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Editorial Office

Faculty of Economics, Universitas Islam Indonesia
Condong Catur, Depok, Sleman, Yogyakarta, Indonesia, 55283
Phone: +6274-881546, 883087, Fax +6274-882589
E-mail: editor.ejem@uii.ac.id

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Global economic policy uncertainty and fiscal responses in emerging markets: A dynamic error-correction panel

Ali Alsubaie

Department of Administration Sciences, King Fahad Security College, Riyadh, Saudi Arabia
Corresponding email: aalsubaie@hotmail.com; subaiea@kfsc.edu.sa

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Abstract

Purpose — This study examines how global economic policy uncertainty (GEPU) shapes government expenditure dynamics in emerging market economies.

Methods — Using an annual panel of 28 emerging economies from 1998 to 2023, this study analyzes short-run fiscal adjustments and long-run equilibrium relationships between global uncertainty and public expenditure. To address challenges arising from mixed integration orders and cross-sectional dependence driven by global shocks, it employs a multistage empirical strategy that combines fixed-effects estimation, Driscoll–Kraay robust inference, and a cross-sectionally augmented dynamic error-correction model (ECM).

Findings — The results provide robust evidence that higher GEPU is associated with higher government expenditure as a share of Gross Domestic Product (GDP). This relationship holds across alternative specifications and persists in the long run, indicating that fiscal responses to uncertainty are not purely transitory. Dynamic estimates reveal a statistically significant error-correction mechanism, confirming a stable long-term relationship among government expenditure, global uncertainty, and domestic economic conditions. Structural factors, particularly urbanization, further shape fiscal outcomes, whereas income per capita enters with a negative sign, though its effect is not consistently statistically significant across specifications.

Implication — The findings have important implications for fiscal sustainability and policy design in an increasingly uncertain global environment.

Originality — By explicitly accounting for non-stationarity and unobserved common global factors, this study contributes to the literature by providing new evidence of how emerging market governments respond to global risks.

Keywords — Global Economic Policy Uncertainty, Government Expenditure, Emerging Markets, Fiscal Policy, Error-Correction Model.

Introduction

Heightened global economic policy uncertainty (GEPU) has become an increasingly salient feature of the international economic landscape. Episodes such as the global financial crisis, trade policy tensions, the COVID-19 pandemic, and recent geopolitical disruptions underscore how uncertainty originating from major economies can spill over into emerging markets. These shocks affect investment, consumption, and capital flows, thereby exerting pressure on domestic

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macroeconomic stability and complicating fiscal policy implementation. In emerging market economies, global uncertainty creates a particularly complex policy environment. On the one hand, heightened uncertainty depresses private demand and amplifies downside risks, potentially necessitating countercyclical fiscal interventions. On the other hand, it can erode fiscal space by weakening revenue bases, increasing borrowing costs, and exposing public finances to external financing constraints. Therefore, understanding how governments balance these competing forces is central to assessing fiscal resilience and sustainability in emerging markets.

Despite a rapidly expanding literature on economic policy uncertainty (EPU) and macroeconomic outcomes, empirical evidence on fiscal responses to global uncertainty remains limited and methodologically fragmented. Much of the existing literature relies on static panel regressions that abstract from non-stationarity, cross-sectional dependence, and long-run dynamics. Given the global nature of uncertainty shocks and the persistence of fiscal variables, such approaches risk producing spurious or incomplete inferences. Moreover, little is known about whether fiscal responses to uncertainty reflect short-run stabilization efforts or deeper long-run adjustments in government expenditure behavior.

Extensive empirical literature has examined the cyclical behavior of fiscal policy, documenting systematic differences between advanced and developing economies. Gavin and Perotti (1997) show that fiscal policy in Latin America has historically been procyclical, with government expenditure rising during economic expansions and contracting during downturns. This pattern contrasts sharply with the countercyclical fiscal behavior typically observed in advanced economies. Subsequent studies attribute this procyclicality to multiple structural and institutional factors. Alesina et al. (1995) highlight political economic constraints and fiscal rigidities that limit governments' ability to smooth expenditure over the business cycle. Extending this line of reasoning, Talvi and Végh (2005) argue that volatility in tax bases and revenue collection amplifies fiscal procyclicality, particularly in economies with weak automatic stabilizers. More recent evidence indicates that fiscal behavior is shaped by both macroeconomic conditions and institutional quality. Alesina et al. (2008) show that weak institutions, credit constraints, and political distortions systematically bias fiscal policies toward procyclicality in emerging markets, while Woo (2009) finds that political polarization exacerbates fiscal instability and increases the sensitivity of government expenditure to economic shocks.

In parallel with the fiscal policy literature, a growing body of research has examined the macroeconomic consequences of economic policy uncertainty. The seminal contribution by Baker et al. (2016) introduced a systematic measure of EPU and demonstrated its strong association with declines in output, investment, and employment. The theoretical foundations of uncertainty effects trace to Bernanke (1983), who highlights the role of irreversibility and option value in delaying economic decisions under uncertainty. Building on this insight, Bloom (2009) formalizes uncertainty shocks as drivers of business cycle fluctuations, showing that heightened uncertainty leads to sharp, temporary contractions in economic activity. Microeconomic evidence further reinforces these findings. Bloom et al. (2007) show that firm-level investment and hiring decisions are highly sensitive to uncertainty, thereby providing micro foundations for aggregate effects. Complementing these approaches, Jurado et al. (2015) developed an alternative measure capturing latent macroeconomic uncertainty and confirmed its predictive power for economic activity. While this literature convincingly establishes uncertainty as a macroeconomic force, it primarily focuses on private-sector outcomes, leaving the fiscal response to uncertainty comparatively underexplored.

An expanding literature has begun to examine the interaction between uncertainty and fiscal policy. Using structural models, Born and Pfeifer (2014) show that policy uncertainty affects macroeconomic dynamics through expectations and fiscal transmission channels, suggesting that fiscal authorities may respond strategically to uncertainty to stabilize economic outcomes. Empirical studies further highlight nonlinear and state-dependent effects of uncertainty. Caggiano et al. (2014) demonstrate that uncertainty shocks have asymmetric effects across business-cycle regimes, implying that fiscal responses may differ between recessions and expansions. Methodologically, Ramey (2011) emphasizes the importance of treating government expenditure as an endogenous policy response rather than as an exogenous driver of economic activity.

Similarly, Ilzetzki et al. (2013) show that fiscal multipliers vary substantially across countries, with particularly strong effects in emerging economies.

Emerging economies are particularly vulnerable to global shocks because of their trade openness, financial integration, and limited policy buffers. Rodrik (1998) argues that larger governments can serve as a form of social insurance against external volatility, implying a positive relationship between exposure to global shocks and public spending. More recently, Eichengreen et al. (2008) document that global financial and policy shocks transmit unevenly across emerging markets, producing synchronized but heterogeneous macroeconomic responses. These spillovers suggest that fiscal outcomes across countries are unlikely to be independent, particularly in periods of heightened global uncertainty.

Macroeconomic panel data also raises several econometric challenges, including persistence, non-stationarity, and cross-sectional dependence. Im et al. (2003) propose panel unit root tests that accommodate heterogeneity across countries, while Pesaran (2004) develops formal tests for cross-sectional dependence in panel models. To address unobserved common factors, Pesaran (2006) introduces the Common Correlated Effects (CCE) estimator, which augments regression models using cross-sectional averages and has become standard in panels subject to global shocks and spillovers. Dynamic extensions, grounded in Pesaran et al. (1999), explicitly model long-run relationships through error-correction mechanisms. For inference, Driscoll and Kraay (1998) propose covariance estimators robust to heteroskedasticity, serial correlation, and cross-sectional dependence, while Nickell (1981) highlights biases that arise in dynamic fixed-effects panels.

Against this background, the present study examines how GEPU shapes government expenditure dynamics in emerging market economies. Using an annual panel of 28 emerging economies from 1998 to 2023, the analysis employs a comprehensive econometric framework that captures both short-run fiscal dynamics and long-run equilibrium relationships while accounting for unobserved global common factors. The empirical strategy combines fixed-effects estimation, robust inference under cross-sectional dependence, and a dynamic error-correction model augmented with cross-sectional averages.

This study contributes to recent literature in several ways. First, it explicitly models government expenditure as an endogenous response to global uncertainty rather than focusing solely on output or private-sector outcomes. Second, it concentrates on emerging markets, where fiscal constraints and exposure to global shocks are particularly pronounced. Third, by employing modern panel techniques that address non-stationarity and global spillovers, the study yields more credible inferences than much of the existing literature that relies on static specifications.

Methods

Data Sources and Sample

We adopted an unbalanced annual panel of 28 emerging economies from 1998 to 2023. The baseline regressions used 672 observations (primarily limited by the availability of financial development data), while other variables had up to 728 observations. The sample composition was determined by the availability of consistent data on government expenditure, macroeconomic indicators, and structural characteristics. The focus on emerging markets is motivated by their increased exposure to global shocks, more constrained fiscal space, and the historically procyclical fiscal behavior documented in the literature.

Government expenditure, gross domestic product (GDP) per capita, inflation, population, trade openness, financial development, and urbanization data were obtained primarily from the World Bank's World Development Indicators (WDI) and related international databases. These sources ensure cross-country comparability and consistency over time. The key explanatory variable, global economic policy uncertainty (GEPU), is drawn from the Global Economic Policy Uncertainty (GEPU) index developed by Baker et al. (2016). This index aggregates policy uncertainty across major economies and captures uncertainty in fiscal, monetary, and regulatory policies. Because the GEPU index reflects global conditions rather than country-specific developments, it is treated as a common external shock that affects all countries in the sample simultaneously. After merging all series and accounting for missing observations, the final dataset contained between 672 and 728 observations, depending on the availability of variables, as shown in the descriptive statistics.

Dependent Variable

The dependent variable in all empirical specifications is government expenditure as a percentage of GDP. This measure captures the relative size of the public sector and is widely used in fiscal policy literature to assess government activity across countries and over time. Expressing government spending relative to GDP facilitates cross-country comparability by scaling expenditure by economic size and helps isolate fiscal responses from pure output fluctuations. This variable varies significantly across countries and over time, reflecting differences in fiscal capacity, institutional development, and macroeconomic conditions.

Key Explanatory Variable: Global Economic Policy Uncertainty

The central explanatory variable is the GEPU index, which captures uncertainty about economic policy decisions in major economies and is constructed from newspaper-based frequency counts, tax code expiration measures, and professional forecaster disagreements.

In the empirical analysis, the GEPU index is mean-centered to facilitate interpretation while preserving its original variation. Accordingly, the GEPU coefficient can be interpreted as the change in government expenditure (in percentage points of GDP) associated with a one-unit increase in the centered global uncertainty index. Because GEPU is common across countries at any given time, it serves as a proxy for global shocks that are plausibly exogenous to individual emerging economies.

Macroeconomic and Structural Control Variables

To isolate the fiscal response to global uncertainty, this study includes a set of macroeconomic and structural control variables commonly used in fiscal policy research. Income per capita, measured as the logarithm of real GDP per capita, controls differences in economic development and fiscal capacity across countries. Higher-income economies may have more diversified revenue bases and different expenditure priorities, which can influence government expenditure behavior. Inflation, measured as the annual percentage change in the consumer price index, reflects macroeconomic instability and nominal pressures that may constrain real government expenditure or prompt fiscal adjustments.

Population size, measured logarithmically, controls for scale effects and demographic pressures on public expenditure. Larger populations may require higher absolute expenditure but lower expenditure ratios because of economies of scale. Trade openness, defined as the sum of exports and imports relative to GDP, reflects exposure to external shocks and global integration. More open economies may be more vulnerable to global uncertainty, potentially amplifying fiscal responses. Financial development is proxied by a standardized index based on domestic credit measures. Greater financial development may ease borrowing constraints and allow governments to smooth expenditure in response to shocks. Urbanization, measured as the share of the population living in urban areas, reflects structural demand for public services, infrastructure, and social spending. All continuous control variables, except inflation and population, are standardized to improve numerical stability and facilitate the comparison of coefficient magnitudes across regressors.

Interaction Term

To examine whether the fiscal response to global uncertainty varies with the level of economic development, the baseline specification includes an interaction term between global uncertainty and GDP per capita, allowing the marginal effect of uncertainty on government expenditure to vary across countries at different income levels, reflecting potential heterogeneity in fiscal capacity, access to credit markets, and institutional quality.

Empirical Strategy

Using an unbalanced annual panel of 28 emerging economies covering 1998 to 2023, this study examines the fiscal response of emerging-market governments to GEPU. The empirical strategy is designed to address three key challenges inherent in macroeconomic panel data: unobserved country-specific heterogeneity, cross-sectional dependence induced by global shocks, and mixed orders of integration across variables.

To address these challenges, the study was conducted in three stages. First, we estimated a baseline fiscal reaction function with country fixed effects to capture time-invariant heterogeneity across economies. Second, we assessed the robustness of the baseline estimates using Driscoll–Kraay standard errors, which are robust to heteroskedasticity, serial correlation, and cross-sectional dependence. Third, we adopted a dynamic error-correction framework augmented with cross-sectional averages to distinguish short-run fiscal dynamics from long-run equilibrium relationships, while explicitly accounting for unobserved global common factors.

This study relies exclusively on secondary, publicly available macroeconomic data obtained from international databases. The study did not involve human participants; therefore, no confidential information or ethical approval was required. The empirical analysis follows standard econometric reporting practices in applied macroeconomic research, ensuring transparency, reproducibility, and methodological rigor.

Baseline Fixed-Effects Model

The baseline specification models government expenditure as a function of GEPU, domestic macroeconomic conditions, and structural characteristics. The estimated equation is as follows:

$$G_{it} = \alpha_i + \beta_1 GEPU_t + \beta_2 GDPpc_{it} + \beta_3 (GEPU_t \times GDPpc_{it}) + \gamma' X_{it} + \varepsilon_{it} \quad (1)$$

where (G_{it}) denotes government expenditure as a percentage of GDP in country (i) at time (t), (α_i) captures country-specific fixed effects, ($GEPU_t$) denotes the GEPU index, and ($GDPpc_{it}$) represents real GDP per capita. The vector (X_{it}) includes inflation, population size (in logarithms), trade openness, financial development, and urbanization.

Year fixed effects are excluded from Equation (1). Because global economic policy uncertainty varies only over time and is common across countries, including time dummies would mechanically absorb the variation in ($GEPU_t$), thereby leaving its coefficient unidentified. Identification, therefore, relies on common global variation in uncertainty over time, combined with cross-country heterogeneity in fiscal responses, while country fixed effects absorb all time-invariant differences across economies.

Equation (1) is estimated using the fixed-effects estimator with standard errors clustered at the country level. Inference is based on country-clustered standard errors, with Driscoll–Kraay standard errors used as a robustness check to account for serial correlation and cross-sectional dependence.

Robust Inference under Cross-Sectional Dependence

To ensure valid statistical inference in the presence of cross-sectional dependence and serial correlation, Equation (1) is re-estimated using Driscoll–Kraay standard errors. This approach preserves the fixed effects point estimates while adjusting the covariance matrix to be robust to general forms of spatial and temporal dependence. Table 4 presents the Driscoll–Kraay results.

Dynamic Error-Correction Model with Cross-Sectional Averages

To distinguish short-run fiscal adjustments from long-run equilibrium relationships and explicitly control for unobserved global common factors, we employ a dynamic error-correction specification augmented with cross-sectional averages, consistent with the CCE framework.
The estimated model is:

$$\Delta G_{it} = \phi G_{it-1} + \theta_1 GEPU_{t-1} + \theta_2 GDPpc_{it-1} + \theta_3 ln(Population_{it-1}) + \delta_1 \Delta GEPU_t + \delta_2 \Delta GDPpc_{it} + \delta_3 \Delta Inflation_{it} + \eta_1 \overline{G}_{t-1} + \eta_2 \overline{GDPpc}_{t-1} + \kappa_1 \Delta \overline{G}_t + \kappa_2 \Delta \overline{GDPpc}_t + \alpha_i + u_{it} \quad (2)$$

where bars denote cross-sectional averages that capture unobserved common factors, and ($\phi < 0$) measures the speed of adjustment toward the long-run equilibrium. Equation (2) is estimated using fixed effects and country-clustered standard errors.

Results and Discussion

Descriptive Statistics and Preliminary Evidence

Table 1 presents descriptive statistics for all variables. Government expenditure averages approximately 13% of GDP across the sample, with considerable cross-country and temporal variation. GEPU shows substantial volatility, reflecting major episodes such as the global financial crisis and the COVID-19 pandemic. Government expenditure averages approximately 13% of GDP, with substantial cross-country and time-series dispersion. Global uncertainty is highly volatile, consistent with major episodes such as the global financial crisis and the COVID-19 pandemic. Descriptive statistics are reported for the maximum available sample by variable, whereas regression estimates use the common sample implied by each specification (e.g., 672 observations in the baseline model owing to the availability of financial development data). Macroeconomic and structural variables also vary substantially, underscoring the heterogeneity of emerging markets and the importance of controlling for country-specific characteristics in empirical analysis.

Table 1. Descriptive Statistics

Variables	Mean	Std. Dev.	Min	Max	Obs.
Government spending (% GDP)	13.31	4.40	4.85	28.91	725
Global EPU (GEPU)	0.00	64.76	−76.16	170.95	728
Log GDP per capita	0.00	1.09	−2.73	2.56	728
Inflation (%)	6.94	8.55	−2.50	84.60	728
Log population	3.74	1.26	1.33	7.27	728
Trade openness	0.00	1.00	−1.62	4.20	728
Financial development	0.00	1.00	−1.77	2.90	672
Urbanization	0.00	1.00	−1.96	1.69	728

Notes: Control variables reported in standardized form have a mean of zero and unit variance. GEPU is mean-centered but retains its original scale.

Table 2. Panel Unit Root Tests (Levels) – IPS

Variable	W-t-bar	p-value	Order
Government expenditure	−1.42	0.077	I(0)
GEPU	0.03	0.512	I(1)
Log GDP per capita	−0.94	0.173	I(1)
Inflation	−2.21	0.014	I(0)
Log population	−0.31	0.379	I(1)
Trade openness	−0.68	0.249	I(1)
Financial development	−0.41	0.341	I(1)
Urbanization	−0.52	0.301	I(1)

To assess the time-series properties of the panel, unit root tests were conducted using the Im–Pesaran–Shin (IPS) and Fisher-type augmented Dickey–Fuller procedures. Tables 2 and 3 show that government expenditure and inflation are stationary at levels, whereas global uncertainty, income per capita, population, trade openness, financial development, and urbanization are non-stationary but become stationary after first difference (Table 4). This mixed order of integration motivates the use of an error-correction framework.

Table 3. Panel Unit Root Tests (Levels) – Fisher ADF

Variable	χ^2 statistic	p-value	Order
Government expenditure	68.41	0.031	I(0)
GEPU	41.22	0.612	I(1)
Log GDP per capita	45.10	0.487	I(1)
Inflation	92.55	0.001	I(0)
Log population	39.67	0.654	I(1)
Trade openness	44.90	0.501	I(1)
Financial development	42.36	0.593	I(1)
Urbanization	40.12	0.637	I(1)

Table 4. Panel Unit Root Tests (First Differencing)

Variable	IPS W-t-bar	Fisher χ^2	Stationary
Δ Government expenditure	−14.11	388.56	Yes
Δ GEPU	−12.81	319.19	Yes
Δ Log GDP per capita	−8.57	215.15	Yes
Δ Inflation	−19.36	613.96	Yes
Δ Log population	−2.07	104.03	Yes
Δ Trade openness	−17.43	519.61	Yes
Δ Financial development	−12.31	337.29	Yes
Δ Urbanization	−3.52	236.11	Yes

Cross-sectional dependence was formally tested using Pesaran's (2004) CD test applied to the residuals from the baseline fixed-effects model. The test strongly rejects the null hypothesis of cross-sectional independence ($CD = 8.91$, $p < 0.001$), indicating the presence of unobserved global factors affecting fiscal outcomes across countries. These diagnostic results underscore the importance of accounting for common global shocks and motivate the use of Driscoll–Kraay inference and a CCE-based dynamic error-correction framework in subsequent estimations.

Baseline Fixed-Effects Estimates

Table 5 presents the baseline fixed effects estimates from Equation (1). GEPU enters with a positive, statistically significant coefficient, indicating that higher global uncertainty is associated with higher government expenditure as a share of GDP in emerging economies. This finding is consistent with the interpretation of fiscal policy as a countercyclical or precautionary tool during increased global risk.

Table 5. Baseline Fixed-Effects Estimates
Dependent variable: Government expenditure (% of GDP)

Variable	Coefficient	Std. Error	t-statistic
Global Economic Policy Uncertainty (GEPU)	0.0088***	0.0023	3.73
Log GDP per capita	−0.261	0.491	−0.53
GEPU × Log GDP per capita	0.0042**	0.0018	2.29
Inflation	−0.0117	0.0107	−1.09
Log population	−7.697***	1.642	−4.69
Trade openness	−0.103	0.373	−0.28
Financial development	0.636	0.576	1.11
Urbanization	3.407***	0.957	3.56
Constant	42.147***	6.131	6.87
Within R^2: 0.305			
Observations: 672			
Countries: 28			

Notes: Country fixed effects are included. Standard errors are clustered at the national level.
*, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively.

Income per capita enters with a negative coefficient, suggesting that higher-income emerging economies tend to maintain smaller public sectors once unobserved heterogeneity is controlled, although this effect is statistically insignificant in the baseline specification. Importantly, the interaction between global uncertainty and income per capita is positive and statistically significant, indicating that fiscal responses to uncertainty are stronger in more developed emerging economies.

Among the control variables, population size is strongly negatively associated with government expenditure ratios, reflecting scale effects on public expenditure. Urbanization has a positive and statistically significant effect, consistent with higher demand for public services and infrastructure in more urbanized societies. Inflation, trade openness, and financial development do not show robust effects in the baseline fixed-effects model.

Robustness to Cross-Sectional Dependence

Table 6 presents fixed-effects estimates with Driscoll–Kraay standard errors. The coefficient on GEPU remains positive and statistically significant, confirming the robustness of the baseline result to cross-sectional dependence and serial correlation. The interaction between uncertainty and income per capita remains positive and significant, further supporting evidence of heterogeneous fiscal responses across emerging economies.

Table 6. Fixed-Effects Estimates with Driscoll–Kraay Standard Errors
Dependent variable: Government expenditure (% of GDP)

Variable	Coefficient	Driscoll–Kraay Std. Error	t-statistic
Global Economic Policy Uncertainty (GEPU)	0.0088***	0.0025	3.55
Log GDP per capita	−0.261	0.161	−1.62
GEPU × Log GDP per capita	0.0042***	0.0012	3.59
Inflation	−0.0117	0.0116	−1.01
Log population	−7.697***	1.029	−7.48
Trade openness	−0.103	0.264	−0.39
Financial development	0.636**	0.288	2.21
Urbanization	3.407***	0.736	4.63
Constant	42.147***	3.802	11.09
Within R^2: 0.305			
Observations: 672			
Countries: 28			

Notes: The estimation uses fixed effects with Driscoll–Kraay standard errors robust to heteroskedasticity, serial correlation, and cross-sectional dependence. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively. The magnitudes and signs of the remaining control variables are largely unchanged, indicating that the baseline findings are not driven by unmodeled global shocks or error correlation.

Dynamic Adjustment and Long-Run Relationships

Table 7 presents the results from the dynamic CCE error-correction model (ECM). The estimation sample differs from Tables 5 and 6 because the ECM uses lags and first differences and excludes financial development owing to data limitations. The coefficient on the lagged level of government expenditure is negative and highly statistically significant (−0.281), confirming a stable long-run equilibrium relationship among government expenditure, GEPU, and domestic economic conditions. The magnitude of the error-correction term indicates that approximately 28% of deviations from long-run fiscal equilibrium are corrected within one year, suggesting a relatively rapid fiscal adjustment process in emerging market economies.

The positive coefficient on lagged GEPU indicates that global uncertainty shifts the expenditure trajectory beyond contemporaneous adjustments, consistent with persistent effects rather than purely short-lived responses. Income per capita enters negatively in the long-run relationship, consistent with static results. Short-run dynamics show that contemporaneous increases in global uncertainty are associated with higher government expenditure growth, whereas changes in income per capita and inflation have contractionary effects on fiscal expansion. The cross-sectional averages absorb unobserved global shocks, ensuring that the estimated coefficients reflect country-specific fiscal behavior rather than synchronized global movements.

Table 7. Dynamic Common Correlated Effects Error-Correction Estimates
Dependent variable: Δ Government Spending (% of GDP)

Variable	Coefficient	Std. Error	t-statistic
Error-correction term (L. Government spending)	−0.281***	0.040	−7.04
L. Global Economic Policy Uncertainty (GEPU)	0.0020**	0.0010	2.06
L. Log GDP per capita	−0.313	0.234	−1.34
L. Log population	−2.909***	1.019	−2.85
Δ GEPU	0.0021*	0.0012	1.86
Δ Log GDP per capita	−2.856***	0.704	−4.05
Δ Inflation	−0.025***	0.006	−3.98
L. Cross-sectional avg. government spending	0.326	0.215	1.52
L. Cross-sectional avg. GDP per capita	0.821**	0.305	2.69
Δ Cross-sectional avg. government spending	0.895***	0.241	3.71
Δ Cross-sectional avg. GDP per capita	3.427*	1.822	1.88
Constant	10.253***	3.187	3.22
Observations: 696			
Countries: 28			

Notes: *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively. Estimators: fixed effects with cross-sectional averages (CCE-ECM); Standard errors: clustered at the country level. Cross-sectional averages in levels and first differences control for unobserved common global factors. The sample differs from Tables 3 and 4 owing to lag/difference construction and variable availability.

A Hausman specification test was conducted to evaluate the appropriateness of the fixed-effects estimator relative to the random-effects alternative. The test strongly rejects the null hypothesis that the random-effects estimator is consistent ($\chi^2 = 41.8$, $p < 0.001$), confirming that the fixed-effects specification is the appropriate model for the empirical analysis. Collectively, evidence from static fixed-effects regressions, Driscoll–Kraay robustness checks, and the dynamic CCE error-correction framework consistently supports the conclusion that GEPU is central to shaping government expenditure dynamics in emerging market economies. Fiscal responses to uncertainty operate through both short-run adjustments and persistent long-run mechanisms, even after accounting for no stationarity, cross-sectional dependence, and unobserved global factors.

Discussion and Policy Implications

The empirical analysis provides consistent, robust evidence that global economic policy uncertainty (GEPU) significantly influences government expenditure dynamics in emerging market economies, both statistically and economically. Across baseline fixed-effects estimates, Driscoll–Kraay robust inference, and the dynamic error-correction framework, heightened global uncertainty is associated with an expansion in government expenditure as a share of GDP. The stability of this finding across alternative specifications indicates that fiscal responses to global uncertainty reflect systematic policy behavior rather than model-specific artifacts or spurious correlations. These results are consistent with a growing literature emphasizing the macroeconomic consequences of policy uncertainty for government decision-making and fiscal outcomes (Baker et al., 2016; Bloom, 2009; Caldara et al., 2020; Mumtaz & Ruch, 2025).

These findings are consistent with earlier evidence that fiscal policy responds systematically to external shocks in emerging markets. For example, Rodrik (1998) argues that government expenditure serves as social insurance against global volatility, while Alesina et al. (2008) emphasize the role of institutional capacity in shaping fiscal responses. More recent empirical studies also show that policy uncertainty can influence macroeconomic policy behavior by altering expectations, risk perceptions, and fiscal policy priorities (Gulen & Ion, 2015; Julio & Yook, 2012; Mumtaz & Ruch, 2025). Unlike studies such as Baker et al. (2016) and Bloom (2009), which focus primarily on output, investment, and private-sector channels, the present analysis demonstrates that uncertainty directly affects public expenditure decisions. By accounting for non-stationarity and unobserved common global factors, this study also extends prior panel evidence that relied on static specifications and potentially overstated short-run effects (Chudik & Pesaran, 2015; Pesaran, 2006).

From an economic perspective, these results align with precautionary and stabilization-oriented views of fiscal policy. Heightened global uncertainty raises downside risks to private investment and consumption, prompting governments to increase public expenditure to stabilize domestic economic activity and support aggregate demand. This mechanism is consistent with theoretical and empirical research showing that uncertainty shocks can amplify macroeconomic fluctuations and trigger countercyclical fiscal responses (Bloom, 2009; Caldara et al., 2020; Fernández-Villaverde et al., 2015; Jurado et al., 2015; Mumtaz & Ruch, 2025). Importantly, the positive association between global uncertainty and government spending persists even after controlling for cross-sectional dependence and unobserved global common factors, suggesting that emerging market governments actively adjust their fiscal policies in response to external risk rather than passively absorbing global shocks.

The interaction between global uncertainty and income per capita further indicates that fiscal responses are not uniform across emerging markets. The positive and statistically significant interaction term implies that higher-income emerging economies have a stronger expenditure response to global uncertainty, reflecting greater fiscal capacity, improved access to financing, and more developed institutional frameworks. In contrast, lower-income emerging markets may face binding fiscal constraints that limit their ability to expand expenditure in response to external uncertainty, even when stabilization needs are substantial. This heterogeneity is consistent with research highlighting the importance of fiscal space, institutional quality, and financial development in shaping macroeconomic policy responses to global shocks (Alesina et al., 2008; Arroyo et al., 2024; Ilzetzki et al., 2013).

Structural factors also play a critical role in shaping fiscal outcomes. Urbanization consistently emerges as a robust positive determinant of government expenditure, reflecting higher demand for public services, infrastructure, and social expenditure in more urbanized economies. In contrast, population size is negatively associated with government expenditure ratios, suggesting scale effects in public service provision. Financial development becomes statistically significant when inference is robust to cross-sectional dependence, indicating that deeper financial systems may enhance governments' ability to smooth expenditure in the presence of uncertainty and external shocks. These findings align with previous studies emphasizing the role of structural transformation, financial development, and demographic dynamics in determining public expenditure patterns (Arroyo et al., 2024; Svirydzenka, 2016).

The dynamic error-correction results provide further insights into the temporal dynamics of fiscal adjustment. The statistically significant negative error-correction term confirms the existence of a stable long-run equilibrium relationship between government expenditure, global uncertainty, and domestic economic conditions. The estimated speed of adjustment indicates that approximately one-quarter to one-third of deviations from long-run fiscal equilibrium are corrected within a year, suggesting relatively rapid fiscal adjustment in emerging markets. These results are consistent with dynamic macroeconomic models in which fiscal policy gradually adjusts toward long-run targets while responding to short-run shocks (Chudik & Pesaran, 2015; Pesaran, 2006; Pesaran et al., 1999).

From a policy perspective, these findings have important implications for fiscal sustainability in emerging markets. While expansionary fiscal responses to global uncertainty may stabilize economic activity in the short run, persistent increases in government expenditure can place pressure on public finances, particularly in economies with limited fiscal space or high debt burdens. Therefore, policymakers should complement countercyclical fiscal interventions with credible medium-term fiscal frameworks that balance stabilization objectives with sustainability concerns. Strengthening fiscal institutions, improving expenditure efficiency, and prioritizing growth-enhancing public investment can help ensure that uncertainty-driven fiscal responses do not undermine long-term fiscal sustainability (Arroyo et al., 2024; Debrun et al., 2008; Frankel et al., 2013).

More broadly, the results suggest that global economic policy uncertainty has become a structural feature of the international economic environment rather than a temporary disturbance. For emerging market economies, this implies that fiscal policy design must increasingly account for recurring global risks and heightened macroeconomic volatility. Developing institutional mechanisms that enable flexible yet disciplined fiscal responses—such as well-designed fiscal rules with escape clauses or stabilization funds—may enhance resilience to future uncertainty shocks and strengthen macroeconomic stability (Frankel et al., 2013; Mumtaz & Ruch, 2025).

This study has certain limitations. First, while the GEPU index captures global policy-related uncertainty, it does not allow identification of specific transmission channels through which uncertainty affects fiscal decisions. Second, the analysis focuses on aggregate government expenditure and does not distinguish between current and capital spending, which may respond differently to uncertainty shocks. Third, data limitations restrict the sample size and prevent the inclusion of certain institutional variables. These limitations suggest promising directions for future research, including disaggregated fiscal analysis, the role of fiscal rules, and the interaction between institutional quality and fiscal responses to global uncertainty.

Overall, the study demonstrates that fiscal policy in emerging markets operates at the intersection of global risk and domestic capacity. By showing that global uncertainty affects government expenditure through both short-run adjustments and long-run equilibrium mechanisms, the analysis advances understanding of how emerging market governments navigate an increasingly uncertain global economic environment.

Conclusion

This study examines how GEPU shapes government expenditure dynamics in emerging market economies from 1998 to 2023. Motivated by emerging markets' growing exposure to global shocks and unresolved debates in fiscal policy literature, the study combines a conventional fixed-effects framework with robust inference techniques and a dynamic ECM that explicitly accounts for non-stationarity and cross-sectional dependence.

The empirical results provide consistent evidence that heightened GEPU is associated with higher government expenditure as a share of GDP in emerging markets. This relationship is robust across static specifications, Driscoll–Kraay-corrected estimates, and a dynamic CCE error-correction framework. Importantly, the dynamic results confirm a stable long-run relationship between government expenditure, global uncertainty, and domestic economic conditions, while also revealing meaningful short-run adjustment dynamics.

By explicitly addressing mixed orders of integration and unobserved global common factors, this study resolves key methodological concerns that constrain the interpretability of earlier studies. The findings indicate that fiscal responses to global uncertainty are not purely transitory but rather reflect persistent adjustments consistent with precautionary and stabilization motives. This insight contributes to the literature by bridging short-run fiscal reactions and long-run equilibrium behavior in emerging markets, an area that has received limited empirical attention.

These findings have several important implications for policy. As global uncertainty is increasingly beyond the control of emerging market policymakers, fiscal authorities face mounting pressure to respond through public expenditure. While such responses may help stabilize domestic economies in the short run, they also raise concerns about fiscal sustainability, particularly in countries with limited fiscal space or constrained access to financial markets. Strengthening fiscal frameworks, improving expenditure efficiency, and building countercyclical buffers can mitigate these risks.

This study has several limitations. First, although the GEPU index captures common shocks, it cannot fully disentangle the transmission channels through which uncertainty affects fiscal decisions. Second, the study explores aggregate government expenditure and does not distinguish between current and capital expenditure, which may respond differently to uncertainty.

These limitations point to promising avenues for future research, including disaggregating fiscal components and examining the interaction among uncertainty, fiscal rules, and institutional quality.

Overall, using a rigorous econometric framework, this study provides new empirical evidence of the fiscal consequences of global uncertainty in emerging markets. By integrating robust static and dynamic approaches, the study advances understanding of how governments adjust fiscal policies amid rising global uncertainty and offers insights relevant to both academic research and policy design.

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Use of AI tools declaration

Artificial intelligence tools were used to assist with language editing and to enhance the clarity and readability of the manuscript. The authors remain fully responsible for the content and conclusions of this study.

Conflict of interest

The authors declare that there are no competing interests related to this manuscript.

References

Alesina, A., Campante, F. R., & Tabellini, G. (2008). Why Is Fiscal Policy Often Procyclical? *Journal of the European Economic Association*, *6*(5), 1006–1036. https://doi.org/10.1162/JEEA.2008.6.5.1006

Alesina, A., Perotti, R., Giavazzi, F., & Kollintzas, T. (1995). Fiscal Expansions and Adjustments in OECD Countries. *Economic Policy*, *10*(21), 205. https://doi.org/10.2307/1344590

Arroyo Marioli, F., Fatas, A., & Vasishtha, G. (2024). Fiscal policy volatility and growth in emerging markets and developing economies. *International Review of Economics & Finance*, *92*, 758–777. https://doi.org/10.1016/j.iref.2024.01.041

Baker, S. R., Bloom, N., & Davis, S. J. (2016). Measuring Economic Policy Uncertainty*. *The Quarterly Journal of Economics*, *131*(4), 1593–1636. https://doi.org/10.1093/qje/qjw024

Bernanke, B. S. (1983). Irreversibility, Uncertainty, and Cyclical Investment. *The Quarterly Journal of Economics*, *98*(1), 85. https://doi.org/10.2307/1885568

Bloom, N. (2009). The Impact of Uncertainty Shocks. *Econometrica*, *77*(3), 623–685. https://doi.org/10.3982/ECTA6248

Bloom, N., Bond, S., & Van Reenen, J. (2007). Uncertainty and Investment Dynamics. *Review of Economic Studies*, *74*(2), 391–415. https://doi.org/10.1111/j.1467-937X.2007.00426.x

Born, B., & Pfeifer, J. (2014). Policy risk and the business cycle. *Journal of Monetary Economics*, *68*, 68–85. https://doi.org/10.1016/j.jmoneco.2014.07.012

Caggiano, G., Castelnuovo, E., & Groshenny, N. (2014). Uncertainty shocks and unemployment dynamics in U.S. recessions. *Journal of Monetary Economics*, *67*, 78–92. https://doi.org/10.1016/j.jmoneco.2014.07.006

Caldara, D., Iacoviello, M., Molligo, P., Prestipino, A., & Raffo, A. (2020). The economic effects of trade policy uncertainty. *Journal of Monetary Economics*, *109*, 38–59. https://doi.org/10.1016/j.jmoneco.2019.11.002

Chudik, A., & Pesaran, M. H. (2015). Common correlated effects estimation of heterogeneous dynamic panel data models with weakly exogenous regressors. *Journal of Econometrics*, *188*(2), 393–420. https://doi.org/10.1016/j.jeconom.2015.03.007

Debrun, X., Moulin, L., Turrini, A., Ayuso-i-Casals, J., & Kumar, M. S. (2008). Tied to the mast? National fiscal rules in the European Union. *Economic Policy*, *23*(54), 297–362. https://doi.org/10.1111/j.1468-0327.2008.00199.x

Driscoll, J. C., & Kraay, A. C. (1998). Consistent Covariance Matrix Estimation with Spatially Dependent Panel Data. *Review of Economics and Statistics*, *80*(4), 549–560. https://doi.org/10.1162/003465398557825

Eichengreen, B., Gupta, P., & Mody, A. (2008). Sudden Stops and IMF-Supported Programs. In *Financial Markets Volatility and Performance in Emerging Markets* (pp. 219–264). University of Chicago Press. https://doi.org/10.7208/chicago/9780226185040.003.0008

Fernández-Villaverde, J., Guerrón-Quintana, P., Kuester, K., & Rubio-Ramírez, J. (2015). Fiscal Volatility Shocks and Economic Activity. *American Economic Review*, *105*(11), 3352–3384. https://doi.org/10.1257/aer.20121236

Frankel, J. A., Vegh, C. A., & Vuletin, G. (2013). On graduation from fiscal procyclicality. *Journal of Development Economics*, *100*(1), 32–47. https://doi.org/10.1016/j.jdeveco.2012.07.001

Gavin, M., & Perotti, R. (1997). Fiscal Policy in Latin America. *NBER Macroeconomics Annual*, *12*, 11–61. https://doi.org/10.1086/654320

Gulen, H., & Ion, M. (2015). Policy Uncertainty and Corporate Investment. *Review of Financial Studies*, hhv050. https://doi.org/10.1093/rfs/hhv050

Ilzetzki, E., Mendoza, E. G., & Végh, C. A. (2013). How big (small?) are fiscal multipliers? *Journal of Monetary Economics*, *60*(2), 239–254. https://doi.org/10.1016/j.jmoneco.2012.10.011

Im, K. S., Pesaran, M. H., & Shin, Y. (2003). Testing for unit roots in heterogeneous panels. *Journal of Econometrics*, *115*(1), 53–74. https://doi.org/10.1016/S0304-4076(03)00092-7

Julio, B., & Yook, Y. (2012). Political Uncertainty and Corporate Investment Cycles. *The Journal of Finance*, *67*(1), 45–83. https://doi.org/10.1111/j.1540-6261.2011.01707.x

Jurado, K., Ludvigson, S. C., & Ng, S. (2015). Measuring Uncertainty. *American Economic Review*, *105*(3), 1177–1216. https://doi.org/10.1257/aer.20131193

Mumtaz, H., & Ruch, F. U. (2025). Policy Uncertainty and Aggregate Fluctuations: Evidence from Emerging and Developed Economies. *The World Bank Economic Review*. https://doi.org/10.1093/wber/lhaf023

Nickell, S. (1981). Biases in Dynamic Models with Fixed Effects. *Econometrica*, *49*(6), 1417. https://doi.org/10.2307/1911408

Pesaran, M. H. (2004). General Diagnostic Tests for Cross Section Dependence in Panels. *SSRN Electronic Journal*. https://doi.org/10.2139/ssrn.572504

Pesaran, M. H. (2006). Estimation and Inference in Large Heterogeneous Panels with a Multifactor Error Structure. *Econometrica*, *74*(4), 967–1012. https://doi.org/10.1111/j.1468-0262.2006.00692.x

Pesaran, M. H., Shin, Y., & Smith, R. P. (1999). Pooled Mean Group Estimation of Dynamic Heterogeneous Panels. *Journal of the American Statistical Association*, *94*(446), 621–634. https://doi.org/10.1080/01621459.1999.10474156

Ramey, V. A. (2011). Identifying Government Spending Shocks: It's all in the Timing*. *The Quarterly Journal of Economics*, *126*(1), 1–50. https://doi.org/10.1093/qje/qjq008

Rodrik, D. (1998). Why do More Open Economies Have Bigger Governments? *Journal of Political Economy*, *106*(5), 997–1032. https://doi.org/10.1086/250038

Svirydzenka, K. (2016). Introducing a New Broad-based Index of Financial Development. *IMF Working Papers*, *16*(05), 1. https://doi.org/10.5089/9781513583709.001

Talvi, E., & Végh, C. A. (2005). Tax base variability and procyclical fiscal policy in developing countries. *Journal of Development Economics*, *78*(1), 156–190. https://doi.org/10.1016/j.jdeveco.2004.07.002

Woo, J. (2009). Why Do More Polarized Countries Run More Procyclical Fiscal Policy? *Review of Economics and Statistics*, *91*(4), 850–870. https://doi.org/10.1162/rest.91.4.850

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Economic uncertainty, monetary uncertainty and money demand in Pakistan: An asymmetrical analysis

Tahir Mukhtar*, Saira Tufail, Zainab Jehan

Department of Economics, Fatima Jinnah Women University, Rawalpindi, Punjab, Pakistan
*Corresponding Author: tahir.mukhtar@fjwu.edu.pk

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Authors' emails:
sairatufail@fjwu.edu.pk
zjehan@fjwu.edu.pk

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Abstract

Purpose — Adopting an asymmetric approach, this study analyses the impact of economic and monetary uncertainties on money demand within an open-economy framework for Pakistan. Its primary objective is to assess whether the positive and negative components of each type of uncertainty deliver a differential impact on money demand.

Methods — The study employs the Nonlinear Autoregressive Distributed Lag (NARDL) framework to examine the long-run and short-run money demand function over the period 1975–2024.

Findings — The results reveal distinct asymmetric effects. Rising economic uncertainty decreases money demand, while a decline in economic uncertainty has a positive but comparatively weaker effect. Conversely, increasing monetary uncertainty increases demand, while a decline in monetary uncertainty reduces demand for money. These findings suggest that for the positive component of economic uncertainty, the substitution effect dominates the precautionary effect; however, as monetary uncertainty increases, the precautionary effect overwhelms the substitution effect. The overall findings also indicate that agents are more sensitive to real sector volatility than to monetary volatility. Moreover, the exchange rate, along with traditional determinants, significantly influences short- and long-run money demand.

Implication — The results suggest that monetary authorities should consider the source and sign of uncertainty shocks to properly anticipate liquidity needs and achieve monetary stability.

Originality — This study is the first of its kind in Pakistan to explore the asymmetric relationship among economic volatility, monetary volatility, and money demand within an open-economy framework.

Keywords — Money Demand; Economic Uncertainty; Monetary Uncertainty; NARDL

Introduction

A thorough understanding of money demand is crucial, as it underpins sound monetary policy and financial stability. It provides a link between monetary policy and macroeconomic outcomes and helps central banks gauge liquidity preferences. Its importance was underscored by the 2007–2008 global financial crisis, which highlighted how macroeconomic uncertainty affects portfolio choices and economic behaviour.

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The conventional theories of money demand focus on the transaction and opportunity-cost motives, which are usually represented by real income and interest rates (Keynes, 1936). Extrapolations of these models also introduce uncertainty as an additional predictor, as volatility in income or prices can be an important determinant of liquidity preferences. Friedman (1984) pioneered this line of inquiry by arguing that volatility in money growth (a proxy for monetary policy uncertainty) increases overall uncertainty, leading people to hold more cash (i.e., reducing velocity). Similarly, Choi and Oh (2003) theoretically showed that output volatility (real uncertainty) induces labour market and income uncertainty, which, in turn, can increase precautionary money holdings.

Empirical studies have largely supported these ideas: for example, Friedman's (1984) hypothesis was tested by Hall and Noble (1987), who confirmed that uncertainty about money growth increases money demand. Atta-Mensah (2004) and Cronin and Kennedy (2007) found that increased economic uncertainty causes agents to prefer safer assets over cash, while studies like Bruggeman et al. (2003), Greiber and Lemke (2005) and Puah (2008) reported that real or monetary volatility tends to raise money demand. Likewise, by constructing indices of macroeconomic uncertainty, Bahmani-Oskooee et al. (2012), Bahmani-Oskooee et al. (2013), Bahmani-Oskooee et al. (2015), and Ozdemir and Saygili (2013) found a significant role of both output and monetary uncertainty on demand for money. Recently, Kurniawan et al. (2022) underscore the significant negative impact of economic uncertainty on money demand in Indonesia. Choudhry (2023) reported that economic policy uncertainty significantly reduces narrow and broad money demand in the UK. Wu et al. (2025), by employing the Toda-Yamamoto long-run causality method across 28 economies, demonstrated a significant role of economic uncertainty and fear factor in driving precautionary money demand. Nosheen et al. (2025) explained the significant role of economic and monetary uncertainty on money demand in selected Asian countries.

A key insight from recent work is that the response of money demand may be asymmetric. That is, positive and negative shocks to uncertainty might have different effects on cash holdings. Bahmani-Oskooee and Maki-Nayeri (2018), Bahmani-Oskooee and Nayeri (2018) provided evidence of such asymmetries in Korea and Australia: during high uncertainty, people hold more money, but when uncertainty falls, the reduction in holdings is smaller. Expectations about future volatility play a role: even when current uncertainty is low, if people expect future instability, they may still hold large cash balances. These findings suggest that money demand should be modelled with separate effects for increases and decreases in uncertainty. Tan et al. (2020) in China and Bahmani-Oskooee and Arize (2020) in Africa documented that the increase in uncertainty tends to have larger impacts than decrease. These works typically conclude that ignoring asymmetry would miss important dynamics of cash holdings. Recently, using the economic policy uncertainty (EPU) index, studies found an asymmetric impact of EPU on money demand, with different magnitudes and strengths. These studies highlighted the importance of incorporating nonlinearity while examining the relationship among real, monetary, overall policy uncertainty, and money demand (For example, Ongan and Gocer (2019) for Japan; Ivanovski and Churchill (2019) for Australia; Bahmani-Oskooee and Nayeri, 2020 for Japan; Hossain and Arwatchanakarn (2020) for New Zealand; Bahmani-Oskooee and Aftab (2022) for China; Nusair et al. (2024) for advanced economies; and Elroukh (2024) for GCC countries).

Pakistan provides a relevant case for exploring these issues. The country has experienced significant financial and policy reforms (Zaidi, 2015; Choudri et al., 2015) – including liberalisation of markets, inflationary episodes, and shifts in exchange-rate policy – all of which can influence money demand. Previous studies of Pakistan's money demand (Ahmad & Munirs, 2002; Azim et al., 2010; Gul & Sajid, 2020) generally reported a stable long-run relationship driven by income and interest rates, while Akbar (2021) examined the significant impact of inflation uncertainty and exchange-rate volatility on money demand. In the context of economic and policy uncertainty, Mukhtar and Jehan (2022) established a significant response of money demand to economic and monetary uncertainty in Pakistan. They identified that economic agents are more responsive to economic uncertainty than to monetary uncertainty. Hence, the existing literature mainly explores the determinants of money demand in Pakistan, whereas only a limited body of literature examines the role of uncertainty in determining it. However, none of these studies tests for asymmetric effects. A review of the extant literature makes three salient conclusions. First, the strong empirical support for Friedman's hypothesis emphasises the importance of including measures of uncertainty in the money demand function. Second, the empirical work that has been done so far on Pakistan is still mostly limited to the estimation of traditional specifications, thus highlighting a major gap. Consequently, there is an unquenchable need to enrich the model with relevant uncertainty proxies to re-examine the operational validity of the demand-for-money framework in Pakistan's economic context. Thirdly, although the asymmetric relationship is acknowledged globally, its specific dynamics in Pakistan remain poorly understood. Pakistan's unique economic conditions, marked by structural reforms, persistent inflationary pressures, and external sector vulnerabilities, require a dedicated asymmetric analysis to model money demand accurately and formulate effective stabilisation policies. To bridge this gap, the present research estimates a money demand function in an open-economy context for Pakistan that incorporates measures of both economic and monetary uncertainty and uses the Nonlinear Autoregressive Distributed Lag (NARDL) framework to capture potential nonlinearity. We show that money demand in Pakistan indeed responds asymmetrically to uncertainty shocks, with different implications for policy.

The rest of the study is divided into the following structure. Section 2 explains the methodological framework, such as the data sources and method of estimation used in the analysis. The empirical results are reported in Section 3. Lastly, section 4 provides conclusions and considers the broader policy implications of the research.

Methods

Model and Data

The standard model of money demand posits a functional relationship in which MD is commonly expressed as a function of a scale variable (i.e., real income) and the opportunity cost of holding money (i.e., the interest rate and inflation rate). In addition, the integration of the domestic financial market with international financial markets highlights the potential for currency substitution, thereby providing a strong theoretical rationale for including the exchange rate in the standard money demand equation (Mundell, 1963). Building upon this conventional approach and following Choi and Oh (2003) and Bahmani-Oskooee et al. (2015), our study extends the conventional monetary framework by incorporating economic uncertainty (proxied by output uncertainty) and monetary uncertainty (measured by money supply uncertainty) as the main regressors. This modification gives the following regression specification:

$$LM_t = \theta_0 + \theta_1 LY_t + \theta_2 R_t + \theta_3 INF_t + \theta_4 LEX_t + \theta_5 VY_t + \theta_6 VM_t + u_t \quad (1)$$

The empirical model, as indicated in Eq. (1), is estimated using the following variables. The dependent variable is the natural logarithm of real money balances (LM), defined as M2 deflated by the GDP deflator. The key determinants are the natural logarithm of real GDP (LY), the scale variable, the interest rate (R), represented by the discount rate, the inflation rate (INF) to measure opportunity cost, and the natural logarithm of the nominal exchange rate (LEX) to account for currency substitution. In this extended specification, the model uses two measures of uncertainty: economic uncertainty (VY), the rolling standard deviation of real GDP growth, and monetary uncertainty (VM), the rolling standard deviation of nominal money (M2) growth. Asymmetric effects are accounted for by decomposing each uncertainty measure into positive and negative partial sums, which indicate the incremental and decremental effects on uncertainty, respectively. Where the subscript t is the time per year, and u symbolises the white-noise error term.

The study uses annual observations from 1975 to 2024. Data on broad money (M2), nominal GDP, and consumer price index—used to construct the inflation rate—are compiled using the World Bank's World Development Indicators database. Data pertaining to discount rate, GDP deflator, and nominal exchange rate are taken from the International Financial Statistics maintained by the International Monetary Fund. The data and variable descriptions are in Table 1.

Table 1. Variables Construction and Data Sources

Variable Name	Operational Definition for Empirical Analysis	Data Source
Monetary Aggregate/Broad Money (M2)	Natural log of Real Money Balance = M2/GDP deflator (LM)	World Development Indicators by World Bank (2025)
Gross Domestic Product (constant US Dollar)	Natural Log of Real GDP (LY)	World Development Indicators by World Bank (2025)
Interest Rate	Discount Rate (% per annum)	International Financial Statistics by IMF (2025)
Inflation Rate	Percentage Change in the Consumer Price Index (Inflation Rate)	World Development Indicators by World Bank (2025)
Exchange Rate	Natural Log of Nominal Exchange Rate (i.e. Pak Rupee per US Dollar)	International Financial Statistics by IMF (2025)
Economic Uncertainty	Rolling Standard Deviation of Real GDP Growth	Authors' Own Construction
Monetary Uncertainty	Rolling Standard Deviation of Nominal Money (M2) Growth.	Authors Construction
GDP Deflator	GDP Deflator	International Financial Statistics by IMF (2025)

Econometric Technique

The study makes use of the NARDL model put forward by Shin et al. (2014) to examine the short-run and long-run reactions of money demand to the changes in VY and VM. Traditional prior research on the topic of money demand has used linear time-series models, such as the Engle-Granger, Johansen Cointegration, and ARDL models. Whereas these econometric methodologies are effective in determining long-run equilibrium relations and short-term dynamics, they have an inherent assumption of being symmetric (linear) of interaction between money demand and its determinants. The assumption of symmetry, however, could be too limiting, especially during economic policy interventions or structural change within the studied sample period (Ibrahim, 2015). In such cases, the reaction of economic agents may differ depending on whether an explanatory variable increases or decreases. As a result, the econometric models that allow nonlinear and asymmetric adjustments provide a more realistic description of economic behaviour. The assumption of symmetry presumes that the influence of a 1 percent rise in explanatory variable(s) like X on the dependent variable like Y is identical to a 1 percent fall in X on Y, while the phenomenon of asymmetry argues that such impacts of increase and decrease are not equal.

