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# ECONOMIC STRUCTURAL CHANGE AND INCOME INEQUALITY IN ASEAN: THE MODERATING ROLE OF FINANCIAL DEVELOPMENT

**Thi Dong Nguyen**
D.Sc. in Economics, Researcher of the Training Department, Academy of Policy and Development, Hanoi, Vietnam;
e-mail: dongnt@apd.edu.vn
ORCID: 0009-0003-9659-6514

## ABSTRACT

This study repositions financial development from a direct determinant to a moderating variable that shapes how economic structural change influences income inequality across ASEAN. Drawing on an unbalanced panel of all ten member states over 2000–2019 (N = 178), the analysis employs a two-way Fixed Effects Model with an interaction term, complemented by System GMM, Panel-Corrected Standard Errors, instrumental-variable estimation, and lagged-variable specifications to address endogeneity. Although the sample window precedes the COVID-19 pandemic and subsequent geopolitical disruptions, the mechanisms under investigation — productivity-gap-driven inequality arising from sectoral reallocation and its attenuation by financial depth — are structural and institutional in nature; their policy relevance therefore extends beyond the estimation period. The principal findings show that structural change on its own significantly widens income inequality ($\beta_1 = 0.875$, $p < 0.01$), yet financial development exerts a countervailing moderating effect: the interaction coefficient (SCI × FD) is $-0.0092$ ($p < 0.05$), indicating that deeper financial systems attenuate the disequalizing pressure of sectoral transformation. A threshold is identified at approximately 95.1 % of GDP in private credit, beyond which the inequality-increasing effect of structural change is fully neutralized. Sub-sample analysis reveals that this moderating mechanism operates robustly only in upper-middle-income and high-income ASEAN economies; it remains statistically insignificant among lower-middle-income members whose financial systems lack the depth and inclusiveness to buffer distributional shocks. By illuminating a "dual paradox" — finance is directly associated with wider inequality yet indirectly mitigates it through moderation — the study provides actionable policy guidance: ASEAN nations pursuing structural transformation must simultaneously build inclusive financial systems, prioritizing broad-based access over mere credit expansion, to ensure that the gains from economic transition are equitably distributed.

**Keywords:** ASEAN, economic structural change, income inequality, financial development, financial inclusion, moderating effect, threshold analysis, panel data, Kuznets curve

**JEL Classification:** O15, O16, O47, G21

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## INTRODUCTION

The ASEAN region has undergone a profound process of economic structural change in the first two decades of the 21st century, characterized by the shift of labor and value-added from agriculture to industry and services. However, the speed and extent of this transition vary significantly among member states — from Singapore, where the financial and technology service sectors play a dominant role, to Cambodia and Myanmar, where agriculture still accounts for a large share of employment (McMillan et al., 2014; Cuyvers, 2019). Parallel to this transition, income inequality in the region has shown complex and divergent trends, raising a critical question: is economic structural change one of the factors shaping income inequality in ASEAN, and if so, what institutional factors moderate that relationship?

The relationship between structural change and inequality was founded on Lewis's (1954) dual-sector model and Kuznets's (1955) inverted U-shaped curve hypothesis, in which the process of labor migration from low-productivity to high-productivity sectors is seen as the mechanism creating income disparity in the early stages of development. Recent empirical studies show that this impact depends on the nature of the labor-receiving sector and the institutional context (Baymul & Sen, 2020; Ghosh et al., 2023), but the evidence remains inconclusive, especially in developing economies in Asia.

Notably, two related streams of research — the first examining the impact of structural change on inequality, and the second assessing the relationship between financial development and income distribution — have largely been studied independently without systematic intersection. Studies on structural change often treat financial development as a control variable to be excluded, rather than a mechanism to be analyzed in depth (Baymul & Sen, 2020). Conversely, studies on finance and inequality (Greenwood & Jovanovic, 1990; Clarke et al., 2006; Jauch & Watzka, 2016; De Haan & Sturm, 2017) typically measure the direct impact of financial development on inequality without considering whether that impact changes depending on the economic structural context. This separate approach leaves a significant gap: we do not fully understand how financial development shapes the way structural change affects income distribution. This gap is particularly concerning in the ASEAN context, where countries are simultaneously promoting both structural change and financial system development at very different paces. Therefore, the question that needs to be asked is: "Does financial development alter, and if so, how does it alter the relationship between structural change and inequality”?

This study addresses the gap outlined above by reconceptualizing financial development — shifting it from a direct explanatory factor to a conditional moderator — in the relationship between structural change and income inequality. Three interrelated objectives guide the analysis: first, estimating the direct distributional effect of economic structural change; second, assessing the direct association between financial development and inequality; and third — the paper's core contribution — testing whether financial development moderates the structural-change–inequality link and, if so, pinpointing the level of financial depth at which the disequalizing effect of sectoral reallocation is neutralized.

This study focuses on the specific context of ASEAN countries, a region with great diversity in economic development levels, ranging from low-income, agriculture-based economies to a global financial center. Based on this diverse context, the study expects to provide new empirical evidence that helps clarify the complex role of financial development in the process of economic structural transformation in ASEAN countries.

## LITERATURE REVIEW

### Economic Structural Change and Income Inequality

The process of economic structural change, understood as the reallocation of labor across sectors, is a major driver shaping income inequality. In the early stages of development, a clear division emerges between the traditional sector (low productivity) and the modern sector (high productivity). Income disparity arises between the group of workers who have successfully integrated into the modern sector and the majority who remain in the traditional sector, causing inequality to tend to increase. However, as the transition continues and the modern sector absorbs a larger share of the labor force, this disparity will gradually narrow, leading to a decline in overall inequality.

Recent studies show that productivity gaps between industries allow workers in expanding sectors to enjoy higher incomes, while workers in contracting sectors face structural unemployment or income decline (Gollin et al., 2014; McMillan et al., 2014). Furthermore, the shift to industry and services increases the skill premium, widening the wage gap between high-skilled and low-skilled labor (Acemoglu & Autor, 2011). In addition, structural change often concentrates in urban centers and industrial zones, exacerbating spatial inequality between urban and rural areas (Young, 2013).

However, empirical evidence on the aggregate direction of these mechanisms remains inconclusive. Baymul & Sen (2020) indicate that labor shifts to manufacturing tend to reduce inequality, while shifts to the service sector increase it. Hartmann et al. (2017) add that distributional outcomes depend significantly on the level of economic complexity and the institutional structure of each country. Herrendorf et al. (2014) also emphasize that different paces and sequences of structural change across economies can lead to very different distributional outcomes. This inconsistency suggests that the relationship between structural change and inequality may be governed by moderating factors that have not been systematically modeled.

### Financial Development and Income Inequality

How financial-system depth relates to income distribution remains a contested empirical and theoretical question, around which two broad perspectives have crystallized.


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One school of thought holds that an expanding financial system narrows income gaps. Foundational models by Galor and Zeira (1993) and Banerjee and Newman (1993) demonstrate that when credit markets are imperfect, collateral requirements lock the poor out of productive investments in education and entrepreneurship, perpetuating inter-generational inequality. Financial deepening — by loosening borrowing constraints and widening access for low-income households — can interrupt this cycle. Cross-country evidence (Beck et al., 2007) and more recent studies (Omar & Inaba, 2020) lend support, especially in developing-country settings. The rapid diffusion of financial technology is expected to reinforce this channel by extending services to previously excluded populations (Demir et al., 2022). For ASEAN, where broadening financial access has long been a central policy goal (Park & Mercado, 2018), this mechanism is directly relevant.

However, another stream of thought argues that this relationship is more complex and may be non-linear. According to the theory of Greenwood & Jovanovic (1990), in the early stages of financial development, transaction costs and entry requirements are high, allowing only the wealthy to benefit. The benefits only begin to spread to lower-income groups when the financial system reaches a sufficiently large threshold of development. This view is also supported by empirical studies showing that the relationship between financial development and inequality follows an inverted U-shaped curve (Kim & Lin, 2011), or that financial development can even increase inequality in certain contexts (Jauch & Watzka, 2016). Literature reviews (e.g., De Haan & Sturm, 2017; Zhang & Naceur, 2019) all conclude that empirical evidence is still inconsistent, and the actual impact depends heavily on the measurement method, national context, and structure of the financial system.

This inconsistency shows that the impact of financial development is not a universal constant but depends on contextual factors, especially the structure of the economy.

**The Moderating Role of Financial Development**

This study argues that the process of economic structural change itself creates pressures that increase inequality, manifested through productivity gaps between sectors, a widening income gap between skilled and unskilled labor, and geographical polarization. However, the extent to which these pressures translate into actual income inequality depends on the financial system's ability to mitigate and redistribute the impacts of the structural shock. A developed financial system can play a moderating role, softening these negative impacts through four main mechanisms:

First, supporting the labor transition process: Access to credit helps workers in contracting industries cover the costs of retraining or relocating to places with new job opportunities. This makes the labor market more flexible and reduces barriers, thereby limiting prolonged structural unemployment (Claessens & Perotti, 2007).

Second, promoting equality in human capital accumulation: An efficient credit market allows low-income households to borrow for their children's education. In the long run, this helps promote equality of opportunity in human capital accumulation, increasing the supply of skilled labor and thereby narrowing the skill-based income gap (Galor & Zeira, 1993).

Third, enhancing risk resilience: Financial instruments such as savings, insurance, and microcredit help households diversify their income sources and protect themselves against adverse shocks (e.g., crop failure, illness), preventing them from falling into chronic poverty (Demirgüç-Kunt & Levine, 2009).

Fourth, promoting efficient capital allocation: A developed financial system is capable of identifying and quickly allocating capital to new, dynamic economic sectors. This not only promotes growth but also creates more high-productivity jobs, allowing more workers to benefit from the structural change process (Seven & Coskun, 2016).

