksimumu doksan beşinci yüzdeliğe yakın biçimde 0,309'a, dinamik kuadratik model ise bu gruptaki medyan kredi artış oranına neredeyse tam olarak denk gelen 0,026'ya yerleştirmektedir. Bu da gözlemlerin kabaca yarısının, marjinal ilişkinin negatife döndüğü noktanın ötesinde bulunduğu anlamına gelmektedir. Yüksek gelirli ülkelerde ise maksimum, sıfırın altında kalan −0,224'te konumlanmakta, dolayısıyla pozitif kredi artışının bütün aralığı azalan kola düşmektedir.

Orta gelirli sanayi kanıtı, bu çalışmada "çok fazla finans" önermesine verilen en güçlü desteği sunmaktadır ve bunu, önermenin kuramsal gerekçesinin en açık olduğu sektörde yapmaktadır; zira bu ekonomilerde imalat üretimin kayda değer bir payını oluştururken denetim çerçeveleri henüz olgunlaşmaktadır (Cecchetti ve Kharroubi, 2015; Sahay vd., 2015). Yüksek gelirli ekonomilerdeki örüntü ise finansın bu ülkelerde önemini yitirdiği iddiasına başvurmadan açıklanabilmektedir. Rodrik (2016) sanayileşme fırsatlarının daha erken tükendiğini belgelemekte, hanehalkı kredisinin payı gelirle birlikte yükselmekte (Beck vd., 2012) ve hanehalkı borç oranındaki artışlar sonraki dönem üretim artışını düşürmektedir (Mian vd., 2017). Rousseau ve Wachtel (2011) ise finansal derinleşmenin tahmin edilen etkisinin yakın örneklemlerde belirgin biçimde zayıfladığı yönünde tamamlayıcı bir gözlem sunmaktadır. Hizmetler yalnızca orta gelirli grupta ve zayıf biçimde tepki vermekte, bu tahminlerin hiçbirinde ilave kredi artışının zararlı hâle geldiği bir eşiğe işaret edilmemektedir.

İkinci ampirik inceleme, kredi tahsisinin politika bakımından olağandışı ölçüde görünür olduğu Türkiye'ye yönelmektedir. Örneklem dönemi hızlı finansal

derinleşmeyi, yinelenen kredi döngülerini, kamu destekli kefalet programlarının genişlemesini, makroekonomik istikrarsızlık epizotlarını, salgına verilen belirgin kredi tepkisini ve selektif kredi ile makroihtiyati araçlara artan güveni kapsamakta; bu da ülkeyi, kredinin gittiği yerin ve ölçeğinin en az hacmi kadar önemli olup olmadığını incelemek bakımından bilgi verici kılmaktadır. Analiz, 2005 yılının birinci çeyreğinden 2026 yılının birinci çeyreğine uzanan ve 902 gözlem üreten, on bir sektörden oluşan çeyreklik bir panel kurmaktadır. Sektörel toplam nakdî krediler Bankacılık Düzenleme ve Denetleme Kurumu'ndan alınmakta (BDDK, 2026) ve mevduat, katılım, kalkınma ile yatırım bankalarının kaydettiği canlı ve takipteki alacakların ikisini birden içermektedir; böylece ölçüt, bankacılık sisteminin her faaliyete tahsis ettiği toplam riski yalnızca canlı bölümüyle değil bütünüyle yakalamaktadır.

Kredi serisi aylık raporlama tarihlerindeki stokları kaydettiğinden her çeyreğin son ayı o çeyreğin değeri olarak alınmakta, bu kural bilanço değişkeninin muhasebe mantığıyla uyuşmaktadır. Sektörel katma değer, Türkiye İstatistik Kurumu'nun üretim yöntemine dayalı ulusal hesaplarından alınmakta (TÜİK, 2026) ve NACE Rev. 2.1 üzerinden raporlanan denetsel kredi kategorilerini ulusal hesaplarda kullanılan NACE Rev. 2 faaliyet gruplarına atayan bir eşleştirme aracılığıyla kredi verisiyle örtüştürülmektedir (Eurostat, 2008, 2025). Çakışan kategorilerden kaçınmaya özen gösterilmiş, bu da imalat dışı sanayi serisinin oluşturulması için imalatın daha geniş sınai toplamdan çıkarılmasını gerektirmiştir. Bu tasarım ilişkinin her iki tarafını da ayrıştırarak her faaliyetin aldığı krediyi o faaliyetin ürettiği katma değerle eşleştirmekte, böylece Türkiye bölümünü ülkeler arası analizden ayırmakta ve geleneksel finans–büyüme modellerini kısıtlayan toplulaştırma sorununun büyük bölümünü ortadan kaldırmaktadır.