Shin et al. (2014) advanced the NARDL model and, thus, endowed it with several substantive advantages over traditional approaches to cointegration. The most important of them is the ability of the model to combine both the testing of asymmetric long-run relations with short-run dynamics into a single framework that is coherent and enhances the strength of cointegration tests in a significant way, particularly with small samples. In addition, the NARDL methodology departs from the rigid assumption that all regressors share the same integration order. The model is robust despite the inclusion of variables that are either I(0) or I(1), which enhances both its flexibility and practical applicability. The model (1) presented in the NARDL framework is given below:

$$\begin{aligned}\Delta LM_t = {} & \gamma_0 + \gamma_1 LM_{t-1} + \gamma_2 LY_{t-1} + \gamma_3 R_{t-1} + \gamma_4 INF_{t-1} + \gamma_5 LEX_{t-1} + \gamma_6 \Delta VY^{+}_{t-1} + \gamma_7 \Delta VY^{-}_{t-1} + \\ & \gamma_8 \Delta VM^{+}_{t-1} + \gamma_9 \Delta VM^{-}_{t-1} + \textstyle\sum_{i=1}^{p} \delta_{1i} \Delta LM_{t-i} + \sum_{i=0}^{q} \varphi_{1i} \Delta LY_{t-1} + \sum_{i=0}^{q} \phi_{1i} \Delta R_{t-i} + \\ & \textstyle\sum_{i=0}^{q} \phi_{1i} \Delta Inf_{t-i} + \sum_{i=0}^{q} \omega_{1i}\, \Delta LEX + \sum_{i=0}^{q} \theta_{1i} \Delta VY^{+}_{t-i} + \\ & \textstyle\sum_{i=0}^{q} \eta_{1i} \Delta VY^{-}_{t-i} + \sum_{i=0}^{q} \pi_{1i} \Delta VM^{+}_{t-i} + \sum_{i=0}^{q} \rho_{1i} \Delta VM^{-}_{t-i} + \upsilon_t \end{aligned} \quad (2)$$

where, VY and VM are divided into their partial sum components in model (2), with the positive and negative partial sums being the focus as:

$$VY_t = VY_0 + VY_t^{+} + VY_t^{-}$$

VY_0 denotes no change in the *VY* series (i.e. a threshold value of zero), $VY_t^{+}$is the partial sum of positive changes in the *VY* series given by:

$$VY_t^+ = \sum_{i=1}^{t} \Delta VY_i^+ = \sum_{i=1}^{t} max(\Delta VY_i^+, 0)$$

VY_t^- is the partial sum of negative changes in the *VY* series calculated by:

$$VY_t^- = \sum_{i=1}^{t} \Delta VY_i^- = \sum_{i=1}^{t} min(\Delta VY_i^-, 0)$$

Similarly, VM is divided into its positive and negative partial sums in model (2) as:

$$VM_t = VM_0 + VM_t^+ + VM_t^-$$

VM_0 indicates no change in the VM_0 series (i.e. threshold value equal to zero), VM_t^+ is the partial sum of positive changes in the *VM* series calculated by:

$$VM_t^+ = \sum_{i=1}^{t} \Delta VM_i^+ = \sum_{i=1}^{t} max(\Delta VM_i^+, 0)$$

VM_t^- is the partial sum of negative changes in the *VM* series given by:

$$VM_t^- = \sum_{i=1}^{t} \Delta VM_i^- = \sum_{i=1}^{t} min(\Delta VM_i^-, 0)$$

Eq. (2) represents the error correction model (ECM) specification. In this chosen formalisation, the coefficients on the first-differenced terms (Δ) capture short-run dynamics, and the parameters on the one-period lagged level variables reflect the long-run relationships.

The long- run impacts of positive and negative changes in VY and VM on money demand are given by $-\frac{\gamma_6}{\gamma_1}, -\frac{\gamma_7}{\gamma_1}, -\frac{\gamma_8}{\gamma_1}, and -\frac{\gamma_9}{\gamma_1}$, , respectively. For evidence of a short-run relationship in model (2), the lagged level variables can be substituted with the lag of the error correction term (ECT_{t-1}). After estimating the parameters of model (2), the null hypothesis of no cointegration is tested using the cointegration test as follows:

$$\gamma_1 = \gamma_2 = \gamma_3 = \gamma_4 = \gamma_5 = \gamma_6 = \gamma_7 = \gamma_8 = \gamma_9 = 0$$

To this end, the estimated F- test value is compared to Pesaran et al. (2001) lower and upper bounds critical values of the F-test statistic. The presence of cointegration between money demand and all of the explanatory variables in model (2) is indicated by the rejection of the null hypothesis. If cointegration between money demand and all explanatory variables is established, the long- run asymmetry inferences in model (2) can be drawn using the Wald test and the following null hypotheses of symmetry:

Null Hypothesis: $-\frac{\gamma_6}{\gamma_1} = -\frac{\gamma_7}{\gamma_1}$

Null Hypothesis: $-\frac{\gamma_8}{\gamma_1} = -\frac{\gamma_9}{\gamma_1}$

In the same way, the short-run asymmetric association is identified using the Wald test on the following null hypotheses of symmetric influence of VY and VM on money demand in Pakistan:

Null Hypothesis: $\sum_{i=0}^{q} \theta_{1i} = \sum_{i=0}^{q} \eta_{1i}$

Null Hypothesis: $\sum_{i=0}^{q} \pi_{1i} = \sum_{i=0}^{q} \rho_{1i}$

This research employs the NARDL model, adhering to the methodological approach as developed by Shin et al. (2014). To this end, the first and essential step is to decide how the variables are to be integrated. This is necessary because the NARDL bounds test assumes that none of the variables is second-order integrated, as any I(2) variable would invalidate the asymptotic distribution of its test statistic for cointegration. In this direction, the Dickey-Fuller Generalised

Least Squares (DF-GLS) unit-root test is used to ensure that all series are either I(0) or I(1), thereby ensuring that the NARDL estimator is consistent. After this, Eq. (2) is estimated in the NARDL framework to check the existence of a nonlinear or asymmetric long-run relationship. As Shin et al. (2014) note, the positive and negative partial sums are also treated as a single regressor in estimation; this way, the normal F-test of the levels relationship can be preserved, even as the linear specification is replaced by a nonlinear, asymmetric specification. Lastly, standard post-estimation diagnostics are conducted to assess the model's statistical fitness of the model.

Results and Discussion

Unit Roots and Cointegration

Our first empirical endeavour is the stationarity check. Such evaluation is essential to the use of the NARDL bounds test, whose accuracy would be undermined when including second-order integrated (I(2)) variables. In this regard, we use the Dickey-Fuller Generalised Least Squares (DF-GLS) unit root test to assess the null hypothesis that each of the individual series contains a unit root. The stationarity test results, presented in Table 2, reveal that LM, LY, R, VM, VY, and LEX are integrated of order 1, while VY and INF are stationary at the level. The fact that the suitability of the NARDL framework to Eq (1) is being supported by the integration properties of the regressors, though of mixed order, the absence of the I(2) series confirms the suitability of the chosen methodology. As a result, we use the NARDL technique to estimate Eq. (1).

Once the optimal lag length is determined via the Schwartz Bayesian Criteria (SBC), cointegration among the model's variables can be tested for and ascertained. For this purpose, F_{PSS} test proposed by Pesaran et al. (2001) is employed. For model (2), this test evaluates the null hypothesis of no cointegration (H_0: $\gamma_1, \gamma_2 = \gamma_3 = \gamma_4 = \gamma_5 = \gamma_6 = \gamma_7 = \gamma_8 = \gamma_9 = 0$). As reported in Table 3, the computed F-statistic exceeds the upper critical bound, compelling the rejection of the null hypothesis, affirming the presence of a cointegrated relationship in model (2).

Table 2. Results of the DF-GLS Stationarity Test

Variable	Test Statistic		Critical Values at 5% Level of Significance	Order of Integration
	At Level	At First Difference		
LM	-1.47	-15.36	-3.19	I(1)
LY	-0.51	-11.56	-3.19	I(1)
R	-2.39	-6.34	-3.19	I(1)
INF	-3.33	-	-3.19	I(0)
LEX	-1.85	-5.99	-3.19	I(1)
VY	-5.43	-	-3.19	I(0)
VM	-2.35	-7.34	-3.19	I(1)

Table 3. Bound Test for Cointegration

Dept.Var.	Ind.Variable	F-Stat Calculated Value	F-Stat Critical Value(at 5% Significance Level)		Outcome
			I(0)	I(1)	
LM	LY,R,INF,LEX, VY_Positive, VY_Negative, VM_Positive, VM_Negative	13.25	2.55	3.68	Cointegration

Long-Run Asymmetry

The NARDL decomposes a given time series into positive and negative partial sums for gauging asymmetries between dependent and independent variables. Pertaining to the objectives of the study, we have disaggregated VY and VM into their positive and negative components to examine the existence of a potential asymmetric relationship with money demand. The existence of a cointegrating relationship, as ascertained in the last step, does not say anything about the nature of the economic association between the regressors. This study assesses whether VY and VM exert an asymmetric influence on money demand, which constitutes its primary empirical objective. To this end, we begin our empirical analysis with the results of model (2), as displayed in Table 4. The Wald test result is reported at the bottom of Table 4, where we tested the null hypothesis that the positive and negative partial sums of changes in VY and VM have equal or symmetric effects on money demand in Pakistan.

Null Hypothesis: $-\frac{\gamma_6}{\gamma_1} = -\frac{\gamma_7}{\gamma_1}$

Null Hypothesis: $-\frac{\gamma_8}{\gamma_1} = -\frac{\gamma_9}{\gamma_1}$

The significant Wald test statistics allow us to reject the null hypothesis of symmetry, thereby establishing that the influence of VY and VM on money demand is asymmetric. This implies that the positive components of both types of uncertainty differ significantly from their negative components in their effects on money demand in Pakistan. Hence, such an outcome calls into question the suitability of the linear ARDL or any other linear cointegration technique for quantifying the role of macroeconomic uncertainty in money demand in the context of Pakistan. Therefore, the study validates the use of the NARDL to examine the long-run relationship among VY, VM, and money demand in Pakistan.

Table 4. NARDL Long-Run Asymmetric Estimates for Money Demand

Dependent Variable: LM Selected Model: NARDL (1,3,1,1,1,2,1)		
Regressor	Coefficient	t-value
LY	0.71***	5.39
R	-0.08***	-5.84
INF	-0.11**	-2.11
LEX	-0.87.***	-6.48
VM_Positive	0.24***	7.92
VM_Negative	0.09***	3.17
VY_Positive	-0.18***	-3.11
VY_Negative	-0.03**	-2.22
Wald Test (H0: Symmetry) LM-VM	Chi-Square (p-value)	43.87(0.00)
Wald Test (H0: Symmetry) LM-VY	Chi-Square (p-value)	17.57 (0.00)

Note: ***, ** and * denote significant at 1%, 5% and 10% levels, respectively.

Having confirmed the long-run asymmetric relationship, we move to the discussion of the long-run parameter estimates presented in the upper section of Table 3. The results showed that the coefficients of the positive and negative components of VY are statistically significant and negative. This implies that a positive change in economic uncertainty reduces money demand, whereas a negative change in uncertainty increases it. The asymmetric effect is further demonstrated by the difference in the magnitudes of the positive and negative economic uncertainty coefficients: a rise in uncertainty significantly reduces money demand (-0.18), whereas a decrease in uncertainty increases money demand by a smaller amount (-0.03).

For VM, which is a proxy for monetary policy uncertainty, both negative and positive components are positively associated with the money demand function. This relationship suggests that money demand rises as VM rises and falls as VM falls, albeit with varying magnitudes (0.24 and 0.09, respectively). The Wald test also verifies that the coefficients are not statistically similar. Theoretically, it is in accordance with Friedman's (1984) hypothesis that exceptional volatility of monetary growth elevates the level of perceived uncertainty, thereby increasing the public's demand for money as a precautionary asset. Empirically, Bahmani-Oskooee & Arize (2020) reported similar results for Algeria, Mauritius, Morocco, Tanzania and Tunisia.

The theoretical impact of uncertainty on money demand is defined by two countervailing channels: a precautionary effect and a substitution effect. An increase in VY/VM may have both a precautionary effect and a substitution effect (moving away from cash-holding behaviour toward less volatile assets). The substitution effect suggests that money demand reduces with an increase in VY/VM. It demonstrates that people are more likely to maintain alternative, less risky assets (such as gold or real estate) rather than cash balances over the long run, and to avoid remaining liquid during periods of economic instability.

On the other hand, the public can become more cautious and motivated to conserve liquidity when both measures of uncertainty rise simultaneously, leading to a precautionary effect and indicating a positive link between uncertainty and money demand. In summary, the theoretical relationship between VY, VM, and money demand is described by two counteracting forces: one a negative substitution effect, which encourages a shift in the demand for money toward alternative assets, and the other a positive precautionary effect, which increases liquidity preference. Consequently, the overall effect of an increase in VY or VM on money demand is unclear and depends on which of the opposing effects is more prominent.

Our results for both VY and VM are very intriguing in several ways, given this theoretical conjecture. First, for the positive component of VY, the substitution effect dominates the precautionary effect, and people tend to hold less money in their asset portfolios. With the increase in VM, the precautionary effect overwhelms the substitution effect, and money holding increases. These results imply that economic agents are treating both uncertainties differently. Existing literature, such as Atta-Mensah (2004), Choi and Oh (2003), Dahmardeh et al. (2011), Bahmani-Oskooee and Baek (2017), and Tan et al. (2020), confirms our result of an increase in VY. Second, the data demonstrate that the positive components of both types of uncertainty have a stronger influence on money demand than their negative counterparts. Asymmetry for VY also exists in public decisions to hold money with positive and negative sums of volatility, in addition to its magnitude. For VM, asymmetry is present only in the magnitude of response (both components are positive), not in the sign of the relationship. Moreover, in the case of Pakistan, VY with both partial sums is more decisive for money demand behaviour than VM, contrary to Bahmani-Oskooee and Nayeri (2020) in many African countries.

The traditional determinants in the model have intuitive coefficients. Real income has a positive elasticity: a 1% increase in real GDP raises money demand by about 0.71%. This Keynesian/Monetarist result reflects more transactions in a growing economy. The interest rate has the expected negative sign but is relatively small. This implies that money demand in Pakistan is fairly insensitive to interest-rate movements, consistent with a relatively stable velocity of money (Gul & Sajid, 2020; Akbar, 2021). Inflation also reduces real money demand, but its coefficient is small. In fact, comparing coefficients suggests that high inflation provides a slightly stronger incentive to economise on cash than a high nominal interest rate does. The conventional interpretation holds: higher inflation erodes the real value of cash, so people prefer real assets. These magnitude results are consistent with recent studies (Murad et al., 2021; Khan et al., 2021) and reinforce the view that economic uncertainty plays a more decisive role than monetary variables in determining cash holdings in Pakistan. Exchange rate enters negatively: a 1% depreciation of the rupee vis-à-vis the US dollar reduces real money balances by about 0.87% in the long run. This suggests a strong currency-substitution effect: when the currency weakens, residents move wealth abroad or into foreign assets, shrinking domestic money holdings (Akbar, 2021). Notably, exchange-rate elasticity (-0.87) slightly exceeds income elasticity (0.71), indicating that exchange-rate changes are a powerful influence on Pakistan's money demand.

Short-Run Dynamics and Stability

The efficacy of monetary policy is contingent upon a rigorous examination of the short-run determinants of money demand. The positive and negative partial sums of VY (Table 5) are significant and negative. They disclose that both increases and decreases in VY have an immediate negative effect on money demand. This synchronised short-run contraction is the dynamical push that propagates into a persistent long-run asymmetry, and it is confirmed that the system lacks short-run symmetry on its way to a long-run asymmetric equilibrium. The short-run outcomes, even with all lags, are very consistent with the long-run outcome. It demonstrates how people, following a lag, adjust their money holdings systematically. This is also the case for VM. The adjustment in money holding toward VM is even more spontaneous. Notably, while each lag of the uncertainty variable enters the model with a statistically significant coefficient in the short run, it increases substantially over time, yielding a much greater effect, as indicated by the long-run coefficients. This finding is crucial in determining the short-run and long-run stances of monetary policy. Moreover, the Wald test rejects the hypothesis of a symmetric relationship in favour of an asymmetric relationship. The short-run results regarding the impact of income, inflation, and interest rates on money demand are largely consistent with the long-run counterparts in terms of the direction of the relationship. However, the exchange rate has emerged as an insignificant determinant of money demand in the short run.

Table 5. NARDL Short-Run Asymmetric Estimates for Money Demand

Dependent Variable: D(LM) Selected Model: NARDL (1,3,1,1,1,2,1)		
Regressor	Coefficient	t-value
Constant	3.44***	12.73
DLM(-1)	0.67***	8.26
DLY	0.17**	2.18
DLY(-1)	0.09***	2.35
DLY(-2)	0.13	1.37
DR	-0.02**	-2.05
DINF	-0.02***	-5.32
DLEX	-0.06***	-7.21
DVM_Positive	0.03***	8.37
DVM_Positive (-1)	0.09*	1.86
DVM_Negative	0.07*	1.79
DVM_Negative (-1)	0.01**	2.26
DVY_Positive	-0.09**	-1.97
DVY_Negative	-0.03***	-5.33
ECT(-1)	-0.74***	-11.38
Wald Test (H0:Symmetry) LM-VM	Chi-Square (p-value)	12.79(0.00)
Wald Test (H0:Symmetry) LM-VY	Chi-Square (p-value)	20.68(0.00)

Note: ***, ** and * denote significant at 1%, 5% and 10% levels, respectively.

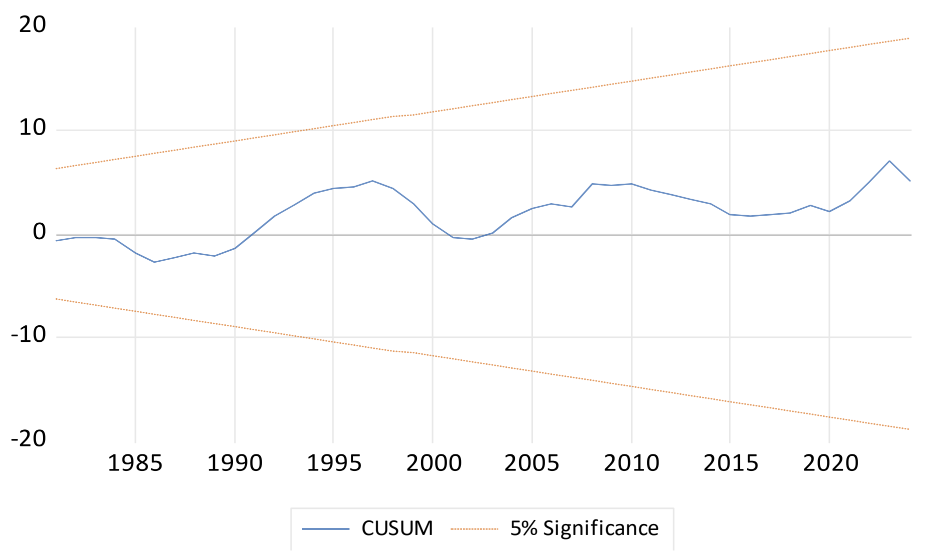

Figure 1. Plot of Cumulative Sum of Recursive Residuals

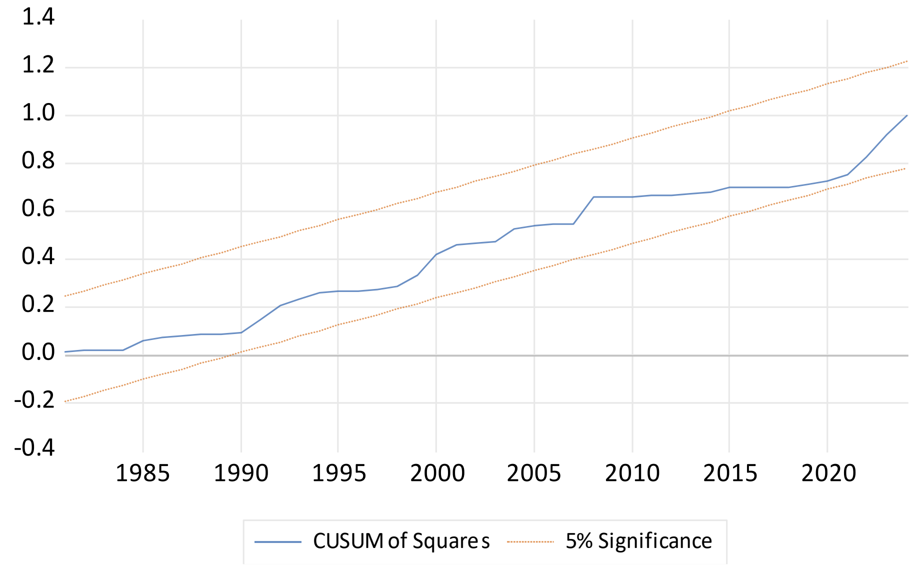

Figure 2. Plot of Cumulative Sum of Squares Recursive Residual

The error-correction term (ECT) exhibits a negative and statistically significant coefficient, providing strong evidence of a stable long-run equilibrium in Pakistan's money demand function. The finding regarding the stability of the money demand model in the long run received additional support from the results of the cumulative sum (CUSUM) and cumulative sum of squares (CUSUMSQ) tests applied to the residuals (Figures 1 and 2). These tests indicated the absence of structural breaks, which proves that the money demand function remained stable during the study period.

Conclusion

This study re-examines Pakistan's demand for money by incorporating macroeconomic uncertainty in a nonlinear framework. Using data from 1975–2024, we estimated an open-economy money-demand function that includes real income, the interest rate, inflation, the exchange rate, and proxies for economic and monetary uncertainty. Employing the NARDL methodology, we find clear evidence of asymmetric long-run effects. Rising economic (output) uncertainty significantly reduces money demand (suggesting a substitution away from cash), whereas increases in monetary (policy) uncertainty raise money demand (indicating a precautionary motive). In both cases, the effects of positive shocks are not offset by the equal effects of negative shocks. These asymmetric responses are novel findings for Pakistan and imply that symmetric models of money demand would miss important dynamics.

In addition, the traditional determinants behave as expected: real income has a strong positive effect, rupee depreciation has a large negative effect, while interest and inflation have modest negative effects in the long run. Overall, once uncertainty and exchange-rate factors are included, the money-demand function is stable and well-specified. The adjustment toward equilibrium (error correction) is rapid, and diagnostic tests indicate the model is free from major specification problems. The estimated elasticities suggest that currency substitution (via the exchange rate) has a larger impact on money demand than income does in the long run, highlighting the openness of Pakistan's economy.

Our findings yield several policy insights. Crucially, the asymmetric effect of uncertainty means that policymakers should distinguish between positive and negative shocks. Real-side (economic) uncertainty has a larger influence on cash holdings than monetary uncertainty in Pakistan. This suggests that macroeconomic policies aimed at stabilising output and expectations can have important effects on liquidity preferences. In contrast, volatility in monetary policy or money growth also matters, but its impact on money demand follows a precautionary pattern. Considering these results, we emphasise two key points for policymakers: first, Stability of money demand: The evidence of a stable long-run money-demand function (even with uncertainty included) implies that monetary aggregates remain useful policy anchors. In practice, this means that M2 growth targets or similar monetary aggregate measures can be reliable guides for policy. When interest-rate tools are constrained (for example, near the zero lower bound), the monetary authority can still monitor the money supply to assess liquidity in Pakistan. Second, the importance of real uncertainty: Since economic uncertainty strongly affects money demand, policymakers should focus on reducing real-sector volatility. Clear and credible fiscal policies, structural reforms, and transparent communication can lower output and inflation uncertainty. By doing so, the monetary authority may prevent excessive precautionary hoarding of cash. In other words, stabilising the real economy helps stabilise money demand.

In summary, our study highlights the need to incorporate macroeconomic uncertainty into analyses of money demand. Accounting for asymmetric effects provides a more complete picture of liquidity behaviour in Pakistan. Policymakers who recognise these dynamics, by maintaining stable economic policies and by tracking uncertainty indicators, will be better positioned to forecast money demand and to design effective stabilisation measures.

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Author contribution

Tahir Mukhtar: Conceptual framework, Data management, Analysis, and finalization of write-up
Saira Tufail: Conceptual framework, Data management, Analysis, and finalization of write-up
Zainab Jehan: Conceptual framework, Data management, Analysis, and finalization of write-up

Use of AI tools declaration

Artificial intelligence tools were used to assist with language editing and to enhance the clarity and readability of the manuscript. The authors remain fully responsible for the content and conclusions of this study.

Conflict of interest

The authors declare that there are no competing interests related to this manuscript.

References

Ahmad, A., & Munirs, M. (2002). An Analysis of Money Demand in Pakistan. *Pakistan Economic and Social Review*, *38*(1), 47–67.

Akbar, M. (2021). Effects of inflation uncertainty and exchange rate volatility on money demand in Pakistan: Bayesian econometric analysis. *International Journal of Finance & Economics*. https://doi.org/10.1002/ijfe.2488

Atta-Mensah, J. (2004). Money Demand and Economic Uncertainty. *Bank of Canada Working Paper*, *2004*(25). https://www.bankofcanada.ca/wp-content/uploads/2010/02/wp04-25.pdf

Azim, P., Ahmed, N., Ullah, S., Zaman, B., & Zakariya, M. (2010). Demand For Money in Pakistan: An ARDL Approach. *Global Journal of Management and Business Research*, *10*(9), 76–80.

Bahmani-Oskooee, M., & Aftab, M. (2022). On the Impact of Policy Uncertainty on the Demand for Money in China: An Asymmetric Analysis. *The Chinese Economy*, *55*, 399–409. https://doi.org/10.1080/10971475.2021.2022837

Bahmani-Oskooee, M., & Arize, A. C. (2020). Asymmetry Cointegration and the J-Curve: New Evidence from Africa. *Journal of Economic Studies*, *47*, 969–984.

Bahmani-Oskooee, M., & Baek, J. (2017). Economic Uncertainty, Monetary Uncertainty and the Korean Demand for Money. *Journal of Economic Policy Reform*, *20*(1), 86–97.

Bahmani-Oskooee, M., Kutan, A. M., & Xi, D. (2013). The Impact of Economic and Monetary Uncertainty on the Demand for Money in Emerging Economies. *Applied Economics*, *45*(23), 3278–3287.

Bahmani-Oskooee, M., & Maki-Nayeri, M. (2018). Asymmetric effects of policy uncertainty on the demand for money in the United States. *Journal of Risk and Financial Management*, *12*(1), 1.

Bahmani-Oskooee, M., & Nayeri, M. (2020). Policy Uncertainty and the Demand for Money in Japan. *Review of Economic Analysis*. https://doi.org/10.15353/rea.v12i1.1692

Bahmani-Oskooee, M., & Nayeri, M. M. (2018). Policy Uncertainty and the demand for money in Australia: An asymmetry analysis. *Australian Economic Papers*, *57*(4), 456–469.

Bahmani-Oskooee, M., Satawatananon, K., & Xi, D. (2015). Economic Uncertainty, Monetary Uncertainty, and the Demand for Money in Thailand. *Global Business and Economics Review*, *17*(4), 467–476.

Bahmani-Oskooee, M., Xi, D., & Wang, Y. (2012). Economic and Monetary Uncertainty and the Demand for Money in China. *Chinese Economy*, *45*(6), 26–37.

Bruggeman, A., Donati, P., & Warne, A. (2003). Is the Demand for Euro Area M3 Stable? *European Central Bank Working Paper Series*, (255). https://ssrn.com/abstract=457521

Choi, W. G., & Oh, S. (2003). A Money Demand Function with Output Uncertainty, Monetary Uncertainty, and Financial Innovations. *Journal of Money, Credit and Banking*, *35*(5), 685–709.

Choudhry, T. (2023). Economic policy uncertainty and the UK demand for money: evidence from the inter-war period. *Journal of Economic Studies*. https://doi.org/10.1108/jes-03-2022-0167

Choudri, E. U., Jan, A., & Malik, H. (2015). Monetary Policy in Pakistan: Effectiveness in Inflation Control and Stabilization. *Working Paper S-37204-PAK-1, International Growth Centre*. https://www.theigc.org/wp-content/uploads/2015/05/Choudhri-et-al-2015-Working-Paper-1.pdf

Cronin, D., & Kennedy, B. (2007). Does Uncertainty Impact Money Growth? A Multivariate GARCH Analysis. *Central Bank and Financial Services Authority of Ireland, Research Technical Paper*, (6/RT/07).

Dahmardeh, N., Pourshahabi, F., & Mahmoudinia, D. (2011). Economic Uncertainty-Money Demand Nexus in Iran Application of the EGARCH Model and the ARDL Approach. *European Journal of Economics, Finance and Administrative Sciences*, *38*, 118–126.

Elroukh, A. (2024). The Asymmetric Impact of Economic Policy Uncertainty on the Demand for Money in the GCC Countries. *International Journal of Economics and Financial Issues*. https://doi.org/10.32479/ijefi.17551

Friedman, B. M. (1984). Lessons from the 1979-82 Monetary Policy Experiment. *The American Economic Review*, *74*(2), 382–387.

Greiber, C., & Lemke, W. (2005). Money Demand and Macroeconomic Uncertainty. *Deutsche Bundesbank Discussion Paper, Series 1: Economic Studies*, (26/2005). https://www.econstor.eu/bitstream/10419/19611/1/200526dkp.pdf

Gul, S., & Sajid, G. M. (2020). Impact of Monetary and Other Economic Uncertainties on Demand for Money: Evidence from Pakistan. *IIIE Journal of Economics and Finance*, *1*(1).

Hall, T. E., & Noble, N. R. (1987). Velocity and the Variability of Money Growth: Evidence from Granger-Causality Tests: Note. *Journal of Money, Credit and Banking*, *19*(1), 112–116.

Hossain, A., & Arwatchanakarn, P. (2020). The effect of economic uncertainty on narrow money demand and its stability in New Zealand: An empirical investigation. *Economic Analysis and Policy*, *68*, 88–100. https://doi.org/10.1016/j.eap.2020.09.002

Ibrahim, M. H. (2015). Oil and food prices in Malaysia: A nonlinear ARDL analysis. *Agricultural and Food Economics*, *3*(2), 1–14.

Ivanovski, K., & Churchill, S. (2019). Economic policy uncertainty and demand for money in Australia. *Applied Economics*, *51*, 4516–4526. https://doi.org/10.1080/00036846.2019.1593936

Keynes, J. M. (1936). *The General Theory of Employment, Interest, and Money.*

Khan, M. A., Adil, M. H., & Husain, S. (2021). The impact of economic uncertainty, stock market uncertainty and monetary uncertainty on money demand in India. *International Journal of Emerging Markets.*

Kurniawan, M., A'yun, I., & Perwithosuci, W. (2022). Money Demand in Indonesia: Does Economic Uncertainty Matter? *Jurnal Ekonomi & Studi Pembangunan.* https://doi.org/10.18196/jesp.v23i2.15876

Mukhtar, T., & Jehan, Z. (2022). Do Economic Uncertainty and Monetary Uncertainty Matter for Money Demand? Evidence from Pakistan. *Journal of Economic Cooperation & Development*, *43*(3), 1–17.

Mundell, R. A. (1963). Capital Mobility and Stabilisation Policy under Fixed and Flexible Exchange Rates. *Canadian Journal of Economics and Political Science*, *29*, 75–485.

Murad, S. M., Salim, R., & Kibria, M. (2021). Asymmetric effects of economic policy uncertainty on the demand for money in India. *Journal of Quantitative Economics*, *19*(3), 451–470.

Nosheen, M., Majeed, R., Saeed, A., & Khan, M. I. (2025). Revisiting Output, Monetary Fluctuations, and Money Demand. *Pakistan Journal of Social Science Review*, *4*(4), 1146–1172.

Nusair, S., Olson, D., & Al-Khasawneh, J. (2024). Asymmetric effects of economic policy uncertainty on demand for money in developed countries. *The Journal of Economic Asymmetries*. https://doi.org/10.1016/j.jeca.2023.e00350

Ongan, S., & Gocer, I. (2019). Monetary policy uncertainties and demand for money for Japan: Nonlinear ARDL approach. *Journal of the Asia Pacific Economy*, *26*, 1–12. https://doi.org/10.1080/13547860.2019.1703880

Ozdemir, K. A., & Saygili, M. (2013). Economic Uncertainty and Money Demand Stability in Turkey. *Journal of Economic Studies*, *40*(3), 314–333.

Pesaran, H. M., Shin, Y., & Smith, R. J. (2001). Bounds testing approach to the analysis of level relationships. *Journal of Applied Econometrics*, *16*(3), 289–326.

Puah, C. H. (2008). Economic Uncertainty and the Demand for Money in Malaysia: A Bounds Testing Estimation. *7th International Conference on Computational Intelligence in Economics and Finance (CIEF 2008).*

Shin, Y., Yu, B., & Greenwood-Nimmo, M. (2014). Modelling asymmetric cointegration and dynamic multipliers in a nonlinear ARDL framework. *Festschrift in Honor of Peter Schmidt*, 281–314.

Tan, W. H., Puah, C. H., Wong, S. S. L., & Chong, M. T. (2020). Economic Uncertainty and the Demand for Broad Money in South Africa. *Business and Economic Research*, *10*(2), 123–133.

Wu, Z., Gan, P., Hadi, F., & Lim, E. (2025). Precautionary Demand for Money and Economic Uncertainty: Evidence from Toda-Yamamoto Long-Run Causality Method. *International Journal of Economics and Financial Issues*. https://doi.org/10.32479/ijefi.18151

Zaidi, A. S. (2015). *Issues in Pakistan Economy: A Political Economy Perspective*.

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Do monetary policy and macroeconomic fundamentals matter under high uncertainty? Evidence from Renminbi exchange rate regimes

Zhang Mengdi[1], Lee Chin[2*], W.N.W. Azman Saini[3]

[1]Graduate School of Studies, Universiti Putra Malaysia, Serdang, Selangor, Malaysia
[2,3]School of Business and Economics, Universiti Putra Malaysia, Serdang, Selangor, Malaysia
[2]Department of Econometrics, Tashkent State University of Economics, Tashkent, Uzbekistan
[*]Corresponding Author: leechin@upm.edu.my

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Authors' emails:
gs63066@student.upm.edu.my
wazman@upm.edu.my

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Abstract

Purpose — This study examines the nonlinear effect of Economic Policy Uncertainty (EPU) on China's Renminbi (RMB) exchange rate.

Methods — Based on the monetary model of exchange rate determination and the behavioural equilibrium exchange rate model, a threshold autoregressive model was estimated with EPU as the threshold variable. Quarterly data from Q1 2005 to Q4 2023 were utilised.

Finding — Under different regimes, the effects of EPU, monetary factors, and macroeconomic factors on the exchange rate are nonlinear. The study also reveals significant differences in the volatility characteristics of exchange rate misalignment under low and high EPU conditions, supporting the hypothesis that EPU has a nonlinear impact on the exchange rate. Furthermore, when EPU is low, external shocks exert a stronger impact on the exchange rate than when EPU is high.

Implications — These results suggest that governments and policymakers can help investors anticipate market shifts by increasing policy transparency and reducing unnecessary policy changes, thereby maintaining economic stability.

Originality — By integrating EPU into a framework combining the monetary model and the BEER model and estimating a nonlinear threshold model, this study provides new evidence on exchange rate dynamics and misalignment under varying EPU regimes.

Keywords — Economic Policy Uncertainty, Exchange Rate, Monetary Model, BEER Model, Threshold Autoregressive Model.

Introduction

The exchange rate (ER), which affects trade balances and macroeconomic stability, is the main channel between a nation's domestic and global markets (Abere, 2025; Cong et al., 2025; Kim et al., 2025). ER movements have been explained by several theoretical models, most of which are based on macroeconomic principles. The complexity of ER fluctuations, especially in reaction to shocks of economic uncertainty, is difficult for these current models to adequately explain (Msomi & Ngalawa, 2024). To develop a more comprehensive framework that accounts for the influence of both EPU[1] and conventional macroeconomic factors on ER fluctuations, this study integrates EPU into two established ER models. The first is the model of the Monetary Model of Exchange

[1] Economic policy uncertainty (EPU) refers to the inability of entities to predict future economic policies or the risks and consequences of policies adopted by the government (Nguyen et al., 2025; M. Zhang & Du, 2025).

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Rate Determination (hereafter MM-ERD). Originally proposed by Dornbusch (1976) and Frenkel (1976), it establishes a relationship between ER and economic factors such as relative currency supply, interest rates, income levels, and inflation differentials. The second model is the Behavioural Equilibrium Exchange Rate (BEER) Model, proposed by Clark and MacDonald (1999), which estimates the equilibrium ER by incorporating other economic variables such as terms of trade (ToT), productivity differences, and foreign assets (Ca' Zorzi & Rubaszek, 2023). Therefore, integrating the two can consider both theory and practicality.

As China's economy has become increasingly integrated into the global financial system, its ER system has undergone several reforms (Table 1). Prior to 1994, China maintained a fixed ER system; following 1994, the country transitioned to a floating ER mechanism. In line with China's globalisation, it has continued to implement periodic ER reforms (e.g., in 2005 and 2015) to achieve ER marketisation. These structural adjustments, coupled with external shocks, motivate this study to investigate the dynamics of the RMB ER in response to EPU. In the context, traditional linear models, such as the Autoregressive Distributed Lag (ARDL) model, fail to capture actual dynamics, whereas nonlinear models, such as the threshold autoregressive (TAR) model, can better reflect asymmetric effects under different EPU regimes (Msomi & Ngalawa, 2024; Tong & Lim, 1980). To capture the nonlinear effect, this study uses a TAR model.

Table 1. China's exchange rate system across different periods

Time	RMB Exchange Rate System
Before 1994	Fixed ER system with strict controls.
1994	Managed float ER system, pegged to the USD.
2005.7.21	ER is based on market supply and demand, adjusted for a portfolio of currencies.
2015.8.11	A basket of currencies serves as the primary benchmark for assessing the ER.
2016	"Counter-cyclical factor + basket currency ER fluctuation" two-factor pricing model.
2017	"Counter-cyclical factor + basket currency ER fluctuation + pro-cyclical factor " three-factor pricing model.

Source: People's Bank of China (PBoC)

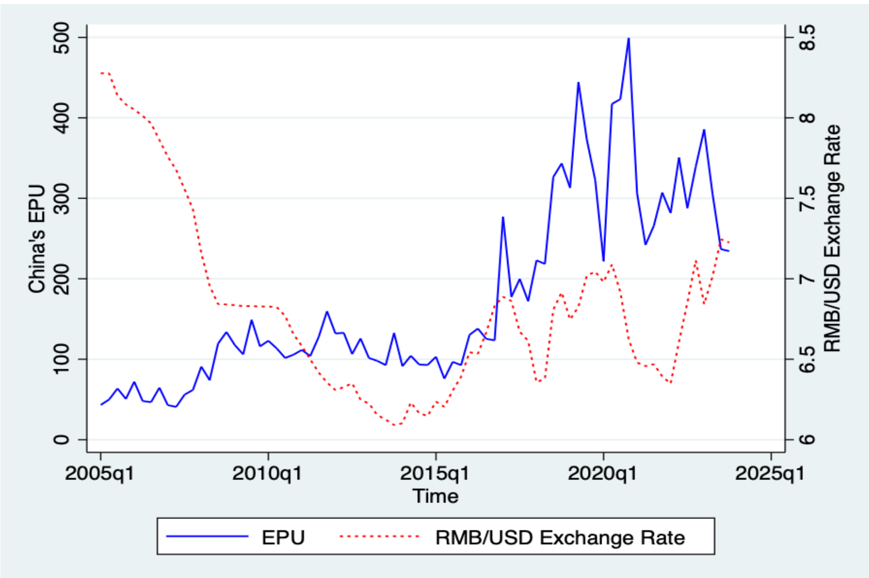

Figure 1. RMB/ USD exchange rate and China's EPU

Due to domestic as well as foreign economic uncertainties, the RMB ER has become more volatile in recent years (Zhang & Huang, 2025). Figure 1 illustrates how important policy events—the 2005 and 2015 ER reforms, the rise of US-China trade tensions in 2018, and the global COVID-19 outbreak in 2020—caused notable fluctuations in the RMB/USD ER, increasing currency valuation uncertainty. EPU, a popular indicator of policy-driven economic unpredictability, can capture these events (Wang, 2024). In other words, the ER varies with the EPU (Figure 1).

ER can be affected by EPU in a few ways. First, it directly affects currency valuations, thereby raising risk premiums in global foreign exchange markets (Glebocki & Saha, 2024). Second, it influences investor expectations, leading to changes in ER adjustments and in speculative behaviour (Baker et al., 2016; Bernanke, 1983; Dixit, 1989; Kim et al., 2025). Thirdly, it affects international capital flows. Increased uncertainty frequently leads to fluctuations in foreign direct investment or capital flight (Goswami et al., 2023; Liu, 2024). The RMB ER continues to move away from its symmetrical point due to the factors noted above.

Thus, we hypothesise that RMB ER misalignment from its equilibrium state can result from changes in the ER system, economic shocks, and other factors (Edwards, 1989; Holtemöller & Mallick, 2013). Enterprise output is affected by this misalignment, prompting investors and practitioners in the global foreign exchange market to reduce risk by "waiting-and-seeing" (X. Wang, 2024). Thus, consistent with (Wen & Usman, 2024b)This study uses the BEER method to investigate the nonlinear effects of various EPU conditions on the RMB ER.

The existing literature supports the view that monetary variables drive ER fluctuations. For example, Chin et al. (2009) confirmed the cointegration relationship between monetary variables and the RMB/USD ER, both before and after the crisis, indicating that the MM-ERD significantly explains how Malaysia's ER is determined. However, many studies indicate that changes in monetary variables alone are insufficient to fully explain ER fluctuations. For example, Afat et al. (2015) found that deviations from PPP prevent the MM-ERD from establishing the predicted relationship between monetary factors and ER. To address these shortcomings, economists developed a series of equilibrium ER models in the 1980s and beyond, including the BEER model. However, the BEER model is essentially an empirical approach (Clark & MacDonald, 2004). As a result, using the BEER model alone to explain ER variation lacks a theoretical basis. To overcome the limitations, this study combines the MM-ERD model with the BEER framework. The former offers a theoretical foundation for the latter, while the latter addresses the former's shortcomings regarding the monetary determinants of ER fluctuations. A comprehensive and inclusive explanation of ER movement is made possible by this integrated approach.

However, previous studies have also provided evidence that traditional economic variables are insufficient to explain ER fluctuations, as they fail to capture policy changes and market expectations (Msomi & Ngalawa, 2024). Incorporating EPU into estimates of ER fluctuations can improve the explanatory power and predictive accuracy of macroeconomic models (Abid, 2020). Meanwhile, many studies have confirmed a direct linear relationship between EPU and ER (Gong et al., 2025). However, there is limited research on the nonlinear effects in this relationship, particularly using a model that combines the MM-ERD and BEER models.

The main contributions of this study are as follows. First, it integrates macroeconomic and monetary factors and considers the impact of the EPU, thereby constructing a more comprehensive ER determination model and enriching the literature on ER modelling. Second, it provides a more accurate analysis of dynamic adjustment processes by using a nonlinear approach to capture the nature of ER fluctuations. Third, it examines RMB ER misalignment under different EPU conditions, generating empirical evidence to support policies that maintain ER equilibrium and mitigate risks in foreign exchange markets. Finally, the findings provide a theoretical basis for investment decisions of foreign exchange market participants and production decisions of international enterprises.

The remaining sections of this paper, in order, review the relevant literature, describe the methodological framework, present the results of data analysis, and lastly, discuss the findings with appropriate policy recommendations.

Methods

Model Specification

According to the MM-ERD, the ER can be explained by four key factors: income, inflation, money supply, and interest rate (all in relative terms) (Chin, Azali, & Matthews, 2007; Chin, Azali, Yusop, et al., 2007; Korap, 2024; MacDonald & Taylor, 1991). The functional relationship of this model is expressed as:

$$e_t = (m_t - m_t^*) - \alpha(y_t - y_t^*) + \theta(r_t - r_t^*) + \lambda(\pi_t - \pi_t^*) \quad (1)$$

Here, e_t represents the RMB/USD ER, while π_t, m_t, y_t, and r_t denote the inflation level, money supply, output level, and interest rate, respectively, in the domestic country. Similarly, π_t^*, m_t^*, y_t^*, r_t^* represent the corresponding variables in the foreign country. In this study, China was the domestic country, while the US represented the foreign country.

As explained previously, to overcome the shortcomings of individual methods, this study analysed the ER by integrating the MM-ERD and the BEER model. Based on previous research, two fundamental macroeconomic variables from the BEER model were selected: ToT (Wen & Usman, 2024a) and foreign exchange reserves (FER) (Abid, 2020; Aizenman et al., 2024). The new model is expressed as follows:

$$e_t = \alpha_1 + \alpha_2(m_t - m_t^*) + \alpha_3(y_t - y_t^*) + \alpha_4(r_t - r_t^*) + \alpha_5(\pi_t - \pi_t^*) + \alpha_6 FER_t + \alpha_7 TOT_t \quad (2)$$

Following (Msomi & Ngalawa, 2024), this study then introduced EPU into the new model to address the limitations of traditional economic variables in reflecting policy changes and market expectations. The final model, which combines the monetary variables with the macroeconomic fundamentals from the BEER model and considers EPU as an influencing factor, is below:

$$e_t = \alpha_1 + \alpha_2(m_t - m_t^*) + \alpha_3(y_t - y_t^*) + \alpha_4(r_t - r_t^*) + \alpha_5(\pi_t - \pi_t^*) + \alpha_6 FER_t + \alpha_7 TOT_t + \alpha_8 EPU_t \quad (3)$$

Measurement of Exchange Rate Misalignment

The degree of temporary ER misalignment refers to the deviation between the actual effective ER and its equilibrium level. Aligning with previous research (Banerjee & Goyal, 2021), this study adopted the BEER approach to calculate exchange rate misalignment (ME) using the following expression:

$$ME_t = \frac{e_t - \hat{e}_t}{\hat{e}_t} \quad (4)$$

Where, e_t is the RMB/USD ER, $\hat{e}$ is the equilibrium ER.

Data and Variables

Table 2 presents the descriptive analysis results of the variables, based on quarterly data from Q1 2005 to Q4 2023. The data's starting point was set as 2005 due to the beginning of China's "managed ER system" that year, meaning that ER data after this period better reflects the impact of EPU.

The dependent variable was the ER, denoted as *e*. Its maximum, minimum, and average values are 8.277, 6.091, and 6.831, respectively (Table 2). Additionally, the independent variable is the China EPU[2] Index. It was developed by (Davis et al., 2019). The index extracts relevant EPU-related information by analysing news media and newspaper reports. Its highest recorded value of 499.200, its minimum of 40.800, and its average of 173.425 (Table 2).

Regarding the control variables, Δm, Δi, Δy, and Δπ respectively represent the differences in money supply, interest rates, output levels, and inflation between China and the US. China's money supply was measured via M2 data obtained from the CEInet Statistics Database, while US M2 data was sourced from the Federal Reserve Economic Data (FRED). Interest rates were derived from the three-month (90-day) Treasury bond yield for China and the three-month U.S. Treasury yield for the U.S., both sourced from the U.S. Federal Reserve. China's and the US's output levels were measured using GDP, with data sourced from the International Financial Statistics. For inflation, the CPI growth rate was the proxy for both countries, with data obtained

[2] The data was obtained from the following website: https://policyuncertainty.com/china_monthly.html.

from the International Monetary Fund. Finally, FER and ToT represent China's FER and ToT (total exports/total imports), acquired from the CEInet Statistics Database.

Table 2. Descriptive statistics

	Variable	Obs	Mean	Std. dev.	Min	Max
Dependent variable	*e*	76	6.831	0.567	6.091	8.277
Independent variables/Threshold variable	EPU	76	173.425	114.210	40.800	499.200
Control variables	Δm	76	6919.395	6956.037	-3291.600	19640.900
	Δi	76	2.006	2.43384	-2.9233	6.113
	Δy	76	1.23E+07	7592378	934001	2.80E+07
	$\Delta\pi$	76	-0.057	1.260	-2.966	2.497
	FER	76	27448.660	9021.847	6418	39853
	ToT	76	1.222	0.101	0.992	1.485

Estimation Methods

Threshold Autoregression (TAR) Model

This study applied the TAR model by Tong and Lim (1980) to examine the EPU-ER relationship using the new integrated MM-ERD–BEER Model. The TAR model allows changes in the dynamics of a time series when a threshold is exceeded, effectively capturing asymmetric, nonlinear, or transitional patterns that linear models may miss. The models used in this study are as follows:

$$e_t = \alpha_1 + \theta_1 X_t + \lambda_1 EPU_t + \epsilon_{1,t}, EPU_t \le q \quad (5)$$

$$e_t = \alpha_2 + \theta_2 X_t + \lambda_2 EPU_t + \epsilon_{2,t}, EPU_t > q \quad (6)$$

Where e_t is the RMB/USD ER, X_t is a set of control variables, such as monetary variables, FER, and ToT; EPU_t is China's EPU, and q is the threshold value. When $\lambda_1 \neq \lambda_2$, there is a threshold. While α_1 and α_2 are the intercepts, $\epsilon_{2,t}$, $\epsilon_{2,t}$ are error items.

Threshold Effect Test

Before estimating the TAR model, it is crucial to verify the existence of threshold effects to ensure that the model accurately captures nonlinear patterns. This study uses the likelihood ratio test to assess the existence and statistical significance of the threshold effect and calculates the F-statistic to determine whether the threshold model fits the data better than the linear model. The formula for this test is as follows:

$$F = \frac{(SSR_{lnear} - SSR_{single})/(k_1 - k_0)}{SSR_{single}/(n - k_1)} \quad (7)$$

Where SSR_{linear} represents the sum of squared residuals of the linear model and SSR_{single} represents the sum of squared residuals of the single-threshold model. n denotes the sample size, k_0 is the total number of parameters in the linear model, and k_1 is the total number of parameters in the threshold model, including the intercept, non-interaction terms, and interaction terms. If the F-statistic is small, the increased fit from the single-threshold model is insignificant, such that the null hypothesis of "non-existent threshold effect" cannot be rejected.

If the F-statistics are substantial, the threshold effect is statistically significant. In this case, the fit between the single- and double-threshold models is further compared using the LR Test Statistic. Its calculation is as follows:

$$F = \frac{(SSR_{single} - SSR_{double})/(k_2 - k_1)}{SSR_{double}/(n - k_2)} \quad 8)$$

Where SSR_{single} represents the sum of squared residuals of the single-threshold model, and SSR_{double} represents the same for the double-threshold model. k_2 is the total number of parameters in the double-threshold model.

Results and Discussion

Unit Root Test

This study employed the Augmented Dickey-Fuller unit root test to examine the stationarity of each time series variable. As shown in Table 3, the series for EX and Δi were not stationary at the level. However, after first-order differencing, both variables become stationary. Therefore, their first-order differences—denoted as dex and dΔi, respectively—were used in subsequent analyses. In addition, Δπ, FER, and ToT reject the unit root under the constant specification at the 1% significance level. Δy rejects the unit root at the 1% significance level, EPU and Δm reject the unit root at the 5% significance level, under the trend specification.

Table 3. Results of the unit root test

Variables	Constant without trend		Constant with the trend
	Level	First dif.	Level
EX	0.089	0.000***	0.867
EPU	0.200	-----	0.023**
Δm	0.977	-----	0.021**
Δi	0.840	0.000***	0.986
Δy	0.828	-----	0.000***
Δπ	0.000***	-----	-----
ToT	0.000***	-----	-----
FER	0.001***	-----	-----

Note: ***, **, * indicate critical values at 1%, 5%, and 10% significance levels, respectively.

Threshold Analysis

This study's main aim was to examine ER changes under different conditions of EPU; accordingly, EPU was set as the threshold variable. Table 4 presents the threshold-effect findings from the likelihood ratio test. For the single-threshold model, the null hypothesis of no threshold effect was rejected at the 1% significance level, confirming the presence of nonlinearity and thus the suitability of the TAR model. Next, Equation (8) was used to test the double-threshold model, and the null hypothesis was not rejected at the 5% significance level. Since the double-threshold effect was excluded, this study adopted a single-threshold TAR model.

Table 4. Threshold effect results

Threshold	F-statistics	P-value	Estimated threshold value
Single	4.130	0 .001	159.745
Double	2.264	0.132	-------
Robustness	chi2(1)	Prob > chi2	
Breusch-Pagan test	1.480	0.223	
Durbin-Watson test	0.090	0.764	

Table 5 summarises the results for the two different EPU states. Regime 1 indicates that the EPU is in a low state. Regime 2 indicates that the EPU is in a high state. The results indicate that in Regime 1, China's EPU has a significant positive impact on the RMB/USD ER. From Table 5, we can see that for every 1% increase in EPU, the RMB/USD ER rises by around 0.1%, indicating RMB depreciation. This finding aligns with Abid (2020), who observed that an increase in EPU leads to currency depreciation. Table 5 also indicates that Δm is significantly correlated with the RMB/USD ER. And it also shows a significant correlation between Δm and the RMB/USD ER. It means that when China's money supply grows faster than the US money supply, the RMB tends to depreciate. This supports the findings of (Chin, Azali, Yusop, et al., 2007), who found that relative money supply directly affects the nominal ER of the peso against the US dollar.

In Regime 1, Δi has a significant negative impact on the RMB/USD ER. It means that rising interest rates led to RMB appreciation. For every 1% increase in the relative interest rates between

China and the US, the RMB/USD ER will decrease by 2.4%. This may be because rising interest rates attract capital inflows, thereby strengthening the domestic currency ER. Similarly, Chin, Azali, and Matthews (2007) and Msomi and Ngalawa (2024) found that, ceteris paribus, interest rate growth and currency appreciation generally move together. Although the same negative influence on ER was observed for Δy at low EPU, this effect was not statistically significant. In other words, higher levels of GDP decrease the RMB/USD ER and appreciate the RMB, but not enough to be statistically significant. This finding supports the work of Korap (2024), who also noted a non-significant yet long-term negative correlation between relative income and the nominal ER.

