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FINANCIAL DEVELOPMENT AND INDUSTRIALISATION IN SUB-SAHARAN AFRICA: AN EMPIRICAL PANEL ASSESSMENT, 2000–2023

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Abstract

Since 2000, Sub-Saharan Africa (SSA) has grown at a steady pace, but this expansion has done little to change the structure of production: manufacturing value added (MVA) has stayed close to 10 per cent of GDP. This paper asks whether, and to what extent, financial development has influenced industrialisation in the region over 2000–2023. We use the IMF multidimensional Financial Development (FD) Index (Svirydzenka, 2016) for a panel of 21 countries and estimate a static model by two-stage least squares (2SLS) and a dynamic model by two-step system generalised method of moments (GMM-SYS). In both cases the FD coefficient is positive and significant at the 1 per cent level (0.379 under 2SLS and 0.250 under GMM-SYS). Because MVA is expressed in per cent of GDP and the index lies between 0 and 1, these estimates imply that a 0.1-point rise in the index, roughly one sample standard deviation, is associated with an increase in MVA of about 0.038 percentage points of GDP under 2SLS and 0.025 points under GMM-SYS, other things equal. The association is therefore statistically robust but modest in size. Human capital and domestic investment are positively associated with MVA. The coefficients on natural-resource rents and trade openness change sign between the static and the dynamic specification, and no statistically significant quadratic relationship between financial development and MVA is detected.

Keywords: *financial development; industrialisation; manufacturing value added; Sub-Saharan Africa; system GMM.*

1. Introduction

Since the early 2000s, Sub-Saharan Africa has combined a return to growth with very limited structural change. Real GDP expanded by 3.5 to 5.0 per cent a year on average between 2000 and 2019, before the COVID-19 pandemic and the war in Ukraine interrupted that trend (African Development Bank, 2020). Over the same period, the share of manufacturing in regional GDP did not rise and remained close to 10 per cent, about half the level observed in South-East Asia at a similar stage of development (UNIDO, 2022). Rodrik (2016) describes this pattern of growth without industrial transformation as premature deindustrialisation, that is, a decline of manufacturing that begins before the sector has been able to reshape the economy.

The role of finance lies at the heart of this puzzle. In the Schumpeterian tradition revived by King and Levine (1993), financial intermediaries support industrialisation because they screen investment projects and spread risk. McKinnon (1973) and Shaw (1973) add that a gradual liberalisation of the financial sector raises savings and improves the allocation of capital, which speeds up productive accumulation. Financial systems in SSA nevertheless remain shallow. Credit to the private sector averages less than 30 per cent of GDP, compared with nearly 100 per cent in emerging Asia (Sahay et al., 2015; Ibrahim & Alagidede, 2018). In addition, a large part of available credit goes to short-term trade finance, household consumption and real estate, so that the long-term funding required by manufacturing firms remains scarce.

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This paper therefore examines whether, and through which channels, financial development has affected industrialisation in SSA between 2000 and 2023. The question is worth revisiting for three reasons. First, most earlier studies (Beck et al., 2000; Demetriades & Hussein, 1996; Ibrahim & Alagidede, 2018) rely on single indicators such as M2/GDP or private credit/GDP, which capture depth but ignore access and efficiency. We use instead the IMF composite index (Svirydzenka, 2016), which combines depth, access and efficiency for both financial institutions and financial markets. Second, the 2000–2023 period includes a series of large shocks, among them the global financial crisis, the commodity price collapse of 2014–2016, the pandemic and the war in Ukraine, which makes it a demanding setting for testing the finance–industry relationship. Third, we combine a static instrumental-variable estimator (2SLS) with the dynamic system estimator of Blundell and Bond, so that both the endogeneity of financial development and the persistence of manufacturing output are taken into account.

The existing evidence is mixed, and the sources of disagreement are informative. Kinda and Sawadogo (2023), using data on 35 SSA countries for 1996–2018, find that finance hinders industrialisation when financial depth is low and supports it only once a certain level has been reached. Lo and Cissokho (2023), in a sample of more than one hundred countries, emphasise institutions: the link between finance and manufacturing is strong where institutions function well and weak elsewhere, which is the situation of most SSA countries. For 25 countries, Asongu and Odhiambo (2022) show that broader access to credit increases value added in services but has no effect on agriculture or manufacturing. Nwani et al. (2023), on the other hand, report that deeper financial institutions raise manufacturing value added and partly offset the adverse effect of resource dependence, while Asngar (2025) finds a positive overall effect alongside a negative contribution of monetary expansion. Shikur (2024) observes that financial development benefits African manufacturing in the short run but harms it in the long run. These differences largely reflect three methodological decisions that every study in this area has to make: the measure of financial development, the measure of industrialisation, and the treatment of endogeneity and non-linearity.

The measurement of industrialisation also deserves attention, since views on African manufacturing have changed in recent years. The premature deindustrialisation argument associated with Rodrik (2016) has been questioned by Nguimkeu and Zeufack (2024). Using data on 45 SSA countries from 1960 to 2016, they find no robust sign of deindustrialisation across the region: the share of manufacturing in output is stagnant, but its share in employment rises slowly, and only Southern Africa shows a modest decline that may reflect a more mature economic structure. Drawing on the GGDC/UNU-WIDER Economic Transformation Database, Kruse et al. (2023) report a revival of manufacturing in low-income economies of Asia and SSA, much of it driven by small and often informal firms. Firm-level studies add further detail. Diao et al. (2025), using panel data from Tanzania and Ethiopia, show that large firms are productive but create few jobs, whereas small firms absorb labour with little productivity growth. Manufacturing value added, the dependent variable in this paper, therefore captures an important part of industrialisation but not all of it, and our coefficients should be interpreted with this limitation in mind.

The paper makes three contributions. It uses a multidimensional measure of financial development for 21 SSA economies; it compares static and dynamic estimates of the finance–manufacturing relationship and, within the dynamic model, derives the implied long-run effect from the coefficient on the lagged dependent variable; and it reports Westerlund (2007) panel cointegration tests as a complementary check, given that the series do not all share the same order of integration.

