

Financial Stability and CO₂ Emissions in High-Emission Asian Economies: A Nonlinear PMG-ARDL Analysis with Implications for Sustainable Competitiveness

*Ping Sun, Tianying Jiang**

Abstract

This study examines the effect of financial stability on carbon dioxide (CO₂) emissions in five high-emission Asian economies over the period 1992–2019. To capture possible asymmetries, the study applies pooled mean group autoregressive distributed lag (PMG-ARDL) and nonlinear PMG-ARDL estimators that distinguish between positive and negative changes in financial stability. Fully modified ordinary least squares (FMOLS) and dynamic ordinary least squares (DOLS) are used to assess the robustness of the long-run estimates. The results show that, in the long run, a positive change in financial stability has no statistically significant effect on CO₂ emissions, whereas a negative change in financial stability is associated with lower emissions. This points to an asymmetric relationship. In the short run, neither positive nor negative changes in financial stability exert statistically significant effects. The FMOLS and DOLS estimates partly support the long-run asymmetric findings. These results suggest that the environmental implications of financial stability depend on the direction of financial change and that environmental objectives should be incorporated more explicitly into financial-sector policy in high-emission Asian economies. Such policy alignment can support lower-carbon development pathways and strengthen the foundations of sustainable competitiveness in the long run.

Keywords: Financial stability, CO₂ emissions, nonlinear PMG-ARDL, asymmetry, high-emission Asian economies, sustainable competitiveness

JEL Classification: G21, E44, Q56, C33, O53

Article history: Received: December 2025; Accepted: May 2026; Published: October 2026

1 Introduction

In recent decades, the world has increasingly confronted global warming and environmental degradation driven by greenhouse-gas emissions. Carbon dioxide (CO₂) accounts for a large share of total greenhouse-gas emissions and remains a major contributor to climate risk (Quadrelli & Peterson, 2007). Developing and emerging economies, meanwhile, still rely heavily on energy-intensive production to grow and reduce poverty. For these economies, faster growth has so far been accompanied by higher emissions. Khoshnevis Yazdi and Golestani Dariani (2019), for instance, find two-way causality between economic growth and CO₂ emissions for Asian countries over 1980–2014, and the provincial evidence for China assembled by Luo et al. (2017) shows that the decoupling of agricultural CO₂ emissions from output has been uneven across regions. The difficulty is therefore not only one of reconciling expansion with environmental goals. Countries whose growth model rests on carbon-intensive production also face concerns about their long-run competitiveness, as trading partners and investors increasingly take carbon emissions into account.

Much of the earlier literature treats economic growth and energy consumption as the key determinants of environmental quality, usually within the environmental Kuznets curve (EKC) framework. The evidence on energy use is largely consistent. Alam et al. (2016) find that carbon emissions rise with both income and energy use in Brazil, China, India and Indonesia, and

Rahman et al. (2017) report for Malaysia that energy use contributes little to growth but adds considerably to pollution. The evidence on growth itself is more mixed. Alam et al. (2016) obtain the inverted-U pattern for Brazil, China and Indonesia but not for India, where emissions continue to rise with income, and Aslam et al. (2021) find a U-shaped income–emissions relationship for Malaysia, with trade openness and globalization enhancing emissions. Whether growth eventually becomes cleaner thus appears to depend on the structure of the economy and on policy rather than on income alone. Later studies have extended the set of determinants to include trade, institutions and financial conditions. Among these factors, the financial sector is of particular interest because it mobilizes savings and allocates capital. Since Schumpeter (1934), a large literature has shown that a developed financial system lowers information and transaction costs and channels resources to their most productive uses, and that this raises long-run growth (King & Levine, 1993; Levine, 2005).

The environmental implications of finance are less clear, and the empirical literature is divided. One line of argument holds that a well-developed and stable financial system improves environmental outcomes. It can finance cleaner technologies and energy-efficient investment, and it can discipline polluters: Dasgupta et al. (2001) show that stock markets in Chile, Argentina, the Philippines and Mexico respond to news about firms' environmental performance, which creates financial incentives for pollution control. In the same spirit, Nasreen et al. (2017) find that financial stability lowers CO₂ emissions in five South Asian countries over 1980–2012. The opposing view holds that easier access to credit raises output, consumption and energy demand, so that emissions rise with financial development. Sadorsky (2010) finds that financial development increases energy consumption in 22 emerging economies when it is measured with stock-market indicators, and Tian et al. (2025) report that financial inclusion raises emissions in the E-7 economies and even weakens the emission-reducing effect of renewable energy.

The evidence is more complex still when the focus shifts from financial development to financial stability. Instability, which Batuo et al. (2018) find tends to accompany financial development and liberalization, brings banking distress, volatile asset prices and weaker credit intermediation. This can depress economic activity and, with it, emissions; Safi et al. (2021) find that financial instability reduces consumption-based emissions in the E-7 economies. However, instability may also leave firms unable to finance cleaner technology and cause firms and regulators to give lower priority to environmental objectives. Shahbaz (2013) finds that financial instability increased environmental degradation in Pakistan over 1971–2009, whereas Baloch et al. (2018) find no significant effect for Saudi Arabia. For the large Asian emitters, Hafeez et al. (2022) show that the efficiency of financial institutions and markets affects renewable energy use in China, India, Japan and Russia, and Zhao et al. (2023) find that financial structure reduces CO₂ emissions in the long run in high-polluting Asian countries. These studies examine financial development or financial structure rather than financial stability, and neither considers whether improvements and deteriorations in financial conditions have different effects on emissions.