Table 5. TAR model and linear regression results

Parameter estimation	Regime 1 (EPU<= 159.745）	Regime 2 (EPU> 159.745）	Linear Regression
_cons	0.002	6.625	-0.254
	(0.152)	(2.409)	(0.201)
EPU	0.001*	-0.001	-0.001
	(0.001)	(0.001)	(0.001)
Δm	0.001**	-0.001	-0.001
	(0.001)	(0.001)	(9.30E-06)
dΔi	-0.024*	-0.082	-0.065***
	(0.014)	(0.051)	(0.020)
Δy	-3.00E-09	4.38E-08**	1.56E-08*
	(1.18E-08)	(1.76E-08)	(9.23E-09)
Δ*π*	0.001	-0.046	0.003
	(0.008)	(0.032)	(0.012)
FER	-7.43E-06**	-0.000198**	2.00E-06
	(2.82E-06)	(0.001)	(2.48E-06)
ToT	0.030	-0.890	0.091
	(0.111)	(0.599)	(0.159)

Note: Standard errors are in parentheses; ***, **, * indicate critical values at 1%, 5%, and 10% significance levels, respectively.

Next, Table 5 shows that Δπ has no significant effect on the RMB/USD ER. This contradicts the Sticky Price Monetary Model's theoretical proposition that the relative inflation rate between China and the US is associated with RMB depreciation. Chin et al. (2009) reported a similar outcome on the ringgit–dollar nominal ER, attributing it to price rigidity. The impact of inflation on ER typically has a considerable lag; its short-term effects are not statistically significant. The results further indicate that, under Regime 1, China's FER has a significant negative impact on the RMB/USD ER and promotes RMB appreciation. Aizenman et al. (2024) also reached the same conclusion: a higher FER will strengthen the ER of the domestic currency. Finally, at Regime 1, Table 5 shows that improvements in China's ToT led to a depreciation of the RMB, although this result is not statistically significant. This result may support Clark and MacDonald (1999) view that improvements in ToT generally lead to an appreciation of the domestic currency.

Meanwhile, Regime 2 shows that interest rates and FER have a significant negative impact on ER, consistent with the situation under low EOU. Although inflation and the ToT have opposite effects on ER under high EPU, these effects are not statistically significant. In Regime 2, the negative impact of relative inflation on ER was not significant. This can be attributed to China's effective control over inflation expectations. According to Wang & Wang (2015), China's "expectation management" policy in 2009 enhanced the central bank's ability to control changes in inflation expectations, thereby strengthening market confidence in the RMB and promoting its appreciation. The advancement of the ER system also appears to have contributed to the RMB's appreciation. For ToT, its improvement appears to lead to RMB appreciation. Although not significant, this result agrees with the findings of Abid (2020), when the ToT of Mexico, Brazil, and India improved, their currencies also appreciated.

In conclusion, many variables in Regime 1—such as EPU, money supply, interest rates, FER, and inflation—generally follow BEER theory and conventional monetary models. However, in Regime 2, the signs of most variables become insignificant or their directions "reverse". The linear regression results, shown in Table 5, indicate that EPU has no significant impact on ER. This suggests that if EPU exceeds a certain threshold, the linear relationships predicted by the classical monetary model and the BEER framework no longer fully hold. Therefore, the use of the TAR model can better capture ER changes under varying EPU conditions and, consequently, facilitate timely interventions to maintain ER stability.

Exchange Rate Misalignment

In the second part of the analysis, this study investigated the degree of misalignment of the RMB ER, calculated as the difference between the equilibrium ER (estimated through the BEER approach) and the actual ER. Figure 2 plots the different levels of ER misalignment under the two EPU regimes, where the solid blue line represents low EPU and the red dotted line represents high EPU. Each data point corresponds to the misalignment from the regime (Regime 1: EPU <= 159.745; Regime 2: EPU > 159.745) to which the observation belongs. The results show that the volatility of ER misalignment varies significantly under EPU conditions, supporting the hypothesis that EPU has a nonlinear effect on ER misalignment.

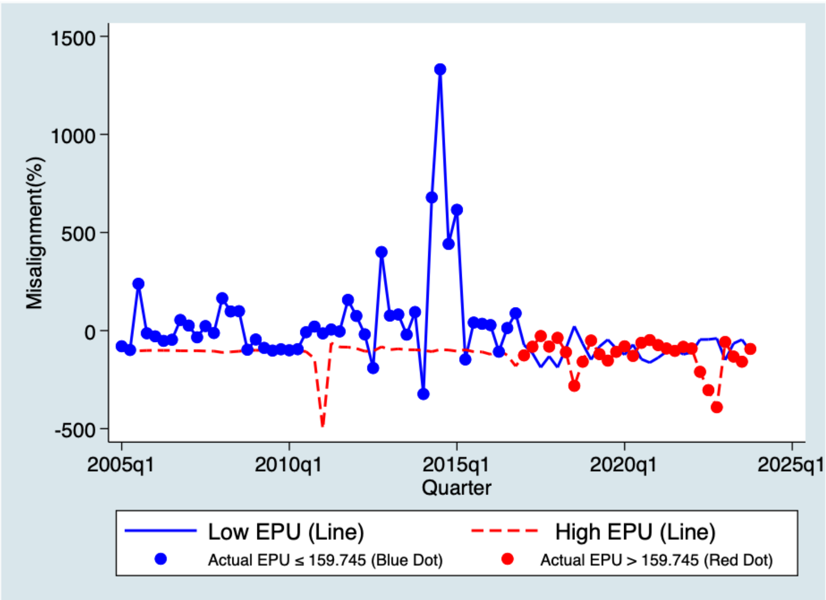

Figure 2. ER misalignment under different EPU states.

Note: Misalignment is measured as the difference between the actual RMB/USD ER and its estimated equilibrium ER. Each dot represents the degree of deviation for the quarter. It is colour-coded according to the EPU regimes. Blue dots indicate that the EPU for the quarter is in Regime 1, while red dots indicate that the EPU for the quarter is in Regime 2. Regimes are identified using the TAR model and the BEER-based equilibrium estimation described in Section 3. A positive deviation (above zero) suggests an undervalued RMB, while a negative deviation (below zero) indicates an overvalued RMB.

Figure 2 shows that periods with higher EPU are largely concentrated after 2017.-In 2018, the US–China trade war broke out, exerting depreciation pressure on the RMB. On August 3, 2018, the PBoC announced that the foreign exchange risk reserve ratio for forward foreign exchange sales would be raised from 0% to 20%. This measure elevated costs for businesses and financial institutions engaging in forward foreign exchange purchases, thereby mitigating excessive speculative activities. Subsequently, on August 24, 2018, the PBoC reinstated the counter-cyclical factor to mitigate excessive RMB depreciation. Then, from 2022 to 2023, the US Federal Reserve significantly raised the federal funds rate, leading to a rise in US interest rates[3]. In China, the PBoC dropped the reserve requirement ratio and interest rates several times in 2023, including two reserve requirement ratio cuts in March and September[4] and loan prime rate cuts in June and August[5], aiming to support domestic economic recovery after COVID-19. The RMB faces continued depreciation pressure. As shown in Figure 2, when the EPU is high, the RMB ER is undervalued most of the time.

However, Figure 2 shows that before 2017, the EPU was low most of the time. And the RMB ER was overvalued of times, but the RMB also alternates between undervaluation and overvaluation when EPU is low. In addition, the degree of misalignment is greater than that at Regime 2. This is because, following a series of ER reforms, government intervention decreases in a less uncertain environment, leaving the RMB more vulnerable to market supply and demand (Cui & Si, 2024). In Regime 2, the most significant instance of RMB undervaluation occurred in 2014. On March 15, 2014, the PBoC announced that the daily trading range of the RMB against the USD would be widened from 1% to 2%, increasing the ER's flexibility and improving its representation of market supply and demand (Funke et al., 2015). However, the expanded trading range also triggered market expectations of RMB depreciation. It will lead to increased capital outflow pressure and accelerated RMB depreciation. As a result, the RMB depreciated sharply against the US dollar in 2014, undervaluing itself by approximately 13%, the largest undervaluation under Regime 1.

The RMB is more susceptible to market supply and demand when EPU is low, thus making it more prone to periods of alternating overvaluation and undervaluation. Conversely, when EPU is high, the RMB is undervalued for most of the time. Therefore, the difference in the degree of RMB ER misalignment under high and low EPU environments supports the conclusion that there is a non-linear relationship between EPU and ER misalignment. In addition, when EPU is low, external shocks (e.g., international economic events) exert a stronger impact on the ER than when EPU is high.

Conclusion

This research developed a TAR model based on the MM-ERD and the BEER Model to explore the nonlinear association between EPU and the RMB ER. Analysis of quarterly data from Q1 2005 to Q4 2023 revealed that the RMB/USD ER behaves differently across different levels of the threshold variable, EPU. Specifically, at a threshold of 159.745, the EPU environment can be divided into low and high states (Regime 1 and Regime 2). The results show that in Regime 1, most variables, such as EPU, money supply, interest rate, FER, and inflation rate, are consistent with the MM-ERD and BEER models. However, in Regime 2, most variables become insignificant, or their directions reverse. These results indicate that once EPU exceeds a threshold, the linear relationships predicted by the classic monetary model and the BEER framework no longer hold in full.

This study also examined whether the RMB is overvalued or undervalued across different EPU conditions. The findings show that the volatility of ER misalignment differs significantly between low- and high-EPU regimes. When the EPU is low, the ER may be overvalued or undervalued, but it is overvalued most of the time. Conversely, when the EPU is high, the RMB is effectively undervalued. In addition, the degree of misalignment is greater at low EPU than at high EPU. This conclusion supports the hypothesis that EPU has an asymmetric impact on ER misalignment. Given that traditional ER theories (e.g., the MM-ERD and the BEER Model) do not explicitly consider policy uncertainty, this study's examination of EPU's role in ER variation expands extant theories and provides new perspectives for the development of better explanatory ER models.

[3] According to the St. Louis Federal Reserve Bank, between 2022 and 2023, the Federal Reserve implemented a series of interest rate increases, raising the federal funds target rate from 0.00%–0.25% to 5.25%–5.50%, a cumulative increase of 5.25 percentage points.

[4] According to the PBoC, the first reserve requirement ratio reduction was on March 27, 2023; the ratio for financial institutions' RMB deposits was reduced by 0.25 percentage points. The second was on September 15; the PBoC again decreased the ratio for financial institutions' RMB deposits by 0.25 percentage points.

[5] According to the PBoC, on June 20, 2023, the one-year loan prime rate was lowered from 3.65% to 3.55%, and the rate for more than five years was lowered from 4.30% to 4.20%. On August 20, the one-year loan prime rate was further lowered to 3.45%, while the rate for more than five years was maintained at 4.20%.

In fact, the empirical evidence in this paper confirms that EPU has a non-linear impact on ER, providing an important reference for central banks when making intervention decisions in the foreign exchange market. This study also highlights the critical importance of tracking regimes of EPU and ER misalignment for ensuring ER stability. Investors can even use ER misalignment as an indicator of EPU to make informed decisions about holding foreign currency assets. Governments and policymakers should increase policy transparency to improve "expectation management" policies, helping investors better anticipate market fluctuations. Furthermore, unnecessary policy changes should be minimised to maintain a stable economic environment, as predictable markets are crucial for maintaining ER equilibrium.

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Author contribution
All authors contributed to the conception and design of the study, data collection, analysis and interpretation of the results, and the writing of the manuscript. All authors read and approved the final version of the manuscript.

Use of AI tools declaration
Artificial intelligence tools were used to assist with language editing and to enhance the clarity and readability of the manuscript. The authors remain fully responsible for the content and conclusions of this study.

Conflict of interest
The authors declare that there are no competing interests related to this manuscript.

References

Abere, S. S. (2025). Asymmetric impacts of exchange rate and petroleum pump price on economic welfare in Nigeria. *Economic Journal of Emerging Markets*, 219–234. https://doi.org/10.20885/ejem.vol17.iss2.art8

Abid, A. (2020). Economic policy uncertainty and exchange rates in emerging markets: Short and long runs evidence. *Finance Research Letters*, *37*, 101378. https://doi.org/https://doi.org/10.1016/j.frl.2019.101378

Afat, D., Gómez-Puig, M., & Sosvilla-Rivero, S. (2015). The failure of the monetary model of exchange rate determination. *Applied Economics*, *47*(43), 4607–4629. https://doi.org/10.1080/00036846.2015.1031878

Aizenman, J., Ho, S.-H., Huynh, L. D. T., Saadaoui, J., & Uddin, G. S. (2024). Real exchange rate and international reserves in the era of financial integration. *Journal of International Money and Finance*, *141*, 103014. https://doi.org/10.1016/j.jimonfin.2024.103014

Baker, S. R., Bloom, N., & Davis, S. J. (2016). *Measuring economic policy uncertainty*. https://www.policyuncertainty.com/media/EPU_BBD_Mar2016.pdf

Banerjee, K., & Goyal, A. (2021). Behavioural equilibrium real exchange rates and misalignments: Evidence from large emerging markets. *Economic Analysis and Policy*, *70*, 414–436. https://doi.org/https://doi.org/10.1016/j.eap.2021.03.004

Bernanke, B. S. (1983). Irreversibility, uncertainty, and cyclical investment. *The Quarterly Journal of Economics*, *98*(1), 85. https://doi.org/10.2307/1885568

Ca' Zorzi, M., & Rubaszek, M. (2023). How many fundamentals should we include in the behavioral equilibrium exchange rate model? *Economic Modelling*, *118*, 106071. https://doi.org/10.1016/j.econmod.2022.106071

Chin, L., Azali, M., & Matthews, K. G. (2007). The monetary approach to exchange rate determination for Malaysia. *Applied Financial Economics Letters*, *3*(2), 91–94. https://doi.org/10.1080/17446540600993845

Chin, L., Azali, M., Yusop, Z. B., & Yusoff, M. B. (2007). The monetary model of exchange rate: evidence from The Philippines. *Applied Economics Letters*, *14*(13), 993–997. https://doi.org/10.1080/13504850600706032

Chin, L., Habibullah, M. S., & Azali, M. (2009). Tests of different monetary aggregates for the monetary models of the exchange rate in five ASEAN countries. *Applied Economics*, *41*(14), 1771–1783. https://doi.org/10.1080/00036840902845517

Clark, P. B., & MacDonald, R. (1999). Exchange rates and economic fundamentals: a methodological comparison of BEERs and FEERs. In *Equilibrium exchange rates* (pp. 285–322). Springer.

Clark, P. B., & MacDonald, R. (2004). Filtering the BEER: A permanent and transitory decomposition. *Global Finance Journal*, *15*(1), 29–56. https://doi.org/10.1016/j.gfj.2003.10.005

Cong, S., Chin, L., & Kumarusamy, R. (2025). Does trade freedom affect exchange rate movement? A perspective of high-technology trade. *The Journal of International Trade & Economic Development*, *34*(3), 552–572. https://doi.org/10.1080/09638199.2024.2320164

Cui, L., & Si, D.-K. (2024). Does RMB exchange rate regime reform decrease corporate risk-taking? Evidence from China. *Finance Research Letters*, *69*, 106111. https://doi.org/https://doi.org/10.1016/j.frl.2024.106111

Davis, S. J., Liu, D., & Sheng, X. S. (2019). Economic policy uncertainty in China since 1949: The view from mainland newspapers. *Fourth Annual IMF-Atlanta Fed Research Workshop on China's Economy Atlanta*, *19*, 1–37.

Dixit, A. (1989). Entry and exit decisions under uncertainty. *Journal of Political Economy*, *97*(3), 620–638. https://doi.org/10.1086/261619

Dornbusch, R. (1976). Expectations and exchange rate dynamics. *Journal of Political Economy*, *84*(6), 1161–1176. https://doi.org/10.1086/260506

Edwards, S. (1989). Exchange rate misalignment in developing countries. *The World Bank Research Observer*, *4*(1), 3–21. https://doi.org/10.1093/wbro/4.1.3

Frenkel, J. A. (1976). Inflation and the formation of expectations. *Journal of Monetary Economics*, *1*(4), 403–421. https://doi.org/10.1016/0304-3932(76)90011-8

Funke, M., Shu, C., Cheng, X., & Eraslan, S. (2015). Assessing the CNH–CNY pricing differential: Role of fundamentals, contagion and policy. *Journal of International Money and Finance*, *59*, 245–262. https://doi.org/10.1016/j.jimonfin.2015.07.008

Glebocki, H., & Saha, S. (2024). Global uncertainty and exchange rate conditions: Assessing the impact of uncertainty shocks in emerging markets and advanced economies. *Journal of International Financial Markets, Institutions and Money*, *96*, 102060. https://doi.org/10.1016/j.intfin.2024.102060

Gong, Y., He, Z., & Xue, W. (2025). EPU spillovers and exchange rate volatility. *International Review of Financial Analysis*, *97*, 103824. https://doi.org/10.1016/j.irfa.2024.103824

Goswami, M., Pontines, V., & Mohammed, Y. (2023). Portfolio capital flows and the US dollar exchange rate: Viewed from the lens of time and frequency dynamics of connectedness. *International Review of Financial Analysis*, *89*, 102754. https://doi.org/https://doi.org/10.1016/j.irfa.2023.102754

Holtemöller, O., & Mallick, S. (2013). Exchange rate regime, real misalignment and currency crises. *Economic Modelling*, *34*, 5–14. https://doi.org/10.1016/j.econmod.2012.09.017

Kim, H.-E., Lee, W., & Moon, J. (2025). US economic policy uncertainty and China's outward investment diversification. *Finance Research Letters*, *73*, 106620. https://doi.org/10.1016/j.frl.2024.106620

Korap, L. (2024). Impact of asymmetry on exchange rate determination: The role of fundamentals. *Emerging Markets Review*, *63*, 101206. https://doi.org/10.1016/j.ememar.2024.101206

Liu, B. (2024). Economic policy uncertainty, cross-border capital inflows and credit asset allocation risk. *Finance Research Letters*, *68*, 105984. https://doi.org/10.1016/j.frl.2024.105984

MacDonald, R., & Taylor, M. P. (1991). Exchange rates, policy convergence, and the European Monetary System. *The Review of Economics and Statistics*, 553–558.

Msomi, S., & Ngalawa, H. (2024). The monetary model of exchange rate determination for south africa. *Economies*, *12*(8), 206. https://doi.org/10.3390/economies12080206

Nguyen, T. H., Nguyen, T. G., & Nguyen, C. C. (2025). Economic policy uncertainty and foreign direct investment: The role of social connectedness in Vietnam. *Economic Journal of Emerging Markets*, 1–11. https://doi.org/10.20885/ejem.vol17.iss1.art1

Tong, H., & Lim, K. S. (1980). Threshold autoregression, limit cycles and cyclical data. *Journal of the Royal Statistical Society: Series B (Methodological)*, *42*(3), 245–268.

Wang, L., & Wang, X. (2015). Central bank expectation management, inflation volatility and bank risk-taking. *Economic Research Journal*, *50*(10), 34–48.

Wang, X. (2024). Extreme risk spillovers in RMB exchange rates: The role of categorical economic policy uncertainties. *International Review of Economics & Finance*, *94*, 103365. https://doi.org/10.1016/j.iref.2024.05.044

Wen, J., & Usman, M. A. (2024a). An empirical investigation of the relationship between real exchange rate and innovation: Evidence from China. *International Journal of Finance & Economics*, *29*(3), 2991–3006. https://doi.org/https://doi.org/10.1002/ijfe.2814

Wen, J., & Usman, M. A. (2024b). Measuring RMB equilibrium exchange rate and misalignment in China: Application of BEER and PEER approaches. *The Journal of International Trade & Economic Development*, *33*(7), 1323–1344. https://doi.org/10.1080/09638199.2023.2250471

Zhang, M., & Du, W. (2025). Economic policy uncertainty, investor sentiment and IPO underpricing. *International Review of Financial Analysis*, *105*, 104420. https://doi.org/10.1016/j.irfa.2025.104420

Zhang, W., & Huang, Z. (2025). Nonlinear dynamics of rmb exchange rate volatility: a multifractal perspective within the g-expectation framework. *Fractal and Fractional*, *9*(11), 746. https://doi.org/10.3390/fractalfract9110746

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Modelling the impact of Information and Communication Technology (ICT) trade, economic complexity and energy structure on carbon neutrality: Evidence from BRICS

Neharika Singh*, Meenakshi Gupta

School of Economics, Shri Mata Vaishno Devi University, Katra-182320, Jammu and Kashmir, India
*Corresponding Author: 21dse003@smvdu.ac.in

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Author's email:
meenakshi.gupta@smvdu.ac.in

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Abstract

Purpose — This study investigates the impact of Information and Communication Technology (ICT) trade flow components, specifically ICT service exports, ICT goods exports, and ICT goods imports, alongside economic complexity and renewable energy share on carbon emissions.

Methodology — A panel of BRICS countries from 2000 to 2022 is estimated using a second-generation cross-sectional autoregressive distributed lag (CS-ARDL) model that accounts for cross-sectional dependence and slope heterogeneity across countries.

Findings — Gross domestic product per capita and economic complexity are positively associated with carbon emissions. ICT trade flows have heterogeneous effects on emissions. ICT services, exports, and renewable energy consumption significantly reduce carbon emissions. However, ICT goods exports and imports have an insignificant effect on carbon emissions.

Implications — The results suggest that the BRICS countries must emphasise policy measures that promote the export of ICT services, accelerate renewable energy adoption, and promote industrial transformation policies towards sustainable production practices.

Originality — This study focuses on supply-side ICT trade channels and disaggregates them into ICT goods exports, imports, and service exports. Furthermore, this study applies second-generation estimation techniques that are robust to cross-sectional dependence and slope heterogeneity.

Keywords — ICT Trade Flows; Economic Complexity; Renewable Energy; CS-ARDL

Introduction

Balancing economic growth with environmental sustainability remains a significant challenge for emerging economies undergoing structural transformation (Shekhawat et al., 2025). Despite rapid economic and technological advancements characterised by the rapid growth of information and communication technology (ICT) and an evolving production structure marked by rising economic complexity and an increasing renewable energy transition, BRICS economies are experiencing a significant rise in carbon emissions. In recent years, these economies have accounted for about 42% of global carbon emissions (Erkılıç et al., 2025). This raises concerns regarding the environmental effects of their developmental pathways.

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These structural changes affect carbon emissions via different mechanisms. ICT may reduce emissions by facilitating technological innovations and digital solutions, promoting efficiency gains across industries (Kashif et al., 2024), whereas it may increase emissions by increasing overall energy demand (Irfan et al., 2025). Economic complexity may reduce emissions through industrial upgrading towards higher value-added and knowledge-intensive exports, whereas it may increase emissions through energy-intensive exports (Balsalobre-Lorente et al., 2023). Renewable energy reduces emissions through the energy substitution channel by substituting fossil fuels with cleaner energy sources (Shekhawat et al., 2025). While ICT development, an evolving production structure, and energy transition are viewed as catalysts for low-carbon and sustainable growth, their environmental implications yield mixed empirical evidence.

The empirical evidence examining the impact of ICT on carbon emissions remains inconclusive (Appiah-Otoo et al., 2023). Studies that measure ICT using demand-based indicators, such as internet adoption rates and mobile and telephone subscription rates, report mixed findings regarding its impact on carbon emissions. One strand of literature reports that ICT significantly improves environmental quality (Sun et al., 2026; Wen et al., 2025). These studies suggest that ICT reduces emissions by enabling resource optimisation, improves energy efficiency, and directs economic activity towards low-carbon service-based sectors (Zuo & Ren, 2025). Whereas other studies found that the growth of ICT increases electricity demand for digital infrastructure and related services, thereby increasing emissions (Adebayo et al., 2025; Le et al., 2025). The literature primarily focuses on demand-side indicators as proxies for digitalisation. It captures ICT adoption usage intensity and efficiency impacts.

However, these demand-side indicators of ICT do not account for the production and supply aspects of ICT, where energy consumption and emissions may be embedded. Thus, it may underestimate the sector's total environmental impact. However, the empirical literature on ICT focuses primarily on the supply side. Empirical studies have reported heterogeneous effects of ICT trade components on carbon emissions. Mhlanga (2025) found that ICT goods exports significantly increase carbon emissions in the case of BRICS countries. The finding is consistent with the energy-intensive nature of ICT goods production, including digital hardware, telecommunications devices, and electronic equipment, which depend on carbon-based inputs and energy-intensive production methods, thereby increasing carbon emissions (Irfan et al., 2025). Whereas Kashif et al. (2024) found that ICT services exports significantly reduce carbon emissions in the case of OECD countries. ICT services, including digital platforms, software, and cloud solutions, enhance environmental quality through dematerialisation and resource efficiency. It also boosts energy efficiency and facilitates sustainable production practices (Wen et al., 2025). F. Khan et al., (2025) investigated the impact of the ICT trade index (measured by using exports and imports of ICT goods and ICT services export indicators) on carbon emissions in the case of five selected Asian countries. The findings indicate that ICT significantly reduces carbon emissions.

However, studies either use individual ICT trade measures or an aggregated measure, neither of which reflects the significant variation among components of ICT trade associated with different energy intensities and production processes. While ICT services exports represent knowledge-based activities with lower carbon intensity, ICT goods trade is associated with energy-intensive production and embodied carbon emissions (Kim, 2024; Lange et al., 2020). It is therefore important to analyse the varying impact of components of ICT trade on carbon emissions. To address this, the present study examines the impact of disaggregated ICT trade channels, including ICT goods exports, ICT goods imports, and ICT services exports, on carbon emissions. Beyond digitalisation, the study has examined structural transformation through the concept of economic complexity. The Economic Complexity Index (ECI) was developed by Hidalgo & Hausmann (2009). It measures the productive knowledge, skills, and know-how embedded in the economy's production structure, as reflected in its capacity to manufacture and export a wide range of sophisticated goods.

Empirical evidence on the environmental impact of economic complexity remains mixed. Some studies found that economic complexity is associated with increased carbon emissions (Balsalobre-Lorente et al., 2023; Feng et al., 2024). These studies suggest that in the early stages of development, increasing economic complexity indicates a shift from agriculture-based exports towards energy-intensive manufacturing-based exports. This transition increases industrial activity and energy demand, leading to higher fossil fuel consumption and, in turn, higher carbon emissions. Whereas other studies found that economic complexity is associated with reduced carbon emissions (Haciimamoglu, 2025; Tabash et al., 2024). Countries with a greater level of complexity specialise in higher-value-added knowledge-intensive exports. Their accumulation of advanced capabilities facilitates technological innovations and energy-efficient, cleaner production methods, thereby reducing carbon emissions. This contrasting evidence suggests that the environmental impact of economic complexity depends on the prevalent production structure and energy mix. In parallel, the renewable energy transition has been widely regarded as a crucial strategy for reducing carbon emissions. Studies have consistently found that renewable energy significantly reduces carbon emissions (Balsalobre-Lorente et al., 2023; Zuo & Ren, 2025).

Despite a growing body of literature, existing literature largely examines information and communication technology (ICT), economic complexity, and renewable energy separately, with limited attention to their combined influence on carbon emissions. This fragmented approach limits understanding of how ongoing structural transformations influence carbon emissions. In addition, most ICT-related studies rely on general ICT adoption metrics, such as internet access and subscription rates. However, these metrics do not capture the supply-side ICT trade flows, which influence production efficiency, technology diffusion, and carbon emissions differently across economies. Therefore, these limitations highlight the need of an integrated analytical framework that accounts for ICT trade, the complexity of economic systems, and the energy structure to provide a comprehensive understanding of how these interconnected factors influence environmental quality.

In this context, the present study extends the STIRPAT model framework to examine the impact of ICT trade flows (exports of ICT services, exports of ICT goods, and imports of ICT goods), economic complexity, and renewable energy on carbon emissions across BRICS economies. The present study contributes to the existing literature in several ways. First, the study includes supply-side ICT trade variables comprising ICT services exports, ICT goods exports and ICT goods imports to assess their environmental impact, thereby moving beyond demand-side ICT usage indicators and providing fresh insights on the impact of ICT trade components on environmental quality. Second, the study includes economic complexity as a structural transformation variable, providing a novel perspective on the impact of evolving productive knowledge and capability structure on environmental quality. Third, the study includes the share of renewable energy (% of total final energy consumption) to analyse the impact of the economy's prevalent energy mix on environmental quality. The study contributes theoretically by extending the STIRPAT model to include ICT trade, economic complexity, and renewable energy as key technological variables, and by analysing how technological progress, underlying production capabilities, and energy structure influence carbon emissions. Additionally, as a methodological advancement, the study employs a cross-sectional autoregressive distributed lag (CS-ARDL) model that accounts for cross-sectional dependence and slope heterogeneity. This is particularly relevant for highly interconnected economies like BRICS, where cross-sectional dependence and slope heterogeneity are prevalent.

The analysis is essential for understanding the development pathways of BRICS economies in the digital era. From a policy perspective, understanding the environmental trade-off associated with ICT-driven industrialisation is essential, as the BRICS economies strive for sustainable growth and greater integration into the global digital trade network. It provides policy insights on how emerging economies might leverage ICT and technological advancement while ensuring environmental sustainability.

Methods

Data Source and Description

The study is conducted on a sample of BRICS countries, namely Brazil, Russia, India, China, and South Africa, for the period 2000 to 2022, based on data availability. Table 2 provides a brief description of the variables employed and data sources.

Table 1. Description of Variables and Data Sources

Variables	Symbols	Units	Sources
Carbon emissions per capita	COEPC	Tonnes per person	Our World in Data
Urban Population	URB	Per cent of total population	Our World in Data
Gross domestic product per capita	GDPPC	constant 2021 US$	Our World in Data
Exports of ICT Services	ICTSER	Per cent of total services exports	World Development Indicators
Exports of ICT goods	ICTEXP	Per cent of total goods exports	World Development Indicators
Imports of ICT goods	ICTIMP	Per cent of total goods exports	World Development Indicators
Economic complexity index	ECI	Index	Atlas of Economic Complexity
Renewable energy share	RES	Per cent total final energy consumption	Our World in Data

Theoretical Framework and Model Specification

The study analyses the determinants of carbon emissions using the Stochastic Impacts by Regression on Population, Affluence, and Technology (STIRPAT) model framework to determine the role of human activities on environmental quality (Dietz & Rosa, 1997). The model can be described using the following equation:

$$I_{it} = aP_{it}^{b}A_{it}^{c}T_{it}^{d}\varepsilon_{it} \tag{1}$$

where *I* represents environmental impact, *P* represents population, *A* represents affluence, and *T* represents the technology variable; ε_{it} represents the error term; *a* is the constant term; *b*, *c*, and *d* are parameters associated with population, affluence and technology variables; *i* is the cross-section unit, and *t* is the time period.

To estimate equation 1, all variables are transformed to their natural logarithms to account for heteroskedasticity and to allow interpretation of the coefficients in elasticities. For the empirical analysis, the model is transformed into log-linear form:

$$lnI_{it} = \alpha + blnP_{it} + clnA_{it} + dlnT_{it} + \varepsilon_{it} \tag{2}$$

In the study, environmental impact is represented by carbon emissions per capita $LnCOEPC_{it}$. Within the STIRPAT framework, the population effect measures the population concentration. It is represented by using the urban population (% of total population), indicated by $LnURB_{it}$. Higher urbanisation is expected to increase carbon emissions ($\beta_1 > 0$) due to increased energy demand, infrastructure expansion and transportation requirements in BRICS (Singh et al., 2024; Wang et al., 2024). Affluence measures the intensity of economic activity per person. It is represented by using gross domestic product per capita, indicated by $LnGDPPC_{it}$. In BRICS countries, higher economic growth is characterised by rapid industrialisation and energy-intensive economic activities. Economic growth increases overall output production and energy consumption, which in turn increases carbon emissions (Singh et al., 2024; Ullah et al., 2023). Therefore, the study hypothesizes positive relationship between GDPC and COEPC ($\beta_2 > 0$).

The study extends the technological component of the STIRPAT model framework to capture ongoing structural transformations in economies characterised by rapid digitalisation (ICT goods exports, ICT goods imports, and ICT services exports), evolving production structures (economic complexity), and the energy transition (renewable energy share). The digitalisation aspect is represented by disaggregated ICT trade components i.e. ICT services exports (% of total services exports) indicated by $LnICTSER_{it}$, ICT goods exports (% of total goods exports) indicated by $LnICTEXP_{it}$ and ICT goods imports (% of total goods exports) indicated by $LnICTIMP_{it}$.

ICT services exports (% of total services exports) reflect an economy's ability to export knowledge-based digital services. The export of ICT services, such as computer programming, consulting, and information services, supports low-carbon and energy-efficient technologies. It also facilitates industrial structure upgrading as economies move away from heavy industrial production towards cleaner, knowledge-intensive industries (Kashif et al., 2024). They are less energy-intensive and resource-intensive than the manufacturing of ICT goods. Thus, it is hypothesised that the expected sign of ICT services exports on carbon emissions is negative (beta sub 3 less than 0).

ICT goods exports (as % of total exports) represent technology diffusion through the trade channel. However, the production of ICT goods such as computers, phones, semiconductors, and other hardware is energy-intensive. Additionally, the manufacturing and disposal of ICT equipment pose challenges for electronic waste management and resource depletion (Ghulam & Abushammala, 2023). Thus, higher ICT goods exports lead to greater production and, hence, an increase in carbon emissions (Mhlanga, 2025). Thus, it is hypothesised that the expected sign of ICT goods exports with carbon emissions is positive in BRICS $(\beta_4 > 0)$.

ICT goods imports (% of total goods exports) facilitate the adoption of advanced energy-efficient digital technologies within the economy. Hence, it reflects the technique effect. Imports of ICT goods reduce emissions through the adoption of cleaner manufacturing processes, energy-efficiency improvements, and dematerialisation within the economy (A. Khan & Ximei, 2022). However, ICT goods imports may also increase energy consumption and emissions (Irfan et al., 2025). Thus, it is hypothesised that the expected sign of ICT goods imports with carbon emissions is ambiguous $(\beta_5 > 0 \text{ or } \beta_5 < 0)$.

In addition to ICT-based digitalisation, the technology component encompasses structural transformation in production, as measured by the Economic Complexity Index. It is indicated by ECI_{it}. It measures the diversity and sophistication of the production structure, as reflected in the economy's export composition of the economy. A higher economic complexity reflects an economy's ability to produce technologically advanced and knowledge-intensive exports that are less carbon-intensive. A lower economic complexity reflects a production structure characterised by resource- and energy-intensive manufacturing, leading to increased carbon emissions (Haciimamoglu, 2025). In the study, Economic complexity in the BRICS economies is often associated with energy-intensive exports, which leads to a rise in emissions (Balsalobre-Lorente et al., 2023). Therefore, ECI is hypothesised to have a positive impact on emissions $(\beta_2 > 0)$.

Further technology components are extended to include the adoption of renewable energy. The study uses the renewable energy share (% of total final energy consumption) indicated by $LnRES_{it}$. It reflects the transition from a carbon-intensive energy structure towards a more sustainable energy mix. Thus, it represents clean energy technology that reduces carbon emissions, as it has a lower life-cycle carbon footprint (Shekhawat et al., 2025). Thus, it is hypothesised that the expected sign of the Renewable energy share (% of total final energy consumption) and carbon emissions is negative $(\beta_6 < 0)$.

Thus, the extended model is specified as:

$$LnCOEPC_{it} = \alpha_0 + \beta_1 LnURB_{it} + \beta_2 LnGDPPC_{it} + \beta_3 LnICTSER_{it} + \beta_4 LnICTEXP_{it} + \beta_5 LnICTIMP_{it} + \beta_6 ECI_{it} + \beta_7 LnRES_{it} + \varepsilon_{it} \tag{3}$$

In equation (3), α_0 represents the constant term, and $\beta_1; \beta_2; \beta_3; \beta_4; \beta_5; \beta_6; \beta_7$ are the model coefficients to be estimated in the model analysis and ε_{it} represents the error term.

Panel Estimation Technique

The study conducts various pre-diagnostic tests, including the cross-sectional dependence test, slope homogeneity test, and panel unit root test, to select the appropriate methodology for the empirical analysis.

Cross-Sectional Dependence (CSD) test

Cross-sectional dependence arises because of unobserved common factors, global shocks or local spillovers which affect all the individual units (e.g. countries or firms) simultaneously. Hence, the residuals are correlated across individual units in a panel data model. To check the cross-sectional dependence across the panel, the study uses the cross-sectional dependence (CSD) test developed by Pesaran (2004). It can be expressed as,

$$Pesaran\ CSD = \sqrt{\frac{2T}{N(N-1)}}\left(\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\hat{\rho}_{ij}\right) \tag{4}$$

Here, N represents the number of cross-sectional units, T represents the time period, and $\hat{\rho}_{ij}$ represents the pairwise correlation of residuals between unit i and unit j. The hypothesis of the test is that the null hypothesis states that there is no cross-sectional dependence among the cross-sectional units, while the alternative hypothesis states that there is cross-sectional dependence.

Second-Generation Panel Unit-Root Test

To determine the stationarity properties of the variables, the study uses the second-generation Cross-Sectional Augmented Im, Pesaran, and Shin (CIPS) test (Pesaran, 2007). The test statistics of CIPS is calculated as:

$$CIPS(N,T) = \frac{1}{N}\sum_{i=1}^{N} CADF_i(N,T) \tag{5}$$

Here N is the cross-sections, T is the time and $CADF_i(N,T)$ is the Cross-sectionally augmented Dickey-Fuller statistic for i[th] cross-section. It is obtained from the regression:

$$\Delta Y_{i,t} = \alpha_i + \beta_i Y_{i,t-1} + c_i\bar{Y}_{t-1} + \sum_{j=0}^{p} d_{ij}\Delta\bar{Y}_{t-j} + \sum_{j=1}^{p}\phi_{ij}\,\Delta Y_{i,t-j} + \varepsilon_{it} \tag{6}$$

The hypothesis of the test is that the null hypothesis states that there is a unit root present in the variable, while the alternative hypothesis states that there is no unit root present in the variable.

Slope Homogeneity Test

The Slope homogeneity test is used to determine if the slope coefficient is constant across various cross-sectional units within a dataset. In panel data analysis, it is used to choose between the pooled estimation technique, which assumes identical slopes across cross-sectional units in the panel, and the estimation technique that accounts for slope heterogeneity (Pesaran & Yamagata, 2008). It is represented by the following equations:

$$\Delta_{SH} = (N)^{\frac{1}{2}}(2k)^{\frac{1}{2}}\left(\frac{1}{N}\tilde{S} - k\right) \tag{7}$$

$$\tilde{\Delta}_{ASH} = (N)^{\frac{1}{2}}\left(\frac{2k(T-k-1)}{T+1}\right)^{-\frac{1}{2}}\left(\frac{1}{N}\tilde{S} - 2k\right) \tag{8}$$

where Δ_{SH} is delta tilde and $\tilde{\Delta}_{ASH}$ represents the adjusted delta tilde test statistic. $\tilde{S}$ is swamy test statistic, K is the number of independent variables, and N is the number of cross-sectional units. The hypothesis of the test is that the null hypothesis states that the slope coefficients are homogeneous across all units, while the alternative hypothesis states that they are heterogeneous across units.

Cross-Sectionally Augmented Autoregressive-Distributed Lag (CS-ARDL) Model

Finally, to estimate long-run coefficients, the study uses a second-generation cross-sectionally augmented autoregressive distributed lag (CS-ARDL) model (Chudik & Pesaran, 2015). The model addresses the cross-sectional dependence and accounts for slope heterogeneity. The generalised form of the model can be represented by the following equation:

$$y_{it} = \alpha_i + \sum_{l=1}^{p_y}\phi_{il}y_{i,t-l} + \sum_{l=0}^{p_x}\beta'_{il}x_{i,t-l} + \sum_{l=0}^{p_{\bar{z}}}\gamma'_{il}\,\bar{z}_{t-l} + \varepsilon_{it} \tag{9}$$

Where, y_{it} is carbon emissions per capita, $x_{i,t}$ is vector of explanatory variables and $\bar{z}_t$ is cross-sectional averages of both dependent and independent variables. Furthermore, to ensure robustness of the baseline model, the study employs a fully modified ordinary least squares (FMOLS) estimator, which corrects for endogeneity of regressors and serial correlation in cointegrated panels (Phillips & Hansen, 1990).

Results and Discussion

The findings of Pesaran, (2004) test for the presence of cross-sectional dependence in the panel are reported in Table 2. For most of the variables, including COEPC, URB, GDPPC, ICTSER, ICTEXP and RES, the CD test statistic is significant at the 1% level of significance (p-value < 0.01). So, we reject the null hypothesis (H_0) for the test, i.e. the series is cross-sectionally independent. Thus, the CSD test confirms that the countries are cross-sectionally dependent. Because CSD is present, the conventional first-generation panel unit root test may yield biased and inconsistent estimates. Therefore, the study proceeds with a second-generation unit-root test that accounts for CSD.

Table 2. Cross-sectional dependence test

Variable	CD-test	p-value	Average joint T	mean ρ	Mean abs ρ
COEPC	5.68	0.00	23.00	0.37	0.59
URB	14.35	0.00	23.00	0.95	0.95
GDPPC	13.48	0.00	23.00	0.89	0.89
ICTSER	9.19	0.00	23.00	0.61	0.61
ICTEXP	-2.37	0.01	23.00	-0.16	0.30
ICTIMP	1.86	0.06	23.00	0.12	0.43
ECI	-1.95	0.05	23.00	-0.13	0.68
RES	6.80	0.00	23.00	0.45	0.46
H_0: Series is cross-sectionally independent.					

The findings of the second-generation panel CIPS unit-root test are reported in Table 3. At the level form, the variables COEPC, GDPPC, ICTEXP, ECI and RES do not display statistically significant CIPS statistic (p-value > 0.05). These variables are therefore non-stationary in level form, as we fail to reject the null hypothesis (i.e., the presence of a unit root). In the first difference, the variables COEPC, GDPPC, ICTEXP, ECI, and RES show statistically significant CIPS statistics (p-values < 0.05). This indicates that these variables become stationary after first differencing; thus, they are integrated of order I(1). ICTSER and ICTIMP report statistically significant CIPS statistics at the level (p-value < 0.05). Whereas URB shows a statistically significant CIPS statistic at the 10% level of significance. Thus, these variables are integrated of order I (0). Overall, the findings confirm a mixed order of integration across variables.

Table 3. Second-generation CIPS unit root test

Variable	CIPS statistic at the level	CIPS statistic at First-difference	Order of Integration
COEPC	-1.14	-3.34***	I (1)
URB	-2.26*	-	I (0)
GDPPC	-1.93	-2.90***	I (1)
ICTSER	-2.86***	-	I (0)
ICTEXP	-1.54	-3.55***	I (1)
ICTIMP	-2.58***	-	I (0)
ECI	-1.31	-4.79***	I (1)
RES	-1.93	-5.04***	I (1)
H_0: There is a unit root.			

Note: ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

The findings of the Pesaran & Yamagata (2008) slope homogeneity test are reported in Table 4. It indicates that both the delta statistic and the adjusted delta test statistic are statistically significant at the 1% level (p-value < 0.01). Thus, we reject the null hypothesis and confirm that the slope coefficients are heterogeneous across the panel.

The diagnostic test in Table 4 is used to select the appropriate estimation method. The panel shows cross-sectional dependence, slope heterogeneity and mixed order of integration across variables. Conventional panel estimation models that rely on homogeneous slope coefficients and cross-sectional independence may not provide robust estimates. Thus, to account for unobserved common factors, heterogeneous slope coefficients, and variables exhibiting mixed orders of integration, the study uses a second-generation cross-sectionally augmented autoregressive distributed lag (CS-ARDL) model (Chudik & Pesaran, 2015). The findings are reported in Table 5.

Table 4. Slope heterogeneity test

Test	Statistic	p-value
Delta	6.46	0.00
Delta Adj.	8.28	0.00
H_0: Slope coefficients are homogeneous.		

Note: ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

Table 5. Long-Run and Short-Run Estimates from the CS-ARDL Model

LNCOEPC	Coef.	Std. Err.	z	P > |z|	[95% Conf. Interval]	
Short Run Est.						
Mean Group:						
LNURB	0.469	2.27	0.21	0.83	-3.980	4.919
LNGDPC	0.812	0.341	2.38	0.01	0.142	1.482
LNSERV	-0.088	0.040	-2.18	0.02	-0.167	-0.009
LNEXP	-0.001	0.050	-0.02	0.98	-0.100	0.097
LNIMP	0.082	0.065	1.26	0.20	-0.045	0.210
ECI	0.088	0.034	2.56	0.01	0.020	0.155
LNRES	-0.381	0.148	-2.57	0.01	-0.671	-0.090
Adjust. Term						
Mean Group:						
lr_LNCOEPC	-1.04	0.090	-11.52	0.00	-1.21	-0.865
Long Run Est.						
Mean Group:						
lr_LNURB	0.172	1.962	0.09	0.93	-3.673	4.017
lr_LNGDPC	0.689	0.315	2.18	0.02	0.070	1.308
lr_LNSERV	-0.088	0.041	-2.14	0.03	-0.169	-0.007
lr_LNEXP	0.002	0.048	0.05	0.95	-0.091	0.096
lr_LNIMP	0.092	0.071	1.29	0.19	-0.048	0.232
lr_ECI	0.095	0.035	2.70	0.00	0.026	0.164
lr_LNRES	-0.359	0.145	-2.47	0.01	-0.644	-0.074
F (85,25) = 3.88						
Prob > F = 0.00						
R-squared (MG) =0.99						
CD Statistic = -1.26 (p-value = 0.206)						

Note: The CD test statistic indicates no significant cross-sectional dependence in the residuals (p-value > 0.05), thus confirming the reliability of the estimates.

The findings show that GDPPC has a positive and statistically significant effect on COEPC in both the short and long run. The estimated coefficient indicates that a 1-unit increase in GDPPC is associated with a 0.68-unit increase in carbon emissions per capita in the long run. This suggests that an increase in economic growth is associated with rising environmental pressures in the case of BRICS countries. This pattern indicates the structural reliance of these countries on energy-intensive industrialisation, infrastructure development, and continued reliance on fossil fuels, resulting in increased carbon emissions. The finding is consistent with the recent studies by Singh et al., (2024) and Ullah et al., (2023).

Furthermore, the findings reveal the differing impacts of ICT goods exports, ICT goods imports, and ICT services exports on per capita carbon emissions across BRICS countries. While ICT services exports reduce emissions, ICT goods exports and imports show no significant effect.

The findings show that ICT service exports have a negative and statistically significant effect on per capita carbon emissions in both the short and long run. The estimated coefficient indicates that a unit increase in ICT services exports decreases carbon emissions per capita by 0.08 units in the long run. The findings support the hypothesis that the integration of digital services into the economy facilitates the transition towards a less carbon-intensive economy. ICT services such as telecommunications, consultancy, courier and information services, and software solutions promote energy efficiency and reduce emissions. Exports of ICT services shift economic activity towards lower-emission, knowledge-intensive sectors, thereby reducing structural dependence on energy-intensive sectors. Further, ICT services enable industries to adopt digital technologies that enhance production efficiency and reduce energy use per unit of output, representing a technique effect. The findings of this study are consistent with the study by Kashif et al., (2024). Thus, the expansion of ICT services exports in international trade can be a viable strategy to decouple economic growth from emissions, as these services have a comparatively lower carbon footprint than manufacturing ICT goods, indicating a sustainable, technology-driven development pathway.

The findings show that economic complexity has a positive and statistically significant effect on carbon emissions per capita both in the short and long run. The estimated coefficient indicates that with a one-unit increase in economic complexity, carbon emissions per capita increase by 0.09 units in the long run. This implies that rising economic complexity in these countries is associated with resource-intensive manufacturing, which increases emissions. Our results are consistent with the previous studies by Balsalobre-Lorente et al., (2023) and Haciimamoglu, (2025).

The findings show that renewable energy has a negative and statistically significant effect on per capita carbon emissions in both the short and long run. The findings indicate that with a one-unit increase in the renewable energy share, carbon emissions per capita decrease by 0.35 units in the long run. The reason could be that renewable energy resources have lower life cycle emissions than fossil fuels. The finding supports the anticipated environmental advantage of switching towards cleaner energy sources. The results align with the work published by Shekhawat et al., (2025).

Table 6. Robustness Analysis Using the FMOLS Model

Variable	Coefficient	Std. Error	t-statistic	prob
LNURB	-2.72	0.58	-4.67	0.000
LNGDPC	1.520	0.269	5.651	0.000
LNSERV	-0.578	0.084	-6.813	0.000
LNEXP	-0.083	0.120	-0.688	0.492
LNIMP	-0.245	0.342	-0.716	0.475
ECI	0.614	0.274	2.238	0.027
LNRES	-0.169	0.060	-2.787	0.006
R-Squared	0.751			
Adjusted R-Squared	0.736			

Further, the findings of the robustness analysis are provided in Table 6. The findings are broadly consistent with the baseline model. Gross domestic product per capita and economic complexity reveal a positive and significant correlation with emissions. Renewable energy share and ICT services exports reveal a negative and significant correlation with emissions. Whereas ICT goods exports and imports remain statistically insignificant. However, the effect of urbanisation on emissions becomes negative and significant in the robustness model, whereas it was insignificant in the CS-ARDL model, indicating its effect is sensitive to model specification.

Conclusion

The present study examined the impact of individual components of ICT trade (ICT services exports, ICT goods exports and ICT goods imports), economic complexity index (ECI) and renewable energy share (RES) on carbon emissions per capita (COEPC) using the STIRPAT model framework. Methodologically, to account for cross-sectional dependence, heterogeneity of slope coefficients, and mixed order of integration, the panel study used a second-generation cross-sectionally augmented autoregressive distributed lag (CS-ARDL) model for the panel of BRICS over the period 2000-2022. The findings reveal that ICT service exports and renewable energy significantly reduce carbon emissions, indicating their positive role in improving environmental quality. Whereas gross domestic product per capita and economic complexity index significantly increase carbon emissions, indicating their harmful effects on environmental sustainability. However, ICT goods exports and imports have no significant effect on carbon emissions. The findings suggest that increasing the adoption of renewable energy and ICT services can help reduce carbon emissions. However, their impact on emissions reduction remains limited relative to the emissions pressure associated with economic growth and rising economic complexity. Therefore, a more effective decarbonization pathway requires aligning these mitigation channels with structural changes in the production system, such as reducing the carbon intensity of complex industries and integrating renewable energy into high-value-added sectors. Thus, renewable energy and ICT services can be an enabling factor in a more comprehensive transition strategy to achieve carbon neutrality objectives in BRICS countries (Balsalobre-Lorente et al., 2023).

Our findings present the following policy implications for enhancing environmental quality in selected economies through ICT trade and the adoption of renewable energy. First, the findings indicate that exports of ICT services are negatively correlated with carbon emissions, suggesting that the expansion of digital services improves environmental quality. Policies should be formulated to facilitate the growth of ICT service industries by building an enabling environment through investment in digital infrastructure and digital skills development. In addition, promoting ICT-related services, including software development, cloud computing, IT consulting, e-governance, and digital banking, promotes dematerialisation and is less energy-intensive. Thus, exports of ICT services can serve as a low-carbon growth strategy for the BRICS economies. Second, the detrimental effect of economic complexity on environmental quality suggests that countries must integrate industrial upgrading strategies with environmental objectives. Policies should be formulated to channelise investments in the development of green technological capabilities, such as research and development (R&D) in renewable energy technologies and energy-efficient manufacturing methods. Further, the study highlights the positive role of renewable energy in environmental sustainability. Policymakers should prioritise investments in renewable energy infrastructure, enhance renewable energy subsidies, and implement regulatory frameworks that encourage business transition towards sustainability.

In conclusion, our findings suggest that ICT-driven growth, characterised by ICT services exports, can promote sustainability and must be integrated with renewable energy technologies, energy-efficient measures, and innovation-driven industrial upgrading strategies.

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Author contribution

All authors contributed to the conception and design of the study, data collection, analysis and interpretation of the results, and the writing of the manuscript. All authors read and approved the final version of the manuscript.

Use of AI tools declaration

Artificial intelligence tools were used to assist with language editing and to enhance the clarity and readability of the manuscript. The authors remain fully responsible for the content and conclusions of this study.

Conflict of interest
The authors declare that they have no known competing financial interests or personal relationships that could have appeared to influence the work reported in this paper.

Availability of data and material
The dataset analysed during the current study is available from the corresponding author on reasonable request.