Thus, financial development is expected to have a negative moderating effect, meaning it mitigates the inequality-inducing impact of structural change. However, this role is not automatic. If the financial system only serves large corporations and the elite in emerging economic sectors, it could amplify rather than mitigate inequality. This is the possibility that Tsounta et al. (2015) call the "financial concentration effect." The actual direction of financial development's impact — whether it mitigates or amplifies inequality in the context of structural change — becomes an empirical question to be tested.

Synthesizing the above arguments, the paper's analytical framework identifies three relationships: the direct impact of structural change on inequality, the direct impact of financial development on inequality, and the moderating impact of financial development on the relationship between structural change and inequality. Three research hypotheses are stated as follows:

- **H1:** Economic structural change increases income inequality in ASEAN countries, consistent with the first half of the inverted U-shaped curve, as most member states are transitioning from agriculture to industry and services.
- **H2:** Financial development reduces income inequality by easing credit constraints and expanding investment opportunities for low-income groups.

- **H3:** Financial development plays a negative moderating role, whereby in countries with higher levels of financial development, the inequality-increasing impact of structural change will be weaker.

## AIMS AND OBJECTIVES

This study is conducted to fill the aforementioned research gap by repositioning financial development — from a direct explanatory variable to a moderating variable — in the relationship between structural change and income inequality. The paper aims for three objectives:

1. First, to assess the impact of economic structural change on income inequality.
2. Second, to analyze the direct impact of financial development on inequality.
3. Third, which is the core contribution, to test the moderating role of financial development, while identifying the threshold of financial development at which the nature of the relationship between structural change and inequality qualitatively changes.

## METHODS

### Data and Sample

The research sample includes all 10 ASEAN member states — Brunei, Cambodia, Indonesia, Laos, Malaysia, Myanmar, the Philippines, Singapore, Thailand, and Vietnam — over the period 2000–2019. The choice of this period is based on three reasons. First, the first two decades of the 21st century were a period of intense structural change for ASEAN, with the share of agriculture in the region's GDP falling from about 12% in 2000 to below 10% in 2019, while services increased from 45% to over 50% (World Bank, 2021); this was also a period of vibrant financial development, especially after the 1997–1998 Asian financial crisis when member states pushed for banking system reforms and expanded financial inclusion. Second, the 2000–2019 period avoids the unusual structural disruptions caused by the COVID-19 pandemic from 2020, which could distort the long-term relationships between structural change, financial development, and inequality. Third, this is the period for which data from all sources are most available: Penn World Table version 10.01 is updated to 2019 (Feenstra et al., 2015), SWIID version 9.6 has increasingly broad coverage for developing countries after 2000 (Solt, 2020), and the World Bank's World Development Indicators (WDI) provide a relatively complete macroeconomic data series for the ASEAN region during this period.

The period after 2019 has witnessed extraordinary disruptions — the COVID-19 pandemic, intensifying US–China strategic rivalry, the Russia–Ukraine military conflict, and associated supply-chain realignments — each of which has reshaped growth trajectories, fiscal balances, and measured inequality across ASEAN. Retaining the 2000–2019 estimation window, rather than extending it to encompass these events, is justified on three grounds. First, pandemic-era emergency transfers (cash subsidies in Indonesia and the Philippines, wage-support schemes in Singapore and Malaysia) temporarily compressed disposable-income inequality for reasons orthogonal to both structural change and financial depth, introducing transient noise that would obscure the long-run mechanisms this study aims to isolate. Second, at the time of data compilation, standardized Gini estimates in SWIID for several ASEAN members — particularly Myanmar, Laos, and Brunei — had not yet been updated beyond 2019 with the cross-country comparability that the econometric design requires. Third, and most critically, the relationships investigated here are structural and institutional: they describe how the depth and inclusiveness of the financial system condition the distributional consequences of long-run sectoral reallocation, not how economies react to exogenous shocks. Post-2019 events, if anything, reinforce the policy relevance of the findings, because the same financial-system attributes that moderate inequality during structural change — broad credit access, risk-sharing instruments, and efficient capital intermediation — are precisely those that determine an economy's capacity to cushion vulnerable populations during crises. The 2000–2019 estimates therefore constitute a robust pre-shock baseline against which post-pandemic distributional dynamics can be rigorously assessed once comparable data become available.

Data are compiled from four publicly available sources, with specific indicator codes and database versions documented below to ensure full reproducibility. The Gini coefficient of disposable income (variable: gini_disp) and market income Gini (variable: gini_mkt) are drawn from the Standardized World Income Inequality Database version 9.6, accessed via Frederick Solt's repository at https://fsolt.org/swiid/ (Solt, 2020). The SWIID v9.6 file used is dated October 2022; this version provides standardized estimates for all ten ASEAN members from 2000 onward, with the exceptions of Brunei (available from 2006) and Myanmar (from 2008).


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Macroeconomic indicators are collected from the World Development Indicators (WDI) database, November 2021 update, accessed via the World Bank DataBank portal (https://databank.worldbank.org/). The specific WDI indicator codes are: NV.AGR.TOTL.ZS (agriculture value added, % of GDP), NV.IND.TOTL.ZS (industry value added, % of GDP), NV.SRV.TOTL.ZS (services value added, % of GDP), FS.AST.PRVT.GD.ZS (domestic credit to private sector, % of GDP), NY.GDP.PCAP.KD (GDP per capita, constant 2015 USD), SP.URB.TOTL.IN.ZS (urban population, % of total), NE.CON.GOVT.ZS (general government final consumption expenditure, % of GDP), and BX.KLT.DINV.WD.GD.ZS (foreign direct investment, net inflows, % of GDP). All WDI series were downloaded simultaneously for the ten ASEAN countries over 2000–2019 using the DataBank's bulk download function.

The human capital index (variable: hc) is obtained from the Penn World Table version 10.01, released in January 2023, downloaded in Stata format from the Groningen Growth and Development Centre (https://www.rug.nl/ggdc/productivity/pwt/). This index is constructed from average years of schooling from Barro and Lee (2013) and an assumed rate of return based on Mincerian estimates (Feenstra et al., 2015).

The IMF's composite Financial Development Index (variable: FD, aggregate index ranging from 0 to 1), used in the robustness analysis, is accessed from the IMF's Financial Development Database at https://data.imf.org/?sk=f8032e80-b36c-43b1-ac26-493c5b1cd33b (Svirydzenka, 2016), 2022 vintage. This multidimensional index encompasses sub-indices for financial institutions and financial markets across three dimensions: depth, access, and efficiency.

Replication code (Stata 17 .do files) documenting all data cleaning, variable construction, and estimation procedures is available from the author upon reasonable request.

Since the calculation of the structural change index (SCI) requires sectoral share data from year t−1, the year 2000 is used as the base year, and the effective analysis period begins in 2001, creating a maximum of 190 country–year observations (10 countries × 19 years). However, data are not uniformly available across countries. Table 1 below summarizes the data status by country, where the effective period reflects the time frame in which all main variables are simultaneously available.

**Table 1. Summary of research data.**

| Country | Sample Period | Number of observations | Notes |
|---|---|---|---|
| Brunei | 2006–2019 | 14 | Limited SWIID data before 2006 |
| Cambodia | 2001–2019 | 19 | — |
| Indonesia | 2001–2019 | 19 | — |
| Laos | 2001–2019 | 19 | — |
| Malaysia | 2001–2019 | 19 | — |
| Myanmar | 2008–2019 | 12 | Very limited SWIID and private credit data before 2008 |
| Philippines | 2001–2019 | 19 | — |
| Singapore | 2001–2019 | 19 | — |
| Thailand | 2001–2019 | 19 | — |
| Vietnam | 2001–2019 | 19 | — |
| Total | | 178 | |

The final sample is an unbalanced panel of 178 country–year observations. The study uses an unbalanced panel rather than dropping countries with missing data or performing wholesale interpolation because retaining Brunei and Myanmar ensures representation of the entire spectrum of ASEAN's economic development — from a high-income oil economy to a low-income transitioning economy. For isolated missing cases (a single missing year between two years with data) in some control variables from the WDI, linear interpolation is applied and transparently noted. To mitigate the influence of outliers, all continuous variables in the model are winsorized at the 1st and 99th percentiles, replacing values outside these two percentiles with the value at the respective percentile, preserving the sample size while reducing the disproportionate impact of extreme observations.

### Measurement of Variables

*Dependent Variable*: Income inequality ($GINI_{it}$) is measured by the Gini coefficient of disposable income (post-tax, post-transfer) from SWIID version 9.6. This coefficient ranges from 0 to 100, with higher values reflecting greater inequality.

The choice of disposable income Gini over market income Gini is based on the rationale that disposable Gini reflects the actual inequality experienced by the population after accounting for the redistributive effects of taxes and social transfers, and thus more fully captures the distributional consequences of both structural change and related financial policies. SWIID is preferred over other sources (WIID, PovcalNet) for its methodological consistency: through a standardization process based on reference data from the Luxembourg Income Study, SWIID ensures cross-country and cross-time comparability—an essential feature in panel data analysis (Solt, 2020). The market income Gini coefficient from the same source is used in the robustness analysis.

*Independent Variable*: The structural change index ($SCI_{it}$) is constructed from data on the value-added shares of the three economic sectors in GDP — agriculture (including forestry and fishing), industry (including mining, manufacturing, construction, electricity, and water), and services — collected from the WDI according to the ISIC Rev.4 classification. The calculation process involves two steps: first, the value-added share of each sector k in country i's GDP at year t ($S_{kit}$) is recorded; then, the SCI index is calculated using the formula:

$$SCI_{it} = \Sigma_{k=1}^{3} |S_{kit} - S_{ki,t-1},|$$

*where k = 1 (agriculture), k = 2 (industry), k = 3 (services).*

The SCI index measures the intensity of output reallocation between economic sectors in a year: a value of 0 indicates a completely static economic structure, while a larger value reflects a stronger degree of structural change. The theoretical maximum value is 200 (when all output shifts from one sector to another), but in practice, observed values typically range from 0–10 percentage points.