The rest of the paper is organised as follows. Section 2 presents the theoretical background and reviews the empirical literature. Section 3 describes the model, the data, the estimation strategy and the preliminary cointegration and specification tests. Section 4 presents and discusses the results, compares them with previous studies and examines possible non-linearity. Section 5 concludes with policy implications and directions for future research.

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2. Theoretical framework and empirical literature

2.1. Successive reformulations of the notion of financial development

The concept of financial development has been redefined several times, and each redefinition has added a new dimension: material, functional and institutional. The material approach, which goes back to Gurley and Shaw (1960) and Goldsmith (1969), measures financial development by the size of financial assets and liabilities relative to GDP. Shaw (1973) and McKinnon (1973) build on this view and argue that freeing interest rates and ending financial repression raise savings and improve the allocation of capital. The functional approach developed by Levine (1997, 2005) looks at the services provided by the financial system rather than at its size: producing information, pooling and mobilising savings, directing capital to the most productive uses, managing risk, easing transactions and monitoring managers. This perspective supports a Schumpeterian interpretation in which the financial system, by selecting innovative entrepreneurs, accelerates creative destruction and structural change (Aghion & Howitt, 1992, 1998; King & Levine, 1993; Rajan & Zingales, 1998). The integrated approach is the most recent of the three. It is reflected in the IMF index (Svirydzenka, 2016), which combines six sub-indices measuring the depth, access and efficiency of financial institutions and financial markets. We adopt this index in the empirical analysis because it reduces the measurement bias associated with single proxies.

A composite measure is particularly relevant in Africa, where the main weakness of the financial system lies in its structure more than in its overall size. Allen et al. (2014) show that African banks serve governments and large firms reasonably well, while small and medium-sized enterprises depend mostly on retained earnings and informal finance. This inclusion gap is persistent and affects precisely the firms that an expansion of manufacturing would need to reach. Studies that break financial development into its components point in the same direction. For 30 African countries over 2000–2020, Girma and Huseynov (2025) separate depth, efficiency, liquidity and stability, and find that the first two support growth whereas the last two are negatively associated with it; a single proxy would conceal these opposing effects. In a sample of 92 developing countries, Ndoya et al. (2024) find that the impact of financial development on economic complexity depends on country stability and is largest where economic, political and financial risks are low. These results suggest that composite indicators such as the one proposed by Svirydzenka (2016) are needed to study the question properly.

2.2. Industrialisation, structural transformation and productive complexity

Industrialisation is part of the broader process of structural transformation analysed by Lewis (1954), Rostow (1960), Gerschenkron (1962) and Kuznets (1975). More recently, Hidalgo and Hausmann (2009) introduced the concept of economic complexity, according to which industrialisation involves diversification towards more sophisticated products and not only a larger manufacturing share. McMillan and Rodrik (2011) show that structural change in SSA has reduced aggregate productivity, because workers leaving agriculture have moved mainly into low-productivity informal services rather than into manufacturing. This feature of the regional experience explains the focus of the present study, which examines whether finance is associated with the expansion of manufacturing specifically, and not with growth in general.

2.3. Theoretical channels of transmission from finance to industry

The literature identifies four main channels through which finance may affect industrialisation. According to the capital accumulation channel (Greenwood & Jovanovic, 1990; Beck et al., 2000; Levine et al., 2000), deeper financial systems mobilise more savings and lower the cost of capital, which relaxes the financing constraints faced by manufacturing firms in SSA (Bigsten & Söderbom, 2006). The Schumpeterian channel (Schumpeter, 1911; Aghion & Howitt, 1992) concerns the quality of allocation: efficient intermediaries direct funds to innovative projects and select among projects in ways that raise aggregate productivity. The liberalisation channel (McKinnon, 1973; Shaw, 1973) holds that removing interest rate controls and increasing competition among banks improves allocative efficiency. Finally, the intersectoral transfer channel (Lewis, 1954) describes the shift of underused resources from the traditional sector to the modern, capital-intensive sector. These channels can operate together and may strengthen one another where institutions allow (Demetriades & Law, 2006). The reduced-form model estimated in this paper does not test any of these channels separately. They serve to interpret the results, and the estimated coefficients cannot reveal which of them dominates.

A fifth channel, discussed in the macro-financial literature, cannot be tested with our specification but is relevant to the interpretation. Studying 36 developed and developing economies over 1980–2017, Botta et al. (2023) find that periods of large non-FDI capital inflows are followed by a fall in the manufacturing share of output and employment and by a lower economic complexity index, with stronger effects in emerging and developing countries. In this view, finance affects industry through relative prices, the real exchange rate and the composition of demand, in addition to credit allocation, and it can shift resources towards non-tradable activities at the expense of manufacturing. Soumtang Bime et al. (2024) raise a related question for SSA and examine whether financial openness influences economic transformation and productive sophistication in the region. Both studies suggest that the use made of financial resources matters at least as much as their volume.

2.4. Non-linearities, critical thresholds and the "too much finance" hypothesis

The idea of a linear relationship between finance and growth has been challenged, and the challenge itself is now debated. Aghion et al. (2005) argue that countries below a certain level of financial development cannot converge towards the technological frontier. Arcand et al. (2015) go further and report an inverted U-shaped relationship in which financial expansion beyond a private credit ratio of about 100 per cent of GDP starts to reduce growth. This result has since been questioned. Luintel et al. (2024) re-examine fourteen indicators of financial development in roughly seven thousand cross-section and panel regressions and are unable to find a robust inverted U; they conclude that the threshold is more likely a statistical artefact than a structural feature of the data. Demetriades et al. (2024) show that, for 2000–2014, private credit has no significant, or even a negative, association with growth once financial fragility is measured directly instead of being inferred from crises. Evidence for SSA points in a different direction. Using 36 countries over 1980–2015, Machado et al. (2021) estimate a hump-shaped relationship with a turning point at only 6.9 per cent of private credit to GDP, which would place most observations in the region beyond the peak; they attribute this to the diversion of scarce credit towards consumption in weak institutional environments. The quadratic specification reported in Section 4.3 is intended to add evidence on this unresolved question.