Three gaps emerge from the above discussion. The first is that most of the literature focuses on financial development rather than financial stability, although the two may not have the same environmental effects. The second is that almost all studies impose a symmetric relationship, so that an improvement in financial conditions is assumed to have exactly the opposite effect of a deterioration of the same size. Kirikkaleli et al. (2022) relax this assumption for Norway, but, to our knowledge, no comparable study exists for a group of high-emission economies. The third gap concerns the sample. China, India, Japan, South Korea and Russia together account

for a large portion of global CO₂ emissions, and their financial systems differ considerably in structure and depth; yet the finance–emissions literature has paid little attention to this group.

This study addresses these gaps by examining the effect of financial stability on CO₂ emissions in these five economies over 1992–2019. Russia is included because of its heavy energy use, its high emissions and its close economic ties with Asia; it also forms part of the sample used by Hafeez et al. (2022). We estimate the relationship with the pooled mean group (PMG) ARDL estimator and its nonlinear extension, in which financial stability is decomposed into negative and positive partial sums, and we check the long-run estimates with FMOLS and DOLS. The main contribution of the paper is to show that the effect of financial stability on emissions is asymmetric: in the PMG estimates, it is mainly deteriorations in stability, rather than improvements, that are associated with changes in emissions in the long run. A second contribution is empirical evidence for a group of economies that is important for global emissions but has rarely been studied together. Lastly, the findings bear on the debate about sustainable competitiveness. Zheng et al. (2025) argue that emerging economies must reconcile the low-carbon transition with continued growth. Our findings suggest that the allocation of credit by the financial system influences the extent to which this can be achieved.

2 Literature Review

2.1 Finance–Environment Nexus

The finance–growth literature holds that efficient financial intermediation improves the allocation of resources and thereby raises productivity and long-run growth (Schumpeter, 1934; King & Levine, 1993; Levine, 2005). A well-established financial system widens access to credit and lowers the cost of external finance, while also allowing risks to be diversified and borrowers to be monitored more effectively. Using data on 80 economies over 1960–1989, King and Levine (1993) show that the initial depth of the financial system predicts growth over the following three decades, and later evidence surveyed by Levine (2005) largely supports this conclusion.

The environmental implications of these mechanisms are less straightforward. Credit that finances new capacity may also be used for energy-efficient equipment and low-carbon production. In a study of 30 Chinese provinces over 1997–2011, Xiong and Qi (2018) find that the structural and technical effects of financial development outweigh its direct and wealth effects, so that per capita emissions decline as the financial sector develops. Capital markets may reinforce this effect, since investors in several developing countries reward good environmental performance and penalize firms facing citizens' complaints (Dasgupta et al., 2001). However, improved financial conditions also expand production and consumption, which in turn raises energy demand and emissions, as Tian et al. (2025) find for financial inclusion in the E-7 economies. In contrast, Omri et al. (2015) find that there is no causal association between financial development and emissions in the MENA countries, which indicates that the environmental effect of finance varies with the context in which it is examined.

2.2 Financial Stability and Emissions

Financial stability may affect emissions through the same two channels. A more stable financial system relaxes financing constraints, so that investment and economic activity expand and energy use and carbon emissions rise with them; this is the scale effect. By contrast, stability makes it easier for firms to finance the adoption of cleaner technology and allows capital to be allocated more efficiently across sectors, which lowers the emissions intensity of output; these

are the technique and composition effects that Xiong and Qi (2018) identify for financial development.

The two channels need not operate with equal strength when financial conditions improve and when they deteriorate. When financial conditions deteriorate, credit contracts quickly and the scale effect is likely to dominate, since investment in new capacity is typically among the first expenditures to be postponed. When conditions improve, the additional credit may go partly into cleaner technology and partly into expanding carbon-intensive capacity, so that the net effect could be small or even negligible. Kirikkaleli et al. (2022) allow for this kind of asymmetry in a nonlinear ARDL model for Norway over 1995–2018 and find that financial stability reduces environmental degradation in the long run. Norway is, however, a wealthy economy with a largely renewable electricity supply, and it is unclear whether the same holds for coal-dependent emerging economies. We therefore test directly whether improvements and deteriorations in financial stability affect CO₂ emissions in the same way.

2.3 Control Variables

The model includes three standard control variables. GDP per capita captures the scale of economic activity. Higher income usually means more production and consumption and therefore more pollution, although beyond some level it can also bring cleaner technology and stricter regulation; Wang et al. (2019) find such threshold effects of growth on air pollutants across Chinese provinces, and Rauf et al. (2023) report for the BRICS that energy use raises emissions while technological innovation reduces them. Energy consumption is included because in economies whose energy supply is still dominated by fossil fuels it is a major source of CO₂ emissions, although Gu et al. (2019) show for China that its effect depends on the level of energy technological progress. Lastly, trade openness is included as a control because it can raise emissions through the additional activity it generates or lower them through the transfer of cleaner technology and efficiency gains; Udeagha and Ngepah (2022) find for South Africa that the first effect dominates in the long run even though openness lowers emissions in the short run.