A deliberate choice is made to construct SCI from value-added shares rather than the employment shares more commonly used in the structural-change literature (McMillan et al., 2014; Herrendorf et al., 2014). This decision rests on three grounds. First, value-added shares directly capture the productivity dimension of structural change — the reallocation of output across sectors with heterogeneous labor productivities — which is the channel through which structural transformation generates income inequality in the theoretical framework of this study. Employment-based indices measure labor reallocation but do not reflect whether workers move into sectors with genuinely higher productivity or merely into low-productivity informal services (Gollin et al., 2014). Second, in the ASEAN context, employment data suffer from severe measurement limitations: large informal sectors in Cambodia, Laos, Myanmar, and Vietnam are inconsistently captured across national labor-force surveys, creating comparability problems that are less acute for national-accounts-based value-added data reported under standardized UN SNA methodology. Third, disaggregated sectoral employment data from the WDI exhibit substantially more missing observations for ASEAN members than do value-added shares, which would further reduce an already small sample. The study acknowledges that the three-sector classification does not capture intra-sectoral dynamics and that value-added and employment shares may diverge where capital intensity varies sharply across sectors. However, structural growth models from Lewis (1954) to Herrendorf et al. (2014) consistently identify inter-sectoral shifts — between agriculture, industry, and services — as the core axis of transformation that creates productivity gaps and income differentiation, and value-added shares most directly measure the output-reallocation dimension of this process. In the robustness analysis, the value-added shares of industry ($INDVA_{it}$) and services ($SERVA_{it}$) in GDP are used as alternative measures to check if the results are sensitive to the specific direction of change.

*Moderating Variable*: Financial development ($FD_{it}$) is measured by the ratio of domestic credit to the private sector as a percentage of GDP, collected directly from the WDI. This index reflects the extent to which the banking system transforms savings into credit for private enterprises and households and is the most commonly used measure of financial depth in research on finance and inequality. The study recognizes that the credit/GDP ratio only captures one dimension — the depth of financial intermediation through the banking channel — without encompassing access, efficiency, or the development of capital markets and financial technology. However, this choice is theoretically grounded; private sector credit is directly linked to the credit constraint relaxation mechanism in the Galor & Zeira (1993) model — the central channel through which financial development impacts inequality in the paper's analytical framework. At the same time, this is the measure used by the majority of foundational studies in the field, including Beck et al. (2007), Clarke et al. (2006), Jauch & Watzka (2016), as well as Baymul & Sen (2020), facilitating the comparison of results across studies. Regarding its suitability for the ASEAN context, the banking sector dominates financial intermediation in most member countries—even in Singapore and Malaysia, where capital markets are more developed, bank credit remains the main source of financing for small and medium-sized enterprises and households. Therefore, private credit/GDP most closely reflects the financial access mechanism relevant to income distribution in the region. To check the robustness against this single-dimension limitation, the IMF's composite financial development index ($FDX_{it}$) — ranging from 0 to 1 and covering the depth, access, and efficiency of both the banking system and capital markets (Svirydzenka, 2016) — is used in a supplementary analysis.


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*Control Variables* are selected based on theory and empirical studies on the determinants of income inequality. GDP per capita ($\ln GDPPC_{it}$), in constant 2015 USD and taken in natural logarithm, controls for the non-linear relationship between economic development and inequality; its squared form $(\ln GDPPC_{it})^2$ is included to test the inverted U-shaped Kuznets curve hypothesis. The urbanization rate ($URB_{it}$), measured as the percentage of the population living in urban areas, controls for the spatial impact of population concentration, which is often associated with income differentiation in the early stages (Young, 2013). Government final consumption expenditure as a percentage of GDP ($GEXP_{it}$) represents the state's redistributive role through the provision of public goods and social services. The human capital index ($HC_{it}$) from Penn World Table 10.01, constructed based on average years of schooling and Mincerian returns to education (Feenstra et al., 2015), reflects the average skill level of the labor force — a key factor determining the wage gap between skilled and unskilled labor. Net foreign direct investment inflows as a percentage of GDP ($FDI_{it}$) control for the impact of international economic integration, which affects inequality through the channels of job creation, technology transfer, and wage premiums in the foreign-invested enterprise sector (Figini & Görg, 2011).

Table 2 summarizes all variables, notations, measurements, data sources, and expected signs based on the theoretical framework above.

**Table 2. Summary of variables in the model**. Note: The expected signs are based on the theoretical framework and hypotheses H1, H2, and H3 in Section 2. "+/−" indicates that the direction of the effect is theoretically ambiguous.

| Variable | Notation | Measurement | Source | Expected Sign |
|---|---|---|---|---|
| Income Inequality | $GINI_{it}$ | Gini coefficient of disposable income (0–100) | SWIID v9.6 | — |
| Structural Change | $SCI_{it}$ | Sum of absolute changes in the value-added shares of the three sectors (% points) | WDI | + (H1) |
| Financial Development | $FD_{it}$ | Domestic credit to private sector (% of GDP) | WDI | − (H2) |
| Interaction Term | $SCI_{it} \times FD_{it}$ | Product of SCI and FD | Calculated | − (H3) |
| GDP per capita | $\ln GDPPC_{it}$ | Ln(GDP per capita, constant 2015 USD) | WDI | + |
| GDP per capita² | $(\ln GDPPC_{it})^2$ | Square of lnGDPPC | WDI | − |
| Urbanization | $URB_{it}$ | Urban population (% of total population) | WDI | +/− |
| Government expenditure | $GEXP_{it}$ | Government final consumption expenditure (% of GDP) | WDI | − |
| Human Capital | $HC_{it}$ | Human capital index (based on years of schooling and returns to education) | PWT 10.01 | − |
| Foreign Direct Investment | $FDI_{it}$ | Net FDI inflows (% of GDP) | WDI | +/− |

## Econometric Model Specification

*Model 1 — Baseline model, including only the direct impact of structural change on inequality:*

$$GINI_{it} = \alpha + \beta_1 SCI_{it} + \gamma_1 \ln GDPPC_{it} + \gamma_2 (\ln GDPPC_{it})^2 + \gamma_3 URB_{it} + \gamma_4 GEXP_{it} + \gamma_5 HC_{it} + \gamma_6 FDI_{it} + \mu_i + \lambda_t + \varepsilon_{it} \quad (1)$$

*Model 2 — Adding financial development as a direct explanatory variable:*

$$GINI_{it} = \alpha + \beta_1 SCI_{it} + \beta_2 FD_{it} + \gamma_1 \ln GDPPC_{it} + \gamma_2 (\ln GDPPC_{it})^2 + \gamma_3 URB_{it} + \gamma_4 GEXP_{it} + \gamma_5 HC_{it} + \\ + \gamma_6 FDI_{it} + \mu_i + \lambda_t + \varepsilon_{it} \quad (2)$$

*Model 3 — Moderation model, adding the interaction term between structural change and financial development:*

$$GINI_{it} = \alpha + \beta_1 SCI_{it} + \beta_2 FD_{it} + \beta_3 (SCI_{it} \times FD_{it}) + \gamma_1 \ln GDPPC_{it} + \gamma_2 (\ln GDPPC_{it})^2 + \gamma_3 URB_{it} + \\ + \gamma_4 GEXP_{it} + \gamma_5 HC_{it} + \gamma_6 FDI_{it} + \mu_i + \lambda_t + \varepsilon_{it} \quad (3)$$

*Model 4 — Dynamic model, adding the lagged dependent variable to capture the persistence of inequality:*

$$GINI_{it} = \alpha + \delta GINI_{i,t-1} + \beta_1 SCI_{it} + \beta_2 FD_{it} + \beta_3 (SCI_{it} \times FD_{it}) + \gamma_1 \ln GDPPC_{it} + \gamma_2 (\ln GDPPC_{it})^2 + \gamma_3 URB_{it} + \\ + \gamma_4 GEXP_{it} + \gamma_5 HC_{it} + \gamma_6 FDI_{it} + \mu_i + \lambda_t + \varepsilon_{it}, \quad (4)$$

*where $\mu_i$ is the country-fixed effect, capturing unobserved time-invariant characteristics such as institutions, geography, and culture; $\lambda_t$ is the time-fixed effect, controlling for common shocks affecting all countries in the same year (e.g., the 2008–2009 global financial crisis); and $\varepsilon_{it}$ is the random error term.*

The key coefficient in Models 3 and 4 is $\beta_3$—the coefficient of the interaction term ($SCI_{it} \times FD_{it}$). In a model with an interaction, $\beta_1$ no longer represents the average effect of structural change but reflects the marginal effect when financial development is zero. The marginal effect of structural change on inequality depends on the level of financial development:

$$\partial GINI_{it} / \partial SCI_{it} = \beta_1 + \beta_3 \times FD_{it}.$$

The threshold FD* has significant policy implications: it determines the minimum level of financial development required to completely neutralize the inequality-inducing effect of structural change.

**Estimation Strategy**

*Main Estimation Method.* Models (1)–(3) are estimated with two-way fixed effects and Driscoll–Kraay standard errors (Driscoll & Kraay, 1998; Hoechle, 2007). Fixed effects are preferred over random effects on the basis of the Hausman test. Given the panel dimensions (N = 10, T = 19), Driscoll–Kraay standard errors are favored over conventional cluster-robust errors because they remain consistent in the presence of heteroskedasticity, within-panel serial correlation, and — crucially — cross-sectional dependence. The last property is particularly germane to ASEAN, where tightly interwoven trade, investment, and financial linkages make it likely that an economic disturbance in one country spills over to its neighbors (Pesaran, 2004).

*Supplementary Estimation Method.* For robustness, the static models are additionally estimated using Panel-Corrected Standard Errors (PCSE) following Beck and Katz (1995), a method explicitly designed for panels where N is small relative to T. PCSE accounts for panel heteroskedasticity and contemporaneous correlation across units while avoiding the downward bias in standard errors that conventional OLS produces in such settings. The consistency of results between the Driscoll–Kraay and PCSE approaches enhances confidence in the findings by demonstrating that conclusions are not sensitive to the specific variance–covariance correction employed.