2.5. Empirical evidence from Sub-Saharan Africa

Studies on SSA remain dispersed, and their results tend to differ according to the method used more than according to the countries covered. Ibrahim and Alagidede (2018) find that finance has a positive effect on growth in 29 African countries, but that this effect depends on institutional quality. Amoah and Jehu-Appiah (2022) identify a threshold above which financial development supports industrialisation. Avom et al. (2021) highlight the contribution of regional value chains to the industrialisation of SSA, and Asongu and Odhiambo (2019, 2020) underline the complementary role of digital financial inclusion and insurance.

Four groups of studies are directly relevant here. The first deals with non-linearity and its direction. Kinda and Sawadogo (2023) find, for 35 SSA countries, a threshold below which finance harms industrialisation and above which it helps. Unlike the pattern in Arcand et al. (2015), this relationship is U-shaped, and the

binding constraint lies at low rather than high levels of financial development. Abdulai et al. (2024) apply a dynamic threshold model to 41 countries over 2000–2022 and reach a similar conclusion for private investment: finance reduces investment when private credit is below 6.52 per cent of GDP, or when the composite index is below 23.18 per cent, and stimulates it above these levels. Their results point to a problem of insufficient finance in SSA, the reverse of the concern raised in the debate on advanced economies.

The second group concerns institutions. Lo and Cissokho (2023) show that institutional quality determines whether financial development reaches the manufacturing sector, and that weak institutions explain a large part of the gap between SSA and other developing regions. Shikur (2024), studying African manufacturing over 1996–2021, finds that financial development and economic freedom raise manufacturing output in the short run but reduce it in the long run, and links this reversal to the low competitiveness of African producers rather than to finance as such.

The third group takes a sectoral view, and its results are less encouraging. Asongu and Odhiambo (2022) find that financial access increases value added in services but not in manufacturing or agriculture. For 54 African countries over 1990–2022, Asngar (2025) reports a positive overall effect of financial development on structural transformation together with a negative effect of monetary expansion on manufacturing value added. Nwani et al. (2023) obtain a different result for the depth of financial institutions, which raises manufacturing value added and reduces the crowding-out effect of resource dependence; they also confirm that resource dependence lowers the shares of manufacturing and services in output. Talnan (2025), in a study of 30 African countries, reports that financial development lowers the share of manufactured goods in exports. Taken together, these studies suggest that credit in SSA may be directed mainly towards services, trade and consumption rather than towards manufacturing capacity, which is in line with the pattern of financial inclusion described by Allen et al. (2014).

The fourth group looks at outcomes other than value added. Ndoya et al. (2024) find that financial development increases economic complexity only where the economic, political and financial environment is stable. Soumtang Bime et al. (2024) examine whether financial openness affects economic transformation in SSA, and thus whether the composition of capital flows matters in addition to their size. Botta et al. (2023) relate surges in capital inflows to premature deindustrialisation in developing countries more generally.

Relative to this literature, the present study combines several elements that are seldom found together. It uses the multidimensional IMF index instead of the single proxies that remain common in work on SSA; it covers 2000–2023 and therefore includes the shocks after 2020 that earlier samples do not contain; it estimates static and dynamic models on the same data and, in the dynamic case, reports the implied long-run effect; and it tests explicitly for a quadratic relationship, in relation to the threshold literature from Arcand et al. (2015) to Machado et al. (2021) and Kinda and Sawadogo (2023). Section 3 describes how these elements are implemented.

3. Analytical model, data and estimation strategy

3.1. Specification of the structural model

Following Beck and Levine (2004) and Ibrahim and Alagidede (2018), with adjustments to the African context, the baseline model for country i in year t is written as:

$$MVA_{it} = \alpha_0 + \beta_1 FD_{it} + \beta_2 \ln GDPpc_{it} + \beta_3 OPEN_{it} + \beta_4 NRR_{it} + \beta_5 HC_{it} + \beta_6 INV_{it} + \beta_7 CORR_{it} + \beta_8 \ln POP_{it} + \mu_i + \lambda_t + \varepsilon_{it} \quad (1)$$

where μ_i denotes country fixed effects, λ_t common time effects and ε_{it} the idiosyncratic error. To examine the non-linearity hypothesis of Arcand et al. (2015), a second version adds the squared term FD^2_{it}. The dynamic model includes the lagged dependent variable, $MVA_{i,t-1}$, to account for the persistence of manufacturing output (equation 2):

$$MVA_{it} = \alpha_0 + \rho MVA_{i,t-1} + \beta_1 FD_{it} + \ldots + \beta_8 \ln POP_{it} + \mu_i + \lambda_t + \varepsilon_{it} \quad (2)$$

Table 1. Definition and sources of the variables in the industrialisation model

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Variable	Definition	Source
MVA	Manufacturing value added (% of GDP)	WDI, World Bank (2024)
FD	Financial Development Index (0–1)	IMF, Svirydzenka (2016, updated 2024)
lnGDPpc	Log of GDP per capita (constant 2015 USD)	WDI, World Bank (2024)
OPEN	Trade openness ((X + M)/GDP)	WDI, World Bank (2024)
NRR	Natural-resource rents (% of GDP)	WDI, World Bank (2024)
HC	Human capital (public spending on education, % of GDP)	WDI / UNESCO (2024)
INV	Domestic investment (GFCF, % of GDP)	WDI, World Bank (2024)
CORR	Control of corruption (WGI index, −2.5 to +2.5)	WGI – World Bank (2024)
lnPOP	Log of total population	WDI, World Bank (2024)

Source: Authors, from the databases used (WDI, WGI, IMF, UNESCO).

3.2. Sample, period and the nature of the panel

The sample consists of 21 SSA countries chosen because continuous data are available for 2000–2023: Angola, Benin, Botswana, Burkina Faso, Burundi, Cameroon, Côte d'Ivoire, Ethiopia, Gabon, Ghana, Kenya, Madagascar, Mali, Mauritius, Mozambique, Nigeria, Rwanda, Senegal, Tanzania, Uganda and Zambia. The balanced panel contains 504 country-year observations (21 × 24). Gaps in the original series concern mainly public spending on education (HC), for which the UNESCO data are missing in some years, and control of corruption (CORR), since the Worldwide Governance Indicators were not published for 2001. These gaps were filled by linear interpolation within each country series. Values missing at the ends of a series, namely the first years of manufacturing value added for Angola and Madagascar and natural-resource rents for 2022 and 2023, which the WDI do not yet report, were obtained by linear extrapolation. Because the first and second lags of the FD index serve as instruments and the FD series in our data set starts in 2000, the 2SLS estimates cover 2002–2023, which gives 462 observations (21 × 22). In the dynamic model, the lagged dependent variable is not extended before 2000, so the first year serves as the initial condition and the GMM-SYS sample covers 2001–2023, which gives 483 observations (21 × 23).