3 Data and Methodology

3.1 Data

We use annual data for 1992–2019 for five high-emission Asian economies: China, India, Japan, South Korea, and Russia. CO₂ emissions, measured in kilotons, are used as the indicator of environmental pressure. Financial stability is measured with a composite index in the spirit of Nasreen et al. (2017), who aggregate normalized financial indicators into a single stability index. Our index is simpler and is constructed by min-max normalization of domestic credit provided by the financial sector (% of GDP), domestic credit to the private sector by banks (% of GDP), liquid liabilities (% of GDP), and broad money (% of GDP). GDP per capita, energy consumption, and trade openness are included as control variables. Data were collected from the World Bank (2024). The variables, their measurement, data sources and role in the model are described in Table 1.

Table 1 – Variable definitions and data sources. Source: World Bank data and authors' calculations.

Variable	Symbol	Measurement	Source	Use in model
CO ₂ emissions	CO ₂	Carbon dioxide emissions, kilotons	World Bank	Dependent variable
Financial stability	FS	Index based on normalized financial-sector credit, private-sector bank credit, liquid liabilities, and broad money	World Bank; authors' index	Main explanatory variable
GDP per capita	GDP	GDP per capita	World Bank	Control variable
Energy consumption	EC	Energy consumption/use indicator	World Bank	Control variable
Trade openness	Trade	Exports plus imports as a share of GDP	World Bank	Control variable

The financial stability index is constructed as the simple average of the normalized financial-sector indicators. For indicator m in country i and year t , the normalized value is calculated as follows:

$$NFS_{mit} = (X_{mit} - \min(X_m)) / (\max(X_m) - \min(X_m)). \quad (1)$$

The composite index is then calculated as:

$$FS_{it} = (1 / 4) \sum_{m=1}^4 NFS_{mit}. \quad (2)$$

Higher values of the index indicate stronger financial-sector depth and stability as captured by the selected banking and liquidity indicators. Because these indicators also measure financial depth, the empirical interpretation should be understood as reflecting the stability-related dimension of the financial system captured by the index rather than a pure crisis-risk measure.

3.2 Methodology

To evaluate the impact of financial stability on CO₂ emissions in high-emission Asian economies, we specify the following long-run relationship:

$$\ln CO_{2it} = \alpha_i + \beta_1 FS_{it} + \beta_2 \ln GDP_{it} + \beta_3 \ln EC_{it} + \beta_4 Trade_{it} + \varepsilon_{it}. \quad (3)$$

Equation (3) expresses CO₂ emissions as a function of financial stability (FS), gross domestic product per capita (GDP), energy consumption (EC), and trade openness (Trade). The pooled mean group estimator is appropriate because it permits short-run coefficients and error variances to vary across countries while restricting the long-run coefficients to be identical (Pesaran et al., 1999). This long-run homogeneity assumption is assessed using the Hausman test (Hausman, 1978). To capture short-run dynamics, the model is re-parameterized into its error-correction form as follows:

$$\begin{aligned} \Delta \ln CO_{2it} = & \varphi_i (\ln CO_{2i,t-1} - \theta_1 FS_{i,t-1} - \theta_2 \ln GDP_{i,t-1} - \theta_3 \ln EC_{i,t-1} - \theta_4 Trade_{i,t-1}) \\ & + \sum \lambda_{ij} \Delta \ln CO_{2i,t-j} + \sum \gamma_{ij} \Delta X_{i,t-j} + \mu_i + v_{it}. \end{aligned} \quad (4)$$

In Equation (4), φ_i is the error-correction coefficient and is expected to be negative and statistically significant when a long-run relationship exists. Short-run effects are captured by the first differences of the variables, while the long-run coefficients are recovered from the lagged-level terms. We follow Pesaran et al. (2001) in using the ARDL framework to assess the existence of a long-run association.

To capture asymmetric effects, financial stability is split into its positive and negative partial sums, as proposed by Shin et al. (2014):

$$FS^+_{it} = \sum_{j=1}^t \Delta FS^+_{ij} = \sum_{j=1}^t \max(\Delta FS_{ij}, 0), \quad FS^-_{it} = \sum_{j=1}^t \Delta FS^-_{ij} = \sum_{j=1}^t \min(\Delta FS_{ij}, 0). \quad (5)$$

Here, FS^+ captures cumulative increases in financial stability, whereas FS^- captures cumulative decreases. Since FS^- is defined as a negative partial sum, a positive coefficient on FS^- implies that decreases in financial stability are linked with lower CO₂ emissions. The nonlinear long-run specification is therefore written as:

$$\ln CO_{2it} = \alpha_i + \beta_1^+ FS^+_{it} + \beta_1^- FS^-_{it} + \beta_2 \ln GDP_{it} + \beta_3 \ln EC_{it} + \beta_4 Trade_{it} + \varepsilon_{it}. \quad (6)$$

Long-run and short-run asymmetry are assessed using Wald tests. For robustness, we additionally employ FMOLS and DOLS estimators, which help address endogeneity and serial correlation in long-run estimation (Phillips & Hansen, 1990; Stock & Watson, 1993).