*Dynamic Model.* Model (4) incorporates a lagged dependent variable to capture the well-documented persistence of income inequality — Gini coefficients adjust sluggishly because wage structures, asset ownership, and institutional arrangements change only gradually. Estimation follows the System GMM procedure of Blundell and Bond (1998). Because N = 10 is small, the instrument count is managed through the "collapse" option recommended by Roodman (2009), which restricts the lag depth so that the number of instruments remains below the number of groups. Instrument validity is assessed with the Hansen J statistic, and the Arellano–Bond AR(2) test verifies that second-order autocorrelation in the first-differenced residuals — a consistency requirement for GMM — is absent.

*Handling Endogeneity*: The relationship between structural change, financial development, and inequality may be affected by reverse causality (high inequality could slow structural change or hinder financial development) and omitted variable bias (unobserved institutional factors could simultaneously affect all three variables). In addition to country and time fixed effects, the paper applies two additional measures. First, the independent and moderating variables are used in their one-period lagged forms ($SCI_{i,t-1}$, $FD_{i,t-1}$, and $SCI_{i,t-1} \times FD_{i,t-1}$) in a supplementary analysis, based on the logic that past values of explanatory variables are predetermined relative to current inequality, thus mitigating simultaneous causality. Second, an instrumental variable estimation using two-stage least squares (IV-2SLS) is performed, using deeper lags (t−2, t−3) as instruments for $SCI_{it}$ and $FD_{it}$. Instrument relevance is checked with the Kleibergen–Paap rk Wald F statistic, and instrument validity is tested with the Hansen over-identification test.

*Diagnostic Tests*: To select the appropriate estimation method and assess model assumptions, several diagnostic tests are performed prior to estimation: the Hausman test to choose between fixed and random effects; the Pesaran CD test (Pesaran, 2004) for cross-sectional dependence among countries; the modified Wald test for groupwise heteroskedasticity; the Wooldridge (2010) test for first-order autocorrelation in panel data; the Variance Inflation Factor (VIF) to assess multicollinearity — especially important in models with interaction terms, where multicollinearity between constituent variables and the interaction term needs to be controlled (mean-centering is applied if VIF exceeds the threshold of 10); the Im–Pesaran–Shin (IPS) and Fisher–ADF panel unit root tests to verify the stationarity of the series; and the Westerlund (2007) cointegration test if some variables are confirmed to have a unit root, to ensure the estimated relationship reflects a meaningful long-run equilibrium rather than a spurious regression.

To estimate the regression models and perform the tests, the study uses Stata 17 software.


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# RESULTS

## Descriptive Statistics and Correlation Analysis

Table 3 presents the descriptive statistics for the full sample, providing an initial overview of the data's central tendency, dispersion, and range.

**Table 3. Descriptive statistics of model variables.** Note: All continuous variables have been winsorized at the 1st and 99th percentiles. *(Source: author's calculations from SWIID v9.6, WDI (World Bank, 2021), PWT 10.01, and IMF)*

| Variable | N | Mean | Std. Dev. | Min | Max |
|---|---|---|---|---|---|
| GINI | 178 | 37.82 | 3.94 | 29.14 | 46.83 |
| SCI | 178 | 2.87 | 2.19 | 0.21 | 11.36 |
| FD | 178 | 62.38 | 43.17 | 3.82 | 156.38 |
| lnGDPPC | 178 | 8.19 | 1.34 | 6.02 | 11.06 |
| $(lnGDPPC)^2$ | 178 | 68.87 | 22.14 | 36.24 | 122.32 |
| URB | 178 | 44.52 | 21.93 | 17.61 | 100.00 |
| GEXP | 178 | 10.83 | 3.68 | 4.47 | 19.62 |
| HC | 178 | 2.46 | 0.61 | 1.51 | 3.86 |
| FDI | 178 | 5.32 | 5.76 | −2.14 | 27.63 |

Table 3 reveals substantial cross-country heterogeneity that reflects ASEAN's diverse development landscape. The Gini range (29.14–46.83) spans nearly 18 points: the lower end corresponds to economies such as Vietnam and Cambodia, where historically compressed wage structures and large subsistence-agriculture sectors limit measured disposable-income gaps, whereas the upper end reflects countries like Malaysia and the Philippines, where rapid industrialization and urbanization have produced pronounced income stratification. Financial development displays the widest relative dispersion (coefficient of variation ≈ 0.69): private-sector credit ranges from 3.82 % of GDP — characteristic of Cambodia's nascent banking sector — to 156.38 % in Thailand, a level comparable to many OECD economies. This stark variation is analytically advantageous, providing the cross-country leverage needed to estimate how financial depth conditions the distributional effects of structural change. The mean structural change index of 2.87 percentage points, with a maximum of 11.36, indicates that certain country-year episodes experienced abrupt sectoral reallocation — typically associated with rapid industrialization, resource-boom cycles, or policy-driven shifts — while others evolved very gradually. Together, these statistics underscore that ASEAN constitutes a natural laboratory for studying the interplay between structural transformation, financial systems, and inequality.

Table 4A below presents the correlation matrix between the variables.

**Table 4A. Correlation matrix between variables.** Note: Bold values indicate correlation coefficients exceeding the threshold |r| > 0.8. VIF is calculated from the auxiliary regression of each independent variable on the remaining independent variables.

| | 1 | 2 | 3 | 4 | 5 | 6 | 7 | 8 | VIF |
|---|---|---|---|---|---|---|---|---|---|
| (1) GINI | 1.000 | | | | | | | | - |
| (2) SCI | 0.147 | 1.000 | | | | | | | **1.15** |
| (3) FD | −0.306 | −0.081 | 1.000 | | | | | | **2.75** |
| (4) lnGDPPC | −0.178 | −0.162 | 0.758 | 1.000 | | | | | **3.85** |
| (5) URB | −0.103 | −0.127 | 0.623 | **0.836** | 1.000 | | | | **3.52** |
| (6) GEXP | −0.213 | 0.042 | −0.156 | 0.108 | 0.067 | 1.000 | | | **1.09** |
| (7) HC | −0.241 | −0.093 | 0.568 | 0.781 | 0.537 | 0.058 | 1.000 | | **2.98** |
| (8) FDI | 0.047 | 0.114 | 0.354 | 0.442 | 0.291 | −0.118 | 0.189 | 1.000 | **1.34** |

Most pairwise correlation coefficients fall below 0.7, and the maximum VIF of 3.85 (well under the conventional threshold of 10) confirms that multicollinearity does not pose a threat to the regression estimates. From an economic standpoint, the positive bivariate association between SCI and GINI (r = 0.147) offers preliminary — though unconditional — support for hypothesis H1: countries undergoing faster sectoral reallocation tend to exhibit wider income gaps. The negative

correlation between FD and GINI (r = −0.306) is consistent with the credit-constraint-relaxation channel emphasized by Galor and Zeira (1993), yet — as the moderation analysis will demonstrate — this unconditional association masks a more nuanced conditional relationship in which the direction and magnitude of finance's effect depend on the structural-change context. The strong co-movement between FD and lnGDPPC (r = 0.758) mirrors the well-documented co-evolution of financial depth and per-capita income, underscoring the importance of controlling for income level to disentangle the independent contribution of financial development from the broader development process.

### Pre-Estimation Diagnostic Tests

Before proceeding to regression estimation, a battery of diagnostic tests is conducted to justify the modeling choices. Table 4B reports the results.

**Table 4B. Pre-estimation diagnostic tests.** Notes: The Hausman test is based on Model 2 (without interaction term). The Pesaran CD test uses residuals from the fixed-effects specification. VIF for the interaction term is computed after mean-centering SCI and FD. All unit root tests include individual intercepts and trends with one lag selected by the Schwarz criterion.

| Test | Statistic | p-value | Decision |
|---|---|---|---|
| Hausman (FE vs. RE) | $\chi^2(8) = 42.31$ | 0.000 | Reject RE → Use FE |
| Pesaran CD (cross-sectional dependence) | CD = 4.28 | 0.000 | CSD present → Use DK SE |
| Modified Wald (groupwise heteroskedasticity) | $\chi^2(10) = 387.56$ | 0.000 | Heteroskedasticity present |
| Wooldridge AR(1) (serial correlation) | F(1, 9) = 18.73 | 0.002 | AR(1) present |
| Im–Pesaran–Shin unit root (GINI) | t-bar = −2.847 | 0.008 | Stationary |
| Im–Pesaran–Shin unit root (SCI) | t-bar = −4.215 | 0.000 | Stationary |
| Im–Pesaran–Shin unit root (FD) | t-bar = −2.563 | 0.021 | Stationary |
| VIF: SCI × FD (mean-centered) | 1.68 | — | Below threshold of 10 |

The Hausman test decisively rejects the random-effects specification ($\chi^2 = 42.31$, p = 0.000), confirming that unobserved country-specific heterogeneity is correlated with the regressors and that fixed effects are the appropriate estimation framework. The Pesaran CD statistic (CD = 4.28, p = 0.000) detects significant cross-sectional dependence among the ASEAN countries — an expected result given their dense trade, investment, and policy-coordination linkages through the ASEAN Economic Community — motivating the use of Driscoll–Kraay standard errors, which remain consistent under arbitrary forms of spatial and temporal dependence. The modified Wald test ($\chi^2 = 387.56$, p = 0.000) and Wooldridge test (F = 18.73, p = 0.002) confirm the presence of groupwise heteroskedasticity and first-order serial correlation, respectively, further supporting the Driscoll–Kraay correction, which simultaneously addresses heteroskedasticity, autocorrelation, and cross-sectional dependence — over conventional cluster-robust errors that handle only the first two. Im–Pesaran–Shin panel unit root tests reject the null of a unit root for all key variables at the 5 % significance level (GINI: t-bar = −2.847, p = 0.008; SCI: t-bar = −4.215, p = 0.000; FD: t-bar = −2.563, p = 0.021), confirming that the series are stationary in levels and the regressions do not suffer from spurious-correlation problems. Given stationarity, cointegration testing is unnecessary, and the fixed-effects estimates can be interpreted as short-to-medium-run relationships without concern about non-stationary residuals.

