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Unravelling the Nonlinear Dynamics between Foreign Capital Inflows and Trade Openness in Sub-Saharan Africa: Evidence from Panel Cointegration and Causality Tests

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Abstract

This paper proposes using a nonlinear panel cointegration test called the panel smooth transition regression (PSTR) to investigate the relationship between foreign capital inflows and trade openness for 31 Sub-Saharan African (SSA) countries from 1985 to 2023. Furthermore, to analyze the direction of causality between foreign capital inflows and trade openness, the study estimates a nonlinear Panel Vector Error Correction Model. The paper also analyzes the causal relationships in low economic growth and high economic growth regimes using nonlinear causality tests. Besides, the study deals with the cross-sectional dependency problem in both the nonlinear Panel Vector Error Correction Model and the nonlinear panel cointegration test. The empirical findings revealed the existence of bidirectional causality between foreign capital inflows and trade openness in SSA countries. As well, the results showed a positive and significant relationship between foreign capital inflows and trade openness during the 38 years represented by the data. This implies that the two flows are connected and therefore can exert a positive multiplier effect on each other. The study recommends that more steps should be taken to deepen interregional and regional economic integration in Africa, including the implementation of free trade agreements (FTAs), so