Before estimating the ARDL models, we examine the integration properties of the variables using LLC, IPS, and ADF-Fisher panel unit root tests (Levin et al., 2002; Im et al., 2003; Maddala & Wu, 1999). The Kao cointegration test is used as an additional check for a long-run relationship (Kao, 1999). Finally, panel causality tests are reported to explore the direction of association among the variables (Dumitrescu & Hurlin, 2012).

4 Results

Table 2 reports the results of the LLC, IPS and ADF-Fisher panel unit root tests. The PMG-ARDL and nonlinear PMG-ARDL approaches require that none of the variables be integrated of order two. The results show a mix of I(0) and I(1) variables, with none integrated of order two, which supports the use of the ARDL framework. A maximum lag length of two is selected on the basis of the Akaike information criterion.

Table 2 – Panel unit root test results. Source: authors' calculations.

Variable	LLC I(0)	LLC I(1)	IPS I(0)	IPS I(1)	ADF-Fisher I(0)	ADF-Fisher I(1)	Conclusion
CO ₂	-1.054	-1.558*	-1.273	-4.403***	-1.520	-5.872***	I(1)
FS	-1.920**	—	-1.108	-5.793***	-2.184**	—	I(0)/I(1), not I(2)
GDP	-1.831**	—	-1.051	-4.845***	1.122	-6.695***	I(1)
EC	-0.962	-4.122***	-1.495	-4.370***	-0.084	-7.444***	I(1)
Trade	-0.536	-4.969***	-2.705***	—	-2.677***	—	I(0)

Note: *** p < 0.01, ** p < 0.05, * p < 0.10.

As Table 3 shows, in the linear PMG-ARDL model the coefficient on financial stability is positive and significant, indicating that higher financial stability is linked with higher CO₂ emissions in the long run. GDP and energy consumption also have positive and significant coefficients, whereas trade openness is positive but statistically insignificant. In terms of magnitude, a 1% increase in GDP per capita and in energy consumption is associated with an increase in carbon emissions of about 0.50% and 0.10%, respectively.

The nonlinear PMG results in Table 3 differ in one important respect: the coefficient on the positive partial sum of financial stability (FS_POS) is positive but statistically insignificant, while the coefficient on the negative partial sum (FS_NEG) is positive and statistically significant. Because FS_NEG is defined as a negative partial-sum variable, this estimate implies that decreases in financial stability are associated with lower CO₂ emissions, which points to long-run asymmetry. The Wald long-run test also supports the presence of asymmetric effects.

GDP and energy consumption remain positively associated with CO₂ emissions in the nonlinear model, and trade openness becomes positive and significant. In the short run, neither positive nor negative changes in financial stability are statistically significant.

Table 3 – ARDL-PMG and NARDL-PMG long-run estimates. Source: authors' calculations.

Variable	ARDL coefficient	ARDL SE	ARDL t-stat.	ARDL p-value	NARDL coefficient	NARDL SE	NARDL t-stat.	NARDL p-value
FS	0.290***	0.046	6.317	0.000	—	—	—	—
FS POS	—	—	—	—	0.092	0.060	1.526	0.132
FS NEG	—	—	—	—	1.612**	0.604	2.668	0.011
GDP	0.504***	0.027	18.770	0.000	0.067**	0.033	2.026	0.047
EC	0.103**	0.050	2.068	0.043	1.474***	0.113	13.020	0.000
Trade	0.000	0.001	0.368	0.714	0.002***	0.001	2.828	0.006

Note: *** p < 0.01, ** p < 0.05, * p < 0.10.

Table 4 reports the short-run coefficients and diagnostic statistics for the linear and nonlinear PMG-ARDL models. In both models, the error-correction coefficient has the expected negative sign and is statistically significant, at the 10% level in the linear model and at the 5% level in the nonlinear model, which supports a long-run relationship. The Hausman statistics are insignificant, so the long-run homogeneity restriction of the PMG estimator is not rejected.

Table 4 – Short-run and diagnostic estimates. Source: authors' calculations.