To address the mechanical multicollinearity between the interaction term and its constituent variables, mean-centering is applied to both SCI and FD prior to constructing the product term (Aiken & West, 1991). After centering, the VIF of the interaction term SCI_mc * FD_mc drops from 9.42 to 1.68, and all remaining variables record VIF values well below 10, confirming the absence of substantive multicollinearity. All regression results in Tables 6–8 are estimated using mean-centered variables; threshold values and marginal effects are expressed in the original metric for interpretive clarity, as centering leaves the interaction coefficient and threshold calculation algebraically invariant.

### Baseline Model Results: The Impact of Structural Change on Inequality

Table 5 presents the estimation results for Model 1 and Model 2, testing the direct impacts of structural change (SCI) and financial development (FD) on income inequality (GINI). The models are built sequentially: column (1) estimates the impact of SCI with control variables; column (2) adds the squared term of GDP per capita to test the Kuznets hypothesis; and column (3) introduces the financial development variable into the model. All estimations use the two-way fixed effects (TWFE) method with Driscoll–Kraay standard errors to ensure robustness against autocorrelation, heteroskedasticity, and cross-sectional dependence.


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**Table 5. Regression results of the direct impact on income inequality (baseline model).** Notes: Driscoll–Kraay standard errors in parentheses. Statistical significance levels: *** p<0.01, ** p<0.05, * p<0.1.

| *Dependent Variable: Gini Coefficient (GINI)* | | | |
|---|---|---|---|
| **Variable** | **(1)** | **(2)** | **(3)** |
| **SCI** | 0.351** | 0.384*** | **0.412***** |
| | (0.159) | (0.135) | **(0.128)** |
| **FD** | | | **0.025**** |
| | | | **(0.011)** |
| **lnGDPPC** | 0.873 | 5.142** | 4.987** |
| | (1.104) | (2.015) | (2.130) |
| **(lnGDPPC)²** | | -0.281** | **-0.305**** |
| | | (0.124) | **(0.131)** |
| **URB** | 0.048* | 0.051* | 0.045 |
| | (0.029) | (0.028) | (0.030) |
| **GEXP** | -0.215** | -0.233** | -0.208* |
| | (0.098) | (0.101) | (0.112) |
| **HC** | -1.567*** | -1.602*** | -1.814*** |
| | (0.451) | (0.433) | (0.502) |
| **FDI** | 0.031 | 0.029 | 0.017 |
| | (0.045) | (0.042) | (0.048) |
| **Constant** | 35.11*** | 18.24* | 21.05** |
| | (5.67) | (9.88) | (10.12) |
| **Observations** | 178 | 178 | 178 |
| **Number of countries** | 10 | 10 | 10 |
| **R² (within)** | 0.614 | 0.649 | 0.662 |
| **Fixed Effects** | Country & Year | Country & Year | Country & Year |

The results in Table 5 confirm a statistically significant and economically meaningful positive link between structural change and income inequality. In the most complete specification (column 3), a one-percentage-point acceleration in the structural change index is associated with a 0.412-point increase in the Gini coefficient ($p < 0.01$). Given that the sample-mean SCI is 2.87, this implies that the typical annual pace of sectoral reallocation contributes roughly 1.2 Gini points to observed inequality — a non-trivial magnitude relative to the sample standard deviation of 3.94. Economically, this is consistent with the Lewisian transition mechanism: as labor migrates from low-productivity agriculture to higher-productivity industry and services, the resulting inter-sectoral wage differential widens the income distribution before the modern sector absorbs a critical mass of workers.

The positive coefficient of FD (0.025, $p < 0.05$) in column (3) indicates that, on average, a ten-percentage-point increase in private-credit-to-GDP is associated with a 0.25-point rise in the Gini. This aligns with the Greenwood–Jovanovic (1990) prediction that, in economies where financial systems are still maturing, credit expansion tends to benefit asset-rich households and established firms disproportionately, before broader access diffuses the benefits to lower-income segments. In the ASEAN context, this pattern likely reflects a reality in which formal bank credit is still concentrated among urban enterprises and salaried workers, while smallholders, informal-sector participants, and rural populations remain largely unbanked.

The inverted-U relationship between GDP per capita and inequality — positive lnGDPPC (4.987) and negative $(\text{lnGDPPC})^2$ (−0.305), both significant at the 5 % level—validates the Kuznets hypothesis within the sample: inequality initially rises with income growth and subsequently declines once economies pass a turning point, estimated here in the vicinity of the sample mean. The strongly negative human-capital coefficient (−1.814, $p < 0.01$) carries direct policy relevance: each unit increase in the Penn World Table human-capital index is associated with a nearly two-point reduction in the Gini, confirming that investment in education narrows the skill premium that drives much of the income gap during structural transition. Government consumption expenditure (GEXP) exerts the expected inequality-reducing effect (−0.208, $p < 0.10$), reflecting the redistributive role of public spending through social transfers and public-service provision.

**Moderation Model Results: The Role of Financial Development**

Table 6 presents the estimation results for the interaction model (Model 3), where the interaction term SCI × FD is introduced to test this moderating effect. Column (4) presents the main estimation results using the Two-Way Fixed Effects (TWFE) method with Driscoll-Kraay standard errors. Columns (5) and (6) report robustness check results using the Panel-Corrected Standard Errors (PCSE) method and System GMM to ensure the reliability of the findings.

**Table 6. Regression results on the regulatory role of financial development.** Note: Standard errors in parentheses. Statistical significance levels: *** p<0.01, ** p<0.05, * p<0.1. Column (4) uses Driscoll–Kraay standard errors. Column (5) uses PCSE standard errors. Column (6) is a System GMM estimation with collapsed instruments; reports p-values for the Hansen test of instrument validity and the AR(2) test for autocorrelation.

| *Dependent variable: Gini coefficient (GINI)* | | | |
|---|---|---|---|
| **Variable** | **(4) TWFE-DK** | **(5) PCSE** | **(6) System GMM** |
| **SCI** | **0.875*** ** | **0.881*** ** | **0.314** ** |
| | (0.241) | (0.235) | (0.138) |
| **FD** | 0.018 | 0.015 | 0.006 |
| | (0.012) | (0.011) | (0.009) |
| **SCI × FD** | **-0.0092** ** | **-0.0094** ** | **-0.0045*** |
| | (0.0038) | (0.0035) | (0.0024) |
| **$GINI_{i,t-1}$** | | | **0.782*** ** |
| | | | (0.065) |
| **lnGDPPC** | 5.011** | 5.234** | 1.855* |
| | (2.188) | (2.205) | (1.091) |
| **$(lnGDPPC)^2$** | -0.310** | -0.321** | -0.112* |
| | (0.135) | (0.139) | (0.068) |
| **URB** | 0.041 | 0.038 | 0.015 |
| | (0.033) | (0.031) | (0.021) |
| **GEXP** | -0.211* | -0.205* | -0.088* |
| | (0.115) | (0.109) | (0.051) |
| **HC** | -1.790*** | -1.804*** | -0.653** |
| | (0.514) | (0.498) | (0.281) |
| **FDI** | 0.012 | 0.015 | 0.007 |
| | (0.049) | (0.047) | (0.025) |
| **Constant** | 22.84** | 21.98** | 5.12 |
| | (10.55) | (10.21) | (3.76) |
| **Observations** | 178 | 178 | 168 |
| **Number of countries** | 10 | 10 | 10 |
| **$R^2$ (within)** | 0.685 | - | - |
| **Fixed Effects** | Country & Year | Country & Year | Country |
| **Hansen J test (p-value)** | | | 0.285 |
| **AR(2) test (p-value)** | | | 0.411 |

The results from column (4) provide robust evidence for hypothesis H3. The interaction coefficient SCI × FD is −0.0092 ($p < 0.05$), confirming that financial development serves as a negative moderator: it attenuates the inequality-widening pressure generated by structural change. In economic terms, this means that for every additional percentage point of private credit relative to GDP, the marginal disequalizing effect of a given pace of structural change is reduced by 0.0092 Gini points. While this per-unit effect appears modest, its cumulative impact is substantial: moving from Cambodia's credit depth (≈ 4 % of GDP) to Thailand's (≈ 150 %) implies a reduction in the marginal effect of structural change by roughly 1.34 Gini points—more than offsetting the baseline effect entirely.

In the interaction framework, the constituent coefficients acquire conditional interpretations. The SCI coefficient (0.875, $p < 0.01$) now captures the marginal effect of structural change when the financial system is absent (FD = 0) — a theoretical limiting case that reveals the full disequalizing potential of sectoral reallocation without any financial intermediation to facilitate labor mobility, human-capital investment, or risk sharing. That this coefficient is more than double the unconditional estimate in Table 5 (0.412) illustrates how omitting the interaction term averages across heterogeneous country


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contexts, masking the severity of inequality pressures in financially underdeveloped settings. Meanwhile, the FD coefficient (0.018) loses statistical significance, implying that financial depth has negligible relevance for inequality in the absence of structural change (SCI = 0) — a static economy where no sectoral reallocation is taking place. This reinforces the central argument: the primary distributional role of finance in the ASEAN context is not direct but conditional, operating through its capacity to cushion the adjustment costs of economic transformation.

The marginal effect equation ∂GINI / ∂SCI = 0.875 − 0.0092 × FD yields a threshold of FD* = 0.875 / 0.0092 ≈ 95.1 % of GDP. Below this level, structural change continues to widen inequality, albeit with diminishing force as financial depth increases; above it, the financial system is sufficiently deep and inclusive to fully offset — and potentially reverse — the distributional costs of sectoral reallocation. The sample-mean FD of 62.38 % implies that the typical ASEAN economy during 2000–2019 was still well below this threshold: its marginal effect of SCI on Gini remained positive (0.875 − 0.0092 × 62.38 ≈ 0.30), though substantially attenuated relative to the 0.88 benchmark under zero financial development. Only Singapore, Malaysia, and Thailand in the later years of the sample period approached or exceeded the threshold, suggesting that for most ASEAN members, accelerating financial inclusion is a prerequisite for ensuring that structural transformation does not come at an unacceptable distributional cost.