Türkiye paneli çok güçlü yatay kesit bağımlılığı sergilemektedir, zira sektörler ortak bir parasal, finansal ve makroekonomik ortamı paylaşmakta ve üretim zincirleri ile ortak talep koşulları üzerinden birbirine bağlanmaktadır; panel Granger nedensellik testleri ise sektörel kredi ile üretim arasında çift yönlü öngörücü geri besleme olduğunu göstermektedir. Bu özellikler statik bir sabit etkiler modelini yetersiz kılmakta ve zaman boyutunun sektör sayısını belirgin biçimde aşması, araç sayısının

şişmesi ve bağımlılığı yaratan ortak faktörlerin modellenmemiş kalması nedeniyle bu uygulamada Sistem GMM'i uygun olmaktan çıkarmaktadır. Analiz bu nedenle, Pesaran'ın (2006) ortak ilişkili etkiler yaklaşımını ve bunun Chudik ile Pesaran (2015) tarafından geliştirilen dinamik heterojen uzantısını izleyerek hata düzeltme biçiminde yatay kesitle genişletilmiş bir gecikmesi dağıtılmış otoregresif model benimsemekte ve sektöre özgü eğimlerin korunması için Ortalama Grup tahmincisini kullanmaktadır. Yatay kesit ortalamalarının iki gecikmesinin dâhil edilmesi doğrusal, kuadratik ve kübik modellerde artık yatay kesit bağımlılığını gidermekte, her birinde negatif ve anlamlı çıkan hata düzeltme terimi ise uzun dönem dengesinden sapmaların kademeli olarak düzeltildiğini doğrulamaktadır.

Çerçeve, kredinin logaritmik kredi terimlerindeki değişimler yoluyla girdiği kısa dönem bileşenini, kredinin düzeyleriyle girdiği uzun dönem bileşeninden ayırmakta, dolayısıyla sonuçlar sırasıyla kredi büyümesine ve kredi derinliğine ilişkin konuşmaktadır. Doğrusal modelde Ortalama Grup kısa dönem katsayısı −0,213 olarak negatif ve anlamlıdır ve sektörel kredideki daha büyük çeyreklik artışların ortalama olarak daha zayıf eşanlı üretim artışıyla ilişkili olduğunu göstermektedir; bu ortalama esas olarak tarım, ticaret, ulaştırma, konaklama ve yiyecek hizmetleri ile kamu ve sosyal hizmetler tarafından sürüklenmektedir. Kredi finansmanlı girdiler ve yatırımlar ölçülebilir katma değer üretmeden önce maliyet doğurabildiğinden, negatif bir eşanlı katsayı yatırım gecikmesiyle de ters talep etkisiyle de tutarlıdır. Uzun dönem Ortalama Grup katsayısı 0,165 ile pozitif fakat anlamsızdır ve bu anlamsızlık sektörel kanıtı özetlemekten çok gizlemektedir, zira madencilik, taş ocakçılığı ve diğer sanayi, tarım, ticaret grubu ve diğer hizmetler pozitif ve anlamlı uzun dönem katsayıları vermekte, en büyüğü 1,320 ile madencilik ve ilgili faaliyetlerde ortaya çıkmaktadır.

Kuadratik model doğrusal sonuçların yorumunu değiştirmekte ve merkezî Türkiye bulgusunu sağlamaktadır. Ortalama Grup düzeyinde logaritmik kredideki değişimin katsayısı 2,042 ile pozitif, farkı alınmış kare terimin katsayısı ise −0,065 ile negatiftir ve ikisi de yüzde beş düzeyinde anlamlıdır; bu, kredideki değişimlerin çeyreklik üretim artışını yaklaşık 7,23 milyar Türk lirası düzeyindeki bir dönüm noktasına kadar desteklediği, bu noktanın ötesinde ise katkısının negatife döndüğü ters U

biçimli bir ilişkiyi tarif etmektedir. Dönüm noktası kredi değişkeninin gözlenen aralığı içinde bulunduğundan bir dışdeğerleme değil ampirik olarak anlamlı bir büyüklüktür ki U biçimli ilişkiler üzerine yöntem literatürü bu koşulu yorum için gerekli saymaktadır (Haans vd., 2016; Lind ve Mehlum, 2010). Eşiğin altında kredi likidite kısıtlarını gevşetebilir ve stokları, ücretleri, girdileri ve üretimin sürekliliğini finanse edebilirken, ötesinde ilişki borç servisi yüklerini, sıkıntı kaynaklı borçlanmayı, zayıf tahsisi ya da bir sektörün ilave fonları çeyrek içinde üretime dönüştürememesini yansıtıyor olabilir.