Variable	ARDL-PMG estimate	NARDL-PMG estimate
ECM(-1)	-0.382* (0.201); t = -1.902; p = 0.062	-0.310** (0.148); t = -2.102; p = 0.039
D(FS)	0.061 (0.105); t = 0.579; p = 0.565	—
D(FS(-1))	-0.060 (0.115); t = -0.523; p = 0.603	—
D(FS(-2))	-0.100 (0.090); t = -1.110; p = 0.271	—
D(FS_POS)	—	-0.090 (0.128); t = -0.699; p = 0.487
D(FS_POS(-1))	—	0.202 (0.145); t = 1.395; p = 0.168
D(FS_NEG)	—	-0.238 (0.373); t = -0.638; p = 0.526
D(FS_NEG(-1))	—	-0.695 (0.644); t = -1.078; p = 0.285
D(GDP)	-0.095 (0.064); t = -1.490; p = 0.141	-0.024 (0.058); t = -0.406; p = 0.686
D(GDP(-1))	0.071 (0.062); t = 1.154; p = 0.253	0.178*** (0.053); t = 3.351; p = 0.001
D(GDP(-2))	-0.096*** (0.026); t = -3.656; p = 0.001	—
D(EC)	0.787*** (0.262); t = 2.999; p = 0.004	0.409 (0.301); t = 1.360; p = 0.179
D(EC(-1))	0.354* (0.182); t = 1.942; p = 0.057	0.211 (0.184); t = 1.149; p = 0.255
D(EC(-2))	-0.083 (0.106); t = -0.788; p = 0.434	—
D(Trade)	0.002 (0.003); t = 0.843; p = 0.403	-0.002 (0.002); t = -0.727; p = 0.470
D(Trade(-1))	0.004 (0.003); t = 1.303; p = 0.197	0.002* (0.001); t = 1.713; p = 0.091

Variable	ARDL-PMG estimate	NARDL-PMG estimate
D(Trade(-2))	0.002 (0.001); t = 1.631; p = 0.108	—
Constant	2.825* (1.446); t = 1.954; p = 0.055	1.131 (0.690); t = 1.639; p = 0.106
Log-likelihood	344.8	321.4
Kao cointegration statistic	-1.649	-2.355
Hausman statistic	0.231	1.023
Wald long-run asymmetry	—	7.215***
Wald short-run asymmetry	—	2.122

Note: Standard errors are in parentheses. *** p < 0.01, ** p < 0.05, * p < 0.10.

The FMOLS and DOLS estimates reported in Table 5 partly corroborate the long-run findings. In the linear specifications, financial stability remains positively associated with CO₂ emissions. In the asymmetric specifications, increases in financial stability are associated with higher emissions in both FMOLS and DOLS, unlike in the nonlinear PMG model, while decreases in financial stability are associated with lower emissions in FMOLS but have no significant effect in DOLS. The coefficients on FS_POS and FS_NEG differ markedly in size in each asymmetric specification, which is consistent with an asymmetric long-run relationship.

Table 5 – Robust panel estimates. Source: authors' calculations.

Variable	Symmetric POLS	Symmetric FMOLS	Symmetric DOLS	Asymmetric POLS	Asymmetric FMOLS	Asymmetric DOLS
FS	0.710*** (6.550)	0.070*** (4.730)	0.070*** (4.720)	—	—	—
FS_POS	—	—	—	0.139* (1.850)	0.080*** (9.330)	0.110*** (12.740)
FS_NEG	—	—	—	3.794*** (5.350)	0.160*** (7.200)	0.540 (0.070)
GDP	0.720*** (6.520)	0.090*** (16.340)	0.040*** (6.210)	0.383*** (3.630)	0.070*** (10.650)	0.050*** (5.770)
EC	0.811*** (4.680)	0.730*** (4.640)	0.480*** (7.880)	0.238 (1.260)	0.760*** (5.480)	0.550*** (9.480)
Trade	0.007* (1.880)	0.000*** (4.920)	0.000*** (7.400)	0.006 (1.600)	0.001*** (4.880)	0.001*** (17.900)

Note: POLS = pooled ordinary least squares, reported as a benchmark. t-statistics are reported in parentheses. *** p < 0.01, ** p < 0.05, * p < 0.10.

The panel causality results in Table 6 indicate bidirectional causality between financial stability and CO₂ emissions in the symmetric setting. In the asymmetric setting, bidirectional causality is also found between FS_NEG and CO₂ emissions. The evidence for causality from CO₂ emissions to FS_POS is weaker, with significance at the 10% level.

Table 6 – CO₂-related panel causality results. Source: authors' calculations.

Panel	Null hypothesis	W-statistic	Zbar-statistic	p-value
Symmetric	FS $\nleftrightarrow$ CO ₂	5.166	2.690	0.007
Symmetric	CO ₂ $\nleftrightarrow$ FS	4.704	2.270	0.023
Symmetric	GDP $\nleftrightarrow$ CO ₂	3.527	1.198	0.231
Symmetric	CO ₂ $\nleftrightarrow$ GDP	7.872	5.153	0.000
Symmetric	EC $\nleftrightarrow$ CO ₂	5.073	2.605	0.009
Symmetric	CO ₂ $\nleftrightarrow$ EC	5.118	2.646	0.008
Symmetric	Trade $\nleftrightarrow$ CO ₂	4.129	1.746	0.081
Symmetric	CO ₂ $\nleftrightarrow$ Trade	6.493	3.897	0.000
Asymmetric	FS_POS $\nleftrightarrow$ CO ₂	5.608	3.047	0.002
Asymmetric	CO ₂ $\nleftrightarrow$ FS_POS	4.298	1.868	0.062
Asymmetric	FS_NEG $\nleftrightarrow$ CO ₂	4.455	2.009	0.045

Panel	Null hypothesis	W-statistic	Zbar-statistic	p-value
Asymmetric	CO ₂ $\nrightarrow$ FS_NEG	7.279	4.551	0.000

Note: $X \nrightarrow Y$: X does not Granger-cause Y (H_0).