References

Adebayo, T. S., Özkan, O., Uzun Ozsahin, D., Eweade, B. S., & Gyamfi, B. A. (2025). Exploring the role of ICT adoption technologies and renewable energy consumption in achieving a sustainable environment in the United States: an SDGs-based policy framework. *Environmental Sciences Europe*, *37*(1), 20. https://doi.org/10.1186/s12302-025-01059-z

Appiah-Otoo, I., Acheampong, A. O., Song, N., & Chen, X. (2023). The impact of information and communication technology (ICT) on carbon dioxide emissions: Evidence from heterogeneous ICT countries. *Energy & Environment*, *34*(8), 3080–3102. https://doi.org/10.1177/0958305X221118877

Balsalobre-Lorente, D., Contente dos Santos Parente, C., Leitão, N. C., & Cantos-Cantos, J. M. (2023). The influence of economic complexity processes and renewable energy on CO2 emissions of BRICS. What about industry 4.0? *Resources Policy*, *82*, 103547. https://doi.org/10.1016/j.resourpol.2023.103547

Chudik, A., & Pesaran, M. H. (2015). Common correlated effects estimation of heterogeneous dynamic panel data models with weakly exogenous regressors. *Journal of Econometrics*, *188*(2), 393–420. https://doi.org/10.1016/j.jeconom.2015.03.007

Dietz, T., & Rosa, E. A. (1997). Effects of population and affluence on CO $_2$ emissions. *Proceedings of the National Academy of Sciences*, *94*(1), 175–179. https://doi.org/10.1073/pnas.94.1.175

Erkılıç, E., Gazeloğlu, C., & Özgören Ünlü, E. (2025). Renewable energy solution to carbon emissions: BRICS countries in the grip of globalization and economic growth. *Sustainability*, *17*(9), 4117. https://doi.org/10.3390/su17094117

Feng, Q., Usman, M., Saqib, N., & Mentel, U. (2024). Modelling the contribution of green technologies, renewable energy, economic complexity, and human capital in environmental sustainability: Evidence from BRICS countries. *Gondwana Research*, *132*, 168–181. https://doi.org/10.1016/j.gr.2024.04.010

Ghulam, S. T., & Abushammala, H. (2023). Challenges and opportunities in the management of electronic qaste and its impact on human health and environment. *Sustainability*, *15*(3), 1837. https://doi.org/10.3390/su15031837

Haciimamoglu, T. (2025). Investigating the EKC and LCC hypotheses for BRICS countries: the role of economic complexity in environmental degradation. In *Natural Resources Forum*. Oxford, UK: Blackwell Publishing Ltd. https://doi.org/10.1111/1477-8947.70013

Hidalgo, C. A., & Hausmann, R. (2009). The building blocks of economic complexity. *Proceedings of the National Academy of Sciences*, *106*(26), 10570–10575. https://doi.org/10.1073/pnas.0900943106

Irfan, M., Quddus, A., Shahzad, F., & Wang, Y. (2025). Do ICT trade balances and natural resources foster carbon emissions? The role of government effectiveness and green technology innovation. *Structural Change and Economic Dynamics*, *72*, 320–329. https://doi.org/10.1016/j.strueco.2024.12.00 7

Kashif, U., Shi, J., Naseem, S., Dou, S., & Zahid, Z. (2024). ICT service exports and CO2 emissions in OECD countries: the moderating effect of regulatory quality. *Economic Change and Restructuring*, *57*(3), 94. https://doi.org/10.1007/s10644-024-09685-y

Khan, A., & Ximei, W. (2022). Digital economy and environmental sustainability: Do information communication and technology (ICT) and economic complexity matter? *International Journal of Environmental Research and Public Health*, *19*(19), 12301. https://doi.org/10.3390/ijerph191912301

Khan, F., Zahir, S., Rahman, H. U., Raza, A., & Noor, S. (2025). Exploring the role of ICT and education in reducing environmental degradation among Asian countries. *Natural Resources Forum*, *49*(4), 3836–3865. https://doi.org/10.1111/1477-8947.12548

Kim, S. (2024). The Effect of Information and Communication Technology on Electricity Intensity in Korea. *Energies*, *17*(8), 1906. https://doi.org/10.3390/en17081906

Lange, S., Pohl, J., & Santarius, T. (2020). Digitalization and energy consumption. Does ICT reduce energy demand? *Ecological Economics*, *176*, 106760. https://doi.org/10.1016/j.ecolecon.2020.106760

Le, V. L. T., Gupta, R., Pham, K. D., & Nguyen, L. H. (2025). Financial inclusion, ICT development, and CO2 emissions: an ARDL approach. *Sustainable Development*, *33*(5), 7791-7806. https://doi.org/10.1002/sd.3547

Mhlanga, M. (2025). Digitalization and environmental sustainability: Evidence from selected BRICS economies. *SSRN Electronic Journal*. https://doi.org/10.2139/ssrn.5391625

Pesaran, M. H. (2004). General diagnostic tests for cross section dependence in panels. *SSRN Electronic Journal*. https://doi.org/10.2139/ssrn.572504

Pesaran, M. H. (2007). A simple panel unit root test in the presence of cross-section dependence. *Journal of Applied Econometrics*, *22*(2), 265–312. https://doi.org/10.1002/jae.951

Pesaran, M. H, & Yamagata, T. (2008). Testing slope homogeneity in large panels. *Journal of Econometrics*, *142*(1), 50–93. https://doi.org/10.1016/j.jeconom.2007.05.010

Phillips, P. C. B., & Hansen, B. E. (1990). Statistical inference in instrumental variables regression with I(1) Processes. *The Review of Economic Studies*, *57*(1), 99. https://doi.org/10.2307/2297545

Shekhawat, K. K., Kumar, N., Kumar, P., & Sharma, V. (2025). Impact of renewable energy, financial globalization, and technological innovation on environmental sustainability in BRICS. *Discover Sustainability*, *6*(1), 644. https://doi.org/10.1007/s43621-025-01486-1

Singh, G. J., Singh, P. K., & Lal, P. (2024). Dynamic approach to study relationship among carbon dioxide emissions, urbanization, and economic growth in BRICS countries. *Journal of the Knowledge Economy*, *16*(1), 3386–3403. https://doi.org/10.1007/s13132-024-01964-3

Sun, W., Zhang, S., & Uddin, I. (2026). Sustainable development in BRICS economies: Linking digitalization, higher education, and energy efficiency under the N-shaped environmental kuznets curve. *Energy Reports*, *15*, 109048. https://doi.org/10.1016/j.egyr.2026.109048

Tabash, M. I., Farooq, U., Aljughaiman, A. A., Wong, W.-K., & AsadUllah, M. (2024). Does economic complexity help in achieving environmental sustainability? New empirical evidence from N-11 countries. *Heliyon*, *10*(11), e31794. https://doi.org/10.1016/j.heliyon.2024.e31794

Ullah, A., Raza, K., & Mehmood, U. (2023). The impact of economic growth, tourism, natural resources, technological innovation on carbon dioxide emission: evidence from BRICS countries. *Environmental Science and Pollution Research*, *30*(32), 78825–78838. https://doi.org/10.1007/s11356-023-27903-4

Wang, A., Shan, S., Ibrahim, R. L., & Omokanmi, O. J. (2024). A new look at environmental sustainability from the lens of green policies, eco-digitalization, affluence, and urbanization: Empirical insights from BRICS economies. *Energy & Environment*, *35*(8), 4195–4222. https://doi.org/10.1177/0958305X231177736

Wen, J., Khalid, S., Mahmood, H., Alam, K., & Zakaria, M. (2025). Examining the influence of ICT on carbon emissions in emerging economies. *Structural Change and Economic Dynamics*, *74*, 353–360. https://doi.org/10.1016/j.strueco.2025.03.015

Zuo, L., & Ren, Y. (2025). How renewable energy consumption and digitalization contribute to environmental sustainability: Evidence from One Belt One Road countries. *Journal of Environmental Management*, *380*, 124379. https://doi.org/10.1016/j.jenvman.2025.124379

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The role of women in empowering economic prosperity through domestic credit in the Southern African Development Community (SADC)

Magwana Ibrahim Ngollo*, Frank Arbogast Mwombeki, Beny Benjamin Mwenda

College of Business Education, Mbeya campus, Accounts and Marketing Department, Mbeya, Tanzania
Corresponding Author: magwanaibrahim@gmail.com

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Authors' emails:
frank.mwombeki@cbe.ac.tz
beny.mwenda@cbe.ac.tz

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Abstract

Purpose — This study examines the symbiotic relationship between women's representation in parliament and domestic credit, and the direct and indirect pathways by which these factors influence economic prosperity in Southern African Development Community (SADC) countries.

Methods — Panel data from 16 SADC countries over the period 1997 -2022 are analysed using Generalised Method of Moments (GMM) and Generalised Structural Equation Modelling (GSEM), with a focus on the examination of interaction effects and diminishing returns of women's parliamentary representation and domestic credit on economic prosperity.

Findings — The results indicate that women's representation in parliament has a significant positive effect on economic prosperity, with the effect strengthened by domestic credit. Domestic credit also contributes indirectly to economic growth by enhancing women's economic influence.

Implication — These findings provide important insights for policymakers, highlighting the need for a balanced strategy that promotes both women's political participation and financial inclusion while avoiding potential economic imbalances.

Originality — This study contributes to the literature by integrating gender, finance, and economic growth within the SADC context and by uncovering indirect pathways through which domestic credit affects economic prosperity via women's political empowerment.

Keywords — Women in Parliament; Domestic Credit; Economic Prosperity; Financial Inclusion; SADC

Introduction

The symbiotic relationship between women and credit has been a key factor in the longevity and prosperity of societies and empires throughout history (De Nicola, 2022; De Soto, 2019; Harris, 2019; Hudson, 2020; Keller, 2022; Ma et al., 2023). Perhaps one of the most vivid historical examples of the intertwining of the two elements is Queen Isabella's support of Christopher Columbus's voyage, which contributed to the expansion of the Spanish Empire, among other factors (Milam, 2019). When credit and women work in unison, it is possible to achieve a fair and balanced allocation of power, resources and opportunities in society (Perunovic, 2023; Wee, 2021; Yadav, 2023). It might seem easy to bring women and credit together, but certain countries around

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the world face obstacles in advancing gender equality and financial inclusion. This issue has multiple dimensions and impacts both developed and developing countries in diverse ways (Fernández et al., 2021; Ojo, 2023). Even though credit and women have greatly influenced the prosperity of ancient empires, modern African societies have not yet fully utilised these elements to their full potential. Furthermore, although women in Africa make up the vast majority of the population, their contribution to the continent's gross domestic product accounts for only 33% (Moodley et al., 2019).

Despite their little contribution towards GDP, women achieve that level of contribution while facing obstacles in the financial infrastructure. Research has shown that African women lack adequate access to credit, leaving them with limited financing options (Adewusi, 2021; Vandi et al., 2022). Lack of adequate access to credit hinders women's effective participation in enhancing the economic prosperity of the societies they live in. At this juncture, it seems quite plausible to ponder that African economic prosperity can be enhanced by improving women's access to credit and promoting their empowerment across all facets of daily undertakings in society. Improvement of women's access to credit could be essential for harnessing their economic potential (Agarwala et al., 2022). Women can use money obtained through domestic credit to invest in businesses, education, and healthcare, ultimately fostering broader economic growth and stability.

Perhaps one of the key strategies for promoting women is to increase their representation in decision-making roles within the legislative branch of government (Bency, 2018; Gao & Mahutga, 2023). Scholars have shown that higher female representation in parliament improves governance and reduces corruption, both of which are beneficial for economic growth (Gao & Mahutga, 2023; Tusalem, 2022). This idea is particularly relevant in African economies since some of the studies have shown that women in parliament are more likely to advocate for policies that promote financial inclusion, reduce income inequality, and support female entrepreneurship, which are critical for unlocking the economic potential of any society (Goltz et al., 2015; Wang & Naveed, 2021). By encouraging women to take leadership roles, a country can foster creativity in decision-making through diversity of viewpoints and leadership abilities, which may affect economic growth (Arnold & Gammage, 2019; Khushk et al., 2023). Africa is making significant strides in empowering women, especially in sub-Saharan Africa, where the number of women in parliaments has increased in recent years. However, the progress of women legislators varies significantly across the continent (Adams & Wylie, 2020; Mojapelo & Faku, 2020).

Economic prosperity refers to the degree of material economic well-being of the local population (Mueller, 2021). In this study, per capita GDP was used to gauge economic prosperity. GDP per capita was selected as a measure of economic prosperity due to its ability to reflect economic growth, living standards, productivity, and group well-being (Bhide & Khanolkar, 2020; Kaminitz, 2023; Rodriguez-Avi, 2022). Per capita GDP as a measure of economic prosperity has certain drawbacks. There are studies that point out its shortcomings due of factors such as income inequality, variations in the cost of living, and the effects of exchange rates, which are not fully reflected in per capita GDP (Dědeček & Dudzich, 2022; Mohanty, 2019; Tasnim, 2021). Despite its shortcomings, the ultimatum was reached to use it as a measure of economic prosperity due to its ease of data availability and its indirect relationship to living standards, thus providing a dependable indicator of economic progress and material well-being.

Gender development theory is an important framework for understanding economic development, with a specific emphasis on how gender dynamics shape economic outcomes and opportunities (Dow, 2020). For a society to achieve its full economic potential, women must be included in economic activities (Guerrero et al., 2023). However, in many societies, women are not fully utilised in economic activities, thus limiting the potential of economies (Buterin et al., 2023). With gender-based analysis, it is possible to facilitate the creation of economic policies that are better tailored to address the distinct requirements and obstacles faced by women, thereby fostering their utilisation in economic growth (Onaran et al., 2022). Conversely, financial inclusion theory argues that by increasing access to financial services for all members of society, a nation can reduce poverty, boost socio-economic development, and, in turn, stimulate economic growth (Ahmad & Yadav, 2022). Inclusive economic growth and sustainable development are thought to be influenced by financial inclusion (Van et al., 2021). The integration offers a comprehensive framework for analysing the effect of women's empowerment, through politics, on economic prosperity and how that effect can be enhanced by domestic credit (Tripathi & Rajeev, 2023). Increased domestic credit will enable women to consume goods and invest in businesses, thereby contributing to economic growth (Fadil El-Turkey, 2021).

Some of the inconsistencies highlight the regional variations in the effect of women's representation in parliament towards economic growth, such as Mirziyoyeva and Salahodjaev (2023), and Altuzarra et al. (2021), while the former found that female empowerment in the public sector has a positive effect on economic growth, the latter discovered that it boosted economic growth in all developing nations but adversely affected countries in sub-Saharan Africa. Khan et al. (2020) and Murari (2017) found that domestic financing boosts economic growth, whereas Cecchetti and Kharroubi (2019) concluded that increased domestic credit may encourage riskier, less productive investments, ultimately lowering worker productivity and negatively affecting economic growth. Ozili et al. (2023) found that an increase in domestic credit to the private sector in Nigeria increases GDP per capita when the legal system is strong, but not during crises, further suggesting that financial policies, no matter how good, can still be affected by legal frameworks and economic stability. Lastly, prior research highlighted the possibility of the presence of conflicting outputs regarding short-term vs long-term effects. Mohamed (2022) noted that women's participation in parliament boosted Sudan's economy in the short term, but argued that the long-term effect may require additional scrutiny. Given these developments, it was necessary to assess the effect of women's political representation on the economic prosperity of Southern African Development Community (SADC) member states.

This paper contributes to the current body of knowledge by examining a mechanism that has received limited attention in prior research. Previous studies have generally examined the separate impacts of women’s empowerment and domestic credit on economic outcomes; however, our methodology integrates both perspectives to capture their inherent interdependence across various areas and time frames. It incorporates gender and development Theory, financial Inclusion Theory, and the principles of diminishing marginal utility into a cohesive analytical framework. This integrated approach enhances value by connecting the literature on gender studies, financial inclusion, and economic development, while providing context-specific insights for SADC economies that have not been thoroughly examined previously.

Methods

An investigation of the connection between women legislators, domestic credit, and economic prosperity was conducted across 16 SADC countries. SADC was established in 1992 to advance sustainable development and economic integration throughout Southern Africa (SADC, 2022). Its main goals include facilitating commerce, advancing infrastructure development, promoting economic integration, and enhancing peace and security among its member states. The community consists of 16 members: Angola, Botswana, Comoros, Democratic Republic of the Congo, Eswatini, Lesotho, Madagascar, Malawi, Mauritius, Mozambique, Namibia, Seychelles, South Africa, Tanzania, Zambia, and Zimbabwe. The structure of the organisation consists of the SADC Secretariat, the Council of Ministers, the Summit of Heads of State or Government, and standing committees. SADC offers fertile ground for exploring and addressing socio-economic concerns, owing to its dynamic economic integration initiatives and the distinct challenges and opportunities it presents (Muntschick, 2020).

The study period covered the years from 1997 to 2022; these years were selected due to the limited availability of data on the percentage of women in parliament prior to 1997. The data used in this study were obtained from the World Bank database. The annual GDP per capita in current US dollars, used as a proxy for economic prosperity, was the dependent variable. The proportion of seats held by women in national parliaments and domestic credit was an independent variable. The selection of these variables aligns with previous research by various authors (Baskaran et al., 2024; Cecchetti & Kharroubi, 2019; Khan et al., 2020; Mohamed, 2022). A vector of control variables was incorporated to isolate the independent variables and mitigate potential bias from omitted-variable effects, thereby enhancing the coherence of the analysis of the effect of the independent variable on the dependent variable. Inflation rate and the percentage of the rural population with access to electricity were included as control variables in our analysis, given compelling evidence of significant relationships between these variables and gross domestic product, as reported by other researchers (Edward, 2023; Mwale & Allexander, 2022).

Theoretically, economic prosperity as proxied by GDP per capita can be expressed as a function of several main factors, namely:

$$GDPpercapita_{it} = f(perlwomen_{it}, domcred_{it}, ruralelectrper_{it}, inflation_{it}) \quad (1)$$

Based on the discussion presented in this paper and the theoretical framework outlined in Equation 1, and to capture the dynamic nature of the relationship, the following empirical model is specified and estimated using the Generalised Method of Moments (GMM).

$$GDPpercapita_{it} = \beta_0 + \beta_1 L.GDPpercapita_{it-1} + \beta_2 perlwomen_{it} + \beta_3 domcred_{it} + \beta_4 ruralelectrper_{it} + \beta_5 inflation_{it} + \gamma_i + \varepsilon_{it} \quad (2)$$

Equation (2) captures the relationship between economic prosperity, proxied by $GDPpercapita_{it}$, and the key explanatory variables, namely women's representation ($perlwomen_{it}$) and domestic credit ($domcred_{it}$), along with control variables including rural electrification ($ruralelectrper_{it}$) and inflation ($inflation_{it}$). The term $L.GDPpercapita_{it-1}$ represents the one-period lag of GDP per capita, which is incorporated to account for the dynamic models. The parameters β are to be estimated; i denotes the country, t denotes time, represents unobserved effects unique to each country and reflect traits that do not change over time, and ε represents the error term. A detailed description of all variables and their measurements is provided in Table 1.

Table 1. Variable Description

Variable name	Symbol	Measurement	Units
Gross Domestic Product per Capita.	$GDPpercapita$	Natural logarithm of GDP per capita.	US dollars
Proportion of seats held by women in national parliaments.	$perlwomen$	Quotient of the total number of seats occupied by women divided by the total number of seats in parliament.	Percentage
Domestic credit to the private sector.	$domcred$	Quotient of the total value of domestic credit divided by the country's GDP.	Percentage
Percentage of Rural populations with access to electricity.	$ruralelectrper$	A quotient of the number of rural households with electricity divided by the total number of rural households.	Percentage
Inflation	$inflation$	Percentage change in the Consumer Price Index (CPI) over a period of one year, measured via the Laspeyres formula.	Index

The missing data were determined to be missing at random and thus the multiple imputation regression technique was used to create missing values, this technique was used due to its ability to preserve correlations between variables by using observed data to forecast missing values, it conform with maximum likelihood estimate techniques and manage several forms of missing data patterns (Si et al., 2023; Young & Johnson, 2015; Yu et al., 2020). The one-step Generalised Method of Moments estimator and Generalised Structural Equation Modelling (GSEM) were employed to obtain reliable empirical results and address endogeneity in panel data. These techniques were selected for their proven ability to handle endogeneity and multilevel panel data analysis (Carrasco & Nayihouba, 2024; Joo et al., 2022). For the purpose of direct and indirect estimation effects, the GSEM framework presented in Figure 1 was used. This framework sought to determine whether it is women in parliament that indirectly influence GDP per capita through domestic credit, or domestic Credit that indirectly influences GDP per capita through women in parliament.

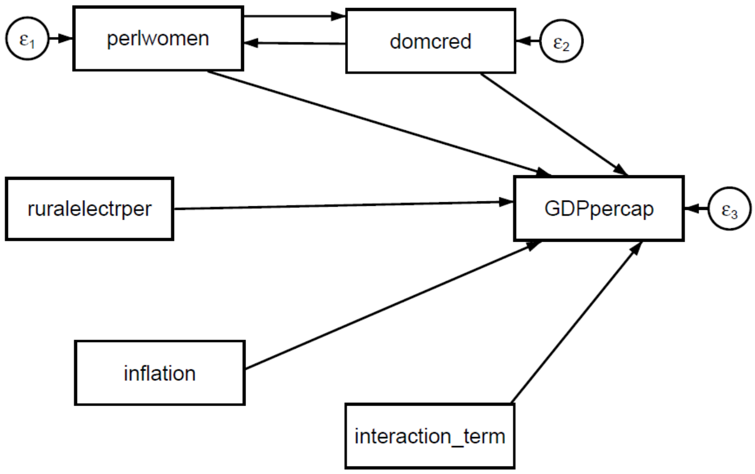

Figure 1. Generalised Structural Equation Model pathways

Results and Discussion

Descriptive statistics of all variables used in this study are presented in Table 2. These statistics provide an initial overview of the data distribution, including the mean, standard deviation, minimum, and maximum. This result indicates the characteristics and variations of the data before further empirical analysis is carried out.

Table 2. Descriptive statistics

Variable	Obs	Mean	Std. Dev.	Min	Max
GDPpercapita	393	3053.56	3481.37	187.86	16851.12
perlwomen	393	20.19	11.92	0.00	46.75
domcred	393	28.37	31.39	1.10	142.42
ruralelectrper	393	30.93	34.30	0.00	100.00
inflation	393	13.96	41.78	-72.00	557.20

From Table 2, it can be observed that the average inflation rate in SADC during the period of the study was approximately 14%, with a standard deviation of 41.776, indicating a large variation in inflation rates across periods. Such a huge variation is a result of certain countries observing an abnormal rate of inflation, such as Zimbabwe. The summary also indicates that some countries experienced a period of deflation, with the lowest inflation rate of -72%. This rate of deflation can be as problematic as high inflation, since it signals economic contraction or depression in certain countries at certain times. The average proportion of women in parliament was roughly 20.2% across regions, with some countries reaching an astounding ratio of almost 47% of legislators being women, while others, during certain periods, had no female members of parliament at all. The standard deviation of 11.19 indicates significant variation in the proportion of women in parliament across SADC member states. Domestic credit to the private sector roughly stood at an average value of 28.4% of GDP; the high standard deviation of 31.393 indicates notable variations throughout the nations in the SADC countries.

The minimum value of 1.09 and the maximum value of 142 further signify the impeccable range of variations of the percentage of GDP issued to the private sector in the form of credit. This could be a result of an extreme financial event such as the 2007 – 2008 financial crisis. The percentage of the rural population with access to electricity stood at an average value of 31%; nevertheless, there is a significant variation, as indicated by the wide standard deviation of 34.299. It can clearly be seen from the minimum and maximum values that, at certain periods, some countries' rural populations had no access to electricity at all, while some achieved a 100% electricity access rate for their entire rural populations, resulting in a large variation. GDP per capita had an average of 3,053.56 US dollars and a standard deviation of 3,481.369 US dollars. As this value is used as a proxy for economic prosperity, it indicates a significant level of economic inequality amongst the SADC member states. The GDP per capita ranged from 187.857 US dollars to a high value of 16,851.12 US dollars. The notable variations in GDP per capita, proportion of women in parliament, domestic credit to the private sector, inflation rates and rural electrification demonstrate the considerable economic and social heterogeneity throughout the SADC area.

Table 3. Variance Inflation Factor (VIF)

	VIF	1/VIF
domcred	1.87	0.54
ruralelectrper	1.69	0.59
perlwomen	1.35	0.74
inflation	1.03	0.97
Mean VIF	1.49	

Table 3 presents the outcome of the variance inflation factor test, which showed tolerable and acceptable value ranges for multicollinearity. Multicollinearity can lead to unstable and inflated coefficients, making it difficult to identify the true effect of each predictor (Shrestha, 2020). Low VIF values indicate a more consistent model with more reliable coefficient estimates. The highest VIF value of 1.8 is below the problematic threshold. The mean VIF of 1.485 indicates very low, tolerable levels of multicollinearity among the independent variables.

Table 4. Generalised Method of Moment Estimation Results

	Dependent Variable: GDPpercapita				
	(1)	(2)	(3)	(4)	(5)
L.GDPpercapita	0.98***	0.92***	0.67***	0.64***	0.66***
	(0.01)	(0.03)	(0.05)	(0.07)	(0.06)
perlwomen	-0.90	44.17*	51.29**	46.10**	18.79
	(5.70)	(23.50)	(21.48)	(22.58)	(22.72)
domcred	2.32	14.13***	16.04	10.79	-17.60***
	(2.69)	(14.13)	(20.22)	(17.38)	(5.79)
perlwomen x domcred		-1.19***	-0.87*	-0.74*	
		(0.40)	(0.50)	(0.44)	
ruralelectrper			49.24***	52.34***	55.53***
			(7.55)	(8.55)	(8.62)
inflation				-11.81	-15.55
				(15.10)	(19.35)
Observations	174	174	174	174	174
Hansen J P-value	0.94	0.86	0.92	0.85	0.87
AR(2)	0.34	0.27	0.12	0.11	0.12

Robust standard errors in parentheses. ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

Table 4 presents the results of the GMM analysis. The Second-order autocorrelation test and the Hansen J-test verify the legitimacy of the instruments used. Due to the small sample size and to avoid model overfitting, the one-step GMM was opted. The outcome of the GMM model (4) with the interaction term and with all control variables indicated that the percentage of women in parliament had a positive effect on the GDP per capita. Each percentage increase in women in parliament resulted in a 46.10 USD increase in GDP per capita; these results were similar to those of (Baskaran et al., 2024; Mirziyoyeva & Salahodjaev, 2023) and contradicted those of (Altuzarra et al., 2021). The results also showed that in the absence of an interaction term, the level of domestic credit has a negative effect on GDP per capita; a percentage increase in the domestic credit led to a reduction of 17.60 USD in per capita GDP. These results are similar to those of (Cecchetti & Kharroubi, 2019) and contradict those of (Murari, 2017).

These results are fascinating, and they reveal how inclusive governance can transform financial development into economic growth. When examined independently, domestic credit seems to impact GDP per capita negatively, which could be the result of misallocation of resources, inadequate institutions, and restricted absorptive ability that hinder the productive utilisation of credit (Altuzarra et al., 2021). However, when women's representation in parliament is incorporated as an interaction term, the impact of domestic credit transitions from negative to positive, highlighting the influence of female legislators in enhancing governance, mitigating corruption, and promoting policies that allocate financial resources to education, health, and family welfare (Mirziyoyeva & Salahodjaev, 2023). From the results, it is plausible to argue that women's political participation of women increases credit productivity and transforms hindrances to growth into catalysts for progress. These findings align with gender and development theory, which emphasises that gender inequality hinders economic advancement and that the empowerment of women in decision-making bodies enhances resource distribution and sustainable development results (Silva & Klasen, 2021).

The influence of women in parliament on GDP per capita is rendered insignificant in the absence of the interaction term, suggesting that the predominance of female lawmakers alone is not a necessary condition for economic growth (Dahlum et al., 2022; Khorsheed, 2020). The presence of women in parliament must be coupled with financial inclusion to yield growth (Wani et al., 2024). Augmenting domestic credit increases the financial resources available to women, enabling them to allocate these funds towards productive endeavours that elevate production and GDP per capita. An increased proportion of women in parliament indicates a policy landscape that is likely to be inclusive and conducive to women's financial access. The mitigation of the adverse impact of domestic credit through the interaction term further emphasises that the interdependent relationship between women and credit ultimately ensures that financial resources are utilised productively, benefiting the whole economy.

However, on closer inspection, these results also suggest a diminishing symbiotic effect of women in parliament on domestic credit, as indicated by the negative coefficient of the interaction term. To our knowledge, this is the first evidence that the interaction between women's parliamentary representation and domestic credit exhibits a diminishing symbiotic effect. This observation is in line with the theory of diminishing returns. In this case, we argue that a combination of women in parliament and domestic credit will positively affect GDP per capita up to a certain level, after which their unitary contribution to per Capita GDP will begin to decline, as the negative interaction term offsets the marginal gains from their joint increase. This suggests that there should be an optimum level and balance of the combination of women in parliament and domestic credit. The joint growth of the two items beyond a certain threshold will yield a smaller and even negative contribution to GDP per capita.

The results of the GMM table also show that the effect of women in parliament is significant when the control variables are included, indicating that for female legislators to have any impact on GDP per capita, the existence of favourable economic conditions and suitable infrastructure is necessary. Similar results have been observed by various authors (Khorsheed, 2020; Mirziyoyeva & Salahodjaev, 2023). Domestic credit is significant and has a positive effect in the model without controls, but with an interaction term, suggesting that domestic credit can be adequately used by women, ceteris paribus, and hence stimulate the effect on a country's economic prosperity. Enhanced credit and women's empowerment are crucial for enhanced per capita GDP. Rural electrification positively affects GDP per capita across all models in which it appears, and the presence of an interaction term between women in parliament and domestic credit enhances the effect of rural electrification on the model containing only control variables. Greater representation of women in parliament strengthens the legitimacy of gender-responsive policymaking, ultimately leading to more effective gender-focused interventions in the presence of robust infrastructure, such as rural electrification (Amuakwa-Mensah & Surry, 2022).

The results of the Generalised Structural Modelling are presented in Table 5. The GSEM results indicate that women in parliament have a positive direct effect on GDP per capita, whereas the level of domestic credit has no significant direct effect. In the GSEM model without the interaction term, a percentage increase of women's representation in parliament contributed to an increase of 38.09 USD increase in per capita GDP. The presence of the interaction term enhances the effect of women in parliament by raising their contribution towards per capita GDP from 38.09 to 53.82 USD. These outcomes are similar to those of one-step GMM, indicating robustness. The effect of the interaction term is negative and significant, albeit small, with every unit increase in the interaction term leading to a reduction of -0.54 USD in GDP per capita. These results are also consistent with the results of one-step GMM. Similar arguments, as presented above when discussing the alignment of one-step GMM results with the theories used in this model, continue to hold for the GSEM model. Credit availability enables women to leverage their role in the labour force by utilising capital for entrepreneurship, business expansion, and productivity gains, thereby amplifying their overall contribution to the economy (Balasubramanian & Kuppusamy, 2020).

Table 5. Generalised Structural Equation Modelling Estimation Results

	I	II
Direct effect		
Perlwomen	38.090***	53.833***
	(11.092)	(11.812)
Domcred	-1.175	15.741
	(5.892)	(10.948)
Ruralelectrper	79.133***	77.404***
	(6.558)	(6.669)
Inflation	-1.148	-0.974
	(2.191)	(2.347)
perlwomen x domcred		-0.548**
		(0.253)
Indirect effects		
perlwomen	0.788	11.255*
Pathway : perlwomen → domcred → per capita GDP	(3.292)	(6.704)
Domcred	2.539***	4.440***
Pathway : domcred → perlwomen → per capita GDP	(0.823)	(1.611)

Notes: Robust standard errors in parentheses. ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

The GSEM findings demonstrate the indirect mechanism by which domestic credit fosters economic growth in conjunction with women's parliamentary representation. In the baseline model, devoid of the interaction term, domestic credit demonstrated a robust positive indirect effect on GDP per capita: each 1 percentage point increase in credit availability was associated with a USD 2.54 increase in per capita GDP, whereas the independent indirect effect of women in parliament was negligible. Upon the introduction of the interaction term, the indirect effect of credit increased to USD 4.40, signifying that the efficacy of credit in enhancing GDP is augmented in environments where women hold parliamentary positions. The presence of women in parliament facilitates the conversion of credit availability into economic growth. These findings emphasise that credit availability is essential for prosperity, but its effect is much enhanced when women are empowered to shape decisions.

These findings are novel, as they shift the focus from the well-studied direct benefits of women's parliamentary representation or domestic credit on growth to their indirect and interaction effects. Even though Global evidence suggests that women's political empowerment can improve economic outcomes when combined with enabling factors, many of them focus on enablers such as civil liberties, civil society participation, and institutional quality (Al-Qahtani et al., 2020; Dahlum et al., 2022). Studies on how credit availability amplifies the effect of women's representation on growth are scarce, especially in Sub-Saharan Africa. Previous research in the region has examined the influence of women's representation and the availability of domestic credit on economic growth, without assessing whether institutional representation is the factor that enables the efficient transfer of credit into economic prosperity (Altuzarra et al., 2021; Jemiluyi & Yinusa, 2021; Mba Fokwa, 2025). Our findings reveal an underexplored mechanism whereby credit availability and women's legislative presence boost growth, offering a new perspective on how gender and finance shape development trajectories in Sub-Saharan Africa.

Conclusion

Women and credit will forever remain an essential component of the prosperity of any society. SADC countries should strive to gain all the possible benefits from the combination of these variables. There is potential to uncover new avenues for utilising women and credit to enhance economic prosperity by providing women with access to credit and promoting their inclusion in the political arena, which affects the economies of those countries. It is essential to note that the utilisation of both factors should be undertaken with a sound understanding of the decreasing returns and potential economic imbalances that may result from the interaction of the said elements. There is a need to ensure that the combined path taken brings economic success to the region in a way that is both sustainable and inclusive. The approach should be well-balanced and tailored to the specific circumstances of the nation.

The aftermath of the findings of this study can be quite handy for policymakers in the SADC region and beyond. The study has demonstrated that improving women's representation in parliament is a key step toward fully exploiting their economic potential. It has also demonstrated that their economic potential can be leveraged through domestic credit, thereby enhancing their effect on per capita GDP. On the other hand, it also highlights the need for policymakers to be cognizant of the decreasing returns associated with these factors. A well-rounded strategy should be employed amongst SADC countries that takes into account and carefully controls the interplay between the level of domestic credit and women's representation in parliament. The strategies to be employed must be context-specific, since the exact combination and scale of factors may vary depending on the economic and social conditions prevailing in each country in the SADC region.

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Author contribution
All authors contributed to the conception and design of the study, data collection, analysis and interpretation of the results, and the writing of the manuscript. All authors read and approved the final version of the manuscript.

Use of AI tools declaration
Artificial intelligence tools were used to assist with language editing and to enhance the clarity and readability of the manuscript. The authors remain fully responsible for the content and conclusions of this study.

Conflict of interest
The authors declare that there are no competing interests related to this manuscript.

References

Adams, M., & Wylie, K. (2020). Transnational organizing, the boosting effect, and women's legislative caucuses in Africa. *Politics, Groups, and Identities*, *8*(3), 615–626. https://doi.org/10.1080/21565503.2020.1752441

Adewusi, A. O. (2021). Modelling loan accessibility among women for economic development. *International Journal of Innovative Research and Development*, *10*(3). https://doi.org/10.24940/ijird/2021/v10/i3/JAN21032

Agarwala, V., Maity, S., & Sahu, T. N. (2022). Female entrepreneurship, employability and empowerment: impact of the mudra loan scheme. *Journal of Developmental Entrepreneurship*, *27*(01), 2250005. https://doi.org/10.1142/S1084946722500054

Ahmad, F. M., & Yadav, D. K. (2022). Financial Theories and Their Relevance in Financial Inclusion. In *Financial Inclusion Schemes in India* (pp. 21–35). Springer Nature Singapore. https://doi.org/10.1007/978-981-19-1316-7_2

Al-Qahtani, M. M. Z., Alkhateeb, T. T. Y., Mahmood, H., Abdalla, M. A. Z., & Qaralleh, T. J. O. T. (2020). The Role of the Academic and Political Empowerment of Women in Economic, Social and Managerial Empowerment: The Case of Saudi Arabia. *Economies*, *8*(2). https://doi.org/10.3390/economies8020045

Altuzarra, A., Gálvez-Gálvez, C., & González-Flores, A. (2021). Is Gender Inequality a Barrier to Economic Growth? A Panel Data Analysis of Developing Countries. *Sustainability*, *13*(1). https://doi.org/10.3390/su13010367

Amuakwa-Mensah, S., & Surry, Y. (2022). Association between rural electrification and agricultural output: Evidence from Sub-Saharan Africa. *World Development Perspectives*, *25*, 100392. https://doi.org/https://doi.org/10.1016/j.wdp.2021.100392

Arnold, J., & Gammage, S. (2019). Gender and financial inclusion: the critical role for holistic programming. *Development in Practice*, *29*(8), 965–973. https://doi.org/10.1080/09614524.2019.1651251

Balasubramanian, S. A., & Kuppusamy, T. (2020). Does female labour force participation contribute to better financial inclusion? Evidence from cross-country analysis. *Journal of Economic and Administrative Sciences*, *37*(4), 643–658. https://doi.org/10.1108/JEAS-08-2020-0151

Baskaran, T., Bhalotra, S., Min, B., & Uppal, Y. (2024). Women legislators and economic performance. *Journal of Economic Growth*, *29*(2), 151–214. https://doi.org/10.1007/s10887-023-09236-6

Bency, B. P. B. (2018). Gender equality through women empowerment strategy-a conceptual view. In *Gender Parity and Women Empowerment–Challenges and Way Forward* (pp. 23–30). BONFRING. https://doi.org/10.9756/BP2018.1012/05

Bhide, N., & Khanolkar, S. (2020). A comprehensive study on the factors impacting the gdp (Per capita) of major economies around the globe using regression analysis. *International Journal of Computer Applications*, *175*(10), 26–30. https://doi.org/10.5120/ijca2020920560

Buterin, V., Fajdetić, B., & Funarić, B. (2023). Understanding the macroeconomic effects of female participation in the labour market. *Economies*, *11*(11), 280. https://doi.org/10.3390/economies11110280

Carrasco, M., & Nayihouba, A. (2024). Regularized estimation of dynamic panel models. *Econometric Theory*, *40*(2), 360–418. https://doi.org/10.1017/S0266466622000469

Cecchetti, S. G., & Kharroubi, E. (2019). Why does credit growth crowd out real economic growth? *The Manchester School*, *87*(S1), 1–28. https://doi.org/10.1111/manc.12295

Kaminitz, S. C. (2023). The significance of GDP: a new take on a century-old question. *Journal of Economic Methodology*, *30*(1), 1–14. https://doi.org/10.1080/1350178X.2023.2167228

Dahlum, S., Knutsen, C. H., & Mechkova, V. (2022). Women's political empowerment and economic growth. *World Development*, *156*, 105822. https://doi.org/10.1016/j.worlddev.2022.105822

De Nicola, B. (2022). Elite women in the mongol empire. In *The Mongol World* (pp. 422–439). Routledge. https://www.taylorfrancis.com/chapters/edit/10.4324/9781315165172-35/elite-women-mongol-empire-bruno-de-nicola

De Soto, P. (2019). Network analysis to model and analyse roman transport and mobility. In P. Verhagen, J. Joyce, & M. R. Groenhuijzen (Eds.), *Finding the Limits of the Limes* (pp. 271–289). Springer International Publishing. https://doi.org/10.1007/978-3-030-04576-0_13

Dědeček, R., & Dudzich, V. (2022). Exploring the limitations of GDP per capita as an indicator of economic development: a cross-country perspective. *Review of Economic Perspectives*, *22*(3), 193–217. https://doi.org/10.2478/revecp-2022-0009

Dow, S. (2020). Gender and the future of macroeconomics: an evolutionary approach. *Review of Evolutionary Political Economy*, *1*(1), 55–66. https://doi.org/10.1007/s43253-020-00001-8

Edward, A. S. (2023). Impact of rural and urban electricity access and economic growth in Ghana: Does line transmission losses matter? *Energy Economics Letters*, *10*(1), 44–68. https://doi.org/10.55493/5049.v10i1.4732

Fadil El-Turkey, E. (2021). Social and economic empowerment of women with breadwinners. *International Journal of Modern Agriculture and Environment*, *1*(1), 79–116. https://doi.org/10.21608/ijmae.2023.215963.1016

Fernández, R., Isakova, A., Luna, F., & Rambousek, B. (2021). Gender equality and inclusive growth. *IMF Working Papers*, *21*(59). https://doi.org/10.5089/9781513571164.001

Gao, M., & Mahutga, M. C. (2023). Women's political representation and corruption. *Sociology of Development*, *9*(3), 217–241. https://doi.org/10.1525/sod.2022.0022

Goltz, S., Buche, M. W., & Pathak, S. (2015). Political Empowerment, Rule of Law, and Women's Entry into Entrepreneurship. *Journal of Small Business Management*, *53*(3), 605–626. https://doi.org/10.1111/jsbm.12177

Guerrero, M. L., Albán, R. A., Fiorella, S., Arias, B., Lizzeth, K., & Tantajulca, V. (2023). *Empowering Women for Household Economic Growth in Latin America: A Systematic Review.* https://doi.org/https://doi.org/10.34190/icgr.6.1.1155

Harris, W. V. (2019). Credit-money in the roman economy. *Klio*, *101*(1), 158–189. https://doi.org/10.1515/klio-2019-0006

Hudson, M. (2020). Origins of money and interest: palatial credit, not barter. In S. Battilossi, Y. Cassis, & K. Yago (Eds.), *Handbook of the History of Money and Currency* (pp. 45–65). Springer Singapore. http://link.springer.com/10.1007/978-981-13-0596-2_1

Jemiluyi, O. O., & Yinusa, D. O. (2021). Female economic participation and economic growth: an empirical evaluation of the nexus for sub-saharan african region. *Journal of Economics, Finance and Accounting Studies*, *3*(1), 72–80. https://doi.org/10.32996/jefas.2021.3.1.7

Joo, Y. S., Kim, J., Lee, J., & Chung, I.-J. (2022). Fine particulate matter and depressive symptoms in children: A mediation model of physical activity and a moderation model of family poverty. *SSM - Population Health*, *17*, 101015. https://doi.org/https://doi.org/10.1016/j.ssmph.2021.101015

Keller, K. (2022). Hidden figures: the holy roman empire as a "realm of ladies." *Central European History*, *55*(3), 339–354. https://doi.org/10.1017/S0008938922000024

Khan, S. M., Sherazi, H., & Liaqat, S. (2020). Impact of money supply and domestic credit on economic well-being: A case of Pakistan. *European Online Journal of Natural and Social Sciences*, *9*(3), pp–618.

Khorsheed, E. (2020). The Impact of Women Parliamentarians on Economic Growth: Modelling & Statistical Analysis of Empirical Global Data. *International Journal of Statistics and Probability*, *9*(3), 23. https://doi.org/10.5539/ijsp.v9n3p23

Khushk, A., Zengtian, Z., & Hui, Y. (2023). Role of female leadership in corporate innovation: a systematic literature review. *Gender in Management: An International Journal*, *38*(3), 287–304. https://doi.org/10.1108/GM-01-2022-0028

Ma, M., Lu, Y., Dong, G., Ren, L., Min, R., Kang, L., Zhu, Z., Li, X., Li, B., Yang, Z., Cili, N., Liu, R., Gao, Y., & Chen, F. (2023). Understanding the transport networks complex between south asia, southeast asia and china during the late neolithic and bronze age. *The Holocene*, *33*(2), 147–158. https://doi.org/10.1177/09596836221131698

Mba Fokwa, A. (2025). Financial access and women's political representation in Sub-Saharan Africa: the role of ethnic fragmentation. *Review of Behavioral Finance*, *17*(2), 249–271. https://doi.org/10.1108/RBF-09-2024-0258

Milam, J. (2019). Venture capital— patronage to apprenticeship to profession. *The Journal of Private Equity*, *22*(4), 19–32. https://doi.org/10.3905/jpe.2019.22.4.019

Mirziyoyeva, Z., & Salahodjaev, R. (2023). Does representation of women in parliament promote economic growth? Considering evidence from Europe and Central Asia. *Frontiers in Political Science*, *Volume 5-2023*. https://doi.org/10.3389/fpos.2023.1120287

Mohamed, E. S. E. (2022). Female human capital and economic growth in sudan: empirical evidence for women's empowerment. *Merits*, *2*(3), 187–209. https://doi.org/10.3390/merits2030014

Mohanty, A. T. (2019). Measuring gdp and economic growth: an economic barometer. *International Journal of Tax Economics and Management*, *2*(1), 42–58. https://doi.org/10.35935/tax/21.5842

Mojapelo, J. T., & Faku, M. E. (2020). Bridging the Gender Gap and Promoting Representation of Women in the National Parliament: A Comparative Analysis Study of South Africa and Lesotho. *Gender Questions*, *7*(1), 15 pages. https://doi.org/10.25159/2412-8457/6458

Moodley, L., Kuyoro, M., Holt, T., Leke, A., Madgavkar, A., Krishnan, M., & Akintayo, F. (2019). The power of parity: Advancing women's equality in Africa. *McKinsey Global Institute*.

Mueller, J. T. (2021). Defining dependence: the natural resource community typology*. *Rural Sociology*, *86*(2), 260–300. https://doi.org/https://doi.org/10.1111/ruso.12357

Muntschick, J. (2020). Regional economic integration in the southern african development community (Sadc): analysing the dynamics and performance. *Vestnik RUDN. International Relations*, *20*(2), 333–346. https://journals.rudn.ru/international-relations/article/view/23974

Murari, K. (2017). Financial development–economic growth nexus: evidence from south asian middle-income countries. *Global Business Review*, *18*(4), 924–935. https://doi.org/10.1177/0972150917692245

Mwale, B. J., & Allexander, M. (2022). Electricity consumption and economic growth in southern african development community region. *Journal of Economics and Sustainable Development*. https://doi.org/10.7176/JESD/13-24-02

Ojo, T. A. (2023). Gendered narratives on financial inclusion: global and continental context. In *The Politics of Financial Inclusion of Women in South Africa* (pp. 95–127). Springer Nature Singapore. https://doi.org/10.1007/978-981-99-1847-8_4

Onaran, Ö., Oyvat, C., & Fotopoulou, E. (2022). Gendering Macroeconomic Analysis and Development Policy: A Theoretical Model. *Feminist Economics*, *28*(3), 23–55. https://doi.org/10.1080/13545701.2022.2033294

Ozili, P. K., Oladipo, O., & Iorember, P. T. (2023). Effect of abnormal increase in credit supply on economic growth in Nigeria. *African Journal of Economic and Management Studies*, *14*(4), 583–599. https://doi.org/10.1108/AJEMS-02-2022-0036

Perunovic, A. (2023). Credit, debt and money as social institutions of trust. *Filozofija i Drustvo*, *34*(2), 273–288. https://doi.org/10.2298/FID2302273P

Rodriguez-Avi, J. (2022). A probabilistic model for the distribution of gdp per capita in nuts 3 zones of europe. *Studies of Applied Economics*, *40*(1). https://doi.org/10.25115/eea.v40i1.5326

Shrestha, N. (2020). Detecting multicollinearity in regression analysis. *American Journal of Applied Mathematics and Statistics*, *8*(2), 39–42. https://doi.org/10.12691/ajams-8-2-1

Si, Y., Heeringa, S., Johnson, D., Little, R. J. A., Liu, W., Pfeffer, F., & Raghunathan, T. (2023). Multiple imputation with massive data: an application to the panel study of income dynamics. *Journal of Survey Statistics and Methodology*, *11*(1), 260–283. https://doi.org/10.1093/jssam/smab038

Silva, M. S., & Klasen, S. (2021). Gender inequality as a barrier to economic growth: a review of the theoretical literature. *Review of Economics of the Household*, *19*(3), 581–614. https://doi.org/10.1007/s11150-020-09535-6

Tasnim, S. (2021). Use of Shape Restricted Regression Methods for Fitting Model of Per Capita GDP: A Global Economic Scenario of 2018. *International Journal of Statistics and Probability*, *10*(4), 52. https://doi.org/10.5539/ijsp.v10n4p52

Tripathi, S., & Rajeev, M. (2023). Gender-inclusive development through fintech: studying gender-based digital financial inclusion in a cross-country setting. *Sustainability*, *15*(13), 10253. https://doi.org/10.3390/su151310253

Tusalem, R. F. (2022). Gendered governance: examining the relationship between women legislators and effective public goods provision from a global sample. *Polity*, *54*(2), 298–333. https://doi.org/10.1086/718579

Van, L. T.-H., Vo, A. T., Nguyen, N. T., & Vo, D. H. (2021). Financial inclusion and economic growth: an international evidence. *Emerging Markets Finance and Trade*, *57*(1), 239–263. https://doi.org/10.1080/1540496X.2019.1697672

Vandi, N. J., Mshelia, S. I., Michael, A., & Kwaji, B. H. (2022). Analysis of women crop farmers' access and utilisation of agricultural credit in yola south local government area of adamawa state, nigeria. *Journal of Agriculture and Applied Biotechnology*, *1*(1), 1–6.

Wang, C., & Naveed, A. (2021). Can Women Empowerment Explain Cross-Country Differences in Inequality? A Global Perspective. *Social Indicators Research*, *158*(2), 667–697. https://doi.org/10.1007/s11205-021-02721-8

Wani, I. U., Alharthi, M., Khanday, I. N., Subhan, M., & Al-Faryan, M. A. S. (2024). Exploring the complementary interaction between financial inclusion and gender equality on economic growth: fresh evidence from developing countries. *Cogent Economics & Finance*, *12*(1), 2365585. https://doi.org/10.1080/23322039.2024.2365585

Wee, L. (2021). Money, power and inequality among married couples in urban Sarawak. *SHS Web of Conferences*, *124*, 2002. https://doi.org/10.1051/shsconf/202112402002

Yadav, S. (2023). Gender equality as a tool for sustainable development. *International Journal of Teaching, Learning and Education (IJTLE)*, *2*(2), 1–3.

Young, R., & Johnson, D. R. (2015). Handling missing values in longitudinal panel data with multiple imputation. *Journal of Marriage and Family*, *77*(1), 277–294. https://doi.org/10.1111/jomf.12144

Yu, L., Liu, L., & Peace, K. E. (2020). Regression multiple imputation for missing data analysis. *Statistical Methods in Medical Research*, *29*(9), 2647–2664. https://doi.org/10.1177/0962280220908613

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Public debt dynamics and real exchange rate volatility: Evidence from South Africa

Yonela Tyani*, Chuma Mbaleki, Asive Nodo

Department of Business Management and Economics, Faculty of Economics and Financial Science, Walter Sisulu University, Mthatha, South Africa
*Corresponding Author: tyaniyonela@gmail.com

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Authors' emails:
cmbaleki@wsu.ac.za
anodo@wsu.ac.za

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Abstract

Purpose — This paper examines the relationship between South Africa's rising public debt and real exchange rate (RER) volatility. Over the past two decades, the country has experienced an alarming increase in external and domestic debt levels, accompanied by episodes of exchange rate instability and deteriorating economic performance.

Method — Using annual data from 2000 to 2024, we estimate an ARDL model to assess the nexus between debt and exchange rate volatility.

Findings — The results suggest that public debt is a significant driver of exchange rate volatility and rand depreciation. However, interest rates, inflation, and trade openness are key factors responsible for significant fluctuations in the exchange rate in South Africa. Similar results are obtained in both the short-run and long-run estimation.

Implications — The paper recommends firm government controls designed to prevent sharp capital movements (inflows or outflows) that could destabilise the rand. This can be achieved by maintaining a favourable trade balance, targeting inflation, and adjusting monetary policy. Secondly, the government can diversify the composition of its debt currency. This helps reduce volatility in debt-servicing payments and further stabilises government budgets and fiscal planning.

Originality — There is little empirical literature on the direct relationship between public debt and real exchange rate volatility in South Africa. This study aims to fill the gap by providing a novel empirical assessment of the long- and short-run dynamics between rising public debt and real exchange rate volatility.

Keywords — Public Debt, Real Exchange Rate Volatility, South Africa, Economic Growth, ARDL Model

Introduction

Despite its evident signs of resistance, South Africa has faced a deteriorating macroeconomic situation over the past decade. From the persistent accumulation of debt stock level, widening deficit and exchange rate volatility. Between 2008 and 2023, public debt increased from approximately 27% to over 70% of GDP, while debt-servicing costs grew as a share of total government liabilities. The country's domestic currency, the rand, is amongst the world's most volatile currencies, frequently responding to global risk sentiment, commodity price shocks, and domestic fiscal dynamics.

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Public debt refers to a government's overall debt to domestic and international creditors, including residents, foreign governments, and domestic and foreign financial institutions (Usman, 2025). On the other hand, the real exchange rate is the relative value of a nation's goods and services compared to those in foreign markets (Mbali, 2021). Dynamics in public debt can destabilise the value of a domestic currency, thereby increasing the cost of servicing debt (IMF, 2025). Governments with a high public debt burden face greater uncertainty, lower investor confidence, and higher borrowing costs (Koçak, 2025). These variations increase import costs, contribute to inflationary pressures, reduce the competitiveness of domestic firms involved in global trade**,** and erode purchasing power (Khomo & Aziakpono, 2020). Although many factors drive exchange rate volatility in South Africa, the role of public debt as a single driver has received less attention. Empirical evidence suggests that fiscal stress increases sovereign risk premia and triggers exchange-rate instability (Manguzvane & Biyase, 2023). Yet a comprehensive analysis explicitly focusing on the South African case remains limited.

Rising public debt in developing countries has significant implications for macroeconomic stability. A study by Olaoye et al. (2022) shows that high debt accumulation tends to suppress economic growth and worsen macroeconomic stability, especially when it exceeds optimal levels. These impacts are further complicated when linked to exchange rate dynamics, as demonstrated by Kuranga et al. (2025). He found a two-way relationship between public debt and the exchange rate. Exchange rate depreciation increases the debt burden, while increased debt can depress the exchange rate by increasing economic uncertainty.

Moreover, as public debt increases, interest rates rise, leading to capital inflows that contribute to real exchange rate appreciation. These findings are confirmed in empirical literature. Specifically, Sang et al. (2025) demonstrate that fiscal expansions financed by debt lead to an appreciation of the real exchange rate. This case applies to both advanced and emerging markets. Similarly, Donnat, (2025) demonstrates that higher public debt increases domestic interest rates, attracting capital inflows, which, in turn, leads to real exchange rate appreciation. Nyoro and Njaramba (2025) further demonstrate that high public debt is associated with real exchange rate depreciation and volatility. Ormaechea (2020) demonstrates that public debt, predominantly denominated in foreign currency, leads to exchange-rate depreciation, particularly in emerging markets. These studies hint at a negative relationship between public debt and the real exchange rate, an important aspect for South Africa, which is characterised by high public debt that has since been exacerbated by loans from the IMF and the World Bank.

A larger pool of empirical evidence suggests that macroeconomic variables have a significant impact on the real exchange rate (David, 2024). Changes in macroeconomic factors such as GDP, interest rates, inflation, external debt, and the balance of trade are major causes of exchange rate fluctuations (Ghauri et al., 2024). On the contrary, Armah et al. (2023) show that interest rates and inflation are unrelated to exchange rates. However, disaggregated studies targeting individual variables yielded mixed outcomes. For example, Özen et al. (2020) found that inflation and interest rates have a significant impact on exchange rates in both developed and emerging economies. Fluctuations in exchange rates and inflation substantially affect economic growth in the Southern African Development Community (SADC), particularly given the region's historical connections and economic clustering. Empirical literature demonstrates that exchange rate volatility and inflation are inversely correlated with economic growth. Olamide et al. (2022) demonstrate that exchange rate volatility worsens the inflation growth relationship in SADC economies. Patel and Mah (2018) and Rapetti (2020) found that real exchange rate appreciation is negatively and significantly associated with exports and economic growth in the long run. In contrast, money supply and foreign direct investment exert a positive and significant influence on the real exchange rate.