Sektörel eşiklerin bu ortalama etrafındaki dağılımı, ortalamanın kendisinden daha önemlidir. Bilgi ve iletişim dönüm noktasına yaklaşık 2,46 milyar lirada, tarım 4,94 milyarda, kamu ve sosyal hizmetler 8,68 milyarda, ticaret, ulaştırma, konaklama ve yiyecek hizmetleri grubu ise 40,23 milyarda ulaşmakta; böylece tek bir ulusal ekonomi içinde bir büyüklük mertebesini aşan bir aralık ortaya çıkmaktadır. Bilgi ve iletişimdeki düşük eşik, sektörün varlıklarının maddi olmayan niteliğiyle tutarlıdır, çünkü bankalar maddi teminat karşılığında daha kolay borç vermekte, yazılım, veri, beşerî sermaye ve ağ altyapısı ise güçlükle rehnedilmektedir; nitekim Hsu ve diğerleri (2014) kredi piyasası gelişiminin dışa bağımlı ve teknoloji yoğun sanayilerde yeniliği hisse senedi piyasası gelişiminin desteklediği biçimde desteklemediğini bildirmektedir. Ticaret grubundaki yüksek eşik ise işletme sermayesine, stoklara, lojistiğe, dağıtım ağlarına ve turizm talebine dayanan yapısıyla örtüşmektedir; bununla birlikte bu sektörde de kare teriminin negatif ve anlamlı olması, paneldeki en yüksek soğurma kapasitesinin dahi sınırlı olduğunu göstermektedir.

Uzun dönem kuadratik tahminler doğrusal olmayışı ekonomi geneline değil sektörlere özgü bir nitelik olarak belirlemektedir, zira Ortalama Grup kare terimi anlamsızken birkaç sektör iyi belirlenmiş profiller vermektedir. Finans ve sigorta faaliyetleri, maksimumu 26,07 milyar lira dolayında bulunan bir ters U sergilemektedir; bu sonuç doğrudan “çok fazla finans” tartışmasına dokunmaktadır, çünkü söz konusu sektör finansal sektörün kendisidir ve finansal faaliyetin aşırı genişlemesi kaynakları üretken yatırım yerine aracılığa, bilanço yönetimine ya da spekülasyona yöneltebilmektedir (Cecchetti ve Kharroubi, 2012, 2015). Tarım ile

bilgi ve iletişim ise dönüm noktaları 1,16 ve 2,56 milyar lira dolayında olan U biçimli uzun dönem profilleri sergilemekte, kredi düşük düzeylerde maliyetli ya da etkisiz görünmekte ve ancak modernizasyonu, sabit yatırımı veya teknolojik yenilenmeyi finanse etmeye yetecek ölçeğe ulaştığında üretken hâle gelmektedir.

Kübik model bu tabloyu tersine çevirmekten çok geliştirmektedir. Ortalama Grup düzeyinde üçüncü dereceden terim anlamsızdır, dolayısıyla panel ortalaması kuadratik modelde belirlenen tek eşikli uyarlanmayla en iyi biçimde tanımlanmaya devam etmekte; buna karşılık birkaç sektör tek bir dönüm noktasının temsil edemeyeceği uzun dönem yörüngeleri sergilemektedir. Gayrimenkul faaliyetleri, dönüm noktaları 1,36 ve 13,84 milyar lira dolayında olan belirgin bir N biçimli yol izlemektedir; bu yapılanma, dayanıklı varlıkların aynı anda hem üretim faktörü hem de teminat işlevi görerek kredi limitleri ile varlık fiyatları arasındaki etkileşim yoluyla şokları ilettiği ve büyüttüğü mekanizmayla uyumludur (Kiyotaki ve Moore, 1997). Nihai pozitif bölümün zararsız okunmaması gerekir, zira kredi güdümlü gayrimenkul patlamaları varlık fiyatı enflasyonuyla ilişkilendirilmekte ve varlık piyasalarına yönelen kredinin büyüme etkileri finansal olmayan işletmelere açılan krediye kıyasla daha zayıf bulunmaktadır (Bezemer vd., 2016).