With 21 of the 48 countries of SSA, the sample covers fewer than half of the countries of the region (about 44 per cent), although its share of the regional population is considerably larger. It includes several of the largest economies, such as Nigeria, Ethiopia, Kenya, Angola, Ghana and Côte d'Ivoire, but it does not include South Africa, the most industrially diversified economy of SSA, nor the Democratic Republic of the Congo. The sample also reflects the institutional diversity of the region: members of the CFA franc zone and countries outside it, a small island economy (Mauritius) and large continental ones, oil exporters and agricultural exporters. Some gaps should be stated from the outset because they limit the scope of our conclusions. The mineral-rich economies of Central Africa and the fragile states of the Sahel and the Horn of Africa are only partly represented, and the countries excluded for lack of continuous data tend to be those with the weakest statistical systems and, in practice, the least developed financial infrastructure. The results should therefore be read as applying to a group of SSA economies with relatively complete data, and not to the region as a whole.

3.3. The dual estimation strategy

Two complementary estimators are used. In the baseline, two-stage least squares (2SLS) deals with the contemporaneous endogeneity of the FD index, which can arise from reverse causality, since industrial growth raises the demand for financial services, or from omitted institutional factors. The model is estimated with xtivreg2 using the within (fixed-effects) transformation, which removes the country effects, and it includes year dummies to capture shocks common to all countries. The excluded instruments actually retained are the first and second lags of the FD index (t−1 and t−2). Legal origin (La Porta et al., 1998), which was initially considered because it shapes financial structure (Demirgüç-Kunt & Levine, 2008), does not vary over time within a country; it is therefore perfectly collinear with the country fixed effects and is not part of the instrument set. With two excluded instruments and one endogenous regressor, the first-stage F-statistic has two numerator degrees of freedom and the Hansen J statistic has one degree of freedom, as shown in Tables 3 and 8. For the dynamic model, the within estimator is biased when a lagged dependent variable is included and the time dimension is short; Nickell (1981) shows that this bias is of order 1/T, which is not negligible with 24 years. The difference GMM estimator of Arellano and Bond (1991) eliminates the country effects by first-differencing and uses earlier levels as instruments for the differenced lag. The system estimator of Arellano and Bover (1995) and Blundell and Bond (1998) adds the equation in levels, which improves efficiency when the series is highly persistent, as manufacturing value added is. We use the two-step system GMM estimator (Blundell & Bond, 1998; Roodman, 2009a) as a robustness check. In line with Roodman (2009b), the instrument matrix is collapsed and the lags are limited to laglimits(2 4), so that the number of instruments (17) remains below the number of countries (21). Standard errors include the Windmeijer (2005) finite-sample correction.

The composition of the instrument set calls for some justification. In the 2SLS model, identification rests entirely on the lagged values of the FD index. These lags are correlated with the current index because financial structures change slowly, and they are assumed to be uncorrelated with current shocks to manufacturing value added. This exclusion restriction cannot be tested directly, although the Hansen test reported below provides some indication. If it holds only approximately, the 2SLS estimate will be biased in one direction, while the GMM-SYS estimate, which relies on internal instruments and on assumptions about initial conditions, may be biased in another. The fact that the two estimates have the same sign and a similar order of magnitude therefore gives more reassurance than either of them taken alone, although it does not remove the need for caution. A similar reasoning underlies recent work on SSA; Abdulai et al. (2024), for instance, combine system GMM with dynamic threshold estimation instead of relying on a single estimator.

3.4. Preliminary tests: stationarity, cointegration and specification

Seven tests precede the estimation: the Levin et al. (2002) panel unit-root test, the Westerlund (2007) panel cointegration test, the Pesaran (2004) cross-sectional dependence test, the Hausman (1978) specification test, the Durbin–Wu–Hausman endogeneity test, the Hansen (1982) J test of over-identifying restrictions and the Arellano–Bond test for serial correlation in the first-differenced residuals. Since the LLC test shows that some variables are I(0) and others I(1), a panel cointegration test is also reported. Tables 2 and 3 present the results.

Table 2. Westerlund (2007) panel cointegration tests

Statistic	Value	p-value	Decision (H_0: no cointegration)
Gt	−2.847	0.012	Rejected at 5%
Ga	−7.124	0.008	Rejected at 1%
Pt	−9.521	0.021	Rejected at 5%
Pa	−6.302	0.015	Rejected at 5%

Source: Authors' calculations under Stata 16.

All four Westerlund (2007) statistics, the two group-mean statistics (Gt, Ga) and the two panel statistics (Pt, Pa), reject the null hypothesis of no cointegration at the 5 per cent level or better. The test is applied to manufacturing value added, the FD index and the two variables found to be I(1) by the LLC test,

domestic investment (INV) and trade openness (OPEN). The p-values reported in Table 2 are the asymptotic ones. Since the Pesaran CD test points to cross-sectional dependence, which these p-values do not take into account, the bootstrap p-values proposed by Westerlund (2007) would be more reliable, and the results should be read with this reservation in mind. Subject to these assumptions, the results are consistent with a long-run relationship between manufacturing value added, the FD index, investment and openness, and they reduce the concern that the level estimates reflect a spurious regression. They do not, by themselves, establish that the level specification is valid. It should also be noted that the dependent variable is I(0) according to the LLC test and that only INV and OPEN are I(1), so the cointegration results are best viewed as a supplementary check on the long-run relationship rather than as the basis of the specification.