Exchange rate fluctuations continue to pose a serious threat to economic growth, as ongoing instability undermines competitiveness and complicates trade-related policy decisions (David, 2024). Whilst international trade is found to negatively affect the exchange rate by empirical results (Yussif et al., 2022), suggesting that greater trade exposure may increase currency volatility, a positive correlation is also found. For example, Moloi (2023) and Mehtiyev et al. (2021) show that, in certain situations, exchange rate volatility can promote trade by creating profit opportunities for both importers and exporters. Nicita (2013) found that fluctuations in exchange rates generally have little impact on international trade flows, with the notable exception of situations involving currency unions or fixed exchange rate systems, where volatility can play a more significant role.

This study contributes to the literature by estimating the impact of public debt on real exchange rate volatility in South Africa using the ARDL model. South Africa offers an excellent case study due to direct participation in global capital markets, relatively high levels of external debt, and well-developed financial markets (Mbaleki, 2024). Although the relationship between public debt and real exchange rate volatility has attracted significant global attention, empirical evidence from developing economies, such as South Africa, remains relatively limited. Motivated by these gaps, this study examines the relationship between public debt dynamics and real exchange rate volatility in South Africa, utilising time-series data spanning from 2000 to 2024. The data is obtained from the South African Reserve Bank (SARB) online statistical query and the World Bank. The study also controls for macroeconomic variables, namely: inflation, real interest rates, foreign direct investment (FDI), and the trade balance.

Methods

Empirical Framework

Nuru and Gereziher (2021) examined how innovations in public expenditure affect the volatility of the real exchange rate in South Africa, utilising a vector autoregressive impulse response model. The baseline Linear Programming model can be defined in the following manner:

$$x_{t+h} = \alpha_h + \varphi_h(L)y_{t-1} + \beta_h\ innovation_t + linear\ trend + \varepsilon_{t+h} \tag{1}$$

In this framework, x denotes government expenditure and exchange rate variations. Y encompasses the lagged values of government spending, exchange rate instability, real GDP, and public income. The expression φh(L) represents a polynomial in the lag operator, with the intercept βh reflecting the response of x at time t + h. The model incorporates a stationary (αh). We calibrate from this model by including public debt and controlling for macroeconomic variables such that:

$$RER_t = \beta_0 + \beta_1 PD_t + \beta_2 INF_t + \beta_3 RINTR_t + \beta_4 EXP_t + \beta_5 IMP_t + \beta_6 FDI_t + \varepsilon_t \tag{2}$$

Where RER_t denotes the real exchange rate in time t, based on the real effective exchange rate adjusted for trade weights. β_1 to β_6 Represents coefficients of the independent variables. INF_t Signifies inflation, measured by the yearly variation in the consumer price index. $RINTR_t$ Indicates real interest rates represented as a percentage of investment. EXP_tcorresponds to exports as a share of GDP, IMP_t While it reflects imports. FDI_t Represents Foreign Direct Investment based on net inflows (% of GDP). ε_t Encompasses all other factors influencing real exchange rate volatility that are not accounted for in the equation.

The study employs the ARDL approach to investigate the impact of public debt on real exchange rate volatility in the South African context. The chosen approach is more efficient since it captures both I(0) and I(1) variables, presents both long-run and short-run elasticities. Furthermore, ADRL corrects endogeneity and autocorrelation simultaneously, as stated by Pesaran et al. (2001). The specified ARDL model is as follows:

$$RER_t = \alpha_0 + \sum_{i=1}^{p} \psi_i RER_{t-i} + \sum_{i=0}^{q_1} \beta_{1,i} PD_{t-i} + \sum_{i=0}^{q_2} \beta_{2,i} INF_{t-i} + \sum_{i=0}^{q_3} \beta_{3,i} RINTR_{t-i} + \sum_{i=0}^{q_4} \beta_{4,i} EXP_{t-i} + \sum_{i=0}^{q_5} \beta_{5,i} IMP_{t-i} + \sum_{i=0}^{q_6} \beta_{6,i} FDI_{t-i} + \epsilon_t \tag{3}$$

Where p , represents the ideal count of lags for RER, q_1 to q_6 represent the ideal lags for every independent variable. α_0: The constant component. ψ_i: The short-term coefficients for the lagged dependent variable (RER). $\beta_{1,i}$ to $\beta_{6,i}$: The short-term coefficients for both current and lagged independent variables. ϵ_t: The residual term.

This study conducts a cointegration analysis using the ARDL Bounds test to examine the long-term relationship among the variables (Pesaran et al., 2001). The test relies on the F-statistic. In this examination, the null hypothesis of no cointegration is rejected if the calculated F-statistic exceeds the upper-bound critical value, suggesting a long-term relationship between the variables. Conversely, the null hypothesis is not rejected if the F-statistic is below the lower bound critical value, which indicates that there is no long-term connection between the variables. If the F-statistic falls between the lower and upper critical values, the test fails to yield a clear result and requires additional analysis to assess cointegration. Once the cointegration is identified, an Error Correction Model (ECM) is presented to show the speed at which the system reverts to equilibrium. The re-specified Error Correction Model (ECM) format of this ARDL model:

$$\Delta RER_t = \alpha_0 + \sum_{i=1}^{p-1} \psi_i^* \Delta RER_{t-i} + \sum_{i=0}^{q_1-1} \beta_{1,i}^* \Delta PD_{t-i} + \sum_{i=0}^{q_2-1} \beta_{2,i}^* \Delta INF_{t-i} + \sum_{i=0}^{q_3-1} \beta_{3,i}^* \Delta RINTR_{t-i} + \sum_{i=0}^{q_4-1} \beta_{4,i}^* \Delta EXP_{t-i} + \sum_{i=0}^{q_5-1} \beta_{5,i}^* \Delta IMP_{t-i} + \sum_{i=0}^{q_6-1} \beta_{6,i}^* \Delta FDI_{t-i} + \lambda ECT_{t-1} + \epsilon_t \quad (4)$$

Where ΔRER_t: The short-run changes in the real exchange rate. ΔPD_{t-i} to ΔFDI_{t-i}The first-differenced independent variables represent their short-run changes. ECT_{t-1}: The error correction term is the lagged residual from the long-run cointegrating regression. λ: The coefficient of the error correction term. ψ_i^* and $\beta_{1,i}^*$ to $\beta_{6,i}^*$: Short-run coefficients on the first-differenced variables.

The long-run relationship is embedded in the ARDL model and can be recovered by solving for the dependent variable (RER_t) when all variables are at their steady-state (or long-run) values. The long-run coefficients are the ratio of the cumulative coefficients of the independent variables to the cumulative coefficient of the lagged dependent variable.

$$RER_t = \alpha_0^{**} + \beta_1^{**} PD_t + \beta_{2,}^{**} INF_t + \beta_{3,}^{**} RINTR_t + \beta_4^{**} EXP_t + \beta_5^{**} IMP_t + \beta_6^{**} FDI_t \ \epsilon_t \quad (5)$$

The long-run coefficient for the j-th independent variable. α_0^{**}: The long-run intercept. ϵ_t: The long-run error term.

The estimation process for the ARDL model involves a series of steps, including testing for stationarity and establishing a long-run association using the bounds test. The process also involves running the long- and short-run regressions, as well as diagnostic tests and model stability checks. To test for stationarity, both informal and formal unit root tests are used.

Data Sources

The study uses time-series data from 2000 to 2024. The data has been obtained from the South African Reserve Bank (SARB) and the World Bank online database. The summary of the variables is presented in Table 1. This research examines the impact of public debt on real exchange rates in South Africa, with the real interest rate as the dependent variable and public debt as an explanatory variable, alongside inflation, real interest rates, foreign direct investment, exports, and imports.

Table 1. Variable description

Variable	Symbol	Source
Public debt	PDT	SARB
Real exchange rate	$RERT$	World Bank
Real interest rate	$RINTRT$	World Bank
Inflation	$INFT$	World Bank
Foreign direct investment	$FDIT$	World Bank
Imports	$IMPT$	Word Bank
Exports	$EXPT$	World Bank

Firstly, unit root tests for stationarity and integration are conducted using graphical analysis, and formal unit root tests are utilised to evaluate the stationarity of the model variables by applying the Augmented Dickey-Fuller and Phillips-Perron tests to assess whether a sequence of data over time series exhibits stationarity or non-stationarity. Secondly, the study determined the suitable lag length using the Schwarz Criterion (SC). The SC is a commonly employed standard for selecting lag length in econometric models, derived from the Bayesian information criterion (BIC). The ARDL bound test for cointegration is presented and scrutinised. Thirdly, after identifying the cointegration, an error correction model is presented. Both short-run and long-run estimates are interpreted. Lastly, diagnostic checks are utilised to check the model's stability.

This study conducted various tests to assess the model's reliability, including the Jarque-Bera test for normality of residuals, White's test, the ARCH test for heteroskedasticity, and the Breusch-Godfrey test for autocorrelation. The CUSUM and CUSUM Q tests are utilised to evaluate the stability and specification of the model.

Results and discussion

Table 2. Correlation matrix

	RERT	PDT	INFT	RINTRT	FDIT	EXPT	IMPT
RERT	1.000						
PDT	-0.700	1.000					
INFT	-0.433	-0.089	1.000				
RINTRT	0.152	-0.060	-0.012	1.000			
FDIT	-0.133	0.171	0.150	-0.381	1.000		
EXPT	-0.750	0.625	0.519	-0.034	0.287	1.000	
IMPT	-0.516	0.245	0.438	0.111	-0.056	0.787	1.000

The study utilises a correlation matrix to assess the relationships among variables, with the findings presented in Table 2. The results reveal that public debt, foreign direct investment, exports, and imports are negatively correlated with real exchange rates. In contrast, inflation and real interest rates are positively correlated with real exchange rates.

The summary statistics for each variable used in the study are presented in Table 3 below. This table provides insights into normality, measures of dispersion, central tendency, and other relevant findings.

Table 3. Descriptive Statistics

	RERT	PDT	RINTRT	INFT	FDIT	EXPT	IMPT
Mean	85.317	43.868	4.381	5.244	1.714	27.732	26.910
Median	80.826	38.045	4.546	5.338	1.064	27.575	26.938
Maximum	104.604	76.355	8.047	9.909	9.660	33.371	33.719
Minimum	69.832	24.044	0.472	-0.692	0.205	22.757	21.808
Std.Dev	11.496	16.444	1.727	2.146	1.998	2.834	3.362
Skewness	0.340	0.732	0.144	-0.281	2.851	0.349	0.222
Kurtosis	1.679	2.252	2.991	4.526	11.409	2.560	2.106
Jarque-Bera	2.299	2.819	0.086	2.759	107.558	0.711	1.039
Probability	0.316	0.244	0.957	0.251	0.000	0.701	0.594
Sum sq.Dev.	3171.949	6490.244	71.641	110.568	95.841	192.801	271.394
Observations	25	25	25	25	25	25	25

Table 4. Unit root test

Test	ADF		PP		Order of integration
Variable	Level	First difference	Level	First difference	ADF & PP
RERT	0.296	0.001	0.241	0.001	I(1)
PDT	0.999	0.038	0.996	0.043	I(1)
RINTR	0.058	0.004	0.092	0.000	I(0)
INF	0.719	0.000	0.330	0.000	I(1)
FDIT	0.001	0.000	0.001	0.000	I(0)
EXPT	0.214	0.000	0.214	0.000	I(1)
MPT	0.065	0.001	0.092	0.000	I(1)

Table 4 provides an overview of the Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests conducted at the level and after differencing. The results indicate that foreign direct investment is stationary in levels. In contrast, all other independent variables included in this study become stationary after first differencing. This study has a mixed order of integration, which includes I(0) and I(1) variables, enabling the use of an ARDL model.

Once the order of integration is identified through the unit root test, this research chooses the maximum lag of 1 based on the Schwarz Criterion (SC). Commonly referred to as the Bayesian Information Criterion (BIC), it favours simpler models, which are essential for minimising over-parameterisation and improving the efficiency of estimates in limited samples. The ARDL model analyses the long- and short-run relationships between public debt and real exchange rates in South Africa, and an ARDL bound test is utilised to detect cointegration.

Table 5. Lag length selection

Lag	LogL	LR	FPE	AIC	SC	HQ
0	-401.093	N/A	1387446.	34.007	34.351	34.098
1	-300.931	133.548*	23494.73*	29.744*	32.493*	30.473*

The research selects the Schwarz Criterion (SC) to determine lag, demonstrating that lag 1 is the best choice. The SC is a commonly employed standard for selecting lag length in econometric models, derived from the Bayesian information criterion (BIC). This is especially beneficial for analysing economic time series, as extended lags can lead to overfitting and subpar performance on out-of-sample data (Kilian & Lütkepohl, 2017; Lütkepohl & Schlaak, 2019).

Table 6. Bounds test for Co-integration

T statistic	Value	K
F Statistic	12.118	6
Critical value bounds (Actual sample size = 24)		
Significance	I(0) Bound	I(1) Bound
10%	2.12	3.23
5%	2.45	3.61
2.5%	2.75	3.99
1%	3.15	4.43

Table 6 shows that the computed F-statistic (12.118) exceeds the upper critical threshold, indicating the presence of long-run cointegration among the variables. Therefore, the null hypothesis of no cointegration is rejected. This aligns with research by (Georgescu et al., 2024), which similarly identified evidence of co-integration among macroeconomic variables.

Table 7: Results from the long-run estimation

Dependent Variable: Real exchange rates				
Variable	Coefficient	Standard errors	t-statistic	p-value
PDT	-1.478	0.374	-3.943	0.001
RINTRT	-3.032	1.574	-1.925	0.074
INFT	-8.793	2.383	-3.689	0.002
FDIT	-0.057	0.949	-0.060	0.952
LEXPT	9.934	3.886	2.556	0.022
IMPT	-4.866	1.835	-2.652	0.018

Table 7 presents empirical results of the long-term relationship between public debt (PDT), real interest rate (RINTRT), inflation (INFT), foreign direct investment (FDIT), exports (EXPT), imports (IMPT), and the real exchange rate in South Africa. The results show that an increase in public debt results in a negative exchange rate in the long run. Essentially, a 1% increase in public debt leads to a -1.47 Rand depreciation. The results are statistically significant. The channel through which this relationship operates is currency depreciation and capital flows. Debt sustainability theories suggest that as debt increases, the country is at greater risk; as a result, investors are likely to channel their investments elsewhere in fear of default and inflationary financing. Therefore, as investors lose confidence in a specific economy, the currency depreciates. On the other hand, capital outflow increases volatility. This is supported by Nyoro and Njaramba (2025). Specifically, they find that elevated public debt levels can lead to a decline in a nation's currency's value.

Additionally, an increase in inflation (INFT) results in depreciation of the domestic currency. Specifically, a 1% increase in inflation results in a -3.0 depreciation in the rand. These results are statistically significant. This is supported by basic macroeconomic expectations in which rising inflation is poised to erode the rand's purchasing power, leading to its depreciation (Magubane, 2025). Another mechanism by which this relationship operates is through higher import costs, diminished investor trust, and slower economic growth.

The results show a negative, statistically significant relationship between imports and the exchange rate, whereas exports are positively related to it. A rise in imports leads to currency depreciation. Rising imports may lead to rand depreciation due to heightened demand for foreign currencies (Matlasedi, 2017). Conversely, an increase in exports results in an appreciation in domestic currency. This is due to increased demand for the rand in the foreign exchange market and, subsequently, to an improvement in the current account balance. Higher demand for the rand from increased exports can boost the currency (Ganelli & Rankin, 2020). Improvements in exports attract foreign investors, increase capital inflows, and strengthen the rand. It is clear, therefore, that Trade openness is one of the major drivers of the exchange rate fluctuations. Stronger exports relative to imports result in rand appreciation, whereas a negative trade balance causes rand depreciation.

Table 8. Short-run estimates and ECM

Dependent Variable: Real exchange rates				
Variable	Coefficient	Standard errors	t-statistic	p-value
C	30.992	21.028	1.473	0.162
PDT	-0.715	0.130	-5.490	0.000
RINTRT(-1)	-1.467	0.060	-2.439	0.028
INFT	-4.255	0.700	-6.073	0.000
FDIT	-0.027	0.458	-0.061	0.952
EXPT(-1)	4.807	1.308	3.675	0.002
IMPT	-2.355	0.572	-4.115	0.001
D(RINTRT)	1.093	0.424	2.577	0.021
D(EXPT)	2.591	1.048	2.471	0.026
CointEq(-1)	-0.483	0.043	-11.008	0.000

Table 8 presents the findings from a short-run estimation and error-correction model. The findings from the short-run estimation and error-correction model reveal a negative, statistically significant relationship between public debt, real interest rates, inflation, imports, and the real exchange rate. However, exports exhibit a positive, statistically significant relationship with the real exchange rate. A rise in public debt leads to a depreciation of the rand. An increase in the real interest rate leads to a decline in the domestic currency. These results are supported by Goonawardhana and Dissanayake (2023). They state that a rise in real interest rates leads to a decline in the real exchange rate, weakening the rand, as higher interest rates might attract foreign investment.

Furthermore, rising inflation reduces the real exchange rate, contributing to a decline in the rand's value. Elevated inflation may erode the currency's purchasing power and lower its value relative to other currencies (Priyatna et al., 2025). An increase in imports leads to a decline in the real exchange rate, resulting in a depreciation of the rand. Increased imports can create higher demand for foreign currency, which, in turn, weakens the rand (Mohr, 2020). Conversely, an increase in exports results in an appreciation of the domestic currency. This aligns with a study by

Kandil (2015), which found that an increase in exports leads to a rise in the real exchange rate, strengthening the rand.

The CointEq(-1) coefficient (-0.484) implies that approximately 48% of the divergence from the long-term equilibrium is adjusted in each period. The negative ECT coefficient of -0.483 confirms a stable long-run relationship where exchange rate volatility corrects toward equilibrium following a debt shock. This value indicates that approximately 48.4% of any divergence from the long-term trend is adjusted within a single period. The relatively high speed of adjustment suggests the system is responsive, clearing nearly half of the equilibrium error almost immediately. Consequently, it would take roughly two periods for the full impact of a shock to dissipate and for the exchange rate to stabilise. This result highlights a robust corrective mechanism that prevents permanent deviations from the established economic fundamentals. The coefficient on the error term is both statistically significant and negative, indicating that the model adjusts for deviations from the long-term equilibrium.

Table 9. Diagnostic tests

Test	Null Hypothesis	test-Statistic	p-value	Decision
Jarque-Bera	Residuals are normally distributed.	4.286	0.117	Residuals are normally distributed.
Arch test	The time series is homoscedastic.	0.115	0.910	The time series data presented in the study exhibit homoscedasticity.
Breusch-Godfrey serial correlation LM test	There is no serial correlation in the residuals.	0.406	0.467	There is no serial correlation in the residuals of the time series data used in this research paper.

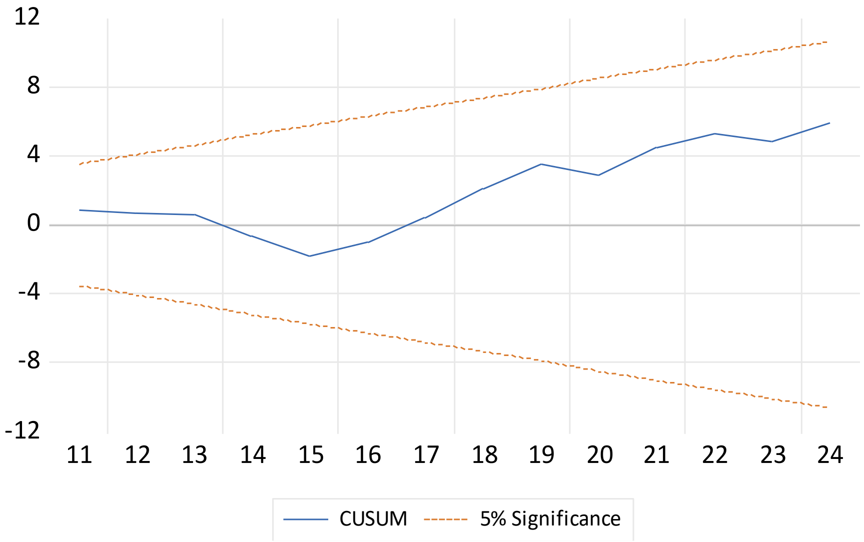

Figure 1. CUSUM Test

We use the Cumulative Sum of Recursive Residuals (CUSUM) and the Cumulative Sum of Squares (CUSUM Q) tests to assess structural stability and specification. Figure 1 shows that the CUSUM chart indicates the line remains within the 5% significance level, indicating that the model's parameters are consistent over time. This uniformity indicates that the model can be dependably utilised for policy guidance and projections.

Based on the CUSUM of squares plot, the line stays within the 5% significance level, indicating that the model's parameters are consistent over time. This consistency implies that the model can be reliably used for policy recommendations and predictions. The lack of substantial deviations from the expected path suggests that the relationships among the model variables are robust and do not often drive significant structural breaks.

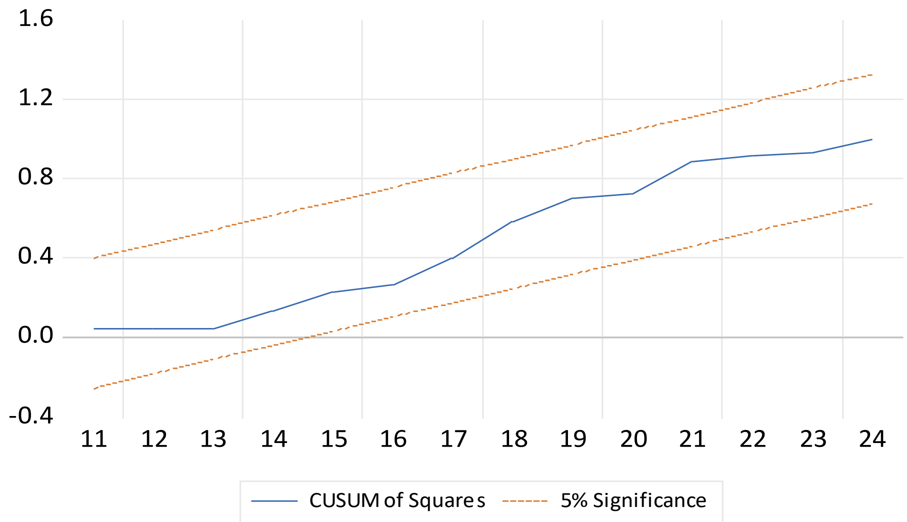

Figure 2. CUSUM of Squares

Conclusion

The study examined the relationship between public debt and real exchange rate volatility in South Africa using time-series data spanning from 2000 to 2024. South Africa is a unique case study in this instance as a country with a growing public debt, a volatile exchange rate, and a heavily import-reliant developing economy. The findings show that public debt is a significant driver of exchange rate volatility. This is the case in both the short- and long-run. This finding supports debt sustainability theories suggesting that rising public debt beyond sustainable levels scares investors due to anticipated default or inflationary financing. This may result in capital outflows, currency depreciation, and increasing exchange rate volatility. Furthermore, the study finds that interest rates, inflation, and trade openness are key macroeconomic factors driving significant fluctuations in South Africa's exchange rate. These fluctuations are more prevalent in the long run than in the short run.

Lastly, the paper recommends firm government controls designed to prevent sudden capital movements (inflows or outflows) that could destabilise the rand. This can be achieved by maintaining a favourable trade balance, targeting inflation, and adjusting monetary policy. Secondly, the government can diversify the composition of its debt currency. This helps reduce volatility in debt-servicing payments and further stabilises government budgets and fiscal planning.

Area for Future Research

Future research could extend this work by exploring threshold effects (non-linearities) in more detail and by examining the real economic costs (e.g., trade, investment) of volatility generated by public debt. It can also decompose debt to capture the exact effect of debt components on the real exchange rate.

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Author contribution

All authors contributed to the conception and design of the study, data collection, analysis and interpretation of the results, and the writing of the manuscript. All authors read and approved the final version of the manuscript.

Use of AI tools declaration

Artificial intelligence tools were used to assist with language editing and to enhance the clarity and readability of the manuscript. The authors remain fully responsible for the content and conclusions of this study.

Conflict of interest
The authors declare that they have no financial or personal relationships that may have inappropriately influenced them in writing this article. Author A is a Master of Commerce student in Economics at Walter Sisulu University; Authors B and C are Lecturers at Walter Sisulu University. No specific grants were received from any funding agency in the public, commercial, or not-for-profit sectors for this research.

References

Armah, M. K., Ofori, I. K., & Andoh, F. K. (2023). A re-examination of the exchange rate — Interest rate differential relationship in Ghana. *Heliyon*, *9*(4), e14605. https://doi.org/10.1016/j.heliyon.2023.e14605

David, O. S. (2024). Impact of macroeconomic variables on exchange rate in south africa. *International Journal of Development and Economic Sustainability*, *12*(3), 26—35. https://doi.org/10.37745/ijdes.13/vol12n32635

Donnat, G. (2025). *Real exchange rate and external public debt in emerging and developing countries*. http://195.220.190.85/GREDEG-WP-2025-15.pdf

Ganelli, G., & Rankin, N. (2020). Fiscal deficits as a source of boom and bust under a common currency. *Journal of International Money and Finance*, *104*, 102149. https://doi.org/10.1016/j.jimonfin.2020.102149

Georgescu, I. A., Oprea, S.-V., & Bâra, A. (2024). Investigating the relationship between macroeconomic indicators, renewables and pollution across diverse regions in the globalization era. *Applied Energy*, *363*, 123077. https://doi.org/10.1016/j.apenergy.2024.123077

Ghauri, S. P., Ahmed, R. R., Streimikiene, D., Qadir, H., & Hayat, A. (2024). Macroeconomic factors driving exchange rate volatility and economic sustainability: case study of Pakistan. *Amfiteatru Economic*, *26*(66), 612—628. https://doi.org/10.24818/EA/2024/66/612

Goonawardhana, V. U. W., & Dissanayake, D. M. U. H. (2023). The Effect of Financial Development on Economic Growth; Incorporation with Impact of the COVID-19 Pandemic: Evidence from Asian Frontier Financial Markets. *South Asian Journal of Business Insights*, *3*(2), 69—89. https://doi.org/10.4038/sajbi.v3i2.58

IMF. (2025). *GLOBAL FINANCIAL STABILITY REPORT Shifting Ground beneath the Calm*. https://www.imf.org/-/media/files/publications/gfsr/2025/october/english/text.pdf

Kandil, M. (2015). On the benefits of nominal appreciations: Contrasting evidence across developed and developing countries. *Borsa Istanbul Review*, *15*(4), 223—236. https://doi.org/10.1016/j.bir.2015.06.003

Khomo, M. M., & Aziakpono, M. J. (2020). The behaviour of the real effective exchange rate of South Africa: Is there a misalignment? *Cogent Economics & Finance*, *8*(1), 1760710. https://doi.org/10.1080/23322039.2020.1760710

Kilian, L., & Lütkepohl, H. (2017). *Structural vector autoregressive analysis:* (1st ed.). Cambridge University Press. https://doi.org/10.1017/9781108164818

Koçak, M. (2025). Evaluation of public debt management practices in turkey from a historical perspective. *Üçüncü Sektör Sosyal Ekonomi Dergisi*, *60*(2), 1997—2010. https://doi.org/10.63556/tisej.2025.1555

Kuranga, A. O., Afolabi, A. C., Jimoh, M. M., & Ojuolape, M. A. (2025). Public Debt Dynamics and Exchange Rate in Nigeria. *Sri Lankan Journal of Business Economics*, *14*(1), 50—70. https://mgt.sjp.ac.lk/bec/wp-content/uploads/2025/06/SLJBE-Vol.-14-1-May-2025-Artical-04.pdf

Lütkepohl, H., & Schlaak, T. (2019). Bootstrapping impulse responses of structural vector autoregressive models identified through GARCH. *Journal of Economic Dynamics and Control*, *101*, 41—61. https://doi.org/10.1016/j.jedc.2019.01.008

Magubane, K. (2025). Determinants of Financial Stability and Development in South Africa: Insights from a Quantile ARDL Model of the South African Financial Cycle. *Journal of Risk and Financial Management*, *18*(9). https://doi.org/10.3390/jrfm18090495

Manguzvane, M. M., & Biyase, M. (2023). *Exchange rate risk and sovereign debt risk in south africa: a regime dependent approach*. Economic and Well-being Research Group. https://www.uj.ac.za/wp-content/uploads/2022/01/manguvzane-biyase-exchange-rate-risk-and-sovereign-debt-risk-in-south-africa-a-regime-dependent-approach.pdf

Matlasedi, T. N. (2017). The influence of the real effective exchange rate and relative prices on South Africa's import demand function: An ARDL approach. *Cogent Economics & Finance*, *5*(1), 1419778. https://doi.org/10.1080/23322039.2017.1419778

Mbaleki, C. (2024). Decomposing south africa's fiscal dynamics through successive policy regimes in the post-apartheid dispensation. *Journal of Public Administration*, *59*(2), 210—222. https://doi.org/10.53973/jopa.2024.59.2.a4

Mbali, A. (2021). *Public debt and economic growth: empirical evidence from South Africa*. http://hdl.handle.net/11427/35872

Mehtiyev, J., Magda, R., & Vasa, L. (2021). Exchange rate impacts on international trade. *Economic Annals-XXI*, *190*(5-6(2)), 12—22. https://doi.org/10.21003/ea.V190-02

Mohr, P. (2020). *Economics for south african students* (6th ed.). Van Schaik Publishers.

Moloi, V. M. M. (2023). The impact of exchange rate, interest rate, and international trade in selected Southern African Development Community (SADC) countries: an ARDL approach. *Eurasian Journal of Economics and Finance*, *11*(3—4), 144—158. https://doi.org/10.15604/ejef.2023.11.0304.004

Nicita, A. (2013). Exchange rates, international trade and trade policies. *International Economics*, *135—136*, 47—61. https://doi.org/10.1016/j.inteco.2013.10.003

Nuru, N. Y., & Gereziher, H. Y. (2021). *The impacts of public expenditure innovations on real exchange rate volatility in South Africa*.

Nyoro, N., & Njaramba, S. G. (2025). Dynamic relationship between public debt and exchange rate misalignment in kenya. *American Journal of Industrial and Business Management*, *15*(11), 1771—1790. https://doi.org/10.4236/ajibm.2025.1511091

Olamide, E., Ogujiuba, K., & Maredza, A. (2022). Exchange Rate Volatility, Inflation and Economic Growth in Developing Countries: Panel Data Approach for SADC. *Economies*, *10*(3). https://doi.org/10.3390/economies10030067

Olaoye, O. O., Tabash, M. I., Omokanmi, O. J., Ogunjumo, R. A., Ojelade, M. O., & Ishola, J. A. (2022). Macroeconomic implications of escalating stock of public debt: Evidence from sub-Saharan African economies. *African Development Review*, *34*(4), 527—540. https://doi.org/https://doi.org/10.1111/1467-8268.12677

Ormaechea, S. A. (2020). Public debt dynamics and intra-year exchange rate fluctuations. *IMF Working Papers*, *20*(261). https://doi.org/10.5089/9781513561585.001

Özen, E., Özdemir, L., & Grima, S. (2020). The relationship between the exchange rate, interest rate, and inflation: the case of Turkey. *Scientific Annals of Economics and Business*, *67*(2), 259 — 275. https://www.ceeol.com/search/article-detail?id=913840

Patel, D., & Mah, G. (2018). Relationship between Real Exchange Rate and Economic Growth: the case of South Africa. In *Journal of Economics and Behavioral Studies* (Vol. 10, Number 1). https://doi.org/10.22610/jebs.v10i1(J).2098

Pesaran, M. H., Shin, Y., & Smith, R. J. (2001). Bounds testing approaches to the analysis of level relationships. *Journal of Applied Econometrics*, *16*(3), 289—326. https://doi.org/10.1002/jae.616

Priyatna, H. N., Suryadi, I., & Ersanti. (2025). Facing global inflation: economic strategies to strengthen people's purchasing power. *MSJ : Majority Science Journal*, *3*(1), 73—81. https://doi.org/10.61942/msj.v3i1.312

Rapetti, M. (2020). The real exchange rate and economic growth: a survey. *Journal of Globalization and Development*, *11*(2), 20190024. https://doi.org/10.1515/jgd-2019-0024

Sang, D. J., Nganai, S., & Kemboi, I. (2025). *Selected Monetary and Fiscal Factors and Exchange Rate Volatility: An Empirical Analysis of Money Supply, Inflation, Foreign Reserves and External Debt in East Africa Partner States*. https://doi.org/10.51244/IJRSI

Usman, K. (2025). A new analytical framework for investigating the debt-growth nexus in E7 economies. *Journal of Applied Economics*, *28*(1), 2600821. https://doi.org/10.1080/15140326.2025.2600821

Yussif, A.-R. B., Onifade, S. T., Ay, A., Canitez, M., & Bekun, F. V. (2022). Modeling the volatility of exchange rate and international trade in Ghana: empirical evidence from GARCH and EGARCH. *Journal of Economic and Administrative Sciences*, *40*(2), 308—324. https://doi.org/10.1108/JEAS-11-2020-0187

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The effect of economic complexity on income levels across countries: A dynamic panel quantile approach

M. Lokman-Hamdan*, W.N.W. Azman-Saini, Yasmin Bani, Anitha Rosland

School of Business and Economics, Universiti Putra Malaysia, Selangor, Malaysia
*Corresponding author: gs58987@student.upm.edu.my

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Authors' emails:
wazman@upm.edu.my
nor_yasmin@upm.edu.my
anitharosland@upm.edu.my

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Abstract

Purpose — This study investigates the impact of economic complexity on income levels across countries at different stages of economic development, with particular emphasis on how these effects vary across the income distribution.

Method — A dynamic panel quantile regression approach is employed to analyse panel data from 115 countries over the period 1995–2020. GDP per capita is used as a proxy for income, allowing the analysis to capture heterogeneous effects across different quantiles of income distribution. The key control variables include human capital, population, trade openness, institutional quality, and inflation.

Findings — The results reveal significant heterogeneity in the effects of economic complexity across income levels. Economic complexity has a positive and significant impact on income at higher quantiles, indicating that more advanced economies benefit from increased productive capabilities. Conversely, at lower quantiles, the effect is negative, suggesting that less-developed countries are unable to fully capitalise on rising complexity.

Implications — The findings suggest that policy strategies should be tailored to different stages of development. Low-income countries need to enhance skill formation and structural transformation to benefit from complexity, while high-income countries should focus on innovation and diversification. Strengthening human capital and institutional quality is essential to mitigating the effects of inequality.

Originality — This study contributes to the literature by highlighting the heterogeneous effects of economic complexity using a dynamic panel quantile framework, offering new insights into income differences across development levels, an aspect largely overlooked in previous research.

Keywords: Economic Complexity, Income Disparity, Panel Quantile Regression, Within-Group Disparity.

Introduction

Decades of research have focused on differences and disparities in income distribution across countries. In this study, income disparity refers to the differences in income levels across countries, capturing variations between lower- and higher-income economies rather than within-country inequality. In 2024, the World Bank estimated that approximately 692 million people lived below the international poverty line of $2.15 per day (in 2017 purchasing power parity terms). The largest percentage of these individuals lived in the least developed regions, such as Sub-Saharan Africa (67%) and South Asia (21%), while high-income countries accounted for only 1%.

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These disparities are unsurprising, as previous studies have shown that developed countries, which produce a diverse and complex range of products, generally exhibit low or modest income disparities (Amarante, Lanzilotta, & Torres-Pérez, 2024; Hartmann, Guevara, Jara-Figueroa, Aristarán, & Hidalgo, 2017; Lee & Vu, 2020). In contrast, least-developed countries tend to produce less complex products and rely heavily on natural resource exports, often exhibiting high levels of disparity (Hartmann, Jara-Figueroa, Guevara, Simoes, & Hidalgo, 2017; Pham, Truong, & Hoang, 2024). Together, these studies provide important evidence that the diversity and sophistication of a country's productive capabilities influence its income distribution.

Studies on income disparity and economic complexity have proposed a measure of the diversity of knowledge that can be translated into products or services (Hausmann, Hidalgo, Bustos, Coscia, & Simoes, 2014; Hidalgo, 2021). Unlike previous approaches that focused on aggregate output and input factors, economic complexity emphasises the productive capabilities embedded in goods and services through the use of machine learning and network techniques (Hidalgo, 2021).

A country with diverse productive capabilities can develop highly sophisticated industries and manufacture complex products, thereby offering a wide range of job opportunities. These countries also have a greater proportion of skilled workers than unskilled workers (Chu & Hoang, 2020). Countries with a higher complexity index can create more complex products and distribute income more fairly (Hartmann, Guevara, et al., 2017). They are highly diversified and export a large number of complex products. In contrast, widening disparities translate to the quality of human capital available in the country. Disparities persist where productive capabilities depend on unskilled workers and primary sector economies. This also implies limited occupational choices in countries with low complexity (Hartmann, Guevara, et al., 2017). Consequently, this constrains their ability to generate and distribute income fairly (Hartmann, Jara-Figueroa, et al., 2017), leading to high-income disparities and low wages for the majority of the population (Chu & Hoang, 2020).

In addition, changes in productive capabilities create more employment opportunities, thereby increasing demand for skilled workers (Hartmann, Guevara, et al., 2017; Lee & Vu, 2020). This enables workers to bargain for wage increases, resulting in lower income disparities. However, the effect of economic complexity on income across countries remains poorly understood. Several studies have found both negative (Hartmann, Guevara, et al., 2017; Lee & Vu, 2020) and positive (Chu & Hoang, 2020) impacts of economic complexity on income disparities.

Existing research often fails to consider whether economic complexity and disparity vary across different income levels. For example, Hartmann, Guevara, et al. (2017) assume that the inequality-reducing effects of complexity are homogeneous across countries. In contrast, Chu and Hoang (2020) find that disparities tend to increase as economies become more complex. However, as Buchinsky (1994) and Martins and Pereira (2004) argue, these studies largely overlook within-group disparities that can be mitigated using quantile regression. Moreover, for cross-country studies, there is a need for a more refined approach that accounts for income distribution at varying levels of complexity. There is a significant income disparity between developed and least developed countries.

This study investigates the impact of economic complexity on income across countries with different development levels. A panel dataset of 115 countries from 1995 to 2020, consisting of developed, developing, and least developed countries, was used. In contrast to past studies (Amarante et al., 2024; Chu & Hoang, 2020; Hartmann, Guevara, et al., 2017; Hartmann, Jara-Figueroa, et al., 2017; Lee & Wang, 2021; Lee & Vu, 2020; Pham et al., 2024; Sepehrdoust, Tartar, & Gholizadeh, 2022), this study utilizes dynamic panel quantiles to estimate the relationship between economic complexity and income disparity. Panel quantiles were selected because economic complexity may also affect the distribution of income across levels of GDP per capita, and the conventional conditional mean regression approach (e.g., GMM, OLS) may obscure substantial parameter heterogeneity in the association between income and economic complexity. Furthermore, this study uses GDP per capita as a proxy for income disparity between countries, unlike many studies that mainly use the Gini index. We take a different approach, objectively, because studies involving the Gini index in cross-country comparisons require careful consideration. Specifically, the Gini coefficient primarily captures within-country income disparities and may not adequately reflect differences across countries.

Measurement of income disparity using GDP per capita

The Gini index is commonly used as a proxy for studies of income disparity. The use of quantiles to examine the entire within-group income distribution (Buchinsky, 1994) has been rendered moot by the Gini index in methods such as quantile regression that rely on the concept of estimating conditional distributions. For this reason, the Gini index does not have a quantile structure (because it is a relative measure rather than an absolute value) suitable for use as the dependent variable. Hence, quantile regression is typically used to estimate the effect of independent variables at different quantiles of the dependent variable's distribution.

Therefore, in this study, GDP per capita is preferred as a proxy for cross-country income gaps rather than for within-country income inequality, for data and methodological reasons. In this study, GDP per capita is also viewed as an approximation of a country's average income (Nolan, Roser, & Thewissen, 2019). For instance, if the impact is positive for the upper quantiles and negative for the lower quantiles, this is taken as evidence of widening income disparity. This approach, also known as within-group income disparity, can be used to measure differences across relevant quantiles (Buchinsky, 1994; Martins & Pereira, 2004). The application of this method highlights the results of the estimation from the longer tails at the end of the income distribution, especially towards less (or more) complex countries.

Furthermore, this technique shows that the impacts of economic complexity differ across income distributions, or, in this case, between low-, medium-, and high-income countries. Finally, by using quantiles, the magnitude of the increase in complexity of GDP per capita can be observed at different points along the income distribution. This provides a clearer and more meaningful basis for comparisons of various income quantiles. Figure 1 shows a positive relationship between economic complexity and GDP per capita, suggesting that greater complexity is associated with higher income levels.

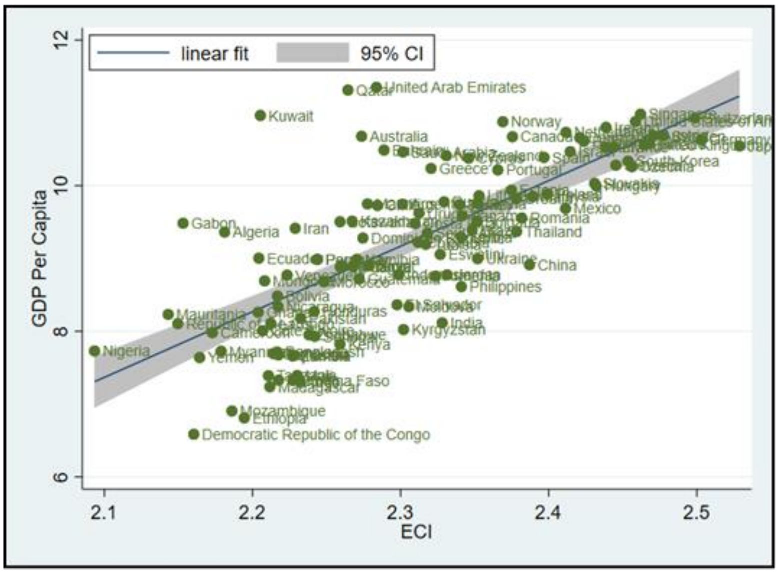

Figure 1. Scatter Plot of Economic Complexity and GDP Per Capita

Methods

This study follows a model proposed by Hartmann, Guevara, et al. (2017), Chu and Hoang, (2020) and Lee and Vu (2020), which stems from the theoretical model of the Kuznets (1955). The effects of economic complexity on income can be expressed in a dynamic form as follows:

$$\mathrm{GDPPC}_{i,t} = \beta_0 + \beta_1 \mathrm{GDPPC}_{i,t-1} + \beta_2 \mathrm{ECI}_{i,t} + \beta_3 \mathrm{HC}_{i,t} + \beta_4 \mathrm{POP}_{i,t} + \beta_5 TO_{i,t} + \beta_6 \mathrm{INS}_{i,t} + \beta_7 \mathrm{INF}_{i,t} + \mu_{i,t} \tag{1}$$

where *i* represents the country, *t* represents time, and μ_t is an error term. *GDPPC*, GDP per capita; ECI, economic complexity index; HC, human capital; POP, total population; TO, trade openness; INS, institutions; and INF, inflation.

Economic complexity, pioneered by Hausmann et al. (2014), refers to the productive capabilities embedded in a country. These capabilities represent the outcome of combining production factors with knowledge, enabling the creation of products or services (Hidalgo & Stojkoski, 2025). In this sense, economic complexity studies the interaction and exchange of knowledge embedded in the economy, weighing and assigning complexity in producing products, thereby contributing to explaining the rate of development, economic growth, and income disparity between countries (Hausmann et al., 2014; Hidalgo, 2021; Hidalgo & Stojkoski, 2025).

By utilizing a method of reflection, the range of products a country can export (diversity) and the number of countries that can make a given product (ubiquity) are iteratively computed. In the concept of economic complexity, a product that can be produced by many is seen as abundant (in terms of productive capabilities). In contrast, products manufactured by a single or a few countries are seen as scarce.

The formulae for deriving the ECI (Hausmann et al., 2014):

$$\text{Diversity} = \mathrm{k}_{c,0} = \textstyle\sum_p M_{cp} \tag{2}$$
$$\text{Ubiquity} = \mathrm{k}_{p,0} = \textstyle\sum_c M_{cp} \tag{3}$$

where M_{cp} is a matrix in which the rows represent different countries and columns represent different products; *c* denotes country, and *p* denotes product. They then jointly and iteratively compute the mean value of diversity and ubiquity to generate a more accurate measure of the number of capabilities available in a country, as follows:

$$\mathrm{k}_{c,N} = \frac{1}{k_{c,0}} \textstyle\sum_p M_{cp} \,.\, k_{p,N-1} \tag{4}$$
$$\mathrm{k}_{p,N} = \frac{1}{k_{c,0}} \textstyle\sum_p M_{cp} \,.\, k_{c,N-1} \tag{5}$$

where *N* denotes the number of iterations. By inserting Equation (5) into Equation (4), we obtain the following equation:

$$\mathrm{k}_{c,N} = \frac{1}{k_{c,0}} \textstyle\sum_p M_{cp} \frac{1}{k_{p,0}} \sum_{c\prime} M_{c\prime p} \,.\, k_{c\prime,N-2} = \sum_{c\prime} k_{c',N-2} \sum \frac{M_{cp}M_{c\prime p}}{k_{c,0}k_{p,0}} \tag{6}$$

This can be rewritten as:

$$\mathrm{k}_{c,N} = \textstyle\sum_{c'} \widetilde{M}_{cc'}\, k_{c',N-2} \tag{7}$$

In which,

$$\widetilde{M}_{cc'} = \sum \frac{M_{cp}M_{c\prime p}}{k_{c,0}k_{p,0}} \tag{8}$$

where $\widetilde{M}_{cc'}$ is a matrix connecting countries exporting similar products. Finally, the ECI is computed as follows:

$$ECI = \frac{\vec{K} - <\vec{K}>}{stdev\ (\vec{K})} \tag{9}$$

where $\vec{K}$ is the eigenvector of $\widetilde{M}_{cc'}$ associated with the second largest eigenvalue, < > represents an average, and stdev denotes the standard deviation.

Data

This study employs an unbalanced panel dataset of 115 countries covering the period 1995–2020, with year selection determined by the availability of economic complexity data. The dependent variable is income, proxied by GDP per capita (constant 2015 US$) from the World Development Indicators (WDI). The independent variable, economic complexity, is represented by the ECI as the primary indicator of productive capabilities. Any changes in productive capabilities are hypothesized to affect income generation, increase complexity, and reduce income disparity in the economy (Hartmann, Guevara, et al., 2017). The data for the control variables are as follows: human capital (HC), population (POP), trade openness (TO), institutions (INS), and inflation (INF). Table 1 summarizes the data used in this study.

Table 1. Data summary

Variable Notation	Variable Name	Variable Description	Measurement Unit	Reference
GDPPC	GDP Per Capita	A proxy to measure income disparity	US$ in 2015	(Buchinsky, 1994; Martins & Pereira, 2004; Nolan et al., 2019)
ECI	Economic Complexity	Economic Complexity Index	Index between -3 and 3.	(Chu & Hoang, 2020; Hartmann, Guevara, et al., 2017; Lee & Vu, 2020)
HC	Human Capital	Human Capital Index	Index between 0 and 5	(Chu & Hoang, 2020; Lee & Vu, 2020)
POP	Population	Population in a country	Sum of the total population	(Lee & Wang, 2021; Lee & Vu, 2020; Morais, Swart, & Jordaan, 2021)
TO	Trade Openness	The sum of exports and imports of goods and services over GDP	Percentage (%) of GDP	(Lee & Wang, 2021; Lee & Vu, 2020; Morais et al., 2021)
INS	Institutions	Institutional Quality	Percentile 0-100	(Chu & Hoang, 2020; Hartmann, Guevara, et al., 2017; Law & Azman-Saini, 2012; Lee & Vu, 2020)
INF	Inflation	Annual percentage change of the consumer price index	Annual percentage (%)	(Glawe & Wagner, 2024; Kamguia, Tadadjeu, Miamo, & Njangang, 2022)

Estimation Techniques

To analyse the effect of economic complexity on income, we utilised quantile regression (Buchinsky, 1994; Koenker & Bassett, 1978; Koenker & Hallock, 2001; Lin, Lee, & Law, 2021; Martins & Pereira, 2004). This technique allows estimation of the conditional quantile function and analysis of the effect of economic complexity on income at different points in the conditional distribution of GDP per capita. It has been chosen because economic complexity may also affect the distribution of income across levels of GDP per capita. Specifically, we are interested in studying the different effects of low- and high-complexity countries at the various quantiles of income distribution.

Furthermore, the dynamic quantile regression approach has been increasingly adopted in empirical growth and inequality studies for its ability to capture heterogeneous effects across the income distribution (Galvao, 2011; Machado & Santos Silva, 2019). It extends the standard quantile framework by incorporating lagged dependent variables, thereby accounting for persistence and dynamic adjustment. Compared to static estimates, our approach provides a more nuanced picture of how economic complexity influences income-disparity segments over time. The advantages of quantile regression include the following: (i) flexibility for modeling data with heterogeneous conditional distributions (Chernozhukov & Hansen, 2008), (ii) median regression is more robust to outliers than ordinary least-squares (OLS) regression (Coad & Rao, 2008), and (iii) it allows for accurate fitting of data with skewed distributions (Kottas & Krnjajić, 2009), which are commonly observed in income datasets.

Additionally, the traditional conditional mean regression approach (such as GMM or OLS) may conceal substantial parameter heterogeneity in the relationship between income and the ECI.

In this case, this technique minimizes $\Sigma_i q|e_i| + \Sigma_i(1-q)|e_i|$, a sum that provides asymmetric penalties $q|e_i|$ for under-prediction and *(1-q)*$|e_i|$ for over-prediction. If ε_i is the model prediction error, OLS minimizes $\Sigma_i e^2{}_i$. In the case of median regression, this technique minimizes $\Sigma_i|e_i|$, which is also known as the least absolute deviations (LAD).

To test the effect of economic complexity on income, the proposed dynamic quantile regression model is as follows:

$$Q_\tau \tau \mid GDPPC_{i,t-1}, X'_{it} = \alpha(\tau)GDPPC_{i,t-1} + X'_{i,t}\beta(\tau) + \mu \tag{10}$$

where X'_{it} is a vector of independent variables; βτ is a k × 1 vector of regression parameters associated with the τ-th percentile. Thus, this study limits the estimations of country-specific effects to be independent of τ across the quantiles. Here, $Q_\tau(GDPPC_{it}|X_{it})$ is τth quantile regression function of $GDPPC$. This estimation recounts projections of quantile functions at the median of the size distribution (50th percentile) and interquartile regressions (0th and 90th percentile) by projecting nine quantile regression functions: Q (0.10), Q (0.20), Q (0.30), Q (0.40), Q (0.50), Q (0.60), Q (0.70), Q (0.80) and Q(0.90). In this framework, the impact of economic complexity on income disparity is assessed by examining the differential effects of economic complexity on GDP per capita across high- and low-income groups.

Results and Discussions

Table 2 provides the summary statistics for the variables used, based on annual data from 115 countries over the 1995–2020 period. The GDPPC ranges from 5.53 (Venezuela in 2019) to 12.02 (Qatar in 2011), with a mean of 9.24. The economic complexity mean is 2.31, ranging between 1.99 (Kyrgyzstan in 1997) and 2.55 (Japan in 2011). The standard deviation of 2.31 (ECI) implies that the ECI scores of the data are clustered around the mean. In contrast, the GDPPC standard deviation of 1.19 indicates a greater spread from the mean. A considerable range is evident across institutions, with a minimum of 0.17 (the Democratic Republic of the Congo in 1998) and a maximum of 4.60 (Finland in 2005).

Table 3 presents the results of the correlation coefficients for all variables used in this analysis. The table shows that two variables (i.e., population and inflation) are negatively correlated with the dependent variable GDPPC, whereas all other variables are positively correlated. It also shows that our main independent variable (i.e., economic complexity) is positively associated with GDP per capita (0.32), which is consistent with the theoretical perspective. In general, all correlation coefficients are below 0.8, indicating no possible problem of collinearity among the variables.

Table 2. Descriptive statistics

Variables	Mean	Std. Dev.	Min.	Max.
GDPPC	9.24	1.19	5.53	12.02
Economic Complexity (ECI)	2.31	0.1	1.99	2.55
Human Capital (HCI)	0.90	0.6	0.05	1.47
Population (POP)	16.50	1.47	13.15	21.07
Trade Openness (TO)	4.27	0.49	2.47	6.08
Institutions (INS)	3.78	0.66	0.17	4.60
Inflation (INF)	4.68	0.15	4.29	7.91

Table 3. Correlation matrix

	GDPPC	ECI	HC	POP	TO	INS	INF
GDPPC	1						
ECI	0.32	1					
HCI	0.76	0.34	1				
POP	-0.16	-0.002	-0.12	1			
TO	0.28	0.13	0.28	-0.61	1		
INS	0.74	0.33	0.58	-0.21	0.25	1	
INF	-0.24	-0.01	-0.13	0.07	-0.11	-0.34	1

To investigate the impact of economic complexity on income, Table 4 presents the results of our main estimates using quantile regression. The table reports the dynamic panel quantile regression results for $\tau = 0.1,\ldots,0.9$. Overall, the effects of economic complexity on income are heterogeneous across quantiles. Quantiles with significant economic complexity coefficients are the 10th, 20th, 40th, 50th, 60th, 70th, 80th, and 90th percentiles. This empirical result demonstrates that the effect is more pronounced at high quantiles, indicating significant positive effects from the 60th to the 90th percentile. At high quantiles, the impact of economic complexity on income is notable at the 90th percentile. This finding seems consistent with the view that more sophisticated capabilities are associated with high-income and developed countries. Thus, improving the productive capabilities of the economy is necessary.