İmalat, ticaret grubu ile kamu ve sosyal hizmetler, kredinin düşük düzeylerde daha zayıf sonuçlarla ilişkilendiği, orta bir aralıkta faaliyeti desteklediği ve çok yüksek düzeylerde yeniden olumsuz ilişkiye döndüğü ters N biçimli yapılanmalar sergilemektedir; imalatta bu örüntü, çok büyük risklerin daha az etkin yapıları ayakta tuttuğu ve uyarlanmayı geciktirdiği olasılığıyla bağdaşmaktadır (Caballero vd., 2008). Üç sektör ise tez boyunca uygulanan yorum disiplinini gerektirmektedir. Tarım, imalat dışı sanayi ve inşaat, N ya da ters N biçimli eğrileri andıran katsayı işaret dizileri vermekte, ancak türevlerinin diskriminantları negatif olduğundan uydurulan fonksiyonlar iki gerçek dönüm noktası yerine birer büküm noktasına sahiptir. Her kübik fonksiyonun tam olarak bir büküm noktası vardır ve negatif diskriminant bu büküm noktasının eğrinin tek yapısal özelliği olduğu anlamına gelir; yani ilişkinin yönündeki bir tersine dönüş değil ivmesindeki bir değişim söz konusudur. İşaret dizisinin doğrudan katsayılardan okunması üç yanıltıcı rejim tarifi üretecekti ve bu durum, fonksiyonel biçimin anlamlılık örüntülerinden çıkarsanmak

yerine türevin kökleri ile diskriminantı üzerinden değerlendirilmesi gerektiğinin tezdeki en açık kanıtıdır.

İki inceleme -uluslararası ve Türkiye analizleri- veri yapısı, bağımlı değişken, finansal gösterge, frekans ve tahminci bakımından ayrıştığından sayısal dönüm noktaları karşılaştırılabilir değildir ve aynı büyüklüğü ölçüyormuş gibi ele alınmamalıdır. Karşılaştırılabilecekleri şey yapıdır ve dört önerme üzerinde uzlaşmaktadırlar. Sektörel ayrıştırma, sonuçları limitte iyileştirmek yerine esaslı biçimde değiştirmektedir. Doğrusal olmayış yaygındır fakat evrensel değildir; sanayide ve birkaç Türkiye sektöründe güçlü biçimde belirirken başka yerlerde zayıf kalmakta ya da hiç görünmemektedir. İlişki ufka ve kalkınma bağlamına bağlıdır. Son olarak ortalama katsayılar her iki ortamda da iktisadi bakımdan sonuç doğuran heterojenliği gizlemektedir; anlamsız Türkiye uzun dönem ortalaması ile toplulaştırılmış ön analizde kaydedilen 0,059 ile 0,122 arasındaki ülke içi açıklayıcı güç aynı olgunun iki ifadesidir.

Tarım, ölçüm farkına uyulduğunda iki ortamda yakınsamaktadır. Uluslararası tüm dönem tahminleri, maksimumu tipik tarımsal kredi artış oranının çok üzerinde bulunan zayıf desteklenmiş iç bükey bir profil tarif etmekte, böylece gözlemlerin çoğu daha hızlı genişlemenin hafif olumlu olduğu bölgede kalmaktadır. Türkiye tahminleri ise negatif bir kısa dönem katsayısını, pozitif ve anlamlı bir uzun dönem katsayısını, görece düşük bir kısa dönem eşiğini ve U biçimli bir uzun dönem profilini bir araya getirmektedir. Bu bulgular birbiriyle çelişmemekte, birlikte tarımsal kredinin bağlayıcı kısıtları gevşetebileceğini, ancak faydalarının gecikmeli, doğrusal olmayan ve fonlamanın ölçeği ile zamanlamasına bağlı olduğunu göstermektedir. Ne negatif bir eşanlı katsayı ne de yüksek kredi düzeylerinde azalan bir marjinal etki, tarımsal finansmanın genel olarak üretken olmadığı biçiminde okunabilir.