To quantify the uncertainty surrounding this threshold estimate, a Monte Carlo simulation is conducted. Drawing 10,000 paired values of ($\beta_1$, $\beta_3$) from their joint normal distribution—parameterized by the point estimates (0.875, −0.0092), standard errors (0.241, 0.0038), and estimated covariance (−0.000320) from the variance–covariance matrix of Model 3 — the empirical distribution of FD* = $-\beta_1/\beta_3$ is constructed. The resulting 95 % confidence interval is [56.3, 183.7], with a median of 93.8 % of GDP. This wide interval confirms that the threshold is estimated with considerable imprecision, as is typical for ratios of correlated coefficients in moderately sized samples. The right-skewed distribution reflects the compounding uncertainty of both the numerator and denominator coefficients, particularly the sensitivity of the ratio to draws where $\beta_3$ approaches zero. The point estimate of 95.1 % should therefore be interpreted as an approximate policy benchmark rather than a precise critical value. Nevertheless, even the lower bound of the confidence interval (56.3 % of GDP) exceeds the private-credit depth of Cambodia (≈ 4 %), Myanmar (≈ 15 %), and Laos (≈ 25 %) throughout the sample period, reinforcing the qualitative conclusion that these economies remain far below the zone where financial development could plausibly neutralize the disequalizing effects of structural change. The upper bound (183.7 %) lies above the observed maximum in the sample (156.4 %), indicating that full certainty about threshold attainment is not achievable even for the region's most financially developed economies within the estimation period.

The robustness checks in columns (5) and (6) corroborate the main finding. The PCSE specification yields a virtually identical interaction coefficient (−0.0094, $p < 0.05$). The dynamic System GMM model in column (6), which explicitly accounts for inequality persistence through the lagged dependent variable (coefficient 0.782, $p < 0.01$), still produces a negative and significant interaction term (−0.0045, $p < 0.10$). The smaller magnitude in the GMM specification is expected, because much of the cross-sectional variation is absorbed by the lagged Gini; what matters is that the sign and significance are preserved. The Hansen test ($p = 0.285$) and AR(2) test ($p = 0.411$) confirm instrument validity and the absence of problematic serial correlation, respectively. The Driscoll–Kraay fixed-effects and PCSE estimators constitute the primary interpretive evidence, as both are methodologically appropriate for panels with small N and moderate T and yield virtually identical interaction coefficients (−0.0092 and −0.0094, respectively). The System GMM result in column (6) serves a complementary role: because the consistency properties of the Blundell–Bond estimator are asymptotic in N rather than T, the theoretical guarantees of the estimator do not strictly apply with only ten cross-sectional units, regardless of the instrument-reduction strategy employed. The collapsed instrument set mitigates instrument proliferation but does not restore large-N asymptotics, and the Hansen J statistic tends toward over-acceptance under weak identification in small-N settings (Roodman, 2009). The GMM estimates should therefore be interpreted as directional evidence — confirming that the sign and approximate magnitude of the moderating effect survive the introduction of dynamics and internal instrumentation—rather than as an independent confirmation of equal inferential standing. That all three approaches point to the same qualitative conclusion — a negative moderating role of financial development — provides reassurance that the finding is not an artifact of any single estimation strategy.

**Robustness Checks and Supplementary Analysis**

*Robustness Checks of the Model*

To test the reliability of the main results (presented in section 4.3), the study performs a series of robustness checks, with the results summarized in Table 7 below.

**Table 7. Robustness checks for the moderation model.** Notes: Driscoll–Kraay standard errors in parentheses. Significance levels: *** p<0.01, ** p<0.05, * p<0.1. All models include the full set of control variables as in Table 6. Column (2) uses the market income Gini as the dependent variable. Column (3) replaces SCI with the industry value-added share (INDVA). Column (4) replaces FD with the IMF's Financial Development Index (FDX); the number of observations is lower due to the unavailability of FDX data for some years. Column (5) uses one-period lagged values for all explanatory variables. Column (6) is an IV-2SLS estimation using second and third lags of endogenous variables as instruments; it reports the Kleibergen-Paap Wald F-statistic for the weak instrument test.

| ***Dependent Variable: Gini Coefficient*** | | | | | | |
|---|---|---|---|---|---|---|
| **Variable** | **(1) Baseline** | **(2) Market Gini** | **(3) Using INDVA** | **(4) Using FDX** | **(5) Lagged Vars.** | **(6) IV-2SLS** |
| **SCI** | **0.875***** | **0.951*** | | **0.913*** | | **1.024** |
| | (0.241) | (0.288) | | (0.255) | | (0.631) |
| **FD** | **0.018** | **0.024*** | **0.015** | | **0.016** | **0.028** |
| | (0.012) | (0.014) | (0.013) | | (0.011) | (0.020) |
| **SCI × FD** | **-0.0092** | **-0.0105** | | | | **-0.0116*** |
| | (0.0038) | (0.0045) | | | | (0.0065) |
| **INDVA** | | | **0.512*** | | | |
| | | | (0.189) | | | |
| **INDVA × FD** | | | **-0.0041** | | | |
| | | | (0.0019) | | | |
| **FDX** | | | | **0.204** | | |
| | | | | (0.158) | | |
| **SCI × FDX** | | | | **-1.152** | | |
| | | | | (0.491) | | |
| **SCI(t-1)** | | | | | **0.859*** | |
| | | | | | (0.230) | |
| **FD(t-1)** | | | | | **0.017** | |
| | | | | | (0.012) | |
| **SCI(t-1) × FD(t-1)** | | | | | **-0.0089** | |
| | | | | | (0.0037) | |
| Control variables | Yes | Yes | Yes | Yes | Yes | Yes |
| Fixed Effects | Country & Year | Country & Year | Country & Year | Country & Year | Country & Year | Country & Year |
| Observations | 178 | 178 | 178 | 165 | 168 | 168 |
| Number of countries | 10 | 10 | 10 | 10 | 10 | 10 |
| $R^2$ (within) | 0.685 | 0.653 | 0.679 | 0.691 | 0.688 | - |
| K-P F-stat | - | - | - | - | - | 14.88 |

Table 7 presents the results of six test models, with column (1) being the baseline model (from Table 6) for easy comparison. Column (2) replaces the dependent variable from disposable income Gini to market income Gini (gini_mkt). The results show that the interaction coefficient SCI × FD remains negative (-0.0105) and is statistically significant at the 5% level, indicating that the moderating role of finance persists even before state redistribution interventions.

Columns (3) and (4) test the sensitivity of the results to the measurement of the main variables. In column (3), the study replaces the composite structural change index (SCI) with a more specific measure, the industry value-added share (INDVA). The interaction coefficient INDVA × FD remains negative and significant, implying that financial development helps mitigate the inequality impact as the economy shifts towards industrialization. In column (4), the study uses the IMF's composite financial development index (FDX), a multidimensional measure covering the depth, access, and efficiency of the financial market. As FDX has a scale from 0 to 1 (instead of the percentage scale of FD), the interaction coefficient SCI × FDX has a significantly larger magnitude (-1.152) and remains highly statistically significant, thereby reinforcing the main result.

Columns (5) and (6) address concerns about endogeneity. Column (5) re-estimates the model with all explanatory variables lagged by one period to minimize the possibility of simultaneous reverse causality. The interaction coefficient SCI(t-1) × FD(t-1) remains negative (-0.0089) and statistically significant, consistent with the baseline result. Column (6) employs IV-2SLS estimation with second and third lags of SCI and FD as instruments. The Kleibergen–Paap F-statistic (14.88)


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exceeds the Stock–Yogo critical value of 10, indicating that the instruments are not weak. The interaction coefficient SCI × FD remains negative and significant at the 10 % level (−0.0116), preserving the qualitative conclusion of a moderating effect. However, an important nuance arises: the direct SCI coefficient (1.024) loses statistical significance ($p > 0.10$), while the interaction term retains it. Two non-mutually-exclusive explanations account for this pattern. First, IV estimation entails a substantial efficiency loss — standard errors roughly triple relative to the fixed-effects baseline — and with only 168 observations the power to detect the direct effect is considerably reduced; the point estimate itself (1.024) remains within the confidence interval of the fixed-effects estimate (0.875), suggesting that imprecision rather than a qualitatively different causal story drives the insignificance. Second, to the extent that OLS estimates of SCI are upward-biased by reverse causality — if, for example, high inequality constrains human-capital accumulation and thereby slows structural change, creating a downward bias, or if governments accelerate industrial policy in response to inequality, creating an upward bias — the IV correction may attenuate the estimated direct effect. Crucially, the survival of the interaction term's significance under instrumentation indicates that the moderating mechanism is robust to endogeneity concerns: even when the causal direct effect of structural change is imprecisely estimated, the conditioning role of financial development in shaping that effect remains detectable. This asymmetry is consistent with the theoretical framework, in which financial depth operates as a structural moderator whose effect is less susceptible to the short-run simultaneity that may contaminate the direct SCI–inequality relationship.

From a financial-economic standpoint, the persistence of the negative interaction coefficient across alternative specifications carries an important implication: the moderating role of finance is not an artifact of how structural change or financial development is measured. Whether one captures structural change as a composite index or as the industry value-added share, whether financial depth is proxied by private credit alone or by the IMF's multidimensional index encompassing market breadth and efficiency, the conclusion remains the same—deeper, more developed financial systems help contain the inequality costs of economic transformation. The IV-2SLS result (column 6) is particularly reassuring from a causal-inference perspective: even after instrumenting to address the possibility that inequality itself shapes the pace of structural change or financial development, the moderating effect survives, albeit with wider confidence intervals reflecting the efficiency cost of instrumentation.

Overall, the main result regarding the negative moderating role of financial development is robust to changes in variable measurement, estimation methods, and efforts to control for endogeneity.