In contrast, the effects of economic complexity on income are negative and significant at lower quantiles, with more substantial effects at the 10th, 20th, and 40th percentiles. Therefore, the positive impact at higher quantiles and the negative impact at the lower quantiles support the idea that economic complexity has widening effects on income disparity. This result aligns with Chu and Hoang (2020) and Hartmann, Guevara, et al. (2017), who found that an increase in economic complexity leads to higher income disparity. The main reason for this widening effect is that much of the workforce in the least-developed countries primarily consists of unskilled workers (Chu & Hoang, 2020; Hartmann, Guevara, et al., 2017; Hartmann, Jara-Figueroa, et al., 2017). However, unlike Amarante et al. (2024), who reported a mitigating effect of complexity on inequality in high-income economies, our results suggest that complexity reinforces disparities when low-income countries lack the absorptive capacity to utilize new capabilities effectively. Thus, improving productive capabilities in lower-quantile countries did not directly increase their incomes or narrow the disparity. It is also possible that low-income countries could only partially absorb this impact. Unequal opportunity distribution benefited only skilled workers and upper-ladder income in the economy. Consequently, this exacerbates disparities in the country.

Interestingly, at several mid-quantiles (30th–50th), the coefficients for ECI appear weak or statistically insignificant, indicating a period of transition among middle-income countries. This pattern suggests that economies in this range are in a structural adjustment phase, benefiting partially from growing productive capabilities but still constrained by limited technological absorption, uneven skill distribution, and institutional inefficiencies. These transitional economies may be developing new industries and capabilities but have not yet achieved the level of diversification and knowledge intensity seen in higher-quantile countries. Consequently, while economic complexity begins to stimulate income growth, its full potential to reduce disparities remains unrealised due to persistent structural constraints.

Across the quantiles, Table 4 shows that trade openness and inflation were positively and significantly associated with GDP per capita in all quantiles. This suggests that, as trade openness and inflation improve, they do not directly influence income disparity, as income tends to increase across all countries regardless of development level. In other words, the effects of trade and inflation on income appear uniform across low- and high-income economies. At the same time, human capital, population, and institutional quality exhibit varying effects across income distribution. This finding is consistent with Chu (2023), which finds that these variables generally have a reducing effect on income disparity, as indicated by their positive impact at lower quantiles and negative impact at higher quantiles. Institutional quality positively affects income per capita across the 10th to 70th percentiles, suggesting that institutional improvements enhance income generation in developing and emerging economies. However, in the 80th and 90th percentiles, the coefficient turns negative, suggesting that further institutional improvements in already high-income countries may yield diminishing returns on income growth.

We then estimate the quantile results simultaneously by testing the equality of the coefficients across quantiles. In doing so, we intend to determine whether the coefficients at lower quantiles differ statistically from those at higher quantiles. Table 5 presents the results for four estimated quantiles: $10^{th} = 90^{th}$, $20^{th} = 80^{th}$, $30^{th} = 70^{th}$, and $40^{th} = 60^{th}$. At the 10^{th} and 90^{th} percentiles, and at the 20^{th} and 80^{th} percentiles, the joint F-statistics results indicate that all coefficients are statistically significant. Thus, the result confirms that the coefficients at lower quantiles are statistically different from those at higher quantiles. However, as we approach the middle quantiles, we see that the F-statistic for the population is not significant at the 30th -70th and 40th - 60th percentiles.

Table 4. Dynamic panel quantile regression

Regressor	Dependent variable: GDPPC								
	Quantile								
	0.1	0.2	0.3	0.4	0.5	0.6	0.7	0.8	0.9
$\ln GDPPC_{t-1}$	0.968*	0.976*	0.978*	0.974*	0.982*	0.987*	0.990*	0.994*	0.995*
	(0.000)	(0.000)	(0.000)	(0.001)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
lnECI	-0.001*	-0.004*	0.000	0.003*	0.000*	0.003*	-0.000*	0.001*	0.002*
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
lnHCI	0.014*	0.016*	0.008*	0.020*	0.012*	0.007*	0.009*	0.003*	-0.005*
	(0.000)	(0.001)	(0.001)	(0.001)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
lnPOP	0.007*	0.003*	0.003*	0.003*	0.003*	0.002*	0.002*	-0.002*	-0.001*
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
lnTO	0.013*	0.009*	0.008*	0.0167*	0.024*	0.017*	0.0186*	0.011*	0.0127*
	(0.000)	(0.000)	(0.002)	(0.001)	(0.001)	(0.000)	(0.000)	(0.000)	(0.001)
lnINS	0.072*	0.038*	0.036*	0.025*	0.014*	0.005*	0.005*	-0.003*	-0.002***
	(0.000)	(0.000)	(0.002)	(0.001)	(0.000)	(0.000)	(0.000)	(0.000)	(0.001)
lnINF	0.015*	0.000	0.017*	0.098*	0.143*	0.216*	0.385*	0.449*	0.534*
	(0.001)	(0.004)	(0.005)	(0.003)	(0.003)	(0.000)	(0.001)	(0.003)	(0.002)
Observations	2,357	2,357	2,357	2,357	2,357	2,357	2,357	2,357	2,357
Number of countries	115	115	115	115	115	115	115	115	115

Note: Standard errors in parentheses. *, **, and *** refer to statistical significance at the 1, 5, and 10% levels, respectively

Table 5. Coefficient differences between lower and upper quantiles

Variables	0.1 = 0.9	0.2 = 0.8	0.3 = 0.7	0.4 = 0.6
	F-stat	*F*-stat	*F*-stat	*F*-stat
ECI	13.03*	15.91*	3.59***	4.93**
HC	8.14*	10.25*	3.23***	2.09**
lnPOP	7.75*	3.50**	0.03	0.01
lnTO	9.35*	2.51**	6.60**	2.73***
lnINS	27.75*	27.83*	17.11*	7.15*
lnINF	74.70*	56.23*	81.22*	25.56*

Note: *, **, and *** indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

Figure 2 graphically illustrates the impact of economic complexity and other control variables on GDP per capita. The horizontal x-axis in the figures shows the quantile scale, and the vertical y-axis shows the estimated coefficients for each variable. As shown in Figure 2(a), past income significantly affects current income, with the effect increasing toward higher quantiles. Figure 2(b) suggests that ECI is significantly negative in low quantiles and positive in medium and high quantiles. This growing effect of economic complexity, which intensifies at higher quantiles, is consistent with Chu and Hoang (2020), who found that disparity increases as countries develop more complex economies. Figure 2(g) shows that lagged GDPPC and inflation have an increasing trend, in contrast to Figures 2(c), (d), and (f), which indicate a decreasing trend for human capital, population, and institutions. Here, we infer that the impact of these variables decreases marginally as quantiles increase. Finally, Figure 2(e) shows sustained fluctuations in the coefficient across quantiles for trade openness.

Robustness Checks

To further validate the dynamic panel quantile results, this study employed a two-step system GMM estimation using a different proxy for income disparity: the Gini index. The Gini index was obtained from the Standardized World Income Inequality Database (SWIID, version 9.5) (Solt, 2020). We utilised the data for the period of 1996 to 2020, and the data were also averaged for 5-year periods in this section. To ensure comparability, the same control variables were retained in both the quantile and GMM estimations. The difference lies only in the dependent variable, GDP per capita (main model) and Gini index (robustness test), which capture different aspects of income disparity. Data for the independent and control variables – ECI, human capital, population, institutions, and inflation – were obtained from previous estimations. However, to successfully investigate the sensitivity of the last result, we selected countries with lower-than-average World GDP per capita in 2020, which was $10,499.6 (constant 2015 US dollars).

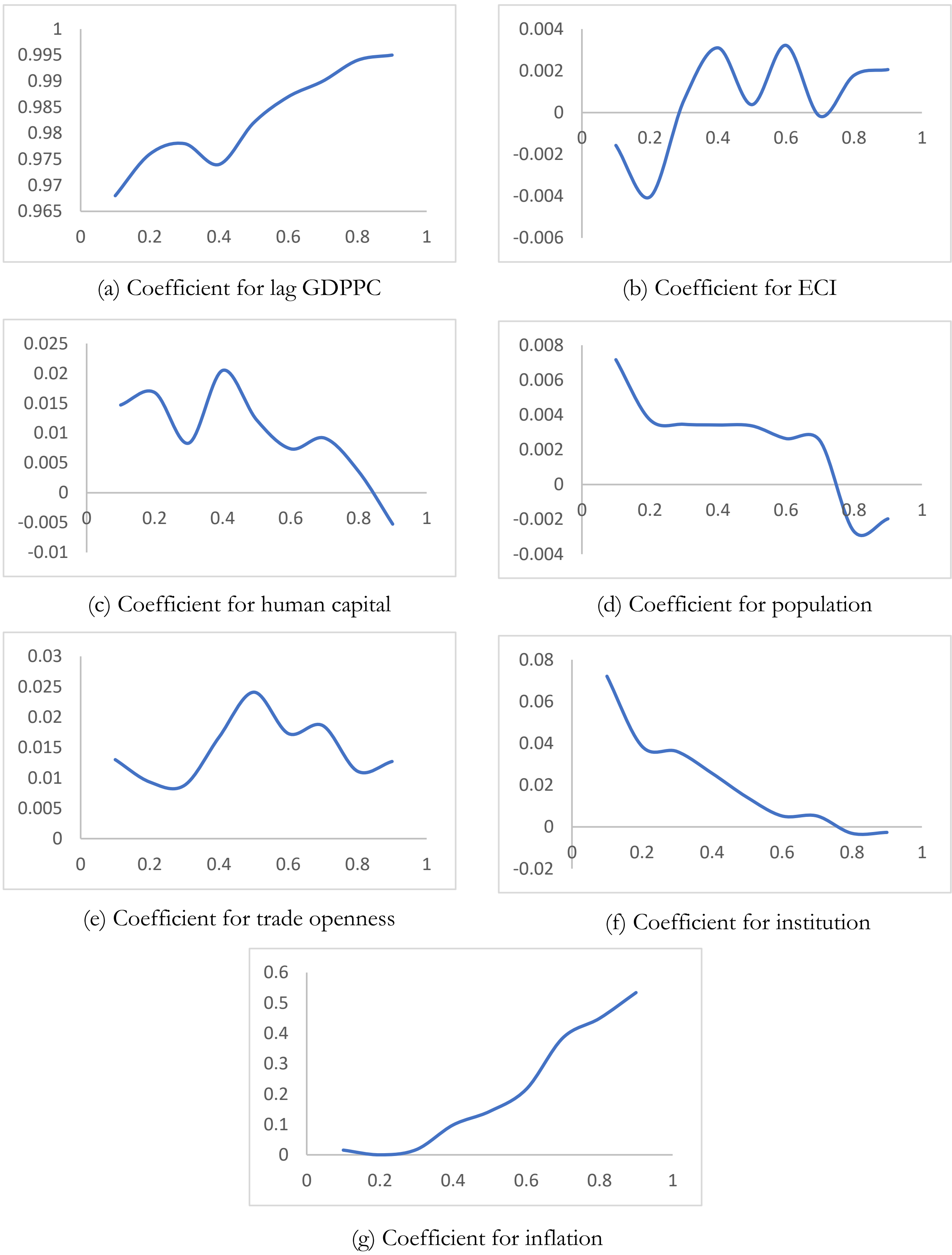

Figure 2. Changes in panel quantiles regression as quantiles vary from 0 to 1

Thus, this sample amounts to 69 countries, representing some middle- and all low-income countries from the quantile estimation. This reduction in sample size may affect the external validity of the robustness check and, if the countries with available Gini data differ systematically from the full panel, introduce sample selection bias. Nevertheless, the direction and significance of the relationship between economic complexity and disparity remain consistent in this subset, offering partial validation of the main findings.

Table 6 shows the robustness test for our estimation, indicating that economic complexity has a statistically significant positive effect on income disparity at the 1% level. The GMM estimates confirm the positive relationship between ECI and Gini, consistent with the upper-quantile results of the dynamic quantile regression, in which economic complexity is associated with higher income disparities. Thus, this result supports previous estimates and shows that increasing productive capabilities in middle- and low-income countries leads to widening disparities in these regions. Furthermore, the persistence of ECI's positive effect across both models suggests that the widening influence of economic complexity on income disparity is not model-specific. However, the stronger effects observed in higher quantiles suggest that the dynamic quantile regression model captures heterogeneity in the income distribution that the GMM approach cannot fully account for.

Additionally, the lagged Gini also affects the current Gini, indicating that this dynamic model is appropriate. Trade openness and inflation appear to positively influence income disparity, whereas human capital and institutions negatively influence the Gini. In addition, the population was found to be insignificant in this estimation. Overall, we can conclude that increasing economic complexity in middle- and low-income countries increases income disparity.

Table 6. System GMM using Gini as a proxy for income disparity without outliers

Regressor	Coefficient	Robust Standard Error
Constant	-0.471*	0.081
Gini $_{t-1}$	0.945*	0.006
ECI	0.084*	0.014
HC	-0.099*	0.035
lnPOP	-0.009	0.049
lnTO	0.089*	0.044
lnINS	-0.008**	0.012
lnINF	0.019**	0.021
Observations	232	
Number of countries	66	
No. of instruments	14	
AR1 (*p*-value)	0.00	
AR2 (*p*-value)	0.413	
Hansen (*p*-value)	0.592	

Note: *, **, and *** refer to statistical significance at the 1, 5, and 10% levels, respectively

Conclusion

This study analyses the effect of economic complexity on income in 115 countries over the 1995–2020 period. The study found significant cross-country impacts using dynamic panel quantile regression. At a high quantile, economic complexity is associated with higher income per capita. Thus, an increase in a country's productive capabilities is linked to higher workers' incomes in developed countries. This study also found that economic complexity negatively impacts income per capita at a low quantile. This implies that an increase in productive capabilities in these countries did not benefit the incomes of low-skilled workers. Thus, the positive impact at higher quantiles and the negative impact at lower quantiles support the idea that economic complexity had a widening effect on income disparity. Meanwhile, increasing human capital, population, and institutional quality reduces income disparity between low- and high-income countries. Finally, increases in trade openness and inflation do not affect income disparity, as income tends to increase across all quantiles.

Based on our findings that economic complexity has a widening effect on income disparity, policy formulation should be strategized to effectively narrow the disparity between high- and low-income countries. The governments of low-income countries should aim to venture into new industries that add value by introducing new productive capabilities in the economy. New productive capabilities not only increase a country's complexity but also contribute to a chain effect in the economy. In high-income countries, policy formation should be structured to foster more complex and diversified activities, especially those that build existing productive capabilities.

Although this study offers useful insights into the relationship between economic complexity and income, several limitations should be noted. First, using GDP per capita as a proxy for income disparity captures differences between countries but not within-country inequality. Second, despite accounting for heterogeneity across income levels, the dynamic panel quantile approach may be subject to endogeneity due to omitted variables, such as labour market structure or capital formation. Lastly, the dataset is unbalanced because not all 115 countries have continuous observations from 1995 to 2020, which may affect the precision of estimates at the lower and upper quantiles. Future research could extend this analysis by incorporating within-country income data to better capture inequality at the individual level. Further studies may also expand the scope to include regional or sectoral data, which could provide a more detailed view of how productive capabilities shape income outcomes across different economic contexts.

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Author contribution

All authors contributed to the conception and design of the study, data collection, analysis and interpretation of the results, and the writing of the manuscript. All authors read and approved the final version of the manuscript.

Use of AI tools declaration

Artificial intelligence tools were used to assist with language editing and to enhance the clarity and readability of the manuscript. The authors remain fully responsible for the content and conclusions of this study.

Conflict of interest

The authors declare that there are no competing interests related to this manuscript.

References

Amarante, V., Lanzilotta, B., & Torres-Pérez, J. (2024). Income inequality and complexity of the productive structure: {New} evidence at the world level. *Economic Analysis and Policy*, *84*, 628–645. https://doi.org/10.1016/j.eap.2024.09.014

Buchinsky, M. (1994). Changes in the US wage structure 1963–1987: Application of quantile regression. *Econometrica*, *62*(2), 405–458.

Chernozhukov, V., & Hansen, C. (2008). Instrumental variable quantile regression: A robust inference approach. *Journal of Econometrics*, *142*(1), 379–398. https://doi.org/10.1016/j.jeconom.2007.06.005

Chu, K. L. (2023). Determinants of economic complexity revisited: Insightful understanding from panel quantile regression. *Journal of Economic and Banking Studies*, *5*, 30–44. https://doi.org/10.59276/jebs.2023.06.2448

Chu, L. K., & Hoang, D. P. (2020). How does economic complexity influence income inequality? {New} evidence from international data. *Economic Analysis and Policy*, *68*, 44–57.

https://doi.org/10.1016/j.eap.2020.08.004

Coad, A., & Rao, R. (2008). Innovation and firm growth in high-tech sectors: A quantile regression approach. *Research Policy*, *37*(4), 633–648. https://doi.org/10.1016/j.respol.2008.01.003

Galvao, A. F. (2011). Quantile regression for dynamic panel data with fixed effects. *Journal of Econometrics*, *164*(1), 142–157. https://doi.org/10.1016/j.jeconom.2011.02.016

Glawe, L., & Wagner, H. (2024). Inflation and inequality: new evidence from a dynamic panel threshold analysis. *International Economics and Economic Policy*, *21*(2), 297–309. https://doi.org/10.1007/s10368-023-00580-x

Hartmann, D., Guevara, M. R., Jara-Figueroa, C., Aristarán, M., & Hidalgo, C. A. (2017). Linking economic complexity, institutions, and income inequality. *World Development*, *93*, 75–93. https://doi.org/10.1016/j.worlddev.2016.12.020

Hartmann, D., Jara-Figueroa, C., Guevara, M., Simoes, A., & Hidalgo, C. A. (2017). The structural constraints of income inequality in Latin America. https://doi.org/10.48550/ARXIV.1701.03770

Hausmann, R., Hidalgo, C. A., Bustos, S., Coscia, M., & Simoes, A. (2014). *The atlas of economic complexity: Mapping paths to prosperity*. Mit Press.

Hidalgo, C. A. (2021). Economic complexity theory and applications. *Nature Reviews Physics*, *3*(2), 92–113. https://doi.org/10.1038/s42254-020-00275-1

Hidalgo, C. A., & Stojkoski, V. (2025). The theory of economic complexity. arXiv. https://doi.org/10.48550/ARXIV.2506.18829

Kamguia, B., Tadadjeu, S., Miamo, C., & Njangang, H. (2022). Does foreign aid impede economic complexity in developing countries? *International Economics*, *169*, 71–88. https://doi.org/10.1016/j.inteco.2021.10.004

Koenker, R., & Bassett, G. (1978). Regression quantiles. *Econometrica*, *46*(1), 33. https://doi.org/10.2307/1913643

Koenker, R., & Hallock, K. F. (2001). Quantile regression. *Journal of Economic Perspectives*, *15*(4), 143–156. https://doi.org/10.1257/jep.15.4.143

Kottas, A., & Krnjajić, M. (2009). Bayesian semiparametric modelling in quantile regression. *Scandinavian Journal of Statistics*, *36*(2), 297–319. https://doi.org/10.1111/j.1467-9469.2008.00626.x

Kuznets, S. (1955). Economic growth and income inequality. *The American Economic Review*, *45*(1), 1–28.

Law, S. H., & Azman-Saini, W. N. W. (2012). Institutional quality, governance, and financial development. *Economics of Governance*, *13*(3), 217–236. https://doi.org/10.1007/s10101-012-0112-z

Lee, C.-C., & Wang, E.-Z. (2021). Economic complexity and income inequality: does country risk matter? *Social Indicators Research*, *154*(1), 35–60. https://doi.org/10.1007/s11205-020-02543-0

Lee, K.-K., & Vu, T. V. (2020). Economic complexity, human capital and income inequality: a cross-country analysis. *The Japanese Economic Review*, *71*(4), 695–718. https://doi.org/10.1007/s42973-019-00026-7

Lin, W. L., Lee, C., & Law, S. H. (2021). Asymmetric effects of corporate sustainability strategy on value creation among global automotive firms: {A} dynamic panel quantile regression approach. *Business Strategy and the Environment*, *30*(2), 931–954. https://doi.org/10.1002/bse.2662

Machado, J. A. F., & Santos Silva, J. M. C. (2019). Quantiles via moments. *Journal of Econometrics*, *213*(1), 145–173. https://doi.org/10.1016/j.jeconom.2019.04.009

Martins, P. S., & Pereira, P. T. (2004). Does education reduce wage inequality? Quantile regression evidence from 16 countries. *Labour Economics*, *11*(3), 355–371. https://doi.org/10.1016/j.labeco.2003.05.003

Morais, M. B., Swart, J., & Jordaan, J. A. (2021). Economic complexity and inequality: does regional productive structure affect income inequality in brazilian states? *Sustainability*, *13*(2), 1006. https://doi.org/10.3390/su13021006

Nolan, B., Roser, M., & Thewissen, S. (2019). Gdp per capita versus median household income: what gives rise to the divergence over time and how does this vary across oecd countries? *Review of Income and Wealth*, *65*(3), 465–494. https://doi.org/10.1111/roiw.12362

Pham, M. H., Truong, H. D. H., & Hoang, D. P. (2024). Economic complexity, shadow economy, and income inequality: fresh evidence from panel data. *International Economic Journal*, *38*(2), 270–292. https://doi.org/10.1080/10168737.2024.2311704

Sepehrdoust, H., Tartar, M., & Gholizadeh, A. (2022). Economic complexity, scientific productivity and income inequality in developing economies. *Economics of Transition and Institutional Change*, *30*(4), 737–752. https://doi.org/10.1111/ecot.12309

Solt, F. (2020). Measuring income inequality across countries and over time: The standardized world income inequality database. *Social Science Quarterly*, *101*(3), 1183–1199.

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Globalisation and growth in Turkiye: Is there a verdict?

Ayuba Napari*, Hasan Vergil, Muhittin Kaplan, Asad Ul Islam Khan

Department of Economics, Ibn Haldun University, Başakşehir/Istanbul, Turkey
*Corresponding author: ayubanapari@ihu.edu.tr/napariayuba@gmail.com

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Authors' emails:
asan.vergil@ihu.edu.tr
muhittin.kaplan@ihu.edu.tr
asad.khan@ihu.edu.tr/asadkhan1983@gamil.com

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Abstract

Purpose — Theoretically, the impact of globalisation is ambiguous, and the empirical evidence inconclusive. This study aims to conclusively determine the effects of the various dimensions of globalisation on Turkey's economic growth.

Method — The study employs the autoregressive distributed lag (ARDL) framework to estimate both short-run and long-run effects of globalisation. Globalisation is measured using the KOF Globalisation Index, disaggregated into de jure, de facto, economic, social, and political dimensions. Human capital is proxied by the Human Development Index (HDI), while physical capital is captured by gross fixed capital formation as a percentage of GDP.

Findings — The results based on the aggregate globalisation index reveal a positive and significant long-run effect of globalisation on Turkey's economic growth. However, neither short-run nor long-run effects are observed when aggregate de jure and de facto globalisation indices are used. Furthermore, economic and social globalisation exert a significant negative impact on GDP growth in the short run, though no long-run effects are detected.

Implications — The findings suggest that while globalisation can support long-term economic growth, its short-term effects—particularly through economic and social channels—may pose adjustment challenges that require appropriate policy responses.

Originality/value — This study contributes to the globalisation–growth literature by providing a comprehensive, disaggregated analysis of globalisation in Turkey using a long time series and the ARDL approach.

Keywords — Globalisation, ARDL, KOF Globalisation Index, Turkiye, Economic Growth

Introduction

The return of President Trump as the 47th President of the United States and the radical, seismic shifts in American foreign policy and bilateral relations have left capitals around the world guessing. President Trump's increasing protectionist rhetoric aligns with Milberg et al. (2024) contention that globalisation has been in decline since at least 2011. This de-globalisation is supposedly characterised by nationalism and protectionism, both of which are fueling a populist political revolution (Morelli & Peluso, 2025).

Despite talk of de-globalisation, Lincicome (2022) forcefully argues that the end of globalisation is not near, with trade in intermediate goods, IT, and service goods still accelerating. Admittedly, the global merchandise trade has been slowing. This, he argued, is a natural and expected evolution, as trade in goods is practically constrained by factors such as transportation costs, tastes,

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and technology. This has nudged globalisation to an evolutionary path, resulting in re-globalisation rather than de-globalisation (Lincicome, 2022). Nonetheless, there seems to be a waning interest in globalisation in the literature. A cursory look at Google N-grams data shows a recession in the popularity of the keywords "Globalisation" and "Globalisation", as depicted in Figure 1. Globalisation and its various versions were hot topics in the early 2000s through the 2010s. Since 2008, however, usage has fluctuated and is now in decline. This, perhaps, is due to the acceptance of globalisation as part of the economic and technological transformational process.

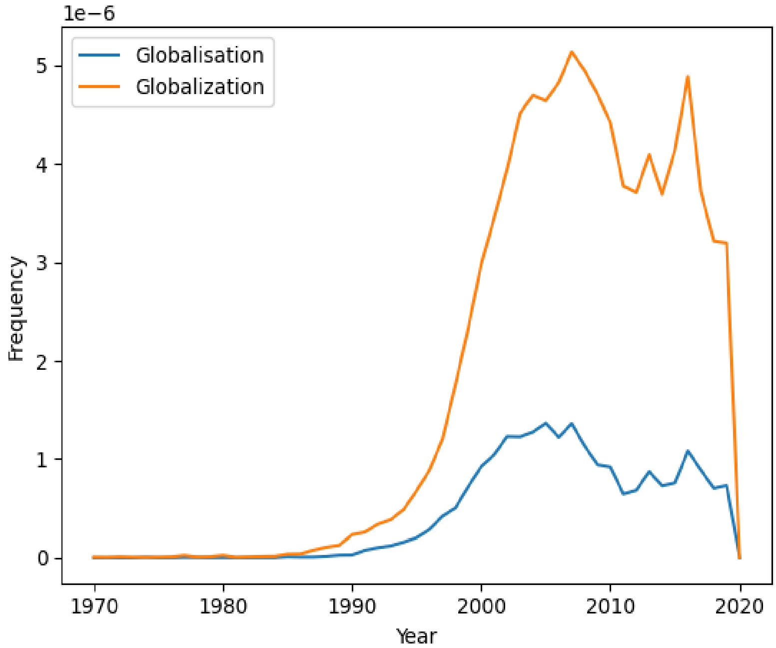

Figure 1. Google N-Gram of Globalisation and Globalisation
Source: Authors' using Data from Google N-grams

Thus, the era of globalisation, as it was known in the first decade of the 21st century, is over, with more questions than answers about its true impact on economies, politics, and society. While some see it as a phenomenon that brings incredible benefits, others criticise its impact on society's cultural and social structures. Economic globalisation is the most debated type, with its effect on economic growth still inconclusive.

The sceptics often trumpet the supposed impact of globalisation on inequalities and environmental pollution. For instance, Goldsmith (1997) makes a definitive claim that Climate change may be the most terrifying problem humanity faces, and economic globalisation accelerates it while, Yay et al. (2016) assert that there is a connection between income disparity and globalisation. The advocates of globalisation, however, see it as economically and, perhaps, socially benign, contending that global economic integration and financial liberalisation that accompany it provide significant benefits for developing countries. Bhagwati (2007), for example, refutes the argument of anti-globalisation advocates, claiming that globalisation, properly managed, can help combat many socio-economic challenges, including poverty.

Beyond the conceptual and moral/social arguments for and against globalisation, many authors have sought to quantitatively gauge the ramifications of globalisation on economies, with contradictory results. Villaverde and Maza (2011) found that globalisation generally promotes economic growth and equality with a semblance of globalisation convergence. Siddharthan (2004) is more cautious after reviewing several studies on the impact of economic globalisation on productivity in India. contending that there is evidence of MNEs having benefited from India's trade and financial liberalisation. Recently, Abubakar (2024) sought to determine the growth impact of globalisation as measured by the KOF Globalisation index in Ghana, Nigeria, Kenya, and South Africa, and concluded a negative but insignificant impact of globalisation on growth. Crafts (2004) stresses that institutional quality is critical in the nexus between globalisation and growth.

This contradictory evidence has also been found for Turkiye. For example, Ojaghlou and Tercan (2024) employed ARDL and DCC-GARCH frameworks, using exports, imports, primary income, remittances, and foreign direct investment as controls, and the KOF index as a gauge of globalisation. They concluded that globalisation has a long-term positive influence on economic growth. However, they did not explain the rationale for the choice of controls. Given that foreign direct investment, real trade flows, income transfers to foreign nationals, and portfolio investment are all used in the calculation of the KOF Economic Globalisation Index (Dreher, 2006), their choice of controls is likely to result in at least multicollinearity.

Çeştepe et al. (2023) also examine the impact of the various dimensions of globalisation on economic growth in Turkiye. Specifically, they used data spanning 1970 to 2018 in an ARDL bounds framework to explore the long- and short-term dynamics among per capita income growth, the KOF globalisation index and its subcomponents, inflation, and the external debt-to-national income ratio. They found that globalisation in the social, political, and economic spheres had a positive impact on per capita income. In contrast, overall globalisation had a negative and significant impact in the long term. In arriving at these conclusions, Çeştepe et al. (2023) modelled overall globalisation as determined by the sub-dimensions of globalisation with inflation and external debt to GDP as controls. The problem with this setup is that the political, social, and economic globalisation are combined to compute the overall globalisation index. As such, using the social, political, and economic globalisation indices, along with the overall globalisation index, as independent variables will result in multicollinearity and conflate the results.

Kılıçarslan and Dumrul (2018) examined the impact of globalisation, in its various forms (de jure and de facto) and types (social, economic, and political), on Turkish economic growth from 1980 to 2015. They found that political globalisation, whether de jure or de facto, decreases economic growth, whereas social globalisation, de facto, increases growth, whereas de jure is inimical to growth. Also, overall economic globalisation was found to have a positive impact on GDP growth, contradicting the findings of Çeştepe et al. (2023) , while the de facto and de jure economic globalisation implied a negative but insignificant impact on growth.

For their part, Bayar and Dabakoğlu (2023) aimed to ascertain if democracy mediates the connection between globalisation and economic growth and found globalisation to have a positive long-term impact on growth, with democratic institutions playing a crucial role. Using human capital and fixed capital formation as control variables, Koyuncu and Sarıtaş (2017) found that while globalisation had a negative, negligible short-term effect on growth in Turkiye, it had a favourable, considerable long-term benefit.

One of the most theoretical and methodologically rigorous studies on Turkiye's globalisation and growth nexus is Utkulu and Özdemir (2004). They underscore the inconclusiveness of endogenous growth models on the trade openness and growth nexus and highlight the inconclusiveness of theory on the direction of causality. They also note the superiority of country-level studies as opposed to cross-country studies, in line with Srinivasan and Bhagwati (2001). In their study, globalisation is measured as the sum of imports and exports as a proportion of GDP. They used real gross domestic investment as a stand-in for physical capital and secondary school enrollment as a stand-in for human capital as controls. They concluded that globalisation, as represented by trade openness, does have a positive impact on economic growth, subject to sound macroeconomic policies and strong investment in both physical and human capital.

Since Utkulu and Özdemir (2004), several globalisation indices have been proposed, one of which is the KOF index. Globalisation indices like the KOF index are more rigorous since they use not just the actual flows of trade but also trade and non-trade restrictions. Also, globalisation indices can capture aspects of globalisation beyond trade, such as financial, political, and social globalisation, which may indirectly affect growth. Utkulu and Özdemir (2004) also used secondary school enrolment as a proxy for human capital development. This, though tolerable, fails to capture the returns to education and the standard of living component of human capital as computed by the United Nations Development Program (UNDP) and even the returns to education as computed by Penn World Tables. Thus, although Utkulu and Özdemir (2004) made significant contributions to conclusively determining the impact of globalisation (trade openness) on economic growth, their study was hindered by data availability problems, which are now surmountable.

The Turkish economy has experienced a series of bursts and recessions since 1970. Figure 2 maps out the oscillations in the GDP per capita of Turkiye from 1970 to 2019.

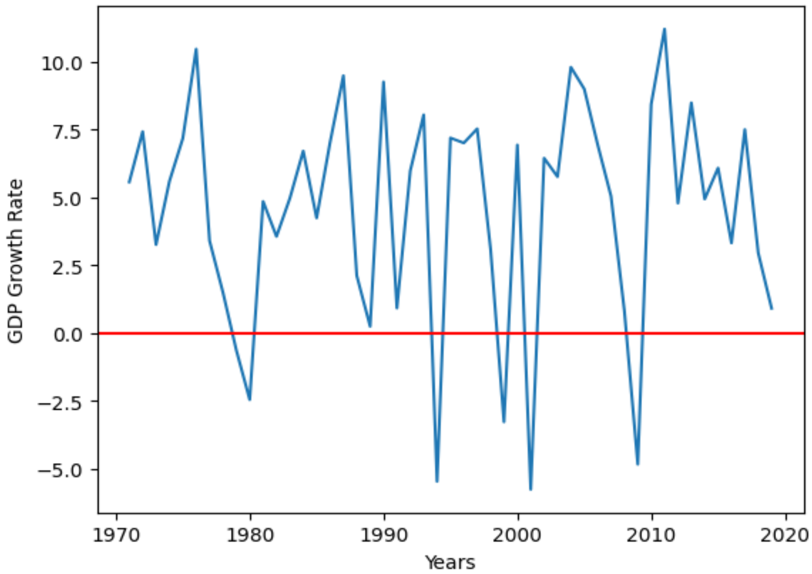

Figure 2. Real GDP Growth (at constant 2017 National Prices Growth)
Source: The authors' use of data from Feenstra et al. (2015)

In the late 1970s, Turkiye struggled to maintain solvency, fight inflation, and restore growth (Singer, 1981), narrowly avoiding recession. From 1978, however, the GDP per capita growth fell below zero, reaching a trough at -2.4% in 1980. The government implemented several measures as part of the 1980 recovery, one of which was to liberalise the economy and pursue an export-driven growth strategy (Ertuğrul & Selçuk, 2001). After the outward, export-led, and globalisation-aligned reforms, the country experienced repeated boom-bust cycles, culminating in the crisis in February 2001. This crisis was mainly precipitated by capital outflows stemming from liberalised capital markets under a crawling peg. The currency immediately fell by almost 30% against the dollar, inflation skyrocketed, government debt nearly doubled, and interest rates rose sharply after the peg was eventually lifted (Dufour & Orhangazi, 2009). This led to a contraction in GDP per capita of over 5.75%.

Dreher (2006) proposes a globalisation index that captures the social, political, and economic dimensions of globalisation. The economic globalisation, as proposed by Dreher (2006) has two types of de facto globalisation and de jure globalisation. The de facto economic globalisation captures actual flows of trade and finance, as well as the extent to which a country employs foreign capital and labour. De jure economic globalisation encompasses restrictions on trade through explicit and implicit import barriers, tariffs, taxes, and capital controls. The de facto and de jure economic globalisation indices are aggregated with equal weighting to obtain the economic globalisation index.

The number of embassies, involvement in UN Security Council missions, and membership in international organisations are weighted at 34%, 34%, and 32%, respectively, to determine the political globalisation index. The social globalisation index is measured using different measures of cultural proximity, information flows, and personal contact. The three dimensions are then aggregated to create the overall globalisation index. The original indices, as proposed by Dreher (2006) have since been disseminated by the KOF Swiss Economic Institute as the KOF Index of Globalisation.

The KOF index has gone through some revisions. One such revision is by Gygli et al. (2019), who distinguish between de jure and de facto globalisation, which is not limited to the economic dimension but has expanded to include de facto and de jure political and social globalisation. Also, the aggregation is designed as a time-varying weighting of the various measures used to compute the indices. The appealing property of the KOF globalisation index is its comprehensiveness. Rather than focusing solely on trade and economics, the index also captures the subtler underlying forces of globalisation. The appetite for imports and exports is likely to be influenced by the extent of exposure to foreign social and political contacts. The trend of Türkiye's various globalisation measures from 1970 to 2021 is plotted in Figure 3. As depicted in the plot, Türkiye has seen a phenomenal increase in all dimensions of globalisation since 1980.

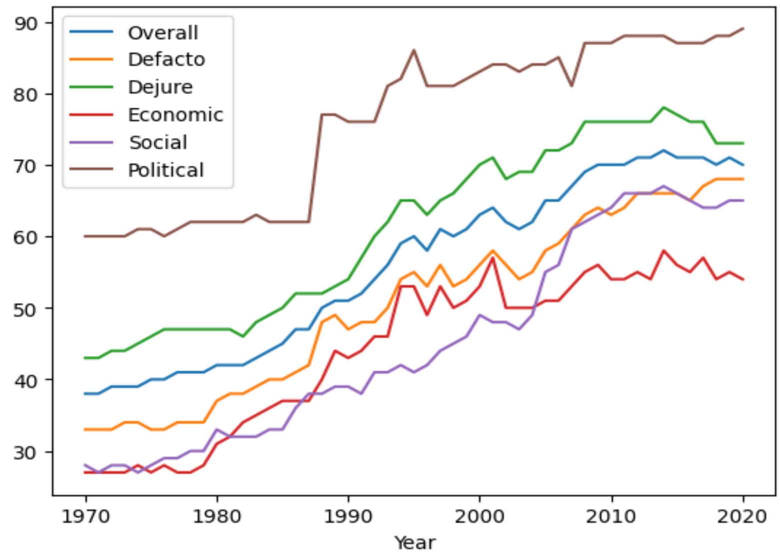

Figure 3. Turkish Globalisation Indices
Source: KOF Globalisation Index

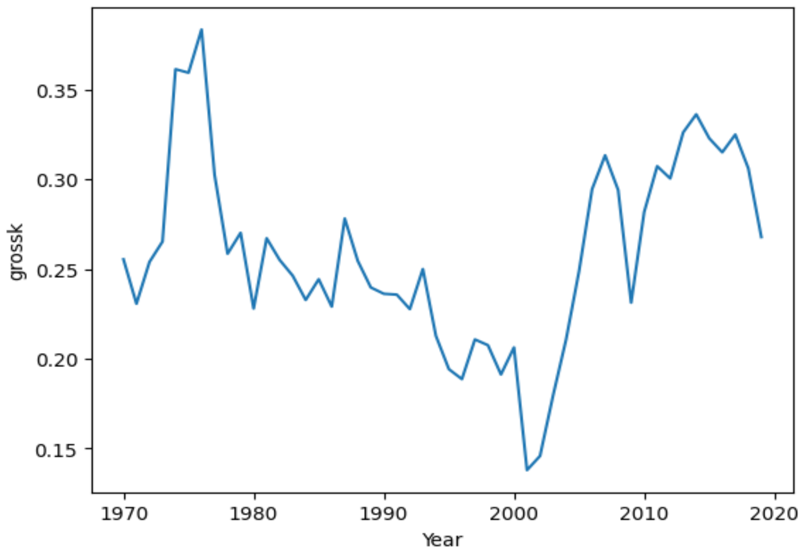

Figure 4. Share of Gross Capital Formation at Current PPPs(grossk)
Source: The authors' use of data from Feenstra et al. (2015).

Over 25% of gross capital formation, the share rose to a staggering 38% in 1976. It has since oscillated, falling to a minimum of 13.8% in 2001. It has since hovered above 21% since 2004.

As the literature indicates, there appears to be disagreement about the overall impact of globalisation on growth, both globally and for Türkiye, with some of the studies exhibiting fundamental deficiencies that this study seeks to address. For example, while Ojaghlou and Tercan (2024) and Bayar and Dabakoğlu (2023) contend that a positive long-term impact on growth, Çeştepe et al. (2023) document a significant negative long-term impact of globalisation on growth.

Our contribution is in three parts. With half a century of data, spanning the era of hyper-globalisation, we hope to provide a more comprehensive answer to the impact of globalisation on economic growth in Turkiye. By using a globalisation index that captures both the de facto and de jure openness, we hope to uncover any subtleties that might have been overlooked in the literature. Also, by employing the UNDP's exhaustive Human Capital Index (HCI), we hope to accurately capture the impact of human development on growth that could have been wrongly ascribed to globalisation. It is found that aggregate globalisation has a positive impact on long-term growth. However, the aggregate and sub-indices are found to have a negative and significant impact on short-term growth.

Methods

In this study, we use the functional form in equation 1 to model the relationship between growth and globalisation.

$$lngdp = f(gi, lhdi, lgrossk) \tag{1}$$

Where *lngdp* is the log of GDP, our proxy for economic growth (Feenstra et al., 2015). The *gi* is our globalisation measure, and the *lhdi* and *lgrossk* are the logs of the human development index and gross capital formation, our control variables, respectively.

We use Utkulu and Özdemir (2004) studies by examining how globalisation, as measured by the aggregate economic, political, and social globalisation indices of KOF, impacts growth. Moreover, we employ the aggregate globalisation index (kofg), the de facto (*dfag*) and de jure (*djag*) aggregate globalisation indices, and the economic (*egli*), social (*sgli*), and political (*pgli*) globalisation indices in the estimations. The recent study spans 1970 to 2019, thus covering half a century of data during the hyper-globalisation era of the 2000s and 2010s.

Since 1990, the United Nations Development Program has published the Human Development Index (HDI), which takes into account life expectancy, mean and projected years of schooling, gross national income (GNI), and health, knowledge, and standard of living. To extrapolate the HDI index from 1970 to 1989, we sourced life expectancy at birth and the Atlas method-computed GNI per capita as reported by the World Bank's World Development Indicators (World Bank, 2017). We then employed Bayesian Ridge Regression to train a predictive model used to extrapolate the HDI index for Turkiye.

We utilise the Autoregressive Distributed Lag (ARDL) model, which is especially appropriate for variables of mixed integration, after determining that one of our independent variables, the lhdi, is stationary (Pesaran et al., 2001). Potential long-term level correlations between the variables are tested using the limits testing approach (Pesaran et al., 2001). Specifically, the ARDL model represents the dependent variable as a combination of short-run and long-run impacts of the independent variables, where the short-run includes first-differenced nonstationary variables. In contrast, the long-run includes level nonstationary variables. The ARDL is as in equation (2).

$$lgdp_t = \varphi + \beta_1 lgdp_{t-1} + \beta_2 gi_{t-1} + \beta_3 lgrossk_{t-1} + \beta_4 lhdi_{t-1} + \sum_{i=1}^{p} \alpha_{i1} \Delta lgdp_{t-i} + \sum_{i=1}^{q_1} \alpha_1 \Delta gi_{t-i} + \sum_{i=1}^{q_2} \alpha_1 \Delta lgrossk_{t-i} + \sum_{i=1}^{q_3} \alpha_1 lhdi_{t-i} + \epsilon_t \tag{2}$$

Where p is the optimum lag selection of the dependent variable $lgdp, q_1$ is the optimum lag for the globalisation measure, q_2 is the optimum lag selected for $lgrossk$ and q_3 is the optimum lag selected for $lhdi$. In selecting the optimum lags, we use the Schwarz (1978) information criterion.

The bounds tests based on Pesaran et al. (2001) are then used to test for the existence of a long-run relationship. Specifically, it compares the ARDL model's F-statistics with the critical values reported in Pesaran et al. (2001). If the F-statistics exceed the upper boundaries of the critical values, the null hypothesis of no cointegration is rejected. Table 3 displays the outcomes of the Bounds test.

The ARDL error correction form for the long-run model is presented in Equation 3.

$$lgdp_t = (\rho + \beta_1 lgdp_t + \beta_2 gi_t + \beta_3 lgrossk_t + \beta_4 lhdi_t) + ECT_t \tag{3}$$

The short-run model is presented in Equation 4.

$$lgdp_t = \varphi + \sum_{i=1}^{p} \alpha_{i1} \Delta lgdp_{t-i} + \sum_{i=1}^{q_1} \alpha_1 \Delta gi_{t-i} + \sum_{i=1}^{q_2} \alpha_1 \Delta lgrossk_{t-i} + \sum_{i=1}^{q_3} \alpha_1 lhdi_{t-i} + \theta ECT_{t-1} + \epsilon_t \quad (4)$$

where θ, also known as the error correction term, denotes how quickly the system returns to equilibrium following a shock.

Results and Discussion

The descriptive statistics are presented in Table 1. For the period of the study, real GDP in constant 2017 dollars averaged 917 billion dollars, with a maximum of 2.2 trillion dollars. Within the period, globalisation, as measured by the KOF Globalisation Index (KOFGI), averaged 55.54. The years in which Turkiye scored the least in globalisation were 1970 and 1971, with a globalisation index score of 38. Since then, Turkiye has become more globalised, with its globalisation index peaking at 72% in 2014. Of all the dimensions of globalisation, Turkiye is more globalised politically, with its political globalisation averaging 77.82, with a maximum of 92.4 and a minimum of 59.41. This is higher than the average level of political globalisation in high-income countries. The Human Development Index (*HDI*) in Turkiye between 1970 and 2021 averaged 0.61 with a standard deviation of 0.14, while the share of gross capital formation (*grossk*) averaged 0.23 with a standard deviation of 0.05. For the estimations, all variables are transformed using the natural logarithm function.

Table 1. Descriptive statistics

	gdp	*Kofg*	*dfag*	*djag*	*egli*	*sgli*	*pgli*	*hdi*	*grossk*
Mean	9.17e+05	55.54	50.14	60.94	43.98	44.82	77.82	0.61	0.23
Standard deviation	5.72e+05	12.04	11.99	12.23	10.96	13.88	12.22	0.14	0.05
Min	2.61+05	37.89	32.97	42.80	26.71	27.48	59.41	0.35	0.14
Max	2.24e+06	71.89	68.38	77.54	57.74	66.67	92.40	0.84	0.38
Observation	50	50	50	50	50	50	50	50	50

Stationarity test

The stationarity test is one of the most important considerations in time series modelling. In this study, we use the Augmented Dickey-Fuller test of Dickey and Fuller (1981), and its variants. This is because the existence of a unit root invalidates mean reversion. Also, the presence or absence of unit roots can help determine the data-generating process of a series and whether the trend is deterministic or stochastic (Phillips & Perron, 1988). The ADF test corrects for autocorrelation by including the lagged terms of the series, while the Phillips-Perron test extends the ADF test by correcting for heteroscedasticity. The stationarity results are presented in Table 2.

Table 2. Unit Root Tests

Variable	ADF I(0) Intercept	ADF I(0) Trend	PP I(0) Intercept	PP I(0) Trend	ADF I(1) Intercept	ADF I(1) Trend	PP I(1) Intercept	PP I(1) Trend
Lgdp	-0.13	-2.71	-0.11	-2.83	-6.77***	-6.70***	-6.77***	-6.69***
Lgrossk	-2.16	-2.14	-2.15	-2.144	-6.88***	-6.79***	-6.88***	-6.79***
Lkofo	-1.27	-0.53	-1.29	-0.59	-6.60***	-6.75***	-6.60***	-6.75***
Ldfag	-0.77	-1.78	-0.77	-1.85	-6.69***	-5.44***	-6.76***	-6.73***
Ldiag	-1.55	-0.03	-1.44	-0.42	-5.25***	-5.44***	-5.24***	-5.43***
Legli	-1.57	-1.06	-1.49	-0.94	-7.71***	-5.45***	-7.71***	-7.88***
Lsgli	-0.67	-2.32	-0.50	-2.32	-5.98***	-5.93***	-5.97***	-5.931***
Lpgli	-1.06	-1.94	-1.14	-1.43	-2.51	-2.55	-7.86***	-7.91***
Lhdi	-5.77***	-2.59	-5.26***	-2.44	-4.45***	-6.23***	-4.61***	-6.36***

Notes: *,**and***represent significance at 1%, 5% and 1% levels, respectively

From Table 2, both tests fail to reject the null of a unit root at levels for all variables except the log of the human development index, which is found to be intercept-stationary. Taking the first difference, however, the null of a unit root is rejected at all levels of significance for all the variables except for the *lpgli*. The ADF fails to reject the unit root for the lpgli even after the first differencing. As such, it was dropped as likely to be integrated of the second order.

As shown in Table 3, the estimated F-statistics exceed the upper bound for all models. This implies the existence of a long-run level relationship among our variables.

Table 3. Cointegration Tests

Model	Empirical model	Optimum lag	F-statistic	Lower bound	Upper bound
Model 1	$lngdp = f(lkofg, lhdi, lgrossk)$	1, 3,3,0	15.57***	3.65	4.66
Model 2	$lngdp = f(ldfag, lhdi, lgrossk)$	1,2,2,1	5.64***	3.65	4.66
Model 3	$lngdp = f(ldjag, lhdi, lgrossk)$	1,1,0,1	20.42***	3.65	4.66
Model 4	$lngdp = f(legli, lhdi, lgrossk)$	4,3,3,0	16.55***	3.65	4.66
Model 5	$lngdp = f(lsgli, lhdi, lgrossk)$	2,3,2,1	4.80***	3.65	4.66

Notes: *,**and***represent significance at 1%, 5% and 1% levels, respectively

Table 4. Long-Run Coefficients from the ARDL Models

	Model 1	Model 2	Model 3	Model 4	Model 5
C	12.77***	13.99***	12.33***	15.57***	13.61***
Lkofg	0.66**				
Ldfag		0.48			
Ldjag			0.81		
Legli				0.09	
Lsgli					0.47
Lhdi	1.79***	2.08**	1.78***	2.35***	2.03**
Lgrossk	0.53***	0.63***	0.59***	0.64***	0.40**

Notes: *,**and***represent significance at 1%, 5% and 1% levels, respectively

Table 5. Short Run Coefficients of the ARDL Models

	Model 1	Model 2	Model 3	Model 4	Model 5
$D(LGDP(-1))$				-0.09	0.17*
$D(LGDP(-2))$				-0.10	
$D(LGDP(-3))$				-0.18**	
$D(LKOFG)$	-0.70***				
$D(LKOFG(-1))$	-0.64***				
$D(LKOFG(-2))$	-0.66***				
$D(LDFAG)$		-0.16			
$D(LDFAG(-1))$		-0.23*			
$D(LDJAG)$			-0.32		
$D(LEGLI)$				-0.43***	
$D(LEGLI(-1))$				-0.28***	
$D(LEGLI(-2))$				-0.23***	
$D(LSGLI)$					0.05
$D(LSGLI(-1))$					-0.18
$D(LSGLI(-2))$					-0.39***
$D(LHDI)$	0.98**	1.19***		1.06**	1.52***
$D(LHDI(-1))$	-1.19**	-0.90**		-0.98**	-1.06***
$D(LHDI(-2))$	-0.69*			-0.68	
$D(LGROSSK)$		0.21***	0.23***		0.19***
$ECT(-1)*$	-0.36***	-0.12***	-0.12***	-0.31***	-0.15***
$R-square$	0.99	0.99	0.99	0.99	0.71
$BG\ Test$	0.32	1.44	0.42	0.06	1.09
$Breusch-Pagan$	1.02	0.67	1.33	1.27	1.19

Notes:*,**and***represent significance at 1%, 5% and 1% levels, respectively

Table 4 shows the long-run estimation results. The estimation results show that the human development index ($lhdi$) and gross capital formation ($lgrossk$) are statistically significant and economically meaningful in all models. However, the impact of globalisation on growth measures is only significant in Model 1, namely the aggregate globalisation index (*lkogf*). This finding implies that in the long-run estimation, only total globalisation significantly affects GDP growth, assuming all other factors remain constant. This finding is consistent with the findings of Bayar and Dabakoğlu (2023) and Koyuncu and Sarıtaş (2017), which document the positive long-run impact of globalisation on economic growth. The control variables, HDI, and gross capital formation are also found to have a consistent positive impact on economic growth. This is in line with the theory and evidence documented by Ulas and Keskin (2017) and Tunali and Boru (2019). The short-run model is presented in Equation 4. The results of Model 4 testing indicate a negative and significant short-run impact of economic globalisation on growth. Model 5, which examines the impact of social globalisation, documents a significant but negative impact of social globalisation on growth with a time lag.

Table 5 shows that aggregate globalisation is a disincentive to growth in the short run. The contemporaneous, first, and second lags of aggregate globalisation are found to be negative and statistically significant at the 1% level. The de facto globalisation index indicates a weak yet significant negative impact on growth, with an insignificant contemporaneous term, whereas the de jure index is negative but insignificant. The de jure index captures formal regulatory globalisation, which involves the removal of trade barriers and the establishment of diplomatic relations. In contrast, the de facto index measures the actual impact in terms of trade and openness. This indicates a gap between policy and practice in Turkey's engagement with globalisation. Essentially, the institutional infrastructure required to translate globalisation policies into impactful economic activity needs improvement.

This finding is consistent with the findings of Atiyas (2012), who noted a sectoral difference in Turkish deregulation, resulting in a divergence between de jure rules and de facto implementation. Ucer and Acemoglu (2015) also note that Turkiye has, at various points, adopted the formal architecture of an open, liberalised economy without delivering the institutions that enable firms to translate openness into investment and growth, especially after the collapse of EU accession talks. Also, it is noted in Acemoğlu and Üçer (2021) that Turkish is riddled with institutional frictions that prevent formal trade and investment liberalisation (de jure) from generating the required spillovers, while the de facto trade flows remain insufficient to drive sustained growth

These impacts could be seen as an adjustment phenomenon reflecting the structural vulnerabilities that make short-run openness costly. For example, Turkiye is considered to be considerably dependent on external sources of energy and intermediate goods, and it experiences a rise in the current account deficit during periods of high economic growth (Gür et al., 2019). This dependency means that deeper economic integration, by expanding trade and investment flows, initially increases import bills rather than spurring domestic productivity. Also, the welcoming of foreign capital during the liberalisation phase resulted in the overvaluation of the currency (Gür et al., 2019) and depressed output in the short run (Bilgili et al., 2024), even as globalisation deepens.