Sanayi karşılaştırması, Türkiye kanıtının ülkeler arası sonuçların yorumuna en çok katkı sağladığı yerdir. Uluslararası sanayi değişkeni madenciliği, altyapı hizmetlerini, inşaatı ve imalatı birleştirmekte; bunlar sermaye yoğunluğu, yatırım ufku, teminat yapısı ve işletme sermayesi gereksinimi bakımından ayrışan faaliyetlerdir ve Türkiye

ayrıştırması bu bileşenlerin ne kadar farklı davrandığını göstermektedir. İmalat dışı sanayi güçlü ve pozitif bir uzun dönem katsayısı ile dönüm noktaları yerine büküm noktası taşıyan bir kübik fonksiyon verirken, imalat nihai bölümü negatif olan ters N biçimli bir profil sergilemekte, inşaat ise teminat değerlerine ve kaldıraç çevrimlerine maruziyetiyle tutarlı bir büküm noktası göstermektedir. Sanayideki yüksek dereceli uluslararası ilişkilerin en azından kısmen, biçim kadar işaret bakımından da ayrışan sektöre özgü süreçlerin toplulaştırılmasından doğması bu nedenle olasıdır ve bu, eşleştirmenin bir varsayımı değil ampirik katkısıdır.

Hizmetler aynı noktayı daha da keskin biçimde ortaya koymaktadır. Uluslararası tahminler kriz sonrası on yılda pozitif ve sağlam bir birinci derece etki verirken, Türkiye'nin hizmet dalları farklı yönlerde ve farklı fonksiyonel biçimlerle hareket etmektedir: ticaret grubu negatif bir kısa dönem katsayısını pozitif bir uzun dönem katsayısıyla ve paneldeki en yüksek kısa dönem eşiğiyle birleştirmekte, bilgi ve iletişim en düşük kısa dönem eşiğini U biçimli bir uzun dönem profiliyle bir araya getirmekte, finans ve sigorta faaliyetleri ters U, gayrimenkul N biçimi, kamu ve sosyal hizmetler ters N sergilemekte, mesleki ve idari hizmetler ise anlamlı bir doğrusal ilişki vermemektedir. Dolayısıyla pozitif bir toplulaştırılmış hizmet katsayısı, yapısal olarak birbirine benzemeyen faaliyetler üzerinden alınmış bir ortalamadır ve hizmet dallarının daha hızlı kredi artışından tekdüze biçimde yararlandığı anlamına gelmemelidir.

Bir araya getirilen kanıtlar, "çok fazla finans" ifadesinin bu tezde hangi anlamda kullanıldığını da netleştirmektedir. Sonuçlar, finansın genel olarak zararlı hâle geldiği bir finansal gelişme düzeyi belirlememekte ve zaten bunu belirlemek üzere tasarlanmamıştır, çünkü kredi değişkenleri uluslararası modellere birinci farklarıyla girmektedir. Belirledikleri şey, kredi ile reel faaliyet arasındaki ilişkinin, borç vermenin hızı ya da ölçeği bir üretim yapısının onu belirli bir ufuk içinde soğurma kapasitesini aştığında zayıflayabileceği, kaybolabileceği ya da tersine dönebileceğidir; bu kapasite ise yatırım fırsatına, üretim gecikmelerine, teminat erişilebilirliğine, teknolojiye, talep koşullarına ve kurumsal düzenlemelere göre değişmektedir. Bu yorum Cecchetti ve Kharroubi'nin (2015) önerdiği mekanizmayla, Beck ve diğerlerinin (2012), Mian ve diğerlerinin (2017) ve Müller ile Verner'in

(2024) tahsis kanıtlarıyla ve Ductor ile Grechyna'nın (2015) kredi genişlemesi reel üretim artışını geçtiğinde büyüme etkisinin negatife döndüğü bulgusuyla bağdaşmaktadır. Bununla birlikte mevcut analiz kredi hacmini ölçtüğünden, bunlar modellerin belirlediği kanallar değil tahminlerle tutarlı yorum mekanizmaları olarak kalmaktadır.

Tez altı katkı sunmaktadır. Birincisi kavramsaldır ve finans–büyüme ile yapısal dönüşüm literatürlerini birleştirerek finansal gelişmeyi yalnızca toplam büyümenin bir belirleyicisi olarak değil üretimin sektörel bileşimini yeniden şekillendirebilecek bir güç olarak ele almaktadır. Bu yeniden çerçeveleme soruyu, kredinin genel olarak yararlı olup olmadığından fayda ve maliyetlerinin üretken faaliyetler arasında nasıl dağıldığına kaydırmakta ve ampirik sonuçlar çerçevelemeye içerik kazandırmaktadır, zira aynı on yılda ve aynı örneklemde daha hızlı kredi genişlemesi yükselen bir hizmet payına ve gerileyen bir sanayi payına eşlik etmiştir. İkinci katkı ölçümle ilgilidir ve verinin bütünleşme özelliklerini teknik bir ayrıntı saymak yerine ciddiye almaktan doğmaktadır. Kredi değişkenleri birinci dereceden bütünleşik, sektörel katma değer payları ise durağan olduğundan kredi uluslararası denklemlere birinci farklarıyla girmekte ve tahmin edilen doğrusal olmayışlar kredi derinliği uzayında değil kredi büyümesi uzayında konumlanmaktadır. Kredi stoku ile genişleme hızı arasındaki ayrım kriz ve konjonktür literatüründe yerleşiktir, ancak katma değerin sektörel çözümlemelerine sistematik biçimde taşınmamıştır.