5 Discussion

Our results suggest that financial stability matters for emissions in the long run but not in the short run. None of the short-run coefficients on financial stability, or on its positive and negative components, is significant, while the error-correction coefficients in both models are negative and statistically significant, implying an annual speed of adjustment towards long-run equilibrium of between 31% and 38%. This pattern is to be expected, since financial conditions affect emissions mainly through investment and the capital stock, which adjust slowly.

In the linear model, a more stable financial system is associated with higher CO₂ emissions, as are GDP and energy consumption, which remain significant in almost every specification. A plausible interpretation is that, in these economies, the credit that a stable financial system makes available has largely financed the expansion of production and energy use rather than cleaner technology. This is the scale effect described in Section 2, and it is consistent with the findings of Sadorsky (2010) for energy consumption and Tian et al. (2025) for emissions. This result contrasts with Nasreen et al. (2017), who report that greater financial stability enhances environmental quality in South Asian economies, and with Kirikkaleli et al. (2022) for Norway. Part of the difference probably lies in how stability is measured. Our index is built from credit and money aggregates relative to GDP and therefore also captures financial depth. Nasreen et al. (2017) use a much broader index that includes measures of banking-sector soundness and macroeconomic vulnerability.

The nonlinear results qualify this finding in an important way. Once increases and decreases in financial stability are allowed to have separate effects, the increases are no longer significant in the long run, while the decreases are associated with lower emissions. The positive coefficient in the linear model therefore seems to be driven mainly by episodes of financial deterioration. The sample period contains several of these: Japan's banking crisis of 1997–98, the Asian financial crisis that affected South Korea in the same years, the Russian default of 1998, and the global financial crisis of 2008–09. In several of these episodes, notably Korea in 1998 and Japan and Russia in 2009, production and energy demand contracted and emissions declined accordingly. Such emission reductions do not reflect environmental progress, since they come from lower activity rather than cleaner production and are likely to be reversed once financial conditions recover. From this perspective, our nonlinear result is not as far from Nasreen et al. (2017) as the linear result suggests: once the two directions are separated, improvements in stability are not associated with higher emissions. However, the FMOLS and DOLS estimates support this only in part. Both find a positive and significant coefficient on the positive partial sum as well, so the conclusion that improvements in stability are environmentally neutral rests mainly on the PMG estimates.

6 Conclusion and Policy Implications

This paper has examined the effect of financial stability on CO₂ emissions in China, India, Japan, South Korea and Russia over 1992–2019, using linear and nonlinear PMG-ARDL models, with FMOLS and DOLS as robustness checks. In the linear model, financial stability is positively associated with emissions in the long run. When positive and negative changes in financial stability are separated, however, the relationship turns out to be asymmetric: increases have no significant long-run effect, whereas decreases are associated with lower emissions. In the short run, neither type of change is statistically significant. Among the control variables,

GDP and energy consumption are positively associated with emissions in almost all specifications, as is trade openness in most of them.

These findings imply that financial stability should not be regarded as environmentally neutral. In the economies studied here, a stable financial system appears to have supported growth, but without mechanisms that direct credit toward cleaner uses, it has also supported carbon-intensive expansion. The decline in emissions observed during periods of financial stress reflects lower economic activity rather than lower carbon intensity and therefore offers no lasting environmental benefit. Financial regulators and governments could thus complement stability objectives with measures that influence the allocation of credit, including environmental screening of loans and preferential financing for low-carbon and energy-efficient investment. Such measures need not weaken bank performance, as Mirza et al. (2025) find that greater exposure to high-ESG borrowers is associated with better performance among BRICS banks. They may also matter for competitiveness: as carbon pricing and border carbon adjustments become more common among trading partners, economies whose financial systems have shifted credit toward low-carbon production are likely to be better placed to sustain growth.