Several factors could explain the lagged negative effect of social globalisation on economic growth. First, with social globalisation comes cross-border awareness of foreign goods, foreign earnings, and living standards. Awareness of foreign goods could increase the preference for them and, with a year or two lag, imports, thereby widening the current account deficit and compressing output. Also, social globalisation also exposes the populace to the pay scales and living standards of other countries, which could lead to the classic phenomenon of "brain drain" while globalisation often leads to Cosmopolitization of society, with its negative impact on social trust (Seitova & Kovacs, 2025). This results in a degradation of informal coordination and trust networks that reduce transaction costs, before formal institutional substitutes are in place. Together, these channels could explain the short-run cost of social globalisation materialising, albeit with a delay. These findings possible to pass the verdict that globalisation, in its various forms, overall, de facto, economic and social, is inimical to growth in Turkey in the short run but has a positive impact in the long run with differing levels of significance.

On the control variables, the consistently positive and significant long-run coefficient on HDI across all models confirms that human capital is a robust enabler of growth in Türkiye, a finding aligned with the earlier work of Utkulu and Özdemir (2004) and consistent with the endogenous growth literature. As to whether human capital has a mediating role in transforming short-term globalisation costs into long-term gains, Prasad et al. (2003) contend that the effects of FDI, a key component of globalisation, on growth are dependent on human capital, with countries that meet a certain HDI threshold found to translate FDI into growth much more seamlessly. In the case of Turkiye, the HDI averaged 0.61 over the study period, rising to 0.84 by 2019. The UNDP classifies countries with HDI values above 0.80 as "very high", a threshold Türkiye began approaching only toward the end of the sample. This trajectory is consistent with the finding that globalisation's aggregate effect is positive only in the long run. As HDI has risen over time, Türkiye's capacity to absorb and capitalise on globalisation has progressively improved.

The Error Correction Term (ECT) for all the models is theoretically and statistically significant. For Model 1, the ECT is -0.36. This indicates that approximately 36% of any shock to the model is corrected within a year, and the economy returns to equilibrium within approximately 2.8 years. For Models 2 and 3, the adjustment is much slower, taking approximately 8.3 years. Model 4 implies approximately 3.2 years for full adjustment. These imply that the speed of adjustment varies across models but consistently confirm the existence of a stable long-run equilibrium relationship.

To check for the stability of the model coefficients, the classic CUSUM test by (Page, 1954) is used. Figure 5 presents the CUSUM plot for Model 1.

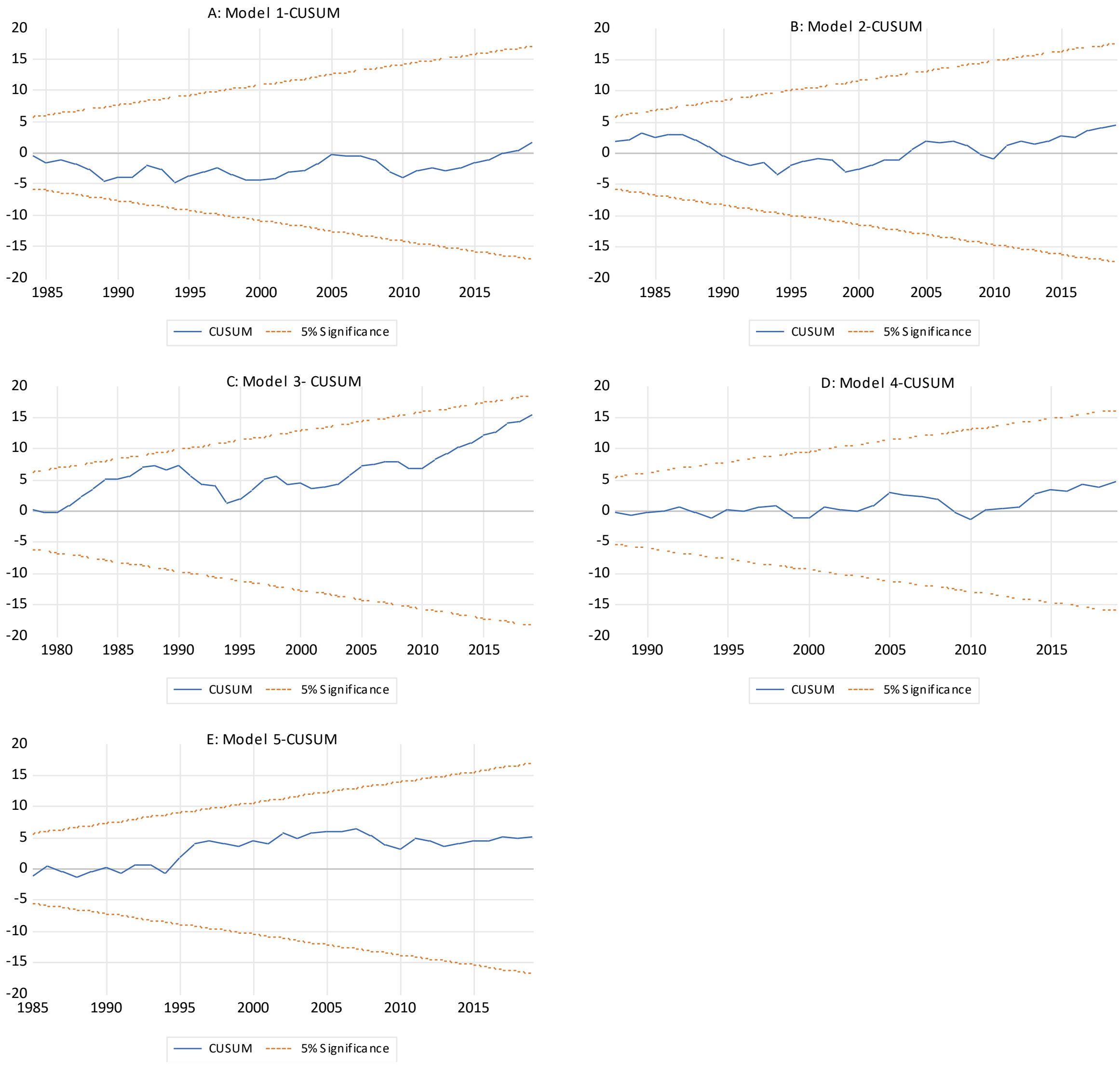

Figure 5. CUSUM Parameter Stability Tests

As depicted in Figure 5 Panel A, Model 1 is stable with the cumulative sum (CUSUM) of recursive residuals remaining within the 5% significance bounds throughout the entire sample period (1970–2019), confirming the structural stability of the estimated parameters. Similar stability is observed across Models 2 to 5 as depicted in Panels B to E.

Conclusion

The empirical analysis using the ARDL bounds approach finds that overall globalisation, as measured by the aggregate KOF globalisation index, is an incentive for long-run GDP growth. However, it is found to be a disincentive to short-run GDP growth in Türkiye. The other measures of globalisation, de facto, de jure, economic, and social globalisation, are found to have no significant impact on GDP growth in the long run. Their impact in the short run, however, is mixed. For example, while the de facto and de jure aggregate globalisation measures are insignificant, the economic and social globalisation is found to have a negative impact on GDP growth in the short run.

A central vulnerability exposed by the short-run results is Turkiye's susceptibility to hot-money flows and exchange-rate volatility. The IMF has recommended restricting foreign currency borrowing without natural hedges, and in the event of tail risks materialising, allowing an orderly depreciation of the exchange rate while letting automatic stabilisers play out IMF (2018). Another approach could be to extend and institutionalise the Reserve Option Mechanism, which has been shown to have a stabilising role on excessive exchange rate movements (Çelik & Oğuş Binatlı, 2022). This would constitute a market-friendly automatic buffer against the sudden stops that amplify the short-run costs of financial globalisation.

To ameliorate the short-term impact of social globalisation, a basic social safety net alongside investments in health and education will go a long way to spur high-quality and shared growth (Acemoğlu & Üçer, 2021). Lastly, undifferentiated, rapid deepening of economic and social globalisation should not be pursued without the domestic preconditions to absorb them. The disaggregated results offer a clear, if nuanced, strategic guide for Türkiye's globalisation policy. Not all dimensions of globalisation are equally growth-oriented, and policymakers should calibrate their engagement accordingly.

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Author contribution
All authors contributed to the conception and design of the study, data collection, analysis and interpretation of the results, and the writing of the manuscript. All authors read and approved the final version of the manuscript.

Use of AI tools declaration
Artificial intelligence tools were used to assist with language editing and to enhance the clarity and readability of the manuscript. The authors remain fully responsible for the content and conclusions of this study.

Conflict of interest
The authors declare that there are no competing interests related to this manuscript.

References

Abubakar, M. (2024). Globalisation and Output Growth Nexus in Sub-Saharan Africa: the Critical Role of Trade Liberalisation. *Journal of the Knowledge Economy*, *15*(1), 2218–2240. https://doi.org/10.1007/s13132-023-01230-y

Acemoğlu, D., & Üçer, E. M. (2021). High-quality versus low-quality growth in Turkey: Causes and consequences. In *Turkish economy at the crossroads: Facing the challenges ahead* (pp. 37–89). World Scientific.

Atiyas, İ. (2012). Economic institutions and institutional change in Turkey during the neoliberal era. *New Perspectives on Turkey*, *47*, 57–81. https://doi.org/10.1017/S0896634600001709

Bayar, İ., & Dabakoğlu, M. (2023). Demokrasi, Küreselleşme ve Ekonomik Büyüme: Türkiye Üzerine Ampirik Bir Analiz. *Turkish Studies-Economics,Finance,Politics*, *Volume 18 Issue 2*(Volume 18 Issue 2), 413–428. https://doi.org/10.7827/turkishstudies.68454

Bhagwati, J. (2007). Why the critics of globalisation are mistaken. *Economist*, *155*(1), 1–21. https://doi.org/10.1007/s10645-006-9040-z

Bilgili, F., Kuşkaya, S., Gençoğlu, P., Ünlü, F., & Magazzino, C. (2024). Exchange rate pass-through at high and low frequencies in Turkey: a continuous wavelet transform model approach. *Journal of Economic Structures*, *13*(1). https://doi.org/10.1186/s40008-024-00341-2

Çelik, M., & Oğuş Binatlı, A. (2022). How Effective Are Macroprudential Policy Instruments? Evidence from Turkey. *Economies*, *10*(4), 76. https://doi.org/10.3390/economies10040076

Çeştepe, H., Ergün Tatar, H., & Erdoğan, S. (2023). The Impact of Globalisation with its Different Aspects on Economic Growth: The Case of Turkey. *Istanbul Journal of Economics/İstanbul İktisat Dergisi*, *73*(2), 717–743. https://doi.org/10.26650/ISTJECON2023-1248345

Crafts, N. (2004). Globalisation and Economic Growth: A Historical Perspective. *The World Economy*, *27*(1), 45–58. https://doi.org/10.1111/j.1467-9701.2004.00587.x

Dickey, D. A., & Fuller, W. A. (1981). Likelihood Ratio Statistics for Autoregressive Time Series with a Unit Root. *Econometrica*, *49*(4), 1057. https://doi.org/10.2307/1912517

Dreher, A. (2006). Does globalisation affect growth? Evidence from a new index of globalisation. *Applied Economics*, *38*(10), 1091–1110. https://doi.org/10.1080/00036840500392078

Dufour, M., & Orhangazi, Ö. (2009). The 2000–2001 Financial Crisis in Turkey: A Crisis for Whom? *Review of Political Economy*, *21*(1), 101–122. https://doi.org/10.1080/09538250802517014

Ertuğrul, A., & Selçuk, F. (2001). A Brief Account of the Turkish Economy, 1980-2000. *Russian & East European Finance and Trade*, *37*(6), 6–30. http://www.jstor.org/stable/27749598

Feenstra, R. C., Inklaar, R., & Timmer, M. P. (2015). The Next Generation of the Penn World Table. *American Economic Review*, *105*(10), 3150–3182. https://doi.org/10.1257/aer.20130954

Goldsmith, E. (1997). Can the environment survive the global economy? *The Ecologist*, *27*, 242+. https://link.gale.com/apps/doc/A20392055/AONE?u=anon~66763e50&amp

Gür, N., Tatlıyer, M., & Dilek, Ş. (2019). The Turkish Economy at the Crossroads: The Political Economy of the 2018 Financial Turbulence. *Insight Turkey*, *21*(4). https://doi.org/10.25253/99.2019214.08

Gygli, S., Haelg, F., Potrafke, N., & Sturm, J.-E. (2019). The KOF Globalisation Index – revisited. *The Review of International Organizations*, *14*(3), 543–574. https://doi.org/10.1007/s11558-019-09344-2

IMF. (2018). Turkey: 2018 Article IV Consultation—Press Release; Staff Report; and Statement by the Executive Director for Turkey. In *IMF Country Report* (Vol. 18, Number 110). International Monetary Fund. https://doi.org/10.5089/9781484348178.002

Kılıçarslan, Z., & Dumrul, Y. (2018). International Journal of Economics and Financial Issues The Impact of Globalisation on Economic Growth: Empirical Evidence from the Turkey. *International Journal of Economics and Financial Issues*, *8*(5), 115–123. http://www.econjournals.com

Koyuncu, C., & Sarıtaş, T. (2017). Analysis of Globalisation and Economic Growth for Turkey with ARDL Model. *Anadolu Üniversitesi Sosyal Bilimler Dergisi*, *17*(2), 51–66. https://doi.org/10.18037/ausbd.417241

Lincicome, S. (2022, June 7). *Globalisation Is Alive, Well, and Changing | Cato Institute*. https://www.cato.org/commentary/globalisation-alive-well-changing

Milberg, W., Liess, T., & Tedesco, M. (2024). Globalisation after De-globalisation. *Social Research*, *91*(3), 939–971. https://doi.org/10.1353/sor.2024.a938583

Morelli, M., & Peluso, E. (2025). *The Populist Geopolitical Order* (Number 03). LISER.

Ojaghlou, M., & Tercan, C. (2024). Impact of liberalization and globalisation on economic growth: Empirical evidence from Türkiye. *JOURNAL OF INTERNATIONAL STUDIES*, *17*(1), 35–51. https://doi.org/10.14254/2071-8330.2024/17-1/3

Page, E. S. (1954). CONTINUOUS INSPECTION SCHEMES. *Biometrika*, *41*(1–2), 100–115. https://doi.org/10.1093/biomet/41.1-2.100

Pesaran, M. H., Shin, Y., & Smith, R. J. (2001). Bounds testing approaches to the analysis of level relationships. *Journal of Applied Econometrics*, *16*(3), 289–326. https://doi.org/10.1002/jae.616

Phillips, P. C. B., & Perron, P. (1988). Testing for a Unit Root in Time Series Regression. *Biometrika*, *75*(2), 335. https://doi.org/10.2307/2336182

Prasad, E., Rogoff, K., Wei, S.-J., Ayhan Kose, M., Bayoumi, T., Berg, A., Clark, P., Edison, H., Husain, A., Jeanne, O., Kumar, M., Mody, A., Morsink, J., Reinhart, C., Robinson, D. J., Sahay, R., Savastano, M., other colleagues Priya Joshi, numerous, Kim, Y., … Wu, Y. (2003). *Effects of Financial Globalisation on Developing Countries: Some Empirical Evidence*. https://www.imf.org/external/np/res/docs/2003/031703.pdf

Schwarz, G. (1978). Estimating the Dimension of a Model. *The Annals of Statistics*, *6*(2), 461–464. http://www.jstor.org/stable/2958889

Seitova, D., & Kovacs, M. (2025). Cosmopolitanism, trust, and democratic support shaping immigrant attitudes: A multilevel analysis across 36 European countries. *International Journal of Intercultural Relations*, *106*, 102178. https://doi.org/10.1016/j.ijintrel.2025.102178

Siddharthan, N. S. (2004). Globalisation: Productivity, Efficiency and Growth: An Overview. *Economic and Political Weekly*, *39*(5), 420–422. http://www.jstor.org/stable/4414574

Singer, M. (1981). Turkey in Crisis. *Current History*, *80*(462), 27–40. http://www.jstor.org/stable/45315937

Srinivasan, T. N., & Bhagwati, J. (2001). Outward-Orientation and Development: Are Revisionists Right? In *Trade, Development and Political Economy* (pp. 3–26). Palgrave Macmillan UK. https://doi.org/10.1057/9780230523685_1

Tunali, C. B., & Boru, F. (2019). The Causality Effects of Manufacturing Sector on Some Macroeconomic Variables in Turkey. *Procedia Computer Science*, *158*, 1109–1113. https://doi.org/10.1016/j.procs.2019.09.233

Ucer, M., & Acemoglu, D. (2015). Why Turkish growth ended: an institutional perspective. *VoxEU*. https://cepr.org/voxeu/columns/why-turkish-growth-ended-institutional-perspective

Ulas, E., & Keskin, B. (2017). *Is There a Relation Between HDI and Economic Performances?* (pp. 61–70). https://doi.org/10.1007/978-3-319-49559-0_6

UNDP. (2018). Human Development Indices and Indicators: Statistical Update 2018. *UNDP (United Nations Development Programme)*. http://report2017.archive.s3-website-us-east-1.amazonaws.com

Utkulu, U., & Özdemir, D. (2004). Does Trade Liberalization Cause a Long Run Economic Growth in Turkey. *Economic Change and Restructuring*, *37*(3–4), 245–266. https://doi.org/10.1007/s10644-005-8080-8

Villaverde, J., & Maza, A. (2011). Globalisation, Growth and Convergence. *The World Economy*, *34*(6), 952–971. https://doi.org/10.1111/j.1467-9701.2011.01359.x

World Bank. (2017). *World Development Indicators 2017*. Washington, DC. https://doi.org/10.1596/26447

Yay, G., Taştan, H., & Oktayer, A. (2016). Globalisation, economic freedom, and wage inequality: A panel data analysis. *Panoeconomicus*, *63*(5), 581–601. https://doi.org/10.2298/PAN130515024Y

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Asymmetric return dynamics and stock price crash risk: Evidence from a quantile regression analysis of an emerging market

Unbreen Arif[1*], Bilal Ahmad[2], Fizza Rizvi[3], Sarah Azhar[4]

[1,4]UE Business School, Division of Management & Administrative Sciences, University of Education, Lower Mall Campus, Lahore, Pakistan
[2,3]Hailey College of Banking & Finance, University of the Punjab, Lahore, Pakistan
*Corresponding Author: Unbreen.arif@ue.edu.pk

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Authors' emails:
bilalahmad@puhcbf.edu.pk
fizza-rizvi@puhcbf.edu.pk
sarah.azhar@ue.edu.pk

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Abstract

Purpose — Understanding extreme downside risk is particularly important in emerging equity markets, where higher market volatility, lower liquidity, and weaker information environments make stock prices more vulnerable to sudden, severe crashes. This study examines downside risk, stock price crash risk, and lower-tail return dynamics using firm-level stock return data for firms listed in the Pakistan Stock Exchange over the period 2014–2024.

Method — Using panel regression and quantile regression techniques, the study investigates the determinants of crash risk and assesses the predictive role of downside risk for future equity returns.

Findings — The results indicate that downside risk is strongly associated with a higher likelihood of extreme negative return realisations, while its effect on average returns remains limited. Quantile-based estimates further show that the impact of downside risk intensifies substantially in the lower tail of the return distribution, highlighting pronounced return asymmetries. These patterns persist across both financial and non-financial firms, although their magnitude varies with market conditions.

Implications — The results carry important implications for investors, regulators, and risk managers concerned with downside protection and the identification of early warning signals in emerging equity markets.

Originality — This study provides new firm-level evidence from an emerging equity market by jointly examining downside risk, crash risk, and return tail behaviour within a unified empirical framework by using quantile regression.

Keywords — Downside Risk, Crash Risk, Return Asymmetry, Quantile Regression, Emerging Markets

Introduction

The stability of emerging equity markets has become a major focus of global financial research, especially as these markets face significant economic and political changes worldwide (Arkol & Azimli, 2024; Bekaert et al., 2023). In asset pricing and risk management, the conventional dependence on mean-variance frameworks has often been inadequate when confronted with *black swan* events, unexpected, non-linear declines in asset valuations that contradict standard distribution assumptions. This event, known as stock price crash risk, is one of the biggest threats to the wealth of investors and the integrity of the market (Ring, 2023). The theoretical motivation for investigating crash risk is deeply rooted in the bad news hoarding hypothesis (Askarzadeh et

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al., 2024). However, there is a physical limit to information suppression. In markets with a lot of information asymmetry and weak institutional oversight, managers often have different reasons to keep bad news from the public. When the accumulated weight of undisclosed negative news reaches a tipping point, it enters the market simultaneously, triggering a catastrophic, rapid price correction (Cheng et al., 2025). Researchers, investors, and regulators are paying increasing attention to these extreme events. Particularly, after a series of financial crises that revealed the limitations of traditional risk measures based on variance and mean returns (Ang et al., 2006; Harvey et al., 2010). Within this broader context, stock price crash risk has emerged as a critical concern in asset pricing and financial stability research (Dierkes et al., 2021; Kelly & Jiang, 2013; Stoja et al., 2025). Empirical studies document that crash risk is closely linked to asymmetric return distributions and negative skewness, implying that investors face disproportionate downside exposure even when average returns appear stable (Boehme et al., 2020; Goetzmann et al., 2022; Habib et al., 2018; Kelly & Jiang, 2013). As a result, understanding the determinants and consequences of crash risk has become central to modern equity market analysis.

While the literature on downside and crash risk has expanded rapidly, much of the empirical evidence is concentrated in developed markets, where information disclosure standards, liquidity conditions, and investor protection mechanisms are relatively strong (Camilleri et al., 2020; Feldman & Kumar, 1995; Station Stereńczak, 2024). These structural features make emerging markets particularly vulnerable to abrupt price collapses, thereby increasing the relevance of downside and tail risks for both investors and regulators (Bae et al., 2006; Fernandes et al., 2024; Zhou et al., 2023). Despite these stylised facts, systematic evidence on the interaction between downside risk, crash risk and tail behaviour in emerging equity markets remains relatively limited. The current literature's predominant focus is on average effects. Traditional linear regression frameworks implicitly assume that risk factors exert uniform influences across the return distribution. However, growing evidence indicates that the pricing and transmission of risk are highly asymmetric, with substantially stronger effects in the lower tail of the distribution (Adrian et al., 2023; Sneller, 2025). Quantile-based approaches provide a more suitable framework for capturing these nonlinear dynamics, particularly when the primary interest lies in extreme downside outcomes rather than central tendencies. Quantile regression provides a granular view of the tail that standard OLS-based panel models typically overlook (Hoque et al., 2024).

A significant gap exists in the current literature regarding the predictive role of downside risk in the context of the emerging economy. Although volatility and systematic risk have been thoroughly examined, the factors influencing firm-level crash risk and the extent to which investors receive sufficient compensation for assuming this risk remain underexplored. The current study also investigates downside risk and crash risk by differentiating financial and non-financial firms. Financial institutions, which have strict capital requirements and are closely monitored by regulators, may have different crash risk profiles from non-financial firms, which are more sensitive to operational leverage and supply chain shocks (Bagh et al., 2025). This study contributes to the literature in several ways. First, this study contributes to the global discourse on the downside risk premium. In efficient markets, investors should theoretically demand higher returns for stocks that exhibit significant downside beta. The current study extends the downside risk and crash risk literature by providing new evidence from an emerging market setting. Second, by jointly examining financial and non-financial firms, the insights offer a more comprehensive view of crash risk across sectors within the same market. Third, the use of quantile regression allows us to move beyond mean-based inference and directly assess how downside risk influences the lower tail of the return distribution. The findings of the study highlight the importance of tail-oriented risk measures for understanding equity market behaviour in emerging economies.

Early asset pricing theories implicitly assumed that investors are primarily concerned with the variance of returns, treating positive and negative deviations symmetrically. However, subsequent research challenged this assumption by demonstrating that investors exhibit asymmetric preferences, placing greater weight on downside outcomes (Arnott & McQuarrie, 2025; Gao et al., 2025). This recognition led to the development of downside risk measures that focus on returns falling below a specified threshold or target (Markowski, 2024). Empirical evidence further shows that negative return realisations carry disproportionate economic and psychological costs, making downside risk a more relevant measure of risk than total volatility (Stoja et al., 2025; Zourai, 2022). Risk type significantly influences financial performance (Suseno & Bamahriz, 2017). This phenomenon is particularly pronounced in emerging markets, where return distributions often exhibit stronger negative skewness and kurtosis than in developed markets, suggesting a higher propensity for extreme negative returns (Barunik & Nevrla, 2018; Calomiris & Mamaysky, 2019). This asymmetry of returns, where large upward movements are not as equally matched by large drawdowns, highlights the importance of considering downside co-moments in asset pricing models, especially during periods of market stress (Markowski, 2024; Xu, 2018). These features imply that average returns fail to fully reflect investors' exposure to extreme losses. While this literature establishes the importance of downside risk, most empirical analyses focus on developed markets. In emerging markets, where liquidity constraints and delayed information dissemination are more prevalent, downside risk is likely to play a more pronounced role in shaping return behaviour. Moreover, existing studies often rely on mean-based estimation techniques, which may obscure the impact of downside risk on extreme return realisations. Therefore, the first hypothesis of the study is

H_1: Firm-level downside risk is positively associated with the likelihood of extreme negative stock returns.

The concept of stock price crash risk is closely tied to theories of information hoarding and asymmetric disclosure. Jin and Myers (2006) argue that managers have incentives to withhold adverse information, leading to the accumulation of undisclosed bad news that is eventually released in a concentrated manner, resulting in stock price crashes. This framework has been widely adopted and empirically validated in subsequent studies, which show that firms with higher opacity and weaker monitoring are more prone to crash events (Andreou et al., 2023; Liu et al., 2024). Salles (2021) reported firm-level risk is not diversifiable in emerging markets. Crash risk has been operationalised using distribution-based measures such as negative conditional skewness and down-to-up volatility, both of which capture the asymmetry of firm-specific returns (Benkraiem et al., 2022; Fiordelisi et al., 2023). Empirical findings suggest that crash risk is not merely a reflection of volatility but represents a distinct risk dimension linked to downside tail behaviour. However, most of this evidence is drawn from developed markets with relatively strong disclosure regimes and institutional frameworks.

Emerging markets differ markedly in this respect. Weaker investor protection, heterogeneous disclosure practices, and episodic liquidity shortages may intensify the buildup and release of negative information, thereby amplifying crash risk. Yet, comparative evidence on crash risk dynamics across financial and non-financial firms within emerging markets remains scarce, which leads to the formation of the hypothesis of the study:

H_2: Higher downside risk is associated with greater stock price crash risk at the firm level.

A central question in asset pricing is whether investors are compensated for bearing downside risk. While traditional models focus on the pricing of systematic volatility, recent studies suggest that downside risk may contain predictive information for future returns, particularly during adverse market conditions (Ang et al., 2006). These findings imply that downside risk may influence expected returns through channels that are not captured by standard risk factors.

However, empirical evidence on downside risk predictability remains mixed. Globalisation alters the risk structure, potentially influencing tail risk and crash probability in financial markets (Said et al., 2027). Some studies find that downside risk commands a risk premium, while others report weak or insignificant effects on average returns (Barunik & Nevrla, 2018; Dai & Harris, 2023). One explanation for these inconsistencies is the reliance on mean-based regression frameworks, which may hide heterogeneous effects across the return distribution.

In emerging markets, downside events are more frequent and severe. So, the predictive role of downside risk may be concentrated in the lower tail rather than the conditional mean. Yet, this possibility has received limited empirical attention, particularly in studies using firm-level data, and this leads to the formulation of the next hypothesis:

H_3: Downside risk has limited explanatory power for average stock returns but significantly predicts lower-tail return outcomes.

Recent advances in financial econometrics emphasise the importance of examining risk-return relationships across the entire return distribution. Quantile regression techniques, introduced by Koenker and Bassett (1978), allow for the capture of heterogeneous effects that vary across different market states. Applications in finance demonstrate that risk factors often exert stronger influences during periods of market stress, particularly in the lower tail of the return distribution (Arnott & McQuarrie, 2025). Tail-focused measures such as value at risk and expected shortfall provide additional insights into extreme downside exposure that cannot be inferred from variance-based metrics alone. While these measures are widely used in risk management, their integration into firm-level return analysis in emerging markets remains relatively underdeveloped. Given the structural characteristics of emerging equity markets, it is reasonable to expect that downside risk exerts a disproportionately stronger effect on extreme negative returns than on central or upper quantiles.

H_4: The impact of downside risk on stock returns is significantly stronger in the lower tail of the return distribution than at the mean.

Financial firms differ from non-financial firms in terms of leverage, regulatory oversight, and sensitivity to market-wide shocks, which may influence their exposure to downside and crash risk. Prior studies suggest that financial firms are particularly vulnerable to tail events due to balance-sheet interconnectedness and maturity mismatches (Aufiero et al., 2025; Ellul et al., 2022; Stoja et al., 2025). However, empirical evidence comparing crash risk dynamics across financial and non-financial firms within the same emerging market remains limited. Understanding whether downside and crash risk behave differently across sectors is crucial for both portfolio allocation and regulatory supervision, especially in markets where financial institutions play a central role in economic stability.

H_5: The magnitude of downside and crash risk effects differs between financial and non-financial firms.

Methods

The study employs a well-balanced panel constructed from monthly stock price data of firms listed on the KSE-100 Index over the period 2014–2024. o. The initial universe includes all companies that were constituents of the index during the sample period. To ensure consistency and reliability of return calculations, firms with insufficient trading history, prolonged trading suspensions, or missing price observations were excluded. After applying these filters, the final sample consists of 85 non-financial firms with continuous return data. Stock prices are used to compute firm-level excess returns, downside risk measures, and crash indicators. Each firm contributes 132 monthly observations, resulting in a balanced panel suitable for panel and quantile regression analysis. The sample covers a broad cross-section of industries, with financial firms (banking and insurance) representing approximately 20% of firm-month observations while non-financial firms comprise the remaining 80%, providing sufficient cross-sectional variation to examine heterogeneity in downside-risk effects across firm types. Stock returns are computed using monthly closing prices, and excess returns are calculated by subtracting the contemporaneous risk-free rate. Market excess returns are measured using the value-weighted market index minus the risk-free rate. This construction follows standard asset-pricing practice (Ang et al., 2006; Fama & French, 2015). Downside risk is measured as the firm-level standard deviation of negative excess returns, capturing exposure to adverse return realisations while excluding upside volatility. Formally, downside risk is computed as:

$$DR_i = \sqrt{\frac{1}{N_i-1}\sum_{t:R_{it}<0}(R_{it} - \bar{R}_{it})^2} \tag{1}$$

where R_{it} denotes the firm i's excess return in month t, and the summation is taken only over periods in which returns are negative. To capture extreme negative return events, a crash indicator is constructed following the tail-event literature (Chen et al., 2001; Kelly & Jiang, 2013). A firm-month observation is classified as a crash if the excess return falls into the bottom 5% of the empirical return distribution. This binary indicator enables the analysis to distinguish between normal negative returns and extreme downside realisations, both of which are central to crash-risk studies. To examine the relationship between downside risk and crash probability, panel logit models are estimated. First, a fixed-effects logit model is employed to control for unobserved firm-specific heterogeneity. However, due to the time-invariant construction of downside risk, this variable is omitted in the fixed-effects specification, reflecting a well-known identification limitation of nonlinear fixed-effects estimators (Wooldridge, 2002). Accordingly, the main crash-risk tests rely on a random-effects logit model, which allows the inclusion of firm-specific characteristics such as downside risk:

$$P(Crash_{it} = 1) = \Lambda(\alpha + \beta_1 DR_i + \beta_2 MKT_t + u_i) \tag{2}$$

This approach is consistent with prior crash-risk studies that incorporate time-invariant firm attributes (Hutton et al., 2009; Kim et al., 2011). To assess whether downside risk is priced in average returns, a fixed-effects panel regression is estimated. The model is as follows:

$$R_{it} = \alpha_i + \beta_1 DR_i + \beta_2 MKT_t + \varepsilon_{it} \tag{3}$$

Given the firm-specific nature of downside risk, the variable is collinear with firm fixed effects and therefore omitted from the estimation. This outcome is informative, indicating that downside risk does not explain within-firm variation in mean returns. This interpretation is consistent with findings in Fama and French (2015), who argue that not all forms of risk command an average return premium.

To capture distributional asymmetries and tail dependence, represented by H_1, H_3, and H_4, a quantile regression is employed in accordance with (Koenker & Bassett, 1978). The baseline quantile regression specification is as follows:

$$Q_\tau(R_{it} \mid X_{it}) = \alpha_\tau + \beta_\tau^{DR} DR_i + \beta_\tau^{MKT} MKT_t \tag{4}$$

Models are estimated at the 10th, 25th, and 50th percentiles. This approach is widely used in tail-risk and asset-pricing research to uncover asymmetric risk–return relationships (Bollerslev et al., 2009; Giglio et al., 2016).

To test whether downside-risk effects differ across firm types (H_5), firms are classified as financial or non-financial based on industry affiliation. Financial firms include banking and insurance companies, consistent with prior studies on systemic and tail risk (Brunnermeier & Oehmke, 2013). Heterogeneity is examined using an interaction-based quantile regression:

$$Q_{0.10}(R_{it}) = \alpha + \beta_1 DR_i + \beta_2 Financial_i + \beta_3 (DR_i \times Financial_i) + \beta_4 MKT_t \tag{5}$$

This specification allows the magnitude of downside-risk effects to differ between financial and non-financial firms while maintaining a unified estimation framework. Standard errors are clustered at the firm level to account for within-firm dependence over time. The consistency of results across fixed-effects, random-effects, and quantile-based estimators mitigates concerns regarding model misspecification or outlier-driven inference.

Results and Discussions

Downside risk is computed as the firm-level standard deviation of negative excess returns. The panel is strongly balanced with 132 monthly observations per firm. The empirical analysis is conducted on a well-balanced panel of firm-level monthly data comprising 11,352 observations from January 2014 to December 2024. The sample spans twelve industries, with financial firms accounting for approximately 20% of firm-month observations. This structure provides sufficient variation to examine downside risk, crash risk, and return distribution asymmetries. Table 1 reports descriptive statistics for the main variables used in the analysis.

Firms exhibit substantial heterogeneity in downside risk exposure. The upper bound (0.255) reflects firms with severe exposure to adverse return realisations. The findings confirm that downside risk is well-defined and economically meaningful.

Table 1. Descriptive Statistics

Variable	Obs.	Mean	Std. Dev.	Min	Max
Excess stock return	11,352	-.0026	.1158	-2.03	1.013
Market excess return	11,352	.0021	.0586	-.269	.147
Downside risk	11,352	0.068	0.027	0.00	0.255
Crash indicator	11,352	-	-	0	1

Table 2 reports quantile regression estimates at the 10th percentile of the return distribution. The coefficient on downside risk is negative, large in magnitude, and highly statistically significant, indicating that firms with greater downside exposure experience substantially more severe losses during extreme negative return realisations. Standard errors are heteroskedasticity-robust. This result provides strong support for *hypothesis 1*, confirming that downside risk is a key determinant of extreme adverse equity outcomes.

Table 2. Quantile Regression Estimates at the 10th Percentile of the Excess Return Distribution

Variable	Coefficient	Std. Error	t-stat	*p*-value
Downside risk	−0.786	0.064	−12.24	0.000
Market excess return	0.864	0.028	30.65	0.000
Constant	−0.059	0.005	−12.43	0.000

Table 3 presents panel logit estimates examining the relationship between downside risk and crash probability. In fixed-effects specifications, downside risk is omitted because it is time-invariant. Random effects are specified at the firm level, allowing the inclusion of firm-specific characteristics that are time-invariant over the sample period. The coefficient on downside risk is positive, indicating that firms with greater exposure to adverse return fluctuations are more likely of experience crashes. However, the estimate is not statistically significant at conventional levels, indicating that the influence of downside risk is not uniform across states of the return distribution. Consistent with the fixed-effects logit estimates, the coefficient on market excess returns is large, negative, and highly statistically significant ($z = -17.68$, $p < 0.01$). This finding implies that deteriorating market conditions substantially increase the likelihood of firm-level crash events. The result underscores the dominant role of systematic market stress in triggering extreme downside outcomes, even after controlling for firm-specific risk characteristics. The estimated variance of the random effects ($\sigma_u = 0.587$) is statistically significant, and the intra-class correlation coefficient ($\rho \approx 0.095$) indicates that approximately 9.5% of the unexplained variation in crash risk is attributable to persistent firm-level heterogeneity. This confirms the relevance of accounting for unobserved firm-specific effects in crash-risk modelling. The market excess returns exert a strong and highly significant effect, highlighting the central role of systematic market downturns in driving extreme negative return realisations. The significance of the random-effects variance further confirms the presence of persistent firm-level heterogeneity in crash risk.

Table 3. Downside Risk and Crash Probability

logit estimates	Variable	Coefficient	Std. Error	z-stat	p-value
Panel A: Fixed-Effects Logit	Downside risk	Omitted			
Market excess return	Market excess return	−17.740***			
Panel B: Random-Effects Logit	Downside risk	12.226	8.174	1.50	0.135
Crash indicator	Market excess return	−17.683	1.406	−12.58	0.000

Notes: entries in *** are significant at 1% significance level.

Table 4. Fixed-Effects Regressions of Average Excess Returns

Variable	Coefficient	Std. Error	*t*-stat	*p*-value
Downside risk	Omitted			
Market excess return	0.947	0.048	19.73	0.000

Table 4 reports fixed-effects regression results for average excess returns. The findings indicate that market excess returns remain the dominant determinant of average return dynamics. The coefficient on downside risk is omitted due to collinearity, reflecting that downside risk is treated as a firm-specific, time-invariant characteristic. This result implies that downside risk does not help explain within-firm variation in average returns over time, consistent with the hypothesis that downside risk is not systematically priced in the conditional mean of returns. In contrast, the coefficient on market excess returns is positive, large, and statistically significant (β = 0.947, t = 19.73, p < 0.05). The within R^2 of 0.23 indicates that approximately 23% of the within-firm variation in excess returns is explained by market movements. The result directly supports that Downside risk is not priced in average returns, as it does not explain within-firm return variation once firm fixed effects are accounted for. Further, systematic market risk dominates the conditional mean; the economic relevance of downside risk therefore lies outside the mean.

Table 5. Quantile Regression across the Return Distribution

Variable	$\tau = 0.10$	$\tau = 0.25$	$\tau = 0.50$
Downside risk	−0.786***	−0.142***	−0.001
Market excess return	0.864***	0.876***	0.896***
Pseudo R^2	0.140	0.131	0.133
Observations	11,103	11,103	11,103

Entries in *** are significant at 1% significance level.

Table 5 shows the test results of H_3 and H_4. It indicates that downside risk significantly predicts returns in the lower tail of the distribution, with strong effects at the 10th percentile and weaker but still significant effects at the 25th percentile. These results strongly support hypothesis 3, which posits that downside risk is not priced in terms of average returns but is highly relevant during adverse return states. Return asymmetry and distributional effects (H_4) test results of quantile regression estimates reveal a clear monotonic decline in the magnitude of downside-risk coefficients from the lower tail toward the centre of the return distribution. This attenuation indicates that downside risk primarily amplifies losses during extreme adverse states. The coefficient on downside risk at the median is economically negligible and statistically insignificant (β = −0.0014, t = −0.04, p = 0.967). This result indicates that downside risk has no discernible effect on median returns, implying that firms with higher downside exposure do not systematically earn higher or lower typical returns. In contrast, market excess returns remain positive and highly statistically significant (β = 0.896, t = 61.90, p < 0.01), confirming that systematic market movements dominate return dynamics around the median, consistent with standard asset pricing models. These findings indicate that downside risk is not priced in typical return outcomes, but instead materialises primarily during adverse states, highlighting pronounced return asymmetries.

Table 6. Quantile Regression Estimates at $\tau = 0.10$ with an Interaction Term

Variable	Coefficient	Std. Error	t-stat	p-value
Downside risk (non-financial firms)	−0.518	0.063	−8.20	0.000
Financial firm dummy	0.090	0.015	6.00	0.000
Downside risk × Financial	−1.110	0.228	−4.87	0.000
Market excess return	0.856	0.026	32.33	0.000

Table 6 reports quantile regression estimates incorporating an interaction between downside risk and a financial-firm indicator. The interaction term is negative and highly statistically significant, indicating that the adverse impact of downside risk on lower-tail returns is significantly stronger for financial firms than for non-financial firms. This finding provides strong support for Hypothesis 5 and suggests that financial institutions are particularly vulnerable to downside shocks during extreme market downturns. Consistent with prior specifications, market excess returns remain positive and highly statistically significant (β = 0.856, t = 32.33, p < 0.01), confirming the dominant role of systematic market conditions in shaping firm-level returns even in the lower tail.

The pseudo-R^2 of 0.146 suggests that the model explains a meaningful portion of the variation in lower-tail returns, which is notable given the inherent volatility of extreme returns. The statistically significant negative interaction term provides strong evidence of heterogeneity, indicating that downside risk has a significantly larger adverse impact on lower-tail returns for financial firms than for non-financial firms.

Table 7. Robustness Check

Variables	Baseline Q10	Financial Control	Bootstrap Q10
Downside risk	-0.779***	-0.694***	-0.779***
Market return	0.886***	0.862***	0.886***

Entries in *** are significant at 1% significance level.

Table 7 presents the robustness checks of the empirical findings. The quantile regression model was re-estimated by including an industry control variable that distinguishes financial firms from non-financial firms. The results indicate that downside risk continues to exhibit a strong negative and statistically significant impact on stock returns in the lower quantile of the return distribution and confirm that the main results are robust to industry-specific effects. Additionally, bootstrap quantile regression was estimated using 200 replications. The results remain consistent with the baseline estimation as downside risk continues to exhibit a strong negative and statistically significant effect on stock returns in the lower quantile of the distribution (-0.779, $p < 0.01$). These findings confirm that the main results are robust to alternative estimation techniques and sampling variation.

Summary of the Hypotheses Testing:

Hypothesis	Description	Result
H_1	Downside risk → extreme negative returns	Supported
H_2	Downside risk → crash probability	Partial support
H_3	Not priced in mean, priced in tail	Supported
H_4	Stronger tail effect than the mean	Strongly supported
H_5	Financial vs. non-financial heterogeneity	Supported

The empirical results provide strong evidence that downside risk is closely linked to extreme negative return realisations, rather than to average return dynamics. These findings are especially relevant for emerging equity markets, where structural vulnerabilities and information frictions magnify tail risk (Arkol & Azimli, 2024; Bekaert et al., 2023). These results align closely with foundational and recent evidence emphasising downside and tail risk in asset pricing. In particular, Ang et al. (2006) and Ang et al. (2009) demonstrate that downside risk and volatility asymmetries become economically meaningful during market downturns. More recent studies also document that tail risk is a dominant feature of equity markets facing heightened macroeconomic and political uncertainty (Ammann et al., 2023; Stoja et al., 2025). By providing firm-level evidence from the PSX, this study extends the downside-risk literature to an emerging market context where non-normal return behaviour and downside asymmetry are especially pronounced (Barunik & Nevrla, 2018; Zhou et al., 2023). In contrast, the weak and statistically imprecise relationship between downside risk and average crash probability is consistent with the crash-risk literature emphasising the role of systematic market stress rather than firm-specific characteristics. Kelly and Jiang (2013) show that crash risk is fundamentally linked to aggregate tail risk, while recent evidence highlights the dominant role of macro-financial shocks in triggering market-wide crashes (Ring, 2023; Stoja et al., 2025). The absence of downside risk effects in fixed effects mean regressions further suggests that downside risk is not rewarded in average returns. These findings reinforce the limitations of mean-variance frameworks and are consistent with the argument that traditional linear models fail to capture the pricing of rare but severe downside events (Bekaert et al., 2023; Ring, 2023). While some studies document downside-related premia under specific conditions (Lee & Yang, 2022). The present evidence suggests that such compensation does not manifest in average returns once firm-specific heterogeneity is controlled. The decline in the magnitude of downside-risk coefficients from the 10th percentile to the median provides compelling evidence of return asymmetry, consistent with behavioural and long-run risk theories. Investors' heightened sensitivity to losses during adverse states, as formalised by Epstein and Zin (1990), implies that downside risk should exert disproportionate effects in the lower tail of the return distribution. Recent empirical work further supports this view by demonstrating that tail risks cannot be adequately captured by mean-based estimation techniques (Sneller, 2025). The quantile-based approaches provide a more suitable framework for analysing downside and crash risk, particularly in emerging markets where extreme outcomes are more frequent (Hoque et al., 2024). The results are also consistent with Giglio and Xiu (2021) that tail risks are not captured by mean-variance frameworks and require distribution-sensitive methods.

A key contribution of this study lies in documenting heterogeneity across firm types. The interaction-based quantile regressions reveal that downside risk has a significantly stronger adverse impact on lower-tail returns for financial firms compared to non-financial firms, consistent with prior research emphasising the heightened vulnerability of financial institutions due to leverage, balance-sheet interconnectedness, and maturity mismatch (Aufiero et al., 2025; Ellul et al., 2022). In emerging markets, where regulatory buffers and crisis-resolution mechanisms are comparatively weaker, these vulnerabilities are likely to amplify the transmission of downside shocks (Bagh et al., 2025). At the same time, our results contrast with those of Bali et al. (2017), that financial firms may benefit from implicit guarantees that dampen downside exposure in developed markets. The divergence likely reflects institutional differences between advanced and emerging economies, where regulatory buffers and safety nets are weaker, making financial firms more vulnerable to downside shocks.

Conclusion

This study examines the role of firm-level downside risk in shaping crash risk and equity return dynamics using a strongly balanced monthly panel from 2014 to 2024. The results provide clear evidence that downside risk is not priced in average returns but plays a critical role in extreme negative return outcomes, particularly in the lower tail of the return distribution. Quantile regression estimates reveal pronounced return asymmetries, with downside risk exerting a strong and economically meaningful impact during adverse states while remaining irrelevant under normal conditions.

Moreover, the analysis uncovers significant heterogeneity between financial and non-financial firms and highlights that financial institutions are particularly exposed to downside risk in extreme market downturns. These findings underscore the importance of moving beyond mean-based models and adopting distribution-sensitive approaches when assessing risk and return relationships.

The findings have several important policy implications. First, regulators and market participants should recognise that downside risk materialises primarily during extreme market conditions, implying that stress testing frameworks should explicitly incorporate downside risk measures rather than relying solely on average volatility metrics. Second, the heightened sensitivity of financial firms to downside risk highlights the need for stronger capital buffers and liquidity requirements, particularly in emerging markets where systemic vulnerabilities are more pronounced. For investors, the results suggest that traditional diversification strategies based on mean returns may be insufficient to mitigate tail risk. Portfolio construction frameworks should therefore incorporate downside risk metrics to better manage exposure to extreme losses.

Despite its contributions, this study has several limitations. First, downside risk is measured using historical return realisations, which may not fully capture forward-looking expectations. Future research could incorporate option-implied downside measures. Second, the analysis focuses on a single emerging market, limiting generalizability. Cross-country studies could examine whether institutional quality moderates the effect of downside risk. Third, while quantile regression captures distributional heterogeneity, it remains static. Future work could employ dynamic tail risk models to explore the temporal evolution of downside risk.

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Author contribution
All authors contributed to the conception and design of the study, data collection, analysis and interpretation of the results, and the writing of the manuscript. All authors read and approved the final version of the manuscript.

Use of AI tools declaration
The authors confirm adherence to all ethical research and publishing standards. This work is original, and all sources have been duly acknowledged. The authors also disclose that ChatGPT was utilized solely for language editing, grammar improvement, and structural refinement of the manuscript.

Conflict of interest
The authors declare that there are no competing interests related to this manuscript.

References

Adrian, T., DeHaven, M., Duarte, F., & Iyer, T. (2023). The Market Price of Risk and Macro-Financial Dynamics. *IMF Working Papers*, *2023*(199), 1. https://doi.org/10.5089/9798400255397.001

Ammann, M., Moerke, M., Prokopczuk, M., & Würsig, C. M. (2023). Commodity tail risks. *Journal of Futures Markets*, *43*(2), 168–197. https://doi.org/10.1002/fut.22381

Andreou, P. C., Lambertides, N., & Magidou, M. (2023). A Critique of the Agency Theory Viewpoint of Stock Price Crash Risk: The Opacity and Overinvestment Channels. *British Journal of Management*, *34*(4), 2158–2185. https://doi.org/10.1111/1467-8551.12693

Ang, A., Chen, J., & Xing, Y. (2006). Downside Risk. *Review of Financial Studies*, *19*(4), 1191–1239. https://doi.org/10.1093/rfs/hhj035

Ang, A., Hodrick, R. J., Xing, Y., & Zhang, X. (2009). High idiosyncratic volatility and low returns: International and further U.S. evidence. *Journal of Financial Economics*, *91*(1), 1–23. https://doi.org/10.1016/j.jfineco.2007.12.005

Arkol, O., & Azimli, A. (2024). Pricing the common stocks in emerging markets: The role of economic policy uncertainty. *Modern Finance*, *2*(1), 31–50. https://doi.org/10.61351/mf.v2i1.93

Arnott, R. D., & McQuarrie, E. F. (2025). *Fear, Not Risk, Explains Asset Pricing.* https://doi.org/10.2139/ssrn.5127501

Askarzadeh, A., Kanaanitorshizi, M., Tabarhosseini, M., & Amiri, D. (2024). International Diversification and Stock-Price Crash Risk. *International Journal of Financial Studies*, *12*(2), 47. https://doi.org/10.3390/ijfs12020047

Aufiero, S., Bartolucci, S., Caccioli, F., & Vivo, P. (2025). *Mapping Microscopic and Systemic Risks in TradFi and DeFi: a literature review.* http://arxiv.org/abs/2508.12007

Bae, K., Lim, C., & Wei, K. C. J. (2006). Corporate Governance and Conditional Skewness in the World's Stock Markets*. *The Journal of Business*, *79*(6), 2999–3028. https://doi.org/10.1086/508006

Bagh, T., Hunjra, A. I., Ntim, C. G., & Naseer, M. M. (2025). Capitalizing on risk: How corporate financial flexibility, investment efficiency, and institutional ownership shape risk-taking dynamics. *International Review of Economics & Finance*, *99*, 104068. https://doi.org/10.1016/j.iref.2025.104068

Bali, T., Engle, R., & Murray, S. (2017). Empirical Asset Pricing: The Cross Section of Stock Returns: An Overview. In *Wiley StatsRef: Statistics Reference Online* (pp. 1–8). Wiley. https://doi.org/10.1002/9781118445112.stat07954

Barunik, J., & Nevrla, M. (2018). Tail Risks, Asset Prices, and Investment Horizons. *SSRN Electronic Journal*. https://doi.org/10.2139/ssrn.3197408

Bekaert, G., Harvey, C. R., & Mondino, T. (2023). Emerging equity markets in a globalized world. *Emerging Markets Review*, *56*, 101034. https://doi.org/10.1016/j.ememar.2023.101034

Benkraiem, R., Goutte, S., Saadi, S., Zhu, H., & Zhu, S. (2022). Investor heterogeneity and negative skewness in stock returns: Evidence from institutional investors. *Journal of International Financial Markets, Institutions and Money*, *81*, 101690. https://doi.org/10.1016/j.intfin.2022.101690

Boehme, R. D., Fotak, V., & May, A. (2020). Seasoned Equity Offerings and Stock Price Crash Risk. *International Journal of Finance & Banking Studies (2147-4486)*, *9*(4), 131–146. https://doi.org/10.20525/ijfbs.v9i4.961

Bollerslev, T., Tauchen, G., & Zhou, H. (2009). Expected Stock Returns and Variance Risk Premia. *Review of Financial Studies*, *22*(11), 4463–4492. https://doi.org/10.1093/rfs/hhp008

Brunnermeier, M. K., & Oehmke, M. (2013). The Maturity Rat Race. *The Journal of Finance*, *68*(2), 483–521. https://doi.org/10.1111/jofi.12005

Calomiris, C. W., & Mamaysky, H. (2019). How news and its context drive risk and returns around the world. *Journal of Financial Economics*, *133*(2), 299–336. https://doi.org/10.1016/j.jfineco.2018.11.009

Camilleri, S. J., Vassallo, S., & Bai, Y. (2020). Predictability in securities price formation: differences between developed and emerging markets. *Journal of Capital Markets Studies*, *4*(2), 145–166. https://doi.org/10.1108/JCMS-07-2020-0025

Chen, J., Hong, H., & Stein, J. C. (2001). Forecasting crashes: trading volume, past returns, and conditional skewness in stock prices. *Journal of Financial Economics*, *61*(3), 345–381. https://doi.org/10.1016/S0304-405X(01)00066-6

Cheng, X., Zheng, Y., & Tu, Y. (2025). Research on the dynamic interaction effects of litigation events, financing constraints, and the risk of corporate stock price crashes. *International Review of Economics & Finance*, *99*, 104008. https://doi.org/10.1016/j.iref.2025.104008

Dai, Y., & Harris, R. D. F. (2023). Average tail risk and aggregate stock returns. *Journal of International Financial Markets, Institutions and Money*, *82*, 101699. https://doi.org/10.1016/j.intfin.2022.101699

Dierkes, M., Hollstein, F., Prokopczuk, M., & Wuersig, C. (2021). Measuring Tail Risk. *SSRN Electronic Journal*. https://doi.org/10.2139/ssrn.3789005

Ellul, A., Jotikasthira, C., Kartasheva, A., Lundblad, C. T., & Wagner, W. (2022). Insurers as Asset Managers and Systemic Risk. *The Review of Financial Studies*, *35*(12), 5483–5534. https://doi.org/10.1093/rfs/hhac056

Epstein, L. G., & Zin, S. E. (1990). 'First-order' risk aversion and the equity premium puzzle. *Journal of Monetary Economics*, *26*(3), 387–407. https://doi.org/10.1016/0304-3932(90)90004-N

Fama, E. F., & French, K. R. (2015). A five-factor asset pricing model. *Journal of Financial Economics*, *116*(1), 1–22. https://doi.org/10.1016/j.jfineco.2014.10.010

Feldman, R. A., & Kumar, M. S. (1995). Emerging Equity Markets: Growth, Benefits, And Policy Concerns. *The World Bank Research Observer*, *10*(2), 181–200. https://doi.org/10.1093/wbro/10.2.181

Fernandes, L. H. S., Fernandes, J. P. V., Silva, J. W. L., Paiva, R. O. A., Pinto, I. M. B. S., & De Araujo, F. H. A. (2024). The (in) efficiency of USA Education Group stocks: before, during and after COVID-19. *Fractals*, *32*(03). https://doi.org/10.1142/S0218348X24500476

Fiordelisi, F., Ricci, O., & Santilli, G. (2023). Environmental engagement and stock price crash risk: Evidence from the European banking industry. *International Review of Financial Analysis*, *88*, 102689. https://doi.org/10.1016/j.irfa.2023.102689

Gao, Y., Hoepner, A. G. F., Prokopczuk, M., Rouxelin, F., & Wuersig, C. (2025). Responsible investing: Upside potential and downside protection? *International Review of Financial Analysis*, *97*, 103754. https://doi.org/10.1016/j.irfa.2024.103754

Giglio, S., Kelly, B., & Pruitt, S. (2016). Systemic risk and the macroeconomy: An empirical evaluation. *Journal of Financial Economics*, *119*(3), 457–471. https://doi.org/10.1016/j.jfineco.2016.01.010

Giglio, S., & Xiu, D. (2021). Asset Pricing with Omitted Factors. *Journal of Political Economy*, *129*(7), 1947–1990. https://doi.org/10.1086/714090

Goetzmann, W., Kim, D., & Shiller, R. (2022). *Crash Narratives*. https://doi.org/10.3386/w30195

Habib, A., Hasan, M. M., & Jiang, H. (2018). Stock price crash risk: review of the empirical literature. *Accounting & Finance*, *58*(S1), 211–251. https://doi.org/10.1111/acfi.12278

Harvey, C. R., Liechty, J. C., Liechty, M. W., & Müller, P. (2010). Portfolio selection with higher moments. *Quantitative Finance*, *10*(5), 469–485. https://doi.org/10.1080/14697681003756877

Hoque, M. E., Billah, M., Kapar, B., & Naeem, M. A. (2024). Quantifying the volatility spillover dynamics between financial stress and US financial sectors: Evidence from QVAR connectedness. *International Review of Financial Analysis*, *95*, 103434. https://doi.org/10.1016/j.irfa.2024.103434

Hutton, A. P., Marcus, A. J., & Tehranian, H. (2009). Opaque financial reports, R2, and crash risk☆. *Journal of Financial Economics*, *94*(1), 67–86. https://doi.org/10.1016/j.jfineco.2008.10.003

Jin, L., & Myers, S. (2006). R2 around the world: New theory and new tests☆. *Journal of Financial Economics*, *79*(2), 257–292. https://doi.org/10.1016/j.jfineco.2004.11.003

Kelly, B. T., & Jiang, H. (2013). Tail Risk and Asset Prices. *SSRN Electronic Journal*. https://doi.org/10.2139/ssrn.2321243

Kim, J.-B., Li, Y., & Zhang, L. (2011). Corporate tax avoidance and stock price crash risk: Firm-level analysis. *Journal of Financial Economics*, *100*(3), 639–662. https://doi.org/10.1016/j.jfineco.2010.07.007

Koenker, R., & Bassett, G. (1978). Regression Quantiles. *Econometrica*, *46*(1), 33. https://doi.org/10.2307/1913643

Lee, K.-H., & Yang, C.-W. (2022). The world price of tail risk. *Pacific-Basin Finance Journal*, *71*, 101696. https://doi.org/10.1016/j.pacfin.2021.101696

Liu, J., Ng, J., Tang, D. Y., & Zhong, R. (2024). Withholding Bad News in the Face of Credit Default Swap Trading: Evidence from Stock Price Crash Risk. *Journal of Financial and Quantitative Analysis*, *59*(2), 557–595. https://doi.org/10.1017/S002210902300008X

Markowski, L. (2024). Conventional and downside CAPM with higher-order moments: Evidence from emerging markets. *Equilibrium. Quarterly Journal of Economics and Economic Policy*, *19*(1), 93–138. https://doi.org/10.24136/eq.2043

Ring, M. A. K. (2023). Entrepreneurial Wealth and Employment: Tracing Out the Effects of a Stock Market Crash. *The Journal of Finance*, *78*(6), 3343–3386. https://doi.org/10.1111/jofi.13280

Said, F. F., Mansur, M., Karim, Z. A., Asri, N. M., & Antoni. (2027). Liberalizing the Malaysian Stock Market. *Jurnal Ekonomi Pembangunan*, *12*(2). https://journal.uii.ac.id/JEP/article/view/508/420

Salles, A. A. de. (2021). Similarity evidence between the country risk and the idiosyncratic risk: An empirical study of the Brazilian case. *Economic Journal of Emerging Markets*, *13*(1), 66–77. https://doi.org/10.20885/ejem.vol13.iss1.art6

Sneller, L. M. (2025). *Sentiment Feedback in Equity Markets: Asymmetries, Retail Heterogeneity, and Structural Calibration*. http://arxiv.org/abs/2509.11970

Stereńczak, S. (2024). Illiquidity and stock returns: the moderating role of investors' holding period in Central and Eastern European markets. *International Journal of Emerging Markets*, *19*(7), 2025–2045. https://doi.org/10.1108/IJOEM-01-2022-0125

Stoja, E., Polanski, A., & Nguyen, L. H. (2025). The taxonomy of tail risk. *Journal of Financial Research*, *48*(2), 701–724. https://doi.org/10.1111/jfir.12423

Suseno, P., & Bamahriz, O. (2017). Examining the impact of bank's risks to Islamic banks' profitability. *Economic Journal of Emerging Markets*, *9*(2), 125–137. https://doi.org/10.20885/ejem.vol9.iss2.art2

Wooldridge, J. M. (2002). *Econometric Analysis of Cross Section and Panel Data* (2nd ed.). The MIT Press.