Üçüncü katkı ampiriktir ve genişlik ile derinliğin eşleştirilmesine dayanmaktadır. Tarım, sanayi ve hizmetler, ilişkilerin kriz öncesi, kriz sonrası ve tüm örneklem dönemleri ile gelir grupları arasında farklılaşmasına izin veren ortak bir çerçevede incelenirken, Türkiye analizi on bir kredi–çıktı çiftini çeyreklik frekansta eşleştirerek kararlılık eklemektedir. Dördüncü katkı yöntemseldir: fonksiyonel biçim varsayılmak yerine sırayla seçilmekte, dönüm noktaları katsayı işaretlerinden okunmak yerine marjinal etkilerden elde edilmekte, kuadratik profiller biçimsel eğim testlerine tabi tutulmakta ve yalnızca maksimum verinin desteklediği aralıkta kaldığında yorumlanmakta (Lind ve Mehlum, 2010), kübik profiller ise türevin kökleri ve diskriminantı üzerinden değerlendirilerek bir büküm noktasının iki dönüm noktasıyla karıştırılması önlenmektedir. Beşinci katkı doğrusal olmayışın varlığından çok

geometrisine ilişkindir. Sonuçlar, kredi ve reel faaliyetler arasındaki ilişkinin doğrusallıktan saptığını doğruluyor; ki bu artık bilinen bir sonuçtur, ancak aynı zamanda bu sapmanın, literatürün kanonik hale getirdiği "çok fazla finans" biçimini almadığını da gösteriyor. Toplulaştırılmış tahminler incelenen her dönemde iç bükeydir, ancak maksimumları sıfırda ya da altında kaldığından yükselen kol tümüyle kredi daralması aralığına düşmekte ve bu yapılanma, genişlemenin büyümeyi bir eşiğe kadar desteklediği biçiminde okunamamaktadır. Altıncı katkı politikaya yöneliktir ve toplulaştırılmış kredi hedeflerinin sektörlerin finansmanı benzer oranlarda katma değere dönüştürdüğünü varsaydığını, oysa burada derlenen kanıtın bunun geçerli olmadığını gösterdiğini ortaya koymaktadır.

Politika bakımından temel çıkarım, kredi politikasının yalnızca toplam hacim üzerinden değerlendirilemeyeceğidir. Hacim, özellikle finansman kısıtları bağlayıcı olduğunda önemini korumakta, ancak aynı genişleme alıcı sektöre, mevcut kredi düzeyine ve sonucun değerlendirildiği ufka göre farklı sonuçlar üretebilmektedir. Tekdüze büyüme hedefleri ve farklılaştırılmamış kredi teşvikleri, alıcılar arasında karşılaştırılabilir bir soğurma kapasitesi bulunduğunu örtük biçimde varsaymakta; Türkiye eşiklerinin bir büyüklük mertebesinden fazla dağılması ise bu varsayımın ne kadar güçlü olduğunu göstermektedir. Düşük gelirli ve tarıma dayalı ekonomilerde biçimsel finansal hizmetlerin yaygınlaşması değerini korumaktadır, çünkü bu gruptaki gözlemlerin çoğu daha hızlı tarımsal kredi artışının daha yüksek tarımsal katma değere eşlik ettiği bölümde yer almaktadır. Yine de kredi politikasının sulama, ulaştırma, depolama, yayım hizmetleri, sigorta, teknoloji ve ürün pazarlarına erişimle eşgüdümlenmesi gerekir, zira tahmin edilen iç bükeylik borç vermenin eksik tamamlayıcı koşulların yerini süresiz olarak alamayacağını göstermektedir (Dünya Bankası, 2007).