### *In-depth Analysis: Sample Heterogeneity*

Next, the study examines whether the results are driven by certain specific countries or if there are differences between groups of countries with different development levels. First, a jackknife analysis, where the model is re-estimated 10 times, each time dropping one country, shows that the interaction coefficient SCI × FD consistently remains negative and statistically significant. This indicates that the results do not depend on any single country.

Table 8 delves deeper into two analyses: excluding high-income economies and splitting the sample by development level.

Column (2) re-estimates the model on a reduced sample of 8 countries, after excluding Singapore and Brunei, two high-income economies with very different financial and economic structures. The results show that the interaction coefficient remains negative (-0.0081) and statistically significant at the 5% level. This demonstrates that the study's findings are not solely a product of the region's most developed financial centers.

More importantly, columns (3) and (4) conduct an analysis on two subgroups: lower-middle-income economies (LMC: Cambodia, Laos, Myanmar, Vietnam) and upper-middle-income and higher economies (UMC: Indonesia, Malaysia, Philippines, Thailand, Brunei, Singapore). The results show a very clear difference.

In the LMC group (column 3), structural change exerts a strong inequality-increasing effect (coefficient 1.103, $p < 0.01$), while the interaction coefficient SCI × FD is not statistically significant. This pattern is consistent with the hypothesis that nascent financial systems lack the depth and breadth to buffer distributional shocks. However, these results must be interpreted with explicit caution: with only four countries, the fixed-effects estimator identifies parameters solely from within-country temporal variation, leaving virtually no cross-sectional degrees of freedom. The estimates are therefore exploratory in nature and should not be cited as definitive evidence of structural heterogeneity between income groups. They are best understood as suggestive of a pattern — that financial moderation requires a minimum institutional threshold — which warrants confirmation with larger cross-country samples that include other developing regions.

Conversely, in the UMC group (column 4), where financial systems are more developed, the moderating role of financial development is more pronounced. The interaction coefficient SCI × FD is -0.0128 and is highly statistically significant (at the 1% level). In these countries, a better-functioning financial system has been able to support labor mobility, promote

human capital accumulation, and allocate capital efficiently, thereby significantly reducing the inequality-inducing impact of structural change.

This analysis not only reinforces the main result but also clarifies an important mechanism: the moderating role of financial development is not an inherent characteristic but is shaped by and dependent on the maturity of the financial system itself.

This heterogeneity carries a sobering policy implication. The countries where financial moderation is most needed—because structural change exerts the strongest disequalizing pressure (SCI coefficient of 1.103 in the LMC group versus 0.798 in the UMC group) — are precisely those where the financial system is least capable of providing it. Cambodia, Laos, and Myanmar are undergoing rapid agricultural-to-industrial transitions, yet their private-credit ratios remain far below the 95.1 % threshold. Without concerted efforts to build inclusive financial infrastructure—mobile banking, agricultural credit programs, microinsurance — these economies risk a prolonged phase of rising inequality that could undermine social cohesion and political support for continued reform. Conversely, the strong moderating effect in the UMC group demonstrates what is achievable: countries like Thailand and Malaysia, where decades of financial-sector reform have broadened credit access, have managed to pursue structural transformation with markedly less distributional friction.

**Table 8. Sample sensitivity and heterogeneity analysis.** Notes: Driscoll–Kraay standard errors in parentheses. Significance levels: *** p<0.01, ** p<0.05, * p<0.1. All models include the full set of control variables and are estimated using the TWFE-DK method. Column (1) is the baseline model on the full sample. Column (2) excludes Singapore (SGP) and Brunei (BRN). Column (3) includes only Lower-Middle Income Countries (LMC): Cambodia, Laos, Myanmar, Vietnam. Column (4) includes Upper-Middle Income and High-Income Countries (UMC): Indonesia, Malaysia, Philippines, Thailand, Brunei, Singapore.

| ***Dependent Variable: Disposable Gini Coefficient*** | | | | |
|---|---|---|---|---|
| **Variable** | **(1) Full Sample** | **(2) Excl. SGP, BRN** | **(3) LMC Group** | **(4) UMC Group** |
| **SCI** | **0.875***** | **0.811***** | **1.103***** | **0.798**** |
| | (0.241) | (0.275) | (0.315) | (0.299) |
| **FD** | **0.018** | **0.014** | **0.031** | **0.016** |
| | (0.012) | (0.011) | (0.025) | (0.014) |
| **SCI × FD** | **-0.0092**** | **-0.0081**** | **-0.0075** | **-0.0128***** |
| | (0.0038) | (0.0039) | (0.0082) | (0.0041) |
| Control variables | Yes | Yes | Yes | Yes |
| Fixed Effects | Country & Year | Country & Year | Country & Year | Country & Year |
| Observations | 178 | 145 | 69 | 109 |
| Number of countries | 10 | 8 | 4 | 6 |
| $R^2$ (within) | 0.685 | 0.692 | 0.715 | 0.681 |

## DISCUSSION

### The Dual Paradox and the Moderating Role of Financial Development

The findings reveal what may be characterized as a "conditional duality" in the distributional role of financial development. In the baseline specification without the interaction term (Table 5, column 3), FD carries a positive and statistically significant coefficient (0.025, $p < 0.05$), suggesting that credit expansion is, on average, associated with wider inequality — consistent with the Greenwood and Jovanovic (1990) prediction that early-stage financial deepening disproportionately benefits asset-rich households. However, once the interaction term is introduced in the full moderation model (Table 6, column 4), the direct FD coefficient becomes statistically insignificant (0.018, $p > 0.10$). This attenuation indicates that the apparent direct effect in the baseline model partially reflects the omitted moderating channel: some of what appears as a direct inequality-increasing association of finance is in fact finance's interaction with ongoing structural change. The instability of the direct coefficient across specifications warrants interpretive caution; the study does not claim that finance unconditionally increases inequality but rather that its net distributional effect is context-dependent. What remains robust across all specifications is the negative interaction term, establishing that the primary distributional relevance of financial depth in ASEAN operates through its capacity to moderate the inequality costs of structural transformation. This phenomenon is particularly relevant in the context of many ASEAN countries, where the process of financial deepening (increasing credit volume) often outpaces comprehensive financial development (expanding access for all).


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On the other hand, and this is the central finding of the paper, when repositioned from a direct explanatory variable to a moderating variable, financial development exhibits a completely different role. The negative and highly statistically significant interaction coefficient ($\beta_3$ = -0.0092) in the moderation model (Table 6) shows that financial development mitigates the inequality-inducing effect of structural change. This finding supports the theoretical framework the study has built, in which a developed financial system acts as an institutional "buffer", helping the economy absorb distributional shocks from the restructuring process. It helps ease labor mobility barriers, supports human capital accumulation, and allocates capital more efficiently, as suggested in the works of Galor & Zeira (1993) and Claessens & Perotti (2007).

This study moves beyond the traditional binary debate on whether financial development is "good" or "bad" for equality. The evidence suggests that the distributional effect of finance is not monolithic but conditional: it depends on whether the economy is undergoing active structural transformation and whether the financial system is deep and inclusive enough to channel resources toward affected populations. The robust finding is not a stable direct effect in either direction, but rather the moderating channel — through which financial depth mitigates the inequality costs of sectoral reallocation. Ignoring this conditional role, as much of the prior literature has done, risks drawing incomplete conclusions about the distributional implications of financial-sector development in transforming economies.

### Marginal Effect Analysis and the "Financial Threshold Map" for ASEAN

The practical significance of the moderating role is most clearly demonstrated through the analysis of the marginal effect. The equation $\partial GINI / \partial SCI = 0.875 - 0.0092 \times FD$ shows that the impact of structural change is not a constant, but varies depending on the depth of the financial system. The coefficient $\beta_1$ = 0.875 carries a profound economic meaning: in a hypothetical scenario with no financial system (FD=0), each percentage point increase in the structural change index would increase the Gini coefficient by nearly 0.88 points. This figure shows the intrinsic power of the restructuring process to create income differentiation without a support mechanism.

More importantly, this analysis allows for the identification of a financial development threshold (FD) of 95.1% of GDP*. This is not just a technical number, but a powerful policy diagnostic tool. It implies that an ASEAN country needs to reach a level of private sector credit of about 95.1% of GDP to completely neutralize the inequality-increasing impact of structural change. Comparing this with reality, the average FD in the research sample is 62.38%, indicating that a typical ASEAN country during the 2000-2019 period was still below this threshold. Their financial systems were able to mitigate, but not strong enough to completely eliminate, the negative impact. For example, in a country with FD = 62.38%, the marginal effect of SCI on Gini is still positive (0.875 - 0.0092 * 62.38 ≈ 0.30). However, this impact has been significantly reduced compared to the 0.88 level without finance. Only economies with the most extensive financial systems like Singapore, Malaysia, and Thailand (in the later years of the period) were able to surpass this threshold, where structural change could even be accompanied by an improvement in equality. Identifying each country's position relative to this threshold has transformed an abstract econometric result into a tangible policy evaluation framework.

### The Role of Finance Depends on the Level of Development

Further analysis on country subgroups (Table 8) provided additional evidence for the moderating mechanism. The results show a clear difference: the moderating role of financial development is almost non-existent in the group of lower-middle-income countries (LMC), but is very strong in the group of upper-middle-income countries and above (UMC).

In LMC countries (Cambodia, Laos, Myanmar, Vietnam), where financial systems are nascent, capital markets are undeveloped, and financial inclusion is limited, the system is not yet "mature" enough to perform the "buffer" role. Although the process of structural change is strong and puts great pressure on inequality (the SCI coefficient is 1.103, very high), the financial system cannot efficiently allocate capital to affected workers or small businesses in new sectors. Therefore, the interaction coefficient is not statistically significant.

Conversely, in UMC countries (Indonesia, Malaysia, Philippines, Thailand, etc.), where the financial system has reached a certain level of depth and efficiency, it can exert its moderating role. The negative and highly statistically significant interaction coefficient (-0.0128) shows that here, a better-functioning financial system has been able to support labor mobility, promote human capital accumulation, and allocate capital efficiently, thereby significantly reducing the inequality-inducing impact of structural change. This analysis confirms that the moderating role of financial development is not an inherent characteristic, but is shaped by and dependent on the maturity level of the financial system itself.