Xu, X. (2018). *Four essays on capital markets and asset allocation* [Université de Lyon]. https://theses.hal.science/tel-01961876

Zhou, Y., Wu, S., Liu, Z., & Rognone, L. (2023). The asymmetric effects of climate risk on higher-moment connectedness among carbon, energy and metals markets. *Nature Communications*, *14*(1), 7157. https://doi.org/10.1038/s41467-023-42925-9

Zourai, H. (2022). On the Effectiveness of Stock Index Futures for Tail Risk Protection. *International Journal of Economics and Financial Issues*, *12*(3), 38–52. https://doi.org/10.32479/ijefi.13011

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The economic consequences of single motherhood on children's cognitive outcomes in Indonesia

Wisnu Setiadi Nugroho[1*], Evi Noor Afifah[2], Andika Ridha Ayu Perdana[3], Zahra Amalia Syarifah[4]

[1,2]Department of Economics, Universitas Gadjah Mada, Yogyakarta, Indonesia
[3]Department of Economics, Universitas Islam Indonesia, Yogyakarta, Indonesia
[4]Department of Sociology, University of California, San Diego, United States
[*]Corresponding author: wisnu.nugroho@ugm.ac.id

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Authors' emails:
evinoor.afifah@ugm.ac.id
andika.ayu@uii.ac.id
zsyarifa@ucsd.edu

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Abstract

Purpose — Single motherhood is widely associated with poorer child outcomes, yet it remains unclear whether these disadvantages stem from family structure itself or from the economic shocks that accompany it. This distinction is particularly important in developing-country contexts, where weak social protection and labor market informality may amplify both channels. We examine how different pathways into single motherhood affect children's cognitive development.

Methods — We use longitudinal data from the Indonesia Family Life Survey (IFLS) and employ Structural Equation Modeling (SEM) to estimate both direct and indirect effects of maternal marital status on children's cognitive outcomes, while controlling for demographic and household characteristics.

Findings — The results show that children in single-mother households, particularly those experiencing divorce, have lower cognitive scores. Poverty plays a key mediating role, as higher poverty levels are associated with worse cognitive outcomes. Households headed by divorced individuals exhibit higher poverty, while the effect of widowhood is smaller and not statistically significant. In addition, larger household size and greater distance from economic centers increase poverty, whereas higher education of the household head and per capita expenditure reduce it.

Implication — The findings suggest that policies targeting single-mother households should address both economic vulnerability and structural constraints, including limited access to services and unequal labor market opportunities.

Originality — This paper contributes to the limited longitudinal literature in developing countries by comparing divorce and widowhood and their roles in perpetuating intergenerational disadvantages through economic and non-economic channels.

Keywords — Single Parenthood; Cognitive Development; Poverty; Family Structure; Indonesia

Introduction

Single mothers across low- and middle-income countries face a distinctive form of socioeconomic vulnerability as they must simultaneously secure income and provide intensive caregiving that often happens in contexts where both markets and states offer limited support. This dual burden is

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especially consequential in settings characterized by high labor informality and fragmented welfare systems. As a result, single motherhood represents a site where broader institutional constraints shape unequal welfare outcomes.

Recent studies attribute this vulnerability to a set of interrelated mechanisms. Gendered labor market structures expose women to wage penalties and unstable employment, particularly following marital dissolution, when caregiving responsibilities intensify, and career interruptions accumulate (Kalil & Ryan, 2020; Rees et al., 2023). These constraints reduce earning capacity and limit re-entry into formal employment. At the same time, the need to combine paid work with caregiving generates persistent job–family strain, often leading single mothers to adopt short-term coping strategies rather than long-term investment. Financial pressures are further compounded by the erosion of social support networks following divorce, which weakens informal insurance mechanisms, especially in developing-country contexts.

The economic strain in these households is further intensified by the dual burden of caregiving and income generation, often leading to significant work–family conflict (Musick & Meier, 2010; Nomaguchi & Milkie, 2020). Evidence suggests that single mothers frequently rely on short-term coping strategies to manage financial hardship and economic instability (Kalil & Ryan, 2020; McKenzie & McKay, 2018), while female-headed households face greater constraints in investing in children's education and development (Cooper & Stewart, 2021; Duncan et al., 2017). Furthermore, divorce reduces both economic and social resources, severely limiting children's developmental opportunities (Agmase, 2021). Household allocation decisions often reflect these precarious financial conditions; while female decision-makers tend to prioritize child-related and health expenditures, they do so under significant constraints (Rees et al., 2023; Ridge & Millar, 2011).

These household-level constraints have well-documented intergenerational consequences. Children in single-parent households tend to experience poorer educational, economic, and psychosocial outcomes, largely due to resource constraints and reduced parental time investment (Lee & McLanahan, 2015; McLanahan et al., 2013; McLanahan & Sandefur, 1997). Children in single-parent families also often experience worse development outcomes, with effects intensifying over time, particularly during early childhood and among boys. In resource-constrained settings, these disadvantages are often amplified as a lack of stable employment restricts households' ability to smooth shocks or invest in children's development. Beyond education, evidence also links parental absence to poorer mental health and higher substance use among children (Annor et al., 2024). Thus, single motherhood can become a channel through which structural inequalities are reproduced across generations.

Despite the issue's salience, their interaction with state support systems remains underexplored in many developing countries. Social protection programs in these contexts are often designed around poverty status or employment categories rather than family structure. This potentially overlooks the specific vulnerabilities associated with single-parent households and raises an important question on the extent to which existing welfare programs can mitigate or inadvertently reproduce the disadvantages faced by single mothers and their children.

Indonesia is particularly useful to examine this question. Over the past two decades, the country has developed an extensive social protection system, including educational transfer, conditional cash transfer, and a universal healthcare system. While these programs provide broad support to low-income households, few are explicitly designed to address the structural vulnerabilities associated with single motherhood. At the same time, rising divorce rates, persistent gender norms, and weak enforcement of child support obligations may further exacerbate these risks.

This paper builds on the broader literature on family structure and inequality by examining how maternal marital status shapes children's well-being in a developing-country context. Using longitudinal data from the Indonesia Family Life Survey (IFLS), we analyze the effects of divorce and widowhood on household poverty and children's cognitive outcomes. Combining Structural Equation Modeling (SEM) with instrumental-variable approaches, the analysis identifies both direct and mediated pathways through which family structure affects child development. In doing so, the paper contributes to a more general understanding of how gender, family structure, and welfare institutions interact to shape intergenerational inequality in developing countries.

Indonesia provides a relevant context due to rising divorce rates and shifting family structures. Divorce cases increased substantially over the past decade, as shown in Table 1, reflecting broader social and economic changes. Qibthiyyah and Utomo (2016) document declining fertility and changing marital patterns, including reduced assortative marriage, indicating evolving household dynamics. Consistent with this trend, Indonesia exhibits relatively high divorce rates compared to other Asian countries (Dommaraju & Jones, 2011). This is reflected in household composition patterns, where female-headed households are more prevalent among younger and older age groups, corresponding to higher rates of divorce and widowhood (Table 2)[1]. Over time, the share of female household heads has also increased, as shown in Table 3.

Table 1. Divorce Cases in Indonesia 2014-2023 (Court Ruling)

Year	Number of Divorce Cases
2014	344,237
2015	347,256
2016	365,633
2017	374,516
2018	408,202
2019	439,002
2020	291,677
2021	447,743
2022	516,344
2023	463,654

Source: Director General of the Religious Courts Administration, Supreme Court of the Republic of Indonesia, and Indonesia Statistic Report

Furthermore, Dommaraju and Jones (2011) show that Indonesia has a higher divorce rate than other Asian countries, and this rate tends to increase each year. The increasing rate of divorces is evident in the household head data from the National Socio-Economic Survey (Susenas). Table 2 shows the household head's gender, age, and place of residence based on the National Socio-Economic Survey (Susenas) in 2016.

Table 2. Household Head, Age, and Living Location

Age	Male (%)			Female (%)		
	Urban	Rural	Total	Urban	Rural	Total
<18	0.06	0.04	0.05	0.4	0.17	0.3
18-25	3.34	2.41	2.89	7.75	1.22	4.65
26-40	34.55	33.76	34.16	11.86	11.93	11.89
41-50	27.92	27.58	27.76	18.4	18.18	18.32
51-60	19.99	20.35	20.18	25.8	25.69	25.73
>60	14.12	15.87	14.96	35.76	42.83	39.09

Source: Susenas 2016, author's calculations.

This study uses IFLS wave 3 and wave 5 data, which provide detailed longitudinal information on household socioeconomic conditions. IFLS enables comparison of children's well-being across divorced and widowed single-mother households in Indonesia. Many divorces are often driven by economic factors and accompanied by weak enforcement of child support obligations, increasing children's vulnerability. Despite these trends, empirical research, especially in Indonesia, remains limited and largely focuses on psychological outcomes. This study extends the literature by examining both current well-being, proxied by consumption-based poverty measures, and future well-being through educational attainment.

We examine the effects of single parenthood on children's well-being and educational outcomes. Key variables include parental marital status, per capita expenditure, access to services,

[1] We are using Susenas 2016 to match the Indonesia Family Life Surveys Data Wave 5

and children's cognitive scores. In the analysis section, we try to distinguish between divorced and widowed single-mother households to capture heterogeneous effects.

Table 3. Percentage of Female Household Head 2012 - 2022

Year	Female Household Head (%)
2012	14.42
2013	14.84
2014	14.73
2015	14.63
2016	15.02
2017	15.17
2018	15.17
2019	15.46
2020	15.82
2021	14.38
2022	12.72

Source: Statistic Bureau of Indonesia, 2012-2022.

Gender differences are also considered. Recent studies show that the effects of family disadvantage may differ by gender, although the patterns are not uniform across outcomes. Evidence suggests that boys tend to be more responsive to family environment in early educational outcomes, while gender differences in longer-term outcomes, such as educational attainment and labor market performance, are less consistent and may even favor girls in certain contexts (Autor et al., 2019; Brenøe & Lundberg, 2018). In addition, children from single-parent or divorced families are more likely to experience adverse educational outcomes, reflecting both economic and non-economic constraints (Mencarini et al., 2019). These findings indicate that gendered responses to family disruption are heterogeneous and context-dependent. To situate these mechanisms, the following section reviews the theoretical and empirical channels linking single parenthood to child outcomes.

The effects of single parenthood on children and household welfare operate through interconnected mechanisms, including economic hardship, reduced parental support, limited access to community resources, parental conflict, and family instability (Cooper & Stewart, 2021; Härkönen et al., 2017; Lee & McLanahan, 2015; McLanahan et al., 2013). These channels jointly shape children's educational, health, and psychosocial outcomes.

Single-mother households face substantially higher poverty risks than two-parent families, reflecting structural economic disadvantages associated with single parenthood (Cooper & Stewart, 2021; McLanahan & Sandefur, 1997). Economic hardship is a key mechanism, as divorce is often followed by a significant decline in household income, particularly for women and their children (Härkönen et al., 2017; Nieuwenhuis & Maldonado, 2018). As a result, children in single-mother households are more likely to experience prolonged exposure to poverty, which reduces investments in health and education (Amato & Patterson, 2017; Shiba & Kondo, 2019). Although income accounts for much of the gap in outcomes, it does not fully explain differences across family types (Magnuson & Duncan, 2006; McLanahan & Sandefur, 1997). Economic vulnerability is further worsened by unstable employment and inconsistent child support (Cancian et al., 2011). The widening inequality reinforces disparities in children's outcomes (Cancian & Meyer, 2018; Cooper & Stewart, 2021; Kalil & Ryan, 2020).

Besides income, single parenthood also often reduces parental support and supervision. Parental involvement is central to child development, yet single parents tend to face higher levels of stress and time constraints (Meier et al., 2016; Nishioka et al., 2020). Empirical evidence suggests that children in single-parent families may receive lower levels of parental monitoring and educational support, which can increase the risk of behavioral and developmental problems (Fomby & Cherlin, 2007; Kalil & Ryan, 2020; McLanahan & Sandefur, 1997). While some children develop resilience, parental absence is also widely associated with emotional distress and adjustment difficulties (Amato, 2000; Härkönen et al., 2017). This pattern is also observed in the Indonesian context (Heri et al., 2022).

Another factor is limited access to community resources. Single-parent families are more likely to face residential and neighborhood disadvantages due to economic constraints, which reduce access to quality institutions and social support (Cooper & Stewart, 2021; Fomby & Cherlin, 2007). These conditions may limit access to education and health services and increase vulnerability during periods of crisis (Cusinato et al., 2020; Moncrief et al., 2014).

Parental conflict and family instability also play critical roles. Exposure to persistent conflict is associated with poorer child outcomes (Amato & Afifi, 2006; Camisasca et al., 2016; Kalmijn, 2016). In some cases, children in high-conflict intact families fare worse than those in divorced families, suggesting that conflict, rather than divorce per se, drives negative effects (Amato & Keith, 1991). However, post-divorce conflict often persists. More broadly, cumulative instability, including residential moves, school changes, and economic shocks, has lasting impacts on mental health and educational outcomes (Fergusson et al., 2007; Hess, 2022; Xu et al., 2008). Family instability may be as detrimental as low income itself (McLanahan & Sandefur, 1997), particularly when combined with repeated transitions (Fomby & Cherlin, 2007; Lee & McLanahan, 2015).

Gendered responses remain mixed. Some evidence suggests that boys tend to exhibit more immediate behavioral problems following family disruption, while girls may experience more internalizing difficulties, such as anxiety or depression. However, overall findings indicate that gender differences are relatively small and vary depending on the outcome and life stage (Autor et al., 2019; Härkönen et al., 2017; McLanahan et al., 2013).

Household characteristics also shape outcomes. Education of the household head is a key determinant of economic well-being, associated with higher productivity and lower poverty (Noble et al., 2015), though its impact varies across contexts. Female-headed households face higher poverty risks due to structural inequalities (Maitra & Vahid, 2006; Okojie, 2002). A larger household size increases the risk of poverty and reduces per capita investment in children.

Spatial factors further influence poverty. Rural and remote areas face persistent disadvantages due to weaker labor markets and limited opportunities (Glasmeier & Farrigan, 2003; Partridge & Rickman, 2008; Swaminathan & Findeis, 2004). Distance from metropolitan centers exacerbates poverty, although local economic growth can mitigate these effects (Fisher, 2004; Partridge & Rickman, 2008).

Method

Based on these mechanisms, this study empirically tests both direct and indirect pathways using a structural framework. This study employs Structural Equation Modeling (SEM) to examine interdependent relationships among variables. SEM allows the estimation of direct and indirect effects within a unified framework. The model follows a mediation structure, where intermediate outcomes link explanatory variables to outcomes:

$$Y_1 = \alpha_0 + \alpha_1 x_1 + \alpha_2 x_2 + \epsilon_1 \quad (1)$$
$$Y_2 = \gamma_0 + \gamma_1 Y_1 + \gamma_2 x_1 + \epsilon_1 \quad (2)$$

Whereas equation (1) is a mediation model of equation (2). Equation (2) with dependent variable Y_2 is explained by equation (1) plus other covariates closely related to the dependent variable in equation (2). The mechanism by which the model works is shown in Figure 1.

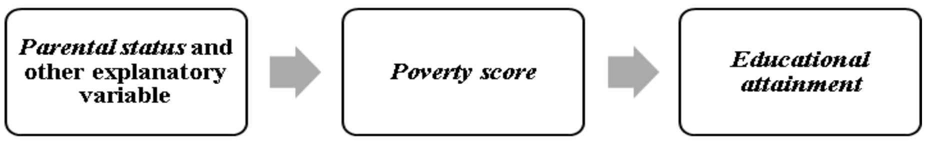

Figure 1. Mechanism Three Main Interest Variables
Source: Authors

We have two different equations (Y1 and Y2). Y1 is the dependent variable in equation (1) but becomes the independent (mediation) variable in equation (2). Here, Y_1 (poverty score) acts as a mediator for Y_2 (cognitive score). Independent variables influence cognitive outcomes both directly and indirectly through poverty.

Table 4. Variables in The Model

Variable	Explanation
Y1	Poverty Score - Constructed based on the multidimensional poverty indicators defined by BPS RI, covering nine household characteristics, including Size of the house, type of floor mat, type of wall, level of education of household head and spouse on average, electricity access, clean water access, toilet access, kitchen and stove type, and asset ownership status. The indicators are summed to form a score ranging from 0 (non-poor) to 9 (very poor).
Y2	Cognitive Score - Derived from the IFLS5 child cognitive test (Book EK), which consists of 17 items assessing basic skills such as pattern recognition and numeracy. The score is calculated as the total number of correct answers and normalized to ensure comparability across modules.
X1	1. Dummy Married (Divorced+Widowed=1) 2. Dummy Married (Divorced=1) 3. Dummy Married (Widowed=1) - Captured using three dummy variables: (1) ever-married but currently single (divorced or widowed), (2) divorced, and (3) widowed, allowing differentiation between types of single-parent households.
X2	Household Size - The number of family members in the household.
X3	Distance from Capital - How far is the household from the nearest *Kecamatan* (Sub-district) Capital?
X4	Distance from Business Center - How far the household is from the nearest business center (traditional market or a group of shops).
X5	Household Head Education - Measures the years of schooling of the household head.
X6	Per Capita Expenditure - Shows the per capita expenditure of each individual in the household.
X7	Age - Explaining the age of respondents when taking the cognitive score test.
X8	Child Status - Explains whether children are adopted (=1) or biological children.
X9	Number of Children in the Household - Shows whether having more children means having fewer resources for each child.
X10	Gender - Child gender.
X11	Teenager (teenager=1) - Showing whether the children are already teenagers (12 years old and above).

Source: IFLS3 and IFLS5 processed by authors

Table 5. Descriptive Variables of All Samples

Variable	Observation	Mean	Standard Deviation	Min	Max
Total Cognitive Score	2,454	8.513855	4.183764	0	17
Sex (=1 Male)	2,454	0.5240424	0.4995234	0	1
Household Size	2,454	6.330073	2.394674	2	16
Distance to Capital (KM)	2,454	21.31842	24.74596	0.5	150
Distance to Business Area (KM)	2,454	3.616809	3.615004	0.2	27
Household Head Education	2,454	3.02608	0.9371464	1	5
Per Capita Expenditure (Rp)	2,454	412874.7	324657.1	54666.67	3824833
Poverty Score	2,454	2.150774	1.344496	0	8
Divorce	2,454	0.0387123	0.1929478	0	1
Share Divorce at Community (%)	2,454	4.299549	5.511272	0	50
Age	2,454	8.72599	5.541036	0	18
Number of Children in the Household	2,454	3.490628	1.84668	1	11
Children by Adoption	2,454	0.0167074	0.128199	0	1
Teenager (Dummy)	2,454	0.3031785	0.4597254	0	1

Source: IFLS3 and IFLS5 processed by authors.

The analysis uses IFLS 3 and IFLS 5 data, covering 2,454 observations. Key variables include poverty score (constructed from nine BPS indicators), cognitive score (based on IFLS child assessments), marital status (divorced and widowed), household size, remoteness (distance to capital and business centers), household head education, and per capita expenditure. Additional controls include age, gender, number of children, adoption status, and a teenage dummy (Table 4). Descriptive statistics in Table 5 show an average cognitive score of 8.51 (range 0–17) and a poverty score of 2.15 (range 0–8). Households average 6.33 members, with a mean per capita expenditure of Rp 412,874.

Divorce is relatively rare (3.87%), while the average child age is 8.73 years. Per capita expenditure averages Rp. 412,874.70, ranging from Rp. 54,666.67 to Rp. 3,824,833. The average age of children in the dataset is 8.73 years, ranging from 0 to 18 years. Households have an average of 3.49 children, ranging from 1 to 11 children. Adoption is rare, with only 2% of children adopted in the dataset. The teenager dummy variable indicates that teenagers make up 30.32% of the observations.

Results and Discussion

Factors Affecting Poverty

Poverty is a multidimensional condition reflecting economic constraints and broader vulnerabilities in living conditions. The results indicate that marital status, household size, spatial accessibility, education, and expenditure significantly shape poverty outcomes. Table 6A shows that family structure is associated with economic vulnerability. Households headed by divorced or widowed individuals exhibit higher poverty scores than married households, with a stronger effect for divorce (0.466) than widowhood (0.145, not significant). This is consistent with evidence that children in single-parent households face higher poverty risks (Rank & Hirschl, 1999), while remarriage reduces such risks (Morrison & Ritualo, 2000), and mothers are more likely than fathers to fall into poverty following separation (Bianchi et al., 1999).

Table 6A. SEM Results – Poverty Model

Variables	(1) Divorced + Widow Non-standardized	(2) Divorced + Widow Standardized	(3) Divorce Non-standardized	(4) Divorce Standardized	(5) Widow Non-standardized	(6) Widow Standardized
Marital status (1=divorced+widowed)	0.235*	0.035*	0.466*	0.036*	0.145	0.018
	(0.129)	(0.019)	(0.248)	(0.019)	(0.150)	(0.019)
Household Size	0.042***	0.075***	0.043***	0.076***	0.042***	0.075***
	(0.011)	(0.019)	(0.011)	(0.019)	(0.011)	(0.019)
Distance District Capital	0.006***	0.122***	0.006***	0.121***	0.006***	0.122***
	(0.001)	(0.020)	(0.001)	(0.020)	(0.001)	(0.020)
Distance from Business Center	0.035***	0.095***	0.034***	0.094***	0.035***	0.094***
	(0.007)	(0.020)	(0.007)	(0.020)	(0.007)	(0.020)
Household head education	-0.372***	-0.257***	-0.373***	-0.257***	-0.372***	-0.257***
	(0.029)	(0.019)	(0.029)	(0.019)	(0.029)	(0.019)
Per Capita Expenditure	-7.62e-07***	-0.178***	-7.62e-07***	-0.178***	-7.60e-07***	-0.177***
	(8.45e-08)	(0.019)	(8.45e-08)	(0.019)	(8.45e-08)	(0.019)
Constant	3.044***	2.292***	3.051***	2.297***	3.050***	2.296***
	(0.119)	(0.090)	(0.119)	(0.090)	(0.119)	(0.090)
R^2 Poverty Score	0.176	0.176	0.176	0.176	0.175	0.175
Observations	2,285	2,285	2,285	2,285	2,285	2,285

Note: Standard errors in parentheses.
*, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively.

Household head education significantly reduces poverty (−0.372), reflecting improved labor market opportunities (Rank & Hirschl, 1999). In contrast, household size increases poverty (0.0422), consistent with resource dilution effects (Becker & Lewis, 1973; Lanjouw & Ravallion, 1995). Spatial factors also matter. A greater distance from district capitals and business centers is associated with higher poverty, reflecting limited access to economic opportunities. Overall, the model explains 17.6 percent of the variation in poverty.

Table 6B shows that poverty significantly reduces children's cognitive scores (−0.0846), consistent with evidence linking economic deprivation to lower cognitive development (Noble et al., 2015). Age is negatively associated with cognitive scores (−0.0161), and adopted children exhibit lower scores (−0.158), likely reflecting early-life disadvantages. Other factors, including gender and number of children, are not statistically significant. Adolescents also perform worse than younger children (−0.0992). Overall, the results highlight a key mechanism by which family structure influences poverty, which in turn negatively affects children's cognitive development.

Table 6B. SEM Results – Cognitive Score Model (natural log)

Variables	(1) Divorced + Widow Unstd	(2) Divorced + Widow Std	(3) Divorce Unstd	(4) Divorce Std	(5) Widow Unstd	(6) Widow Std
Poverty score	-0.085***	-0.231***	-0.085***	-0.231***	-0.085***	-0.231***
	(0.007)	(0.019)	(0.007)	(0.019)	(0.007)	(0.019)
Age	-0.016***	-0.184***	-0.016***	-0.185***	-0.016***	-0.184***
	(0.003)	(0.034)	(0.003)	(0.034)	(0.030)	(0.034)
Child Status (1=Adopted)	-0.158**	-0.042**	-0.158**	-0.042**	-0.158**	-0.042**
	(0.074)	(0.020)	(0.074)	(0.020)	(0.074)	(0.020)
Number of Children in The Household	-0.003	-0.099	-0.003	-0.010	-0.003	-0.010
	(0.005)	(0.020)	(0.005)	(0.020)	(0.006)	(0.020)
Gender (1=Male)	0.002	0.002	0.002	0.002	0.002	0.002
	(0.019)	(0.020)	(0.019)	(0.020)	(0.019)	(0.0196)
Adolescent	-0.099***	-0.094***	-0.099***	-0.094***	-0.099***	-0.094***
	(0.035)	(0.033)	(0.035)	(0.033)	(0.035)	(0.033)
Constant	2.477***	5.103***	2.477***	5.104***	2.477***	5.104***
	(0.030)	(0.082)	(0.030)	(0.082)	(0.030)	(0.082)
R^2 Cognitive Score	0.123	0.123	0.122	0.122	0.123	0.123
Observations	2,285	2,285	2,285	2,285	2,285	2,285

Note: Standard errors in parentheses.
*, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively.

The Factors Affecting Educational Attainment (Cognitive Score)

SEM results show that poverty, age, and child status significantly influence cognitive outcomes. Poverty has a strong negative effect, indicating that economic constraints reduce children's ability to achieve higher cognitive scores. As we mentioned earlier, this relationship operates through multiple channels. Lower-income households face constraints in providing educational inputs, including learning materials, stable environments, and access to quality schools. Limited parental time and resources reduce engagement in children's learning, while unstable conditions weaken school readiness. These findings are like those of Khiem and Kuo (2022), who show that improvements in household conditions enhance educational attainment.

Table 7. Factors Affecting Educational Attainment

	(1)	(2)	(3)	(4)	(5)	(6)
Variables	Total Cognitive Score (Divorce + Widow)	Total Cognitive Score (Divorce)	Total Cognitive Score (Widow)	Log of Total Cognitive Score (Divorce + Widow)	Log of Total Cognitive Score (Divorce)	Log of Total Cognitive Score (Widow)
Poverty Score	-0.568***	-0.568***	-0.568***	-0.084***	-0.084***	-0.084***
	(0.061)	(0.061)	(0.061)	(0.007)	(0.007)	(0.007)
Age	-0.165***	-0.165***	-0.165***	-0.023***	-0.023***	-0.023***
	(0.015)	(0.015)	(0.015)	(0.002)	(0.002)	(0.002)
Child Status (1=Adopted)	-0.845	-0.845	-0.845	-0.152**	-0.152**	-0.152**
	(0.629)	(0.629)	(0.629)	(0.074)	(0.074)	(0.074)
Number of Children in The Household	-0.099**	-0.099**	-0.099**	-0.002	-0.002	-0.002
	(0.045)	(0.045)	(0.045)	(0.005)	(0.005)	(0.005)
Gender (1=Male)	-0.344**	-0.344**	-0.344**	0.004	0.003	0.003
	(0.162)	(0.162)	(0.162)	(0.019)	(0.019)	(0.019)
Constant	11.71***	11.71***	11.71***	2.501***	2.501***	2.501***
	(0.241)	(0.241)	(0.241)	(0.029)	(0.029)	(0.029)
Observations	2,454	2,454	2,454	2,285	2,285	2,285

Note: Standard errors in parentheses.
*, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively.

Family size also constrains cognitive outcomes by diluting per capita resources. The negative association between the number of children and cognitive scores supports the quantity–quality trade-off framework (Becker & Lewis, 1973). Age is negatively associated with cognitive scores, suggesting variation in learning accumulation across cohorts. Gender differences are small and statistically insignificant once socioeconomic. Consistent with prior evidence emphasizing the dominant role of household conditions over gender in shaping cognitive outcomes (Hyde, 2005).

Robustness Check

Given the potential endogeneity of the poverty score and the divorce variables, we conducted a robustness check using instrumental-variables regression. We introduced a new variable called the share of divorce in the community (in percentage). We argue that the percentage of divorce in the community will affect the individual marital status or divorce rates, but will not directly influence the cognitive scores of children. This section includes three different models, each addressing a different type of divorce. The first model estimates the total cognitive score with the divorce variable instrumented. The second model measures the effect of the instrumented divorce variable on the poverty score. Finally, we run the reduced form in the third model, which estimates the total cognitive score, instrumenting both the poverty and divorce variables. All three models will be divided into three types of divorce: divorce (both parents alive), widower (one of the parents died), and a combination of both types.

In general, the equation of this section will be:

$$X_i = p_0 + p_1 Z_i + p_2 \theta_i + v_i \tag{3}$$
$$Y_i = \beta_0 + \beta_1 \widehat{X}_i + \beta_2 \theta_i + u_i \tag{4}$$

Where X is the main interest (divorce and poverty score), Z is the instrumental variable, θ is other control variables, and Y is the variable of interest (educational attainment measured by cognitive score).

Across specifications (Table 8), divorce has a large and statistically significant negative effect on cognitive scores, with stronger effects when divorce is isolated. Household head education consistently improves cognitive outcomes, while larger household size reduces them.

Table 8. Factors Affecting Cognitive Score (Instrumented)

Variables	(1) Total Cognitive Score	(2) Total Cognitive Score	(3) Total Cognitive Score
	Divorce + Widow	Divorce	Widow
Divorce (=1)	-12.38***	-41.96***	-16.57***
	(2.022)	(12.46)	(2.866)
Sex (=1 Male)	-0.248	-0.052	-0.321*
	(0.186)	(0.234)	(0.195)
Household Size	-0.128***	-0.159***	-0.116***
	(0.041)	(0.047)	(0.043)
Distance to District Capital	-0.004	-0.000	-0.006
	(0.004)	(0.007)	(0.004)
Distance to Business Area	-0.039	-0.028	-0.042*
	(0.027)	(0.046)	(0.025)
Household Head Education	1.052***	1.091***	1.039***
	(0.108)	(0.138)	(0.114)
Per Capita Expenditure	2.65e-07	3.64e-07	2.22e-07
	(3.63e-07)	(5.52e-07)	(3.48e-07)
Constant	6.872***	6.638***	6.930***
	(0.458)	(0.561)	(0.478)
Observations	2,454	2,454	2,454

Note: Robust standard errors in parentheses.
*, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively.

Table 9 shows that, once instrumented, divorce does not have a significant direct effect on poverty, whereas household size increases poverty, and education and expenditure reduce it. This suggests that divorce affects poverty indirectly through household characteristics. Table 10 confirms that poverty has a robust negative effect on cognitive scores (approximately 4 points), while divorce continues to exert a strong independent negative effect. These results reinforce the baseline findings and support a causal interpretation. Family disruption and economic conditions jointly shape children's cognitive outcomes. The IV estimates reinforce the causal interpretation of the baseline findings.

Table 9. Factors Affecting Poverty Score (Instrumented)

Variables	(1) Poverty Score	(2) Poverty Score	(3) Poverty Score
	Divorce + Widow	Divorce	Widow
Divorce (=1)	0.703	3.535	0.782
	(0.689)	(2.826)	(0.877)
Sex (=1 Male)	0.0244	0.013	0.027
	(0.038)	(0.039)	(0.038)
Household Size	0.032*	0.034*	0.0315*
	(0.018)	(0.017)	(0.018)
Distance to District Capital	0.004	0.003	0.004
	(0.003)	(0.00)	(0.003)
Distance to Business Area	0.037	0.038	0.037
	(0.024)	(0.023)	(0.024)
Household Head Education	-0.330***	-0.339***	-0.330***
	(0.060)	(0.055)	(0.060)
Per Capita Expenditure	-4.31e-07***	-4.49e-07***	-4.28e-07***
	(1.07e-07)	(1.06e-07)	(1.07e-07)
Constant	2.723***	2.752***	2.726***
	(0.245)	(0.235)	(0.248)
Observations	3,140	3,140	3,140

Note: Robust standard errors in parentheses.
*, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively.

Table 10. Factors Affecting Cognitive Score (Instrumented)

Variables	(1) Total Cognitive Score	(2) Total Cognitive Score	(3) Total Cognitive Score
	Divorce + Widow	Divorce	Widow
Poverty Score	-4.044***	-3.962***	-4.113***
	(0.852)	(0.939)	(0.864)
Divorce (=1)	-11.37***	-34.44***	-15.57***
	(2.550)	(13.24)	(3.469)
Sex (=1 Male)	-0.164	0.001	-0.233
	(0.262)	(0.287)	(0.270)
Household Size	0.060	0.030	0.074
	(0.069)	(0.077)	(0.070)
Distance to District Capital	0.021***	0.023**	0.020***
	(0.008)	(0.009)	(0.008)
Distance to Business Area	0.100**	0.107*	0.0990**
	(0.051)	(0.062)	(0.049)
Household Head Education	-0.585	-0.515	-0.626*
	(0.373)	(0.400)	(0.379)
Per Capita Expenditure	-2.47e-06***	-2.34e-06***	-2.55e-06***
	(6.92e-07)	(8.67e-07)	(6.81e-07)
Constant	19.34***	18.84***	19.62***
	(2.670)	(2.880)	(2.706)
Observations	2,454	2,454	2,454

Note: Robust standard errors in parentheses.
*, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively.

Conclusion

This study shows that single motherhood has a substantial negative effect on children's cognitive outcomes. This is particularly salient among single motherhood through divorce, as it is often followed by a loss of social network. While poverty is an important channel, family disruption exerts an independent and larger effect, showing both economic and non-economic constraints. Household characteristics similarly play a central role, as larger household size and spatial remoteness increase poverty. Meanwhile, education and expenditure reduce it. These findings highlight the importance of economic security and human capital in mitigating intergenerational disadvantage.

These results suggest that policy responses must move beyond short-term relief toward addressing the structural barriers faced by single-parent households. Economic interventions, such as targeted income support and employment opportunities for single mothers, remain essential. Yet, they are insufficient on their own. Investments in early childhood education, improved access to services in remote areas, stronger infrastructure, and academic support are equally important for reducing persistent inequalities, particularly in improving cognitive outcomes. Expanding access to services in remote areas and strengthening infrastructure can further reduce spatial inequalities. Complementary interventions, including counseling and community support, can address non-economic disadvantages.

At the same time, policies must address constraints specific to family structure, including limited enforcement of child support and gendered labor market disadvantages. Family planning programs can mitigate resource constraints associated with larger households, while stronger enforcement of child support and gender-sensitive policies can reduce the economic risks faced by single mothers. Overall, effective policy must integrate income support, human capital investment, and institutional reforms to address both the economic and social dimensions of family disruption. Such an approach is critical to improving children's long-term outcomes.

More broadly, these findings speak to a wider challenge in developing-country contexts where social protection systems are often designed around poverty status while overlooking how family structure shapes vulnerability. By showing that the effects of single motherhood extend beyond income and operate through social and institutional channels, this study highlights the limits of narrowly targeted welfare approaches. Addressing intergenerational inequality, therefore, requires both redistributive capacity and institutional designs that recognize how gender and family disruption structure access to resources and opportunities. In this sense, the case of Indonesia illustrates a more general point: overlooking family structure in policy design risks systematically underserving some of the most vulnerable households.

DATA AVAILABILITY

The supporting data are publicly available on Mendeley Data
https://data.mendeley.com/datasets/s56yk29d7b/1, DOI [10.17632/s56yk29d7b.1]

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Author contribution

All authors contributed to the conception and design of the study, data collection, analysis and interpretation of the results, and the writing of the manuscript. All authors read and approved the final version of the manuscript.

Use of AI tools declaration

Artificial intelligence tools were used to assist with language editing and to enhance the clarity and readability of the manuscript. The authors remain fully responsible for the content and conclusions of this study.

Conflict of interest

The authors declare that they have no known financial or non-financial competing interests that could have influenced the work reported in this paper. The research received no specific grant

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References

Agmase, N. G. (2021). The socio-economic effects of divorce on children in Woreta City Administration, Ethiopia. *Innovare Journal of Social Sciences*, *9*(3), 7–13. https://doi.org/10.22159/ijss.2021.v9i3.41129

Amato, P. R. (2000). The consequences of divorce for adults and children. *Journal of Marriage and Family*, *62*(4), 1269–1287. https://doi.org/10.1111/j.1741-3737.2000.01269.x

Amato, P. R., & Afifi, T. D. (2006). Feeling caught between parents: adult children's relations with parents and subjective well-being. *Journal of Marriage and Family*, *68*(1), 222–235. https://doi.org/10.1111/j.1741-3737.2006.00243.x

Amato, P. R., & Keith, B. (1991). Parental divorce and adult well-being: a meta-analysis. *Journal of Marriage and the Family*, *53*(1), 43. https://doi.org/10.2307/353132

Amato, P. R., & Patterson, S. E. (2017). Single-parent households and mortality among children and youth. *Social Science Research*, *63*, 253–262. https://doi.org/10.1016/j.ssresearch.2016.09.017

Annor, F. B., Amene, E. W., Zhu, L., Stamatakis, C., Picchetti, V., Matthews, S., Miedema, S. S., Brown, C., Thorsen, V. C., Manuel, P., Gilbert, L. K., Kambona, C., Coomer, R., Trika, J., Kamuingona, R., Dube, S. R., & Massetti, G. M. (2024). Parental absence as an adverse childhood experience among young adults in sub-Saharan Africa. *Child Abuse & Neglect*, *150*, 106556. https://doi.org/10.1016/j.chiabu.2023.106556

Autor, D., Figlio, D., Karbownik, K., Roth, J., & Wasserman, M. (2019). Family disadvantage and the gender gap in behavioral and educational outcomes. *American Economic Journal: Applied Economics*, *11*(3), 338–381. https://doi.org/10.1257/app.20170571

Becker, G. S., & Lewis, H. G. (1973). On the interaction between the quantity and quality of children. *Journal of Political Economy*, *81*(2, Part 2), S279–S288.

Bianchi, S. M., Subaiya, L., & Kahn, J. R. (1999). The gender gap in the economic well-being of nonresident fathers and custodial mothers. *Demography*, *36*(2), 195–203. https://doi.org/10.2307/2648108

Brenøe, A. A., & Lundberg, S. (2018). Gender gaps in the effects of childhood family environment: Do they persist into adulthood? *European Economic Review*, *109*, 42–62. https://doi.org/https://doi.org/10.1016/j.euroecorev.2017.04.004

Camisasca, E., Miragoli, S., & Di Blasio, P. (2016). Families with Distinct Levels of Marital Conflict and Child Adjustment: Which Role for Maternal and Paternal Stress? *Journal of Child and Family Studies*, *25*(3), 733–745. https://doi.org/10.1007/s10826-015-0261-0

Cancian, M., & Meyer, D. R. (2018). Reforming policy for single-parent families to reduce child poverty. *RSF: The Russell Sage Foundation Journal of the Social Sciences*, *4*(2), 91–112. https://doi.org/10.7758/RSF.2018.4.2.05

Cancian, M., Meyer, D. R., & Han, E. (2011). Child support: responsible fatherhood and the quid pro quo. *The ANNALS of the American Academy of Political and Social Science*, *635*(1), 140–162. https://doi.org/10.1177/0002716210393640

Cooper, K., & Stewart, K. (2021). Does household income affect children's outcomes? A systematic review of the evidence. *Child Indicators Research*, *14*(3), 981–1005. https://doi.org/10.1007/s12187-020-09782-0

Cusinato, M., Iannattone, S., Spoto, A., Poli, M., Moretti, C., Gatta, M., & Miscioscia, M. (2020). Stress, resilience, and well-being in italian children and their parents during the covid-19 pandemic. *International Journal of Environmental Research and Public Health*, *17*(22), 8297. https://doi.org/10.3390/ijerph17228297

Dommaraju, P., & Jones, G. (2011). Divorce trends in asia. *Asian Journal of Social Science*, *39*(6), 725–750. https://doi.org/10.1163/156853111X619201

Duncan, G. J., Magnuson, K., & Votruba-Drzal, E. (2017). Moving beyond correlations in assessing the consequences of poverty. *Annual Review of Psychology*, *68*(1), 413–434. https://doi.org/10.1146/annurev-psych-010416-044224

Fergusson, D. M., Boden, J. M., & Horwood, L. J. (2007). Exposure to single parenthood in childhood and later mental health, educational, economic, and criminal behavior outcomes. *Archives of General Psychiatry*, *64*(9), 1089. https://doi.org/10.1001/archpsyc.64.9.1089

Fisher, M. (2004). *On the empirical finding of a higher risk of poverty in rural areas: Is rural residence endogenous to poverty?* https://doi.org/https://doi.org/10.22004/ag.econ.18917

Fomby, P., & Cherlin, A. J. (2007). Family instability and child well-being. *American Sociological Review*, *72*(2), 181–204. https://doi.org/10.1177/000312240707200203

Glasmeier, A. K., & Farrigan, T. L. (2003). Poverty, Sustainability, and the Culture of Despair: Can Sustainable Development Strategies Support Poverty Alleviation in America's Most Environmentally Challenged Communities? *The ANNALS of the American Academy of Political and Social Science*, *590*(1), 131–149. https://doi.org/10.1177/0002716203257072

Härkönen, J., Bernardi, F., & Boertien, D. (2017). Family dynamics and child outcomes: an overview of research and open questions. *European Journal of Population*, *33*(2), 163–184. https://doi.org/10.1007/s10680-017-9424-6

Heri, M., Pratama, A. A., & Wijaya, I. G. A. S. (2022). Pengalaman single parent dalam mengasuh anak usia pra-sekolah(6 tahun). *Jurnal Keperawatan Silampari*, *6*(1), 290–296. https://doi.org/10.31539/jks.v6i1.4332

Hess, S. (2022). Effects of inter-parental conflict on children's social well-being and the mediation role of parenting behavior. *Applied Research in Quality of Life*, *17*(4), 2059–2085. https://doi.org/10.1007/s11482-021-10022-y

Hyde, J. S. (2005). The gender similarities hypothesis. *American Psychologist*, *60*(6), 581–592. https://doi.org/10.1037/0003-066X.60.6.581

Kalil, A., & Ryan, R. (2020). Parenting practices and socioeconomic gaps in childhood outcomes. *The Future of Children*, *30*(2020), 29–54. https://doi.org/10.1353/foc.2020.0004

Kalmijn, M. (2016). Father–child contact, interparental conflict, and depressive symptoms among children of divorced parents. *European Sociological Review*, *32*(1), 68–80. https://doi.org/10.1093/esr/jcv095

Khiem, P. H., & Kuo, Y.-C. (2022). Health insurance reform impact on children's educational attainment: evidence from Vietnam. *Review of Economics of the Household*, *20*(4), 1255–1285. https://doi.org/10.1007/s11150-021-09577-4

Lanjouw, P., & Ravallion, M. (1995). Poverty and household size. *The Economic Journal*, *105*(433), 1415. https://doi.org/10.2307/2235108

Lee, D., & McLanahan, S. (2015). Family structure transitions and child development: instability, selection, and population heterogeneity. *American Sociological Review*, *80*(4), 738–763. https://doi.org/10.1177/0003122415592129

Magnuson, K. A., & Duncan, G. J. (2006). The role of family socioeconomic resources in the black–white test score gap among young children. *Developmental Review*, *26*(4), 365–399. https://doi.org/10.1016/j.dr.2006.06.004

Maitra, P., & Vahid, F. (2006). The effect of household characteristics on living standards in South Africa 1993–1998: a quantile regression analysis with sample attrition. *Journal of Applied Econometrics*, *21*(7), 999–1018. https://doi.org/10.1002/jae.853

McKenzie, H., & McKay, F. H. (2018). Thinking outside the box: Strategies used by low-income single mothers to make ends meet. *Australian Journal of Social Issues*, *53*(3), 304–319. https://doi.org/10.1002/ajs4.48

McLanahan, S., & Sandefur, G. (1997). *Growing up with a single parent: what hurts, what helps*. Harvard University Press. https://doi.org/10.2307/j.ctv22tnmnn

McLanahan, S., Tach, L., & Schneider, D. (2013). The causal effects of father absence. *Annual Review of Sociology*, *39*(1), 399–427. https://doi.org/10.1146/annurev-soc-071312-145704

Meier, A., Musick, K., Flood, S., & Dunifon, R. (2016). Mothering experiences: how single parenthood and employment structure the emotional valence of parenting. *Demography*, *53*(3), 649–674. https://doi.org/10.1007/s13524-016-0474-x

Mencarini, L., Pasqua, S., & Romiti, A. (2019). Single-mother families and the gender gap in children's time investment and non-cognitive skills. *Review of Economics of the Household*, *17*(1), 149–176. https://doi.org/10.1007/s11150-017-9385-x

Moncrief, T., Beck, A. F., Simmons, J. M., Huang, B., & Kahn, R. S. (2014). Single parent households and increased child asthma morbidity. *Journal of Asthma*, *51*(3), 260–266. https://doi.org/10.3109/02770903.2013.873806

Morrison, D. R., & Ritualo, A. (2000). Routes to Children's Economic Recovery after Divorce: Are Cohabitation and Remarriage Equivalent? *American Sociological Review*, *65*(4), 560. https://doi.org/10.2307/2657383

Musick, K., & Meier, A. (2010). Are both parents always better than one? Parental conflict and young adult well-being. *Social Science Research*, *39*(5), 814–830. https://doi.org/10.1016/j.ssresearch.2010.03.002

Nieuwenhuis, R., & Maldonado, L. C. (2018). The triple bind of single-parent families: Resources, employment and policies. In *The triple bind of single-parent families* (pp. 1–28). Policy Press.

Nishioka, D., Saito, J., Ueno, K., & Kondo, N. (2020). *Single-parenthood and health conditions among children receiving public assistance in Japan: a cross-sectional study*. https://doi.org/10.21203/rs.3.rs-35089/v1

Noble, K. G., Houston, S. M., Brito, N. H., Bartsch, H., Kan, E., Kuperman, J. M., Akshoomoff, N., Amaral, D. G., Bloss, C. S., Libiger, O., Schork, N. J., Murray, S. S., Casey, B. J., Chang, L., Ernst, T. M., Frazier, J. A., Gruen, J. R., Kennedy, D. N., Van Zijl, P., … Sowell, E. R. (2015). Family income, parental education and brain structure in children and adolescents. *Nature Neuroscience*, *18*(5), 773–778. https://doi.org/10.1038/nn.3983

Nomaguchi, K., & Milkie, M. A. (2020). Parenthood and well-being: a decade in review. *Journal of Marriage and Family*, *82*(1), 198–223. https://doi.org/10.1111/jomf.12646

Okojie, C. E. E. (2002). *Gender and education as determinants of household poverty in Nigeria* (Number 2002/37). https://www.econstor.eu/handle/10419/52915

Partridge, M. D., & Rickman, D. S. (2008). Distance from urban agglomeration economies and rural poverty*. *Journal of Regional Science*, *48*(2), 285–310. https://doi.org/10.1111/j.1467-9787.2008.00552.x

Qibthiyyah, R., & Utomo, A. J. (2016). Family matters: demographic change and social spending in indonesia. *Bulletin of Indonesian Economic Studies*, *52*(2), 133–159. https://doi.org/10.1080/00074918.2016.1211077

Rank, M. R., & Hirschl, T. A. (1999). The economic risk of childhood in america: estimating the probability of poverty across the formative years. *Journal of Marriage and the Family*, *61*(4), 1058. https://doi.org/10.2307/354024

Rees, E., Beeber, S. N., Sampson, R., & Lietz, J. P. (2023). Empowering single parents: navigating socio-economic challenges and fostering resilience in family well-being. *Law and Economics*, *17*(2), 131–150. https://doi.org/10.35335/laweco.v17i2.5

Ridge, T., & Millar, J. (2011). Following families: working lone-mother families and their children. *Social Policy & Administration*, *45*(1), 85–97. https://doi.org/10.1111/j.1467-9515.2010.00755.x

Shiba, K., & Kondo, N. (2019). The Global Financial Crisis and Overweight among Children of Single Parents: A Nationwide 10-Year Birth Cohort Study in Japan. *International Journal of Environmental Research and Public Health*, *16*(6), 1001. https://doi.org/10.3390/ijerph16061001

Swaminathan, H., & Findeis, J. L. (2004). Policy intervention and poverty in rural america. *American Journal of Agricultural Economics*, *86*(5), 1289–1296. https://doi.org/10.1111/j.0002-9092.2004.00679.x

Xu, A., Zhang, J., & Xia, Y. R. (2008). Impacts of parents' divorce on chinese children: a model with academic performance as a mediator. *Marriage & Family Review*, *42*(3), 91–119. https://doi.org/10.1300/J002v42n03_05

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