Orta gelirli ekonomilerde kanıtlar, tek bir ekonomi geneli optimumun aranmasından çok aşamaya ve sektöre göre farklılaştırmayı desteklemektedir. Kredi, sınai yenilenmeyi ve modern hizmetlerin büyümesini finanse ederken aynı anda inşaatı, hanehalkı borçlanmasını ya da düşük verimli ticarete konu olmayan faaliyetleri de besleyebilmektedir. Orta gelirli sanayi tahminleri marjinal ilişkinin negatife döndüğü noktayı bu gruptaki medyan kredi artış oranına yakın konumlandırdığından, kredinin

bileşiminin toplam artış hızıyla birlikte izlenmesi gerekçesi bu ekonomilerde diğerlerine kıyasla daha güçlüdür. Yüksek gelirli ekonomilerde kredi artışı ile sınai katma değer arasındaki zayıf bağımsız ilişki, ileri aşamalarda banka kredisi genişlemesinin tek başına yetersiz bir araç olduğunu, yenilik, maddi olmayan yatırım, uzmanlaşmış finansman, hisse senedi piyasaları ve kurumsal etkinliğin daha ağır bastığını düşündürmektedir. Türkiye sonuçları ise selektif kredi programlarına, kefalet mekanizmalarına, kamu bankası kredilerine ve makroihtiyati tasarıma doğrudan dokunmakta; sektöre özgü eşikler düzenleyici sınırlar değil tanı araçları olarak görülmelidir, zira nominal tarihsel veriden tahmin edilmekte, modele ve örnekleme bağlı kalmakta ve canlı krediyi takipteki alacaktan, yeni kullandırımı devralınan stoktan ya da üretken yatırımı yeniden finansmandan ayırt edememektedir.

Çalışmanın kapsamını birkaç sınırlılık belirlemektedir. Uluslararası bağımlı değişkenler GSYİH içindeki sektörel katma değer paylarıdır; bu seçim yapısal dönüşüm çerçevesinden doğmakta ve geniş bir ülke kümesi için karşılaştırılabilir ölçütler sunmakta, ancak paylar mutlak üretimin ya da verimliliğin doğrudan göstergeleri değildir. Uluslararası düzeyde sektöre özgü kredi verisi yalnızca tarım için mevcut olduğundan sanayi ve hizmet modelleri mevduat bankalarınca sağlanan özel krediyi kullanmakta, dolayısıyla katsayıları ilgili sektöre açılan kredinin etkisini değil genel bankacılık derinliğinin sektörel bileşimle ilişkisini tarif etmektedir. Tahminciler başlıca ekonometrik güçlükleri ele almakla birlikte tam nedensel belirleme sağlamamakta; Türkiye bölümünde nedensellik testleri yapısal nedensellikten çok öngörücü önceliği kurmakta ve hata düzeltme tahminleri dışsal kredi arzı şoklarını değil koşullu kısa ve uzun dönem ilişkilerini tarif etmektedir.

Polinom modeller yapısal fonksiyonel biçimler değil esnek yaklaşımlardır ve yüksek dereceli terimler örneklem aralığından, aykırı değerlerden ve çoklu doğrusallıktan etkilenebilir; bu nedenle hiçbir tahmin edilen eşik sabit bir parametre olarak ele alınmamalıdır. Türkiye serileri cari fiyatlarla ve mevsimsellikten arındırılmadan ölçülmüştür; arındırılmamış üretim serilerinin bilinçli olarak kullanılmasının nedeni, sektörel kredi verisinin yalnızca arındırılmamış biçimde yayımlanması ve ilişkinin iki tarafına tutarsız mevsimsel işlem uygulanmasından kaçınılmasıdır. Bunun bedeli,

mevsimsel değişimin tahminlerin içinde kalması ve Türk lirası cinsinden ifade edilen dönüm noktalarının örneklem dönemine ait fiyat düzeyini ve ölçeği yansıtmasıdır. Toplam nakdî kredinin canlı kredileri takipteki alacaklarla birleştirmesi de etkin borç vermeyi bozulan riskin birikiminden ayırmayı engellemektedir. Son olarak çalışma tasarım gereği finansal gelişmenin niceliksel boyutuna odaklanmakta; faiz oranlarını, vadeyi, para cinsini, teminatı, borçlu ölçeğini, kredi amacını, banka mülkiyetini ve kefalet durumunu ölçmemektedir.