Table 3. Summary of specification and instrument-validity tests

Test	Statistic	p-value	Conclusion
Hausman (fixed vs. random effects)	$\chi^2(9) = 18.79$	0.027	Fixed effects retained
Pesaran CD (cross-sectional dependence)		0.034	Moderate dependence
First-stage 2SLS (instrument strength)	$F(2,411) = 18.50$		Above Stock–Yogo 15% critical value (11.59)
DWH (endogeneity of FD)	$\chi^2 = 6.71$	0.010	Exogeneity of FD rejected; 2SLS preferred
Hansen J (2SLS, over-identification)	$\chi^2(1) = 0.654$	0.419	Not rejected
Arellano–Bond AR(1)		0.003	First-order correlation present (expected)
Arellano–Bond AR(2)		0.936	Not rejected (no second-order correlation)
Hansen J (GMM-SYS, over-identification)		0.811	Not rejected
BIC, final model	5.1866		Benchmark for comparison

Source: Authors' calculations under Stata 16.

Table 3 summarises the specification and instrument tests. The Hausman test ($\chi^2(9) = 18.79$, $p = 0.027$) rejects the null hypothesis that the random-effects estimator is consistent, which indicates that unobserved country characteristics are correlated with the regressors; fixed effects are therefore retained. The Pesaran CD test ($p = 0.034$) points to moderate cross-sectional dependence, which is expected given the common external shocks that affect SSA economies, such as commodity price movements, global financial crises and the pandemic. Driscoll–Kraay standard errors, which are robust to this type of dependence, are not reported in the present tables. The standard errors in Tables 8 and 9 are therefore robust to heteroskedasticity and to serial correlation within countries, but not explicitly to correlation across countries, and the significance levels should be read with this in mind. The first-stage F-statistic (18.50) exceeds the Stock and Yogo (2005) critical value of 11.59 for a maximal IV size of 15 per cent with one endogenous regressor and two excluded instruments, and lies just below the 10 per cent value of 19.93, so weak instruments do not appear to be a serious concern. The Durbin–Wu–Hausman test ($p = 0.010$) rejects the exogeneity of the FD index, which supports the use of 2SLS. The Hansen J test for the 2SLS model ($p = 0.419$) does not reject the over-identifying restriction. For GMM-SYS, the Arellano–Bond tests show first-order serial correlation in the differenced residuals ($p = 0.003$), as expected, but no second-order correlation ($p = 0.936$), which supports the use of lags dated $t-2$ and earlier as instruments. The Hansen J test for the GMM-SYS model ($p = 0.811$) also does not reject the over-identifying restrictions; a high p-value of this kind should nevertheless be read with care, since the test loses power when the number of instruments is large relative to the number of groups (Roodman, 2009b). Overall, the diagnostics reveal no major problem with the specifications, although they cannot prove that the instruments are valid.

4. Empirical results and discussion

4.1. Statistical properties of the series and the architecture of correlations

Table 4. Descriptive statistics of the variables (2000–2023, N = 504)

Variable	Obs.	Mean	Std. dev.	Min	Max
MVA	504	10.915	4.120	3.400	23.400
FD	504	0.146	0.114	0.026	0.593
lnGDPpc	504	6.941	1.029	4.697	9.232
OPEN	504	0.623	0.256	0.206	1.569
INV	504	22.835	9.310	2.000	78.000
HC	504	3.739	1.829	0.350	10.590
NRR	504	12.611	11.274	−0.600	59.700
lnPOP	504	16.720	1.084	14.059	19.244
CORR	504	−0.687	0.535	−1.600	0.800

Source: Authors' calculations under Stata 16. Natural-resource rents for 2022 and 2023, not yet reported in the WDI, are obtained by linear extrapolation of each country series; for countries whose rents are close to zero this produces small negative values, the lowest being −0.600. Since rents cannot be negative by construction, these values result from the extrapolation and not from the source data.

The descriptive statistics already show considerable diversity. Manufacturing value added ranges from 3.4 to 23.4 per cent of GDP across the countries and years of the sample, with a mean of 10.9 per cent. This wide range reflects large differences in industrial structure across countries and supports the inclusion of country fixed effects. The FD index varies even more in relative terms, from 0.026 (Burundi) to 0.593 (Mauritius), and its mean of 0.146 on the 0–1 scale of Svirydzenka (2016) is consistent with the well-known shallowness of financial systems in the region (Sahay et al., 2015). Because the FD index and the private credit ratio are measured on different scales, the 100 per cent credit threshold of Arcand et al. (2015) cannot be applied directly to our data; private credit in the sample countries nonetheless remains far below that level (Sahay et al., 2015).

Comparison with other African studies helps to put these values in context. The mean FD of 0.146 is lower than the mean of 0.209 reported by Girma and Huseynov (2025) for 30 African countries over 2000–2020. It is also below the composite-index threshold of 23.18 per cent estimated by Abdulai et al. (2024), which only the more financially developed countries of our sample exceed. Thresholds expressed in terms of private credit, such as the 6.52 per cent level of Abdulai et al. (2024) or the 6.9 per cent turning point of Machado et al. (2021), cannot be translated directly into values of the composite index. Such differences in measurement are one reason why findings on non-linearity vary across studies, a point taken up in Section 4.3.

Table 5. Linear correlation coefficients (dependent variable: MVA)

Variable	Correlation with MVA
FD	0.029
lnGDPpc	0.004
OPEN	0.137
INV	0.317
HC	0.104
NRR	0.434
lnPOP	−0.207
CORR	−0.128

Source: Authors' calculations under Stata.

The simple correlations in Table 5 may seem surprising at first: the correlation between FD and MVA is very low (0.029), while the correlation between natural-resource rents and MVA is the highest (0.434). This does not contradict the hypothesis that finance supports industry, because bivariate correlations mix direct effects, indirect effects and the influence of omitted factors. In several resource-rich countries of the sample, such as Nigeria, Angola and Gabon, part of manufacturing consists of the primary processing of raw materials, which creates a mechanical link between rents and MVA. A multivariate model that controls for these factors and treats FD as endogenous is needed to estimate its effect.

Table 6. Multicollinearity test (variance inflation factors)

Variable	VIF	Variable	VIF
FD	1.52	HC	1.18
NRR	1.42	lnPOP	1.10
lnGDPpc	1.35	CORR	1.05
OPEN	1.28	Mean VIF	1.29
INV	1.20		

Source: Authors' calculations.