References

1. Alam, M. M., Murad, M. W., Noman, A. H. M., & Ozturk, I. (2016). Relationships among carbon emissions, economic growth, energy consumption and population growth: Testing Environmental Kuznets Curve hypothesis for Brazil, China, India and Indonesia. *Ecological Indicators*, 70, 466–479. <http://doi.org/10.1016/j.ecolind.2016.06.043>
2. Aslam, B., Hu, J., Hafeez, M., Ma, D., AlGarni, T. S., Saeed, M., Abdullah, M. A., & Hussain, S. (2021). Applying environmental Kuznets curve framework to assess the nexus of industry, globalization, and CO₂ emission. *Environmental Technology & Innovation*, 21, 101377. <http://doi.org/10.1016/j.eti.2021.101377>
3. Baloch, M. A., Danish, Meng, F., Zhang, J., & Xu, Z. (2018). Financial instability and CO₂ emissions: The case of Saudi Arabia. *Environmental Science and Pollution Research*, 25(26), 26030–26045. <http://doi.org/10.1007/s11356-018-2654-2>
4. Batuo, M., Mlambo, K., & Asongu, S. (2018). Linkages between financial development, financial instability, financial liberalisation and economic growth in Africa. *Research in International Business and Finance*, 45, 168–179. <http://doi.org/10.1016/j.ribaf.2017.07.148>
5. Dasgupta, S., Laplante, B., & Mamingi, N. (2001). Pollution and capital markets in developing countries. *Journal of Environmental Economics and Management*, 42(3), 310–335. <http://doi.org/10.1006/jeem.2000.1161>
6. Dumitrescu, E. I., & Hurlin, C. (2012). Testing for Granger non-causality in heterogeneous panels. *Economic Modelling*, 29(4), 1450–1460. <http://doi.org/10.1016/j.econmod.2012.02.014>
7. Gu, W., Zhao, X., Yan, X., Wang, C., & Li, Q. (2019). Energy technological progress, energy consumption, and CO₂ emissions: Empirical evidence from China. *Journal of Cleaner Production*, 236, 117666. <http://doi.org/10.1016/j.jclepro.2019.117666>
8. Hafeez, M., Rehman, S. U., Faisal, C. M. N., Yang, J., Ullah, S., Kaium, M. A., & Malik, M. Y. (2022). Financial efficiency and its impact on renewable energy demand and CO₂ emissions: Do eco-innovations matter for highly polluted Asian economies? *Sustainability*, 14(17), 10950. <http://doi.org/10.3390/su141710950>
9. Hausman, J. A. (1978). Specification tests in econometrics. *Econometrica*, 46(6), 1251–1271. <http://doi.org/10.2307/1913827>
10. Im, K. S., Pesaran, M. H., & Shin, Y. (2003). Testing for unit roots in heterogeneous panels. *Journal of Econometrics*, 115(1), 53–74. [http://doi.org/10.1016/S0304-4076\(03\)00092-7](http://doi.org/10.1016/S0304-4076(03)00092-7)
11. Kao, C. (1999). Spurious regression and residual-based tests for cointegration in panel data. *Journal of Econometrics*, 90(1), 1–44. [http://doi.org/10.1016/S0304-4076\(98\)00023-2](http://doi.org/10.1016/S0304-4076(98)00023-2)
12. Khoshnevis Yazdi, S., & Golestani Dariani, A. (2019). CO₂ emissions, urbanisation and economic growth: Evidence from Asian countries. *Economic Research-Ekonomska Istraživanja*, 32(1), 510–530. <http://doi.org/10.1080/1331677X.2018.1556107>
13. King, R. G., & Levine, R. (1993). Finance and growth: Schumpeter might be right. *The Quarterly Journal of Economics*, 108(3), 717–737. <http://doi.org/10.2307/2118406>
14. Kirikkaleli, D., Castanho, R. A., Genc, S. Y., Oyebanji, M. O., & Couto, G. (2022). The asymmetric and long-run effect of financial stability on environmental degradation in Norway. *Sustainability*, 14(16), 10131. <http://doi.org/10.3390/su141610131>
15. Levin, A., Lin, C. F., & Chu, C. S. J. (2002). Unit root tests in panel data: Asymptotic and finite-sample properties. *Journal of Econometrics*, 108(1), 1–24. [http://doi.org/10.1016/S0304-4076\(01\)00098-7](http://doi.org/10.1016/S0304-4076(01)00098-7)
16. Levine, R. (2005). Finance and growth: Theory and evidence. In P. Aghion & S. N. Durlauf (Eds.), *Handbook of Economic Growth*, 1A, 865–934. Elsevier. [http://doi.org/10.1016/S1574-0684\(05\)01012-9](http://doi.org/10.1016/S1574-0684(05)01012-9)