### Relevance of the Findings to the Post-2019 Context

A legitimate concern is whether results derived from the 2000–2019 period retain explanatory power amid the turbulence that has characterized the subsequent years. Three observations suggest they do. First, the moderating mechanism this

study identifies operates through institutional channels — credit accessibility, human-capital financing, and risk-sharing capacity — that evolve slowly and persist across business-cycle fluctuations. The COVID-19 crisis did not fundamentally alter ASEAN's financial architectures; rather, it exposed existing disparities in financial inclusion: economies with broader digital-banking penetration and deeper microfinance networks (e.g., Thailand, Malaysia) mounted faster, more targeted relief efforts than those still reliant on cash-based economies (e.g., Myanmar, Laos). This pattern is fully consistent with the buffering role documented in Tables 6 and 8. Second, the geopolitical realignment — particularly trade diversion from China and accelerated foreign-direct-investment relocation into Vietnam, Indonesia, and the Philippines — has intensified structural change in the region, increasing the urgency of the policy prescription derived here: that financial inclusion must accompany industrial upgrading. Third, the threshold concept (FD* ≈ 95.1 % of GDP) offers a forward-looking benchmark; policymakers can assess whether their country has closed or widened its distance to that threshold during the pandemic period, using the pre-shock estimate as a reference. In sum, while extending the sample to include post-2020 data remains a priority for future research, the structural character of the estimated relationships and the direction of recent economic developments in ASEAN both support the continued relevance of the study's core conclusions.

### Contributions of the Study

This study makes significant contributions in several aspects:

Theoretically, the study has built an integrated analytical framework, connecting two previously independent research streams on structural change–inequality and financial development–inequality. By repositioning financial development as a moderating variable, the paper has clarified the "dual paradox", showing that the impact of finance on inequality is not monolithic but depends on the economic structural context.

Empirically, the study provides evidence of the moderating role of financial development in the context of the diverse economic development levels of ASEAN countries. The identification of a "financial threshold map" and the analysis of heterogeneity among country groups have provided deep and context-relevant insights, going beyond average impact analyses on the full sample.

In terms of policy implications, the research results emphasize the need to distinguish between financial deepening and financial inclusion. Merely expanding credit volume (deepening) can exacerbate inequality (the direct channel). To leverage the positive moderating role (the indirect channel), policies should focus on financial inclusion: expanding access to credit, insurance, and other financial services for groups vulnerable to the structural change process, such as agricultural workers, micro-enterprises, and rural areas. This is a particularly important implication for countries in a strong transition phase like Vietnam, Cambodia, and Myanmar, to ensure that the processes of financial development and structural change complement each other, aiming for sustainable and inclusive growth.

## CONCLUSIONS

This study provides empirical evidence that the relationship between economic structural change, financial development, and income inequality in ASEAN is a complex, non-linear mechanism. The core results suggest that the process of economic structural change itself is a factor that increases inequality in the region's current stage of development. More importantly, the paper has shed light on the "dual paradox" of financial development: while its direct impact on inequality may not be positive, its moderating role is decisive.

The central finding shows that a more developed financial system can significantly mitigate the inequality-inducing impact of structural change. However, this moderating role is not a constant but depends on the development level of the financial system, as demonstrated by the existence of a financial development threshold. It only becomes effective in countries with a more solid financial foundation. By repositioning financial development from an independent explanatory variable to a moderating variable, this study has opened a new perspective, helping to reconcile seemingly contradictory results in previous studies and providing a more robust explanation for the complex interaction between economic transition processes and income distribution outcomes.

Although the estimation window ends in 2019, the structural and institutional nature of the identified mechanisms suggests that the core findings remain pertinent in the post-pandemic landscape. The COVID-19 crisis and ongoing geopolitical realignments have, if anything, accelerated structural change in ASEAN — through supply-chain diversification, digital transformation, and green-energy transitions — while simultaneously exposing the fragility of financial systems that lack inclusive reach. These developments amplify, rather than invalidate, the central policy message: building financial depth without financial breadth leaves the distributional costs of economic transformation unmitigated.


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### Policy Implications

From the research findings, three key groups of policy implications are drawn for ASEAN countries.

First, policymakers need to proactively manage the social impacts of the structural change process. Structural change is a necessary driver for growth, but the research results show that it comes with distributional costs. Therefore, industrial and structural policies must be accompanied by social safety net measures, such as skill retraining programs, unemployment insurance, and mobility support for the labor force in contracting sectors, especially agriculture.

Second, financial development strategy must go beyond the goal of increasing scale (depth) to aim for comprehensiveness and inclusion. The results show that simply expanding credit may not help reduce, and may even increase, inequality. Instead, the top priority is to ensure access to financial services (credit, savings, insurance) for low-income groups, small and medium-sized enterprises, and rural areas that are easily left behind in the transition process. Promoting microfinance, digital banking, and financial literacy programs are effective tools to realize this goal.

Third, there needs to be close and synchronized coordination between structural policy and financial policy. For countries in a strong phase of structural change like Cambodia, Vietnam, and Laos, developing an inclusive financial system is not a secondary or subsequent goal, but must be a parallel strategic priority. Ensuring that the financial system is strong enough to play a "buffer" role will help the benefits of the economic transition be distributed more equitably, creating a foundation for sustainable and inclusive development.

### Limitations and Future Research Directions

Although the research objectives were achieved, several limitations should be acknowledged. First, the small cross-sectional dimension (N = 10) has direct methodological consequences. The asymptotic consistency of the System GMM estimator relies on $N \to \infty$ rather than $T \to \infty$; with only ten countries, the theoretical guarantees of the estimator do not formally hold, the Hansen J-statistic becomes unreliable as a specification test, and the GMM results in this study serve as directional corroboration rather than definitive causal evidence. The Driscoll–Kraay and PCSE fixed-effects estimators, which do not require large-N asymptotics, provide the primary inferential basis. Similarly, the sub-sample analysis for the LMC group (N = 4) is constrained by minimal cross-sectional variation and should be regarded as exploratory. Second, the sample covers 2000–2019 and therefore does not directly capture the distributional consequences of the COVID-19 pandemic, the Russia–Ukraine military conflict, or the intensification of great-power rivalry in the Indo-Pacific. While the structural character of the estimated relationships and the direction of post-2019 developments support the continued validity of the findings, extending the analysis to incorporate post-pandemic data, once reliable standardized inequality estimates become available, is an important priority. Third, measuring structural change through three broad economic sectors may conceal significant intra-sectoral dynamics — for instance, the shift from low-value garment manufacturing to electronics assembly within the industrial sector. Finally, despite the application of fixed effects, lagged variables, and instrumental-variable estimation, residual endogeneity arising from unobserved institutional heterogeneity cannot be entirely excluded.

## ADDITIONAL INFORMATION

### AUTHOR CONTRIBUTIONS

*All authors have contributed equally.*

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*The Authors declare that there is no conflict of interest.*

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*Нгуєн Т. Д.*

## СТРУКТУРНІ ЗМІНИ В ЕКОНОМІЦІ ТА НЕРІВНІСТЬ У ДОХОДАХ У КРАЇНАХ АСЕАН: МОДЕРУЮЧА РОЛЬ ФІНАНСОВОГО РОЗВИТКУ

Це дослідження переосмислює роль фінансового розвитку, розглядаючи його не як прямий визначальний фактор, а як модераторну змінну, що формує вплив економічних структурних змін на нерівність доходів у країнах АСЕАН. На основі незбалансованої панелі даних усіх десяти держав-членів за період 2000–2019 років (N = 178) аналіз використовує двофакторну модель із фіксованими ефектами та інтерактивним членом, доповнену системним методом узагальнених моментів (System GMM), панельними скоригованими стандартними помилками (PCSE), оцінкою за допомогою інструментальних змінних і специфікаціями з лаговими змінними для розв'язання проблеми ендогенності. Хоча вибірковий період передує пандемії COVID-19 і подальшим геополітичним потрясінням, досліджувані механізми — нерівність, зумовлена розривом у продуктивності внаслідок секторального перерозподілу, та її пом'якшення завдяки глибині фінансової системи — мають структурний та інституційний характер; отже, їх політична актуальність виходить за межі періоду оцінки. Основні результати показують, що структурні зміни самі собою значно посилюють нерівність у доходах ($\beta_1 = 0{,}875$, $p < 0{,}01$), однак фінансовий розвиток справляє протилежний пом'якшувальний ефект: коефіцієнт взаємодії (SCI × FD) становить -0,0092 ($p < 0{,}05$), що вказує на те, що глибші фінансові системи послаблюють тиск секторальної трансформації, який призводить до нерівності. Визначено порогове значення на рівні приблизно 95,1% ВВП в приватному кредитуванні, за межами якого ефект структурних змін, що

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посилює нерівність, повністю нейтралізується. Аналіз підвибірок показує, що цей модераторний механізм надійно працює лише в економіках АСЕАН із рівнем доходу, вищим за середній і високим; він залишається статистично незначущим поміж країн-членів із рівнем доходу, нижчим за середній, чиї фінансові системи не мають достатньої глибини та інклюзивності для амортизації розподільчих шоків. Висвітлюючи «подвійний парадокс» — фінанси безпосередньо пов'язані з більшою нерівністю, але опосередковано пом'якшують її через модерацію, — дослідження надає практичні політичні рекомендації: країни АСЕАН, що прагнуть до структурної трансформації, повинні одночасно будувати інклюзивні фінансові системи, надаючи пріоритет широкому доступові, а не простому розширенню кредитування, щоб забезпечити справедливий розподіл вигід від економічного переходу.

**Ключові слова:** АСЕАН, структурні зміни в економіці, нерівність у доходах, фінансовий розвиток, фінансова інклюзія, модеруючий ефект, порогова аналітика, панельні дані, крива Кузнеця

**JEL Класифікація:** O15, O16, O47, G21


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