Table 7. Levin, Lin and Chu (LLC) panel unit-root tests

Variable	LLC stat. (level)	LLC stat. (1st diff.)	Conclusion
MVA	−2.5109***		I(0)
FD	−2.0286**		I(0)
lnGDPpc	−4.8199***		I(0)
HC	−2.9760***		I(0)
NRR	−3.9641***		I(0)
lnPOP	−7.0715***		I(0)
CORR	−2.4882***		I(0)
INV	−1.3972	−16.1726***	I(1)
OPEN	−1.4040	−17.0385***	I(1)

*Source: Authors' calculations. *** significant at 1%; ** at 5%; * at 10%.*

According to the LLC test, seven variables are stationary in levels and two (INV and OPEN) are I(1). The Westerlund tests in Table 2 are consistent with a long-run relationship among the variables, which, under the assumptions discussed in Section 3.4, supports estimation in levels. They do not, however, guarantee its validity.

4.2. Main estimates: 2SLS and GMM-SYS results

Table 8. Two-stage least squares estimates (2SLS), dependent variable: MVA

Variable	Coefficient	z-statistic	p-value
FD	0.3789***	4.42	0.000
lnGDPpc	0.2345***	4.32	0.000
OPEN	0.1988**	2.16	0.031
INV	0.3456***	2.63	0.009
HC	0.2891***	2.78	0.005
NRR	−0.1234**	−2.02	0.044
lnPOP	0.0457**	2.05	0.040
CORR	−0.1766***	−3.08	0.002

Source: Authors' estimates under Stata 16 (xtivreg2, fe), with country fixed effects (within transformation, so no intercept is reported) and year dummies. Excluded instruments: FD(t−1) and FD(t−2). First-stage F(2, 411) = 18.50 (Stock–Yogo critical values for one endogenous regressor and two excluded instruments: 19.93 at 10% and 11.59 at 15% maximal IV

*size); DWH $\chi^2 = 6.71$ ($p = 0.010$); Hansen J $\chi^2(1) = 0.654$ ($p = 0.419$). N = 462 (2002–2023). *** p < 0.01; ** p < 0.05; * p < 0.10.*

Table 9. Two-step system GMM estimates (GMM-SYS), dependent variable: MVA

Variable	Coefficient	z-statistic	p-value
L1.MVA (persistence)	0.4604**	2.54	0.011
FD	0.2496***	3.08	0.002
lnGDPpc	0.0614***	3.42	0.000
OPEN	−0.0365***	−3.24	0.001
INV	0.0329**	2.41	0.023
HC	0.5115***	4.35	0.000
NRR	0.0444**	1.98	0.048
lnPOP	−0.2178**	−2.53	0.032
CORR	0.0723**	2.04	0.041
Constant	4.5565*	2.83	0.067

*Source: Authors' estimates under Stata 16 (xtabond2). Number of instruments = 17 (fewer than the 21 countries); AR(1) p = 0.003; AR(2) p = 0.936; Hansen J p = 0.811; collapse option, laglimits(2 4); Windmeijer correction applied. N = 483 (2001–2023). *** p < 0.01; ** p < 0.05; * p < 0.10.*

The two estimators lead to the same qualitative conclusion: the FD index is positively associated with manufacturing value added, and the coefficient is significant at the 1 per cent level in both cases. The estimate is 0.379 under 2SLS (Table 8) and 0.250 under GMM-SYS (Table 9). The lower value in the dynamic model is consistent with the fact that part of the variation in MVA is absorbed by the lagged dependent variable (L1.MVA = 0.460); in that model, the FD coefficient measures the impact effect, holding past manufacturing output constant. Since MVA is expressed in per cent of GDP and the FD index ranges from 0 to 1, the coefficients must be read in these units. A rise of 0.1 in the index, close to one standard deviation in the sample (0.114), is associated with an increase in MVA of about 0.038 percentage points of GDP under 2SLS (0.379 × 0.1) and 0.025 percentage points under GMM-SYS (0.250 × 0.1), other variables held constant. Under the assumptions of the dynamic model, the implied long-run effect is $\beta/(1 - \rho) = 0.250/(1 - 0.460) \approx 0.46$, so that a permanent increase of 0.1 in the index would raise MVA by about 0.046 percentage points in the long run. Even a large change, such as a move from the sample mean (0.146) to the highest value observed in the panel (0.593), would add only about 0.17 percentage points of GDP to MVA under 2SLS, against a sample mean of 10.9 per cent. The association is therefore statistically robust but modest in economic terms. These are reduced-form coefficients, and they do not reveal which of the channels discussed in Section 2.3 is at work.

The coefficient on lagged MVA (0.460) indicates substantial persistence in the industrial structure of SSA countries: other things equal, each percentage point of manufacturing value added in a given year is associated with about 0.46 points in the following year, which supports the use of a dynamic specification. This persistence is in line with the view that industrial capabilities build up slowly through learning and the accumulation of technical skills, and that they are neither quickly lost after adverse shocks nor quickly gained after favourable ones (Hidalgo & Hausmann, 2009; Rodrik, 2016).

Among the control variables, human capital (HC) has the largest coefficient in the GMM-SYS model (0.512, $p < 0.01$), which is consistent with the role of education in industrialisation emphasised by Lucas (1988) and, for Nigeria, by Adeosun et al. (2023). In our data, HC is measured by public spending on education, so the coefficient reflects an input into skill formation rather than skills as such. Firm-level evidence suggests why skills may matter: Diao et al. (2025) attribute part of Africa's manufacturing puzzle to the gap between capital-intensive large firms and labour-intensive small firms, a gap that a better-trained workforce could help to narrow. Other African studies also point to links between finance and human capital. Kinda and Sawadogo (2023) obtain their thresholds after controlling for human capital, and Nkemgha et al. (2023) estimate joint financial-development and human-capital thresholds beyond which infrastructure investment supports industrialisation. Our model, however, contains no

interaction terms; it does not test whether finance and human capital are complements and shows only that both are positively associated with MVA. Domestic investment (INV) is positive and significant in both models, although the size of its coefficient differs considerably (0.346 under 2SLS and 0.033 under GMM-SYS). This result is consistent with the capital accumulation channel, but here again the model does not test a direct complementarity between investment and financial development. Income per capita (lnGDPpc) also enters positively, which may reflect demand effects and larger domestic markets for manufactured goods.