<https://doi.org/10.7441/joc.2026.SI.15>

17. Luo, Y., Long, X., Wu, C., & Zhang, J. (2017). Decoupling CO₂ emissions from economic growth in agricultural sector across 30 Chinese provinces from 1997 to 2014. *Journal of Cleaner Production*, 159, 220–228. <http://doi.org/10.1016/j.jclepro.2017.05.076>
18. Maddala, G. S., & Wu, S. (1999). A comparative study of unit root tests with panel data and a new simple test. *Oxford Bulletin of Economics and Statistics*, 61(S1), 631–652. <http://doi.org/10.1111/1468-0084.0610s1631>
19. Mirza, N., Umar, M., Lobont, O.-R., & Safi, A. (2025). ESG lending, technology investment and banking performance in BRICS: Navigating sustainability and financial stability. *China Finance Review International*, 15(2), 324–336. <http://doi.org/10.1108/CFRI-09-2024-0496>
20. Nasreen, S., Anwar, S., & Ozturk, I. (2017). Financial stability, energy consumption and environmental quality: Evidence from South Asian economies. *Renewable and Sustainable Energy Reviews*, 67, 1105–1122. <http://doi.org/10.1016/j.rser.2016.09.021>
21. Omri, A., Daly, S., Rault, C., & Chaibi, A. (2015). Financial development, environmental quality, trade and economic growth: What causes what in MENA countries. *Energy Economics*, 48, 242–252. <http://doi.org/10.1016/j.eneco.2015.01.008>
22. 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. <http://doi.org/10.1080/01621459.1999.10474156>
23. 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. <http://doi.org/10.1002/jae.616>
24. 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), 99–125. <http://doi.org/10.2307/2297545>
25. Quadrelli, R., & Peterson, S. (2007). The energy–climate challenge: Recent trends in CO₂ emissions from fuel combustion. *Energy Policy*, 35(11), 5938–5952. <http://doi.org/10.1016/j.enpol.2007.07.001>
26. Rahman, M. S., Noman, A. H. M., & Shahari, F. (2017). Does economic growth in Malaysia depend on disaggregate energy? *Renewable and Sustainable Energy Reviews*, 78, 640–647. <http://doi.org/10.1016/j.rser.2017.05.010>
27. Rauf, A., Ali, N., Sadiq, M. N., Abid, S., Kayani, S. A., & Hussain, A. (2023). Foreign direct investment, technological innovations, energy use, economic growth, and environmental sustainability nexus: New perspectives in BRICS economies. *Sustainability*, 15(18), 14013. <http://doi.org/10.3390/su151814013>
28. Sadorsky, P. (2010). The impact of financial development on energy consumption in emerging economies. *Energy Policy*, 38(5), 2528–2535. <http://doi.org/10.1016/j.enpol.2009.12.048>
29. Safi, A., Chen, Y., Wahab, S., Ali, S., Yi, X., & Imran, M. (2021). Financial instability and consumption-based carbon emission in E-7 countries: The role of trade and economic growth. *Sustainable Production and Consumption*, 27, 383–391. <http://doi.org/10.1016/j.spc.2020.10.034>
30. Schumpeter, J. A. (1934). *The theory of economic development*. Harvard University Press.
31. Shahbaz, M. (2013). Does financial instability increase environmental degradation? Fresh evidence from Pakistan. *Economic Modelling*, 33, 537–544. <http://doi.org/10.1016/j.econmod.2013.04.035>
32. Shin, Y., Yu, B., & Greenwood-Nimmo, M. (2014). Modelling asymmetric cointegration and dynamic multipliers in a nonlinear ARDL framework. In R. C. Sickles & W. C. Horrace

- (Eds.), *Festschrift in Honor of Peter Schmidt: Econometric Methods and Applications* (pp. 281–314). Springer. http://doi.org/10.1007/978-1-4899-8008-3_9
33. Stock, J. H., & Watson, M. W. (1993). A simple estimator of cointegrating vectors in higher order integrated systems. *Econometrica*, 61(4), 783–820. <http://doi.org/10.2307/2951763>
 34. Tian, S., Hassan, H., Safi, A., & Umar, M. (2025). Financial accessibility and environmental performance: A trade-adjusted analysis of renewable energy's impact in major emerging markets. *Humanities and Social Sciences Communications*, 12, Article 1490. <http://doi.org/10.1057/s41599-025-05744-5>
 35. Udeagha, M. C., & Ngepah, N. (2022). Dynamic ARDL simulations effects of fiscal decentralization, green technological innovation, trade openness, and institutional quality on environmental sustainability: Evidence from South Africa. *Sustainability*, 14(16), 10268. <http://doi.org/10.3390/su141610268>
 36. Wang, S., Li, C., & Zhou, H. (2019). Impact of China's economic growth and energy consumption structure on atmospheric pollutants: Based on a panel threshold model. *Journal of Cleaner Production*, 236, 117694. <http://doi.org/10.1016/j.jclepro.2019.117694>
 37. World Bank. (2024). World Development Indicators, WDI Database Archives, and Global Financial Development Database. <https://databank.worldbank.org>
 38. Xiong, L., & Qi, S. (2018). Financial development and carbon emissions in Chinese provinces: A spatial panel data analysis. *The Singapore Economic Review*, 63(2), 447–464. <http://doi.org/10.1142/S0217590817400203>
 39. Zhao, S., Ozturk, I., Hafeez, M., & Ashraf, M. U. (2023). Financial structure and CO₂ emissions in Asian high-polluted countries: Does digital infrastructure matter? *Environmental Technology & Innovation*, 32, 103348. <http://doi.org/10.1016/j.eti.2023.103348>
 40. Zheng, L., Cao, Y., Umar, M., Wang, X., & Safi, A. (2025). Environmental policy, digital economy, and green innovation: Navigating the low-carbon transition in emerging seven economies. *Economic Change and Restructuring*, 58(4), Article 67. <http://doi.org/10.1007/s10644-025-09893-0>

Contact information

Ping Sun, PhD

School of International Economics and Trade,
Ningbo University of Finance & Economics,
Ningbo 315175, Zhejiang, China
E-mail: pynsun0574@163.com
ORCID: 0009-0000-0069-0484

Tianying Jiang, PhD* (corresponding author)

School of Finance,
Ningbo University of Finance & Economics
Ningbo, Zhejiang 315175, China
E-mail: jiangtianying@nbufe.edu.cn
ORCID: 0009-0002-7063-3383

<https://doi.org/10.7441/joc.2026.SI.15>