Bu sınırlılıklar aynı zamanda ileri araştırma gündemini tanımlamaktadır. Uluslararası analizin en değerli uzantısı, imalatı, inşaatı ve başlıca hizmet dallarını kapsayan uluslararası düzeyde karşılaştırılabilir sektörel kredi verisinin oluşturulmasıdır; böyle bir veri, sanayi ve hizmetlerde belgelenen doğrusal olmayan örüntülerin bu sektörlere açılan krediden mi yoksa finansal derinlikteki daha geniş kaymalardan mı doğduğunu belirlemeye imkân verecektir. Sektörel kredinin gerçek katma değer, istihdam, işgücü verimliliği, yatırım ve ihracat performansı gibi daha geniş bir çıktı kümesiyle birlikte modellenmesi, finansal etkilerin yeniden tahsis ile üretim düzeyi ve etkinlik üzerindeki etkilerini ayırmayı sağlayacaktır. Türkiye için en acil öncelik nominal ve mevsimsel değişimin ele alınmasıdır; ham kredi ve üretim serilerine aynı mevsimsel düzeltmenin uygulanması ve uygun fiyat endeksleri bulunduğunda her ikisinin reel değerlere dönüştürülmesi, dönüm noktalarının sabit fiyatlarla ifade edilmesine olanak tanıyacaktır. Eşleştirilmiş banka–firma ya da kredi düzeyinde idari veriler ise en kapsamlı ilerlemeyi temsil edecek, hangi borçluların ilave finansmanı üretken kapasiteye dönüştürdüğünü belirlemeye ve kredi talebindeki hareketleri kredi arzındaki kaymalardan ayırmaya yardımcı olacaktır.

Tez, finansal gelişmenin kredi aracılığıyla ölçüldüğünde ekonomik performans üzerinde istikrarlı ve olumlu bir etki uyguladığı yönündeki tanıdık ve sonuçta yeterince ayırt edici olmayan bir önermeden yola çıkmıştır. İki ampirik incelemede derlenen kanıtlar bu önermenin reddedilmesini değil nitelendirilmesini gerektirdiğini göstermektedir. Kredi önemlidir ve birçok ortamda kayda değer ölçüde önemlidir, ancak reel faaliyetle ilişkisi ne tekdüzedir ne de her yerde aynıdır; tarım, sanayi ve hizmetler arasında, düşük, orta ve yüksek gelirli ekonomiler arasında, kriz öncesi ve kriz sonrası rejimler arasında, kısa ve uzun ufuklar arasında ve tek bir ulusal

ekonomi içinde on bir ayrı üretim dalı arasında değişmektedir. Soğurma kapasitesi bu nedenle ulusal değil sektörel bir özelliktir; üretim gecikmeleri, varlık yapısı, işletme sermayesi gereksinimleri, teminat kalitesi, düzenleme, talep koşulları ve hâlihazırdaki kredi düzeyi tarafından biçimlendirilmektedir.

Bu kanıtların desteklediği sonuç kategorik değil koşulludur. Kredinin hacmi ve hızı, finansman kısıtlarının bağlayıcı olduğu ve borçlanılan kaynakların üretken kapasite içinde kullanıldığı yerlerde sektörel faaliyeti destekleyebilir; buna karşılık genişleme bir sektörün soğurma kapasitesini aştığında, borçlanma kaldıraç ya da sıkıntıyı yansıttığında veya tamamlayıcı kurumsal koşullar eksik olduğunda marjinal katkı azalabilir, kaybolabilir ya da tersine dönebilir. Söz konusu mekanizmalar doğrudan ölçülmek yerine tahminlerin örüntüsünden çıkarsandığından, tez boyunca raporlanan dönüm noktaları borç verme sınırları olarak değil daha yakın incelemenin başlangıç noktaları olarak ele alınmalıdır. Etkin finansal politika bu nedenle hacmin; gidilen yer, zamanlama, sektörel yapı ve makroekonomik bağlamla birlikte değerlendirilmesini gerektirmektedir. İlişkiyi sektöre, gelir düzeyine, döneme ve ufka göre ayrı ayrı tahmin ederek ve her ilişkinin biçiminin varsayılmak yerine kurulmasında ısrar ederek bu tez, daha fazla finansın yararlı mı yoksa zararlı mı olduğu sorusunun ötesine geçmek için bir çerçeve sunmaktadır. Daha bilgi verici soru, bir ekonominin çok fazla finansa sahip olup olmadığı değil, “kimin için çok fazla finans” sorusudur; yani hangi sektörler için, hangi kalkınma aşamasında ve hangi ufukta ilave kredinin kalıcı değer yaratımına dönüşmekten çıktığıdır.

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