Some results differ between the two specifications and must be interpreted with caution. The coefficient on natural-resource rents is negative under 2SLS (−0.123, $p < 0.05$) and positive under GMM-SYS (0.044, $p < 0.05$). This change of sign is a difference between specifications, a static model without lagged MVA and a dynamic model with it, and the comparison of the two estimators does not on its own allow short-run and long-run effects to be separated. The negative 2SLS coefficient is compatible with the Dutch-disease mechanism, in which resource revenues lead to a real appreciation and draw resources away from tradable sectors (Auty, 1993). Within the dynamic model, the long-run effect implied by the GMM-SYS estimates is $\beta/(1 - \rho) = 0.044/(1 - 0.460) \approx 0.08$, which is positive. One possible interpretation is that part of the rents is reinvested in domestic production, but this mechanism is not tested here.

Trade openness shows a similar pattern: its coefficient is positive under 2SLS (0.199) and negative under GMM-SYS (−0.037), with an implied long-run effect in the dynamic model of about −0.07 (−0.037/0.540). As with rents, this is a difference across specifications and not evidence of distinct short-run and long-run effects. The positive static estimate is consistent with access to imported inputs and technology, whereas the negative dynamic estimate is consistent with import competition weakening local manufacturing where production remains concentrated in primary goods. The latter reading is close to Rodrik's (2016) account of premature deindustrialisation and raises the question of how trade liberalisation and structural transformation should be sequenced. Evidence from West Africa adds a further element. For fourteen West African countries over 2002–2022, Bouda et al. (2025) find that openness supports industrial development, that this effect increases with institutional quality, and that it decreases as financial development progresses; they suggest that credit in the sub-region is allocated in a way that shields established firms from import competition. Their negative interaction term and our negative GMM-SYS coefficient both indicate that openness and finance do not necessarily reinforce each other when the financial system lends mainly to non-tradable activities.

The coefficient on control of corruption (CORR) is negative and significant under 2SLS (−0.177, $p < 0.01$) but positive under GMM-SYS (0.072, $p < 0.05$). This instability probably reflects a measurement problem, since the aggregate WGI indicator summarises very different institutional situations in a single score. Studies that examine the separate dimensions of institutional quality would help to clarify this result (Acemoglu et al., 2005; Demetriades & Law, 2006).

4.3. Non-linearity test: the "too much finance" hypothesis in the SSA setting

Table 10. Non-linearity test in the finance–industry relation (quadratic specification)

Variable	Coefficient	z-statistic	p-value
FD	0.4123***	3.89	0.000
FD^2	−0.1780	−1.25	0.213
Control variables	Included		
Implied turning point (for information)	1.16 (outside range)		

Source: Authors' estimates. FD^2 is not statistically significant. The turning point implied by the point estimates, 0.4123/(2 × 0.1780) ≈ 1.16, lies outside both the observed range of the FD index (0.026 to 0.593) and the 0–1 scale of the index; it is reported for information only.

The quadratic specification examines the hypothesis of Arcand et al. (2015) that financial development becomes harmful beyond a certain level. The coefficient on FD^2 is negative (−0.178) but not statistically significant ($p = 0.213$), so the data do not support the presence of a non-linear relationship. For information, the turning point implied by the point estimates, 0.4123/(2 × 0.1780) ≈ 1.16, lies above the highest value observed in the panel (0.593, Mauritius) and even outside the 0–1 range of the index, so it

has no empirical meaning in this sample. The only conclusion that can be drawn is that no statistically significant quadratic relationship between financial development and manufacturing value added is detected over the observed range of the index. This result neither confirms nor rules out a threshold at levels of financial development not observed in the sample.

This result should be set against the threshold literature. Machado et al. (2021) find a turning point at 6.9 per cent of private credit to GDP, Kinda and Sawadogo (2023) find that finance harms industrialisation below a certain threshold, and Luintel et al. (2024) find no robust threshold at all. Our results are closest to the last of these, since we find no significant curvature across the observed range of the FD index. Several differences in research design may explain why the thresholds reported for SSA are not reproduced here. Machado et al. (2021) measure finance by private credit and study aggregate growth, whereas we use a composite index that also covers access and efficiency and study manufacturing value added. A threshold expressed in credit terms and one expressed in terms of the composite index refer to different concepts, and they should not be treated as estimates of the same quantity. The access and efficiency components of the index, which remain low at the bottom of the distribution, may behave differently from credit volume, but our aggregate specification cannot test this. Decomposing the index, as suggested in Section 5, would be one way to address the question.

4.4. Synthesis of results

Overall, the results are consistent with the Schumpeterian view that financial development supports industrialisation, while showing that the association is modest in size in the SSA case. Comparisons with earlier studies must take differences in variables and units into account. Ibrahim and Alagidede (2018) report a coefficient of about 0.12 for the effect of finance on aggregate growth in 29 African countries; because their dependent variable, their measure of finance and hence their units differ from ours, our coefficients of 0.250 and 0.379 cannot be described as larger in any economic sense. The positive association with manufacturing is nonetheless in line with the argument of Rajan and Zingales (1998) that sectors relying on external finance benefit from financial development. Amoah and Jehu-Appiah (2022) identify a threshold beyond which finance and industrialisation reinforce each other; our quadratic test finds no significant non-linearity, so it neither supports nor contradicts their result. Avom et al. (2021) stress the role of regional value chains, and our mixed results for trade openness suggest that the industrial benefits of integration may depend on other conditions, possibly including financial depth, although this is not tested here. The large and positive coefficient on human capital in the GMM-SYS model is consistent with the emphasis that Asongu and Odhiambo (2019) place on complementarities between finance, education and digital inclusion.

Three further comparisons are useful. Lo and Cissokho (2023) show that institutional quality determines whether finance reaches manufacturing. Our results are compatible with theirs: we use a single governance indicator, and the instability of its coefficient across specifications may reflect the conditional relationships they describe. Asongu and Odhiambo (2022) find no effect of financial access alone on manufacturing; the difference with our findings may stem from measurement, since access without depth and efficiency, which is what their indicator captures, appears to favour services more than manufacturing. Nwani et al. (2023) report that resource dependence reduces the share of manufacturing while the depth of financial institutions increases it. Our 2SLS coefficient on rents (−0.123) agrees with the first finding, and the positive coefficient in the GMM-SYS model (0.044) is compatible with their argument that rents can be reinvested in manufacturing. As explained above, however, the change of sign between the static and the dynamic model reflects a difference in specification and should not be read as a decomposition of the resource effect into short-run and long-run components.

The main contribution of the study is to bring together three elements that are rarely combined in this literature: a recent multidimensional index of financial development, a period that includes the shocks after 2020, and the joint use of static (2SLS) and dynamic (GMM-SYS) estimators to address endogeneity and persistence.

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5. Conclusion

This paper has examined the relationship between financial development and manufacturing in a panel of 21 Sub-Saharan African countries over 2000–2023. Using a multidimensional measure of financial development and two estimators, 2SLS and system GMM, it finds that financial development is positively associated with manufacturing value added, with coefficients of 0.379 under 2SLS and 0.250 under GMM-SYS, both significant at the 1 per cent level. In practical terms, a 0.1-point increase in the index, about one sample standard deviation, is associated with an increase in manufacturing value added of between 0.025 and 0.038 percentage points of GDP, and of about 0.046 points in the long run implied by the dynamic model. The relationship is therefore robust in statistical terms but modest in magnitude. Human capital and domestic investment are also positively associated with manufacturing value added. The coefficients on natural-resource rents and trade openness change sign between the static and the dynamic specification, and the negative 2SLS coefficient on rents is compatible with a Dutch-disease effect. The quadratic specification detects no statistically significant non-linearity, so the results do not identify any threshold of financial development for the region.

From a theoretical point of view, the results are consistent with a Schumpeterian interpretation of the link between finance and industry in SSA and support the use of composite measures instead of single indicators. Botta et al. (2023) show that finance can support industry through accumulation while also shifting resources towards non-tradable sectors. Our results are compatible with this dual role, since the positive FD coefficient coexists with a negative 2SLS coefficient on rents and a negative GMM-SYS coefficient on openness, but the model does not test these mechanisms directly. The Schumpeterian view thus seems more useful as an account of how credit is allocated than as a statement about its volume. Empirically, the study provides updated estimates that include the shocks after 2020 and compares static and dynamic estimators on the same data. Methodologically, the combination of cointegration tests, specification diagnostics and two estimators offers an approach that could be applied to other questions of structural transformation.

Several policy implications follow, keeping in mind that the estimated effects are modest and that the transmission channels are not tested directly. First, financial deepening should remain a priority. This involves strengthening the prudential soundness of banking systems (BCEAO, BEAC and national central banks), supporting regional industrial development banks and encouraging long-term lending to manufacturing, for instance through targeted interest subsidies and risk-sharing schemes. Recent evidence on development banks is helpful here. Goga and Bell (2025), studying South Africa's Industrial Development Corporation, show that a development finance institution can reduce the risk of entering new sectors and attract private capital, but that its effect on structural transformation depends on the conditions attached to its loans more than on their volume. Without requirements that link finance to capability building and to linkages along the value chain, long-term loans may have little transformative effect. Rius and Andreoni (2025), in a study of fourteen African development finance institutions, point to a problem on the funding side: domestic institutional investors such as public pension funds, which have long-term liabilities, often hold short-term assets, while development banks that need long-term resources borrow at short maturities. Linking the two could mobilise domestic savings for industrial investment without relying on external capital. Second, digital financial inclusion through mobile money and fintech, which has spread rapidly in SSA (Asongu et al., 2019), could be directed more explicitly towards small and medium-sized manufacturers through suitable products such as working capital loans, equipment finance and partial guarantees. Access alone appears insufficient: Asongu and Odhiambo (2022) find that broader access did not raise manufacturing value added in SSA, and the gap identified by Allen et al. (2014) concerns the use of finance by firms as well as account ownership. Digital platforms reduce the cost of reaching small borrowers, but they do not solve the difficulty of lending to firms that lack collateral, audited accounts or a credit history; partial guarantees and registries for movable collateral are therefore useful complements. Third, public investment in human capital, particularly in technical and vocational training, should be strengthened, given the positive association between education spending and manufacturing value added found in this study.

Four limitations should be noted. First, the sample is restricted to 21 countries because of data availability, and extending it to all 48 SSA countries is left for future work. Second, although the composite index is multidimensional, it does not fully capture the institutional quality of the financial system, such as macroprudential supervision, transparency and resilience to shocks. Third, the analysis uses aggregate manufacturing value added and does not distinguish between branches, for example between heavy industry and agro-processing, or between capital goods and consumer goods. Fourth, because the index aggregates depth, access and efficiency for both institutions and markets, it cannot show which component drives the association with manufacturing, and African evidence suggests that the components may not move in the same direction. Girma and Huseynov (2025) find that depth and efficiency support growth while stability is negatively associated with it, and Talnan (2025) finds that finance reduces the share of manufactures in exports; a single coefficient on the aggregate index averages over such differences. In addition, the reduced-form specification does not test the transmission channels discussed in Section 2.3.

These limitations suggest four directions for further research. The first is to estimate the model separately for the West African Economic and Monetary Union (WAEMU), the Central African Economic and Monetary Community (CEMAC) and non-CFA countries, in order to assess whether monetary arrangements influence the link between finance and industry. The second is to include interactions between finance and institutional quality, so as to test the complementarity suggested by Demetriades and Law (2006). The third is to use disaggregated data from the UNIDO INDSTAT database to identify the manufacturing branches that respond most to financial development. The fourth, which we regard as the most pressing, is to decompose the index along the lines of Girma and Huseynov (2025) and estimate separate effects for depth, access, efficiency and stability, since an aggregate coefficient, however precise, cannot tell central banks or development banks which part of the financial system should be strengthened. Work along these lines would refine the present findings and the policy advice that can be drawn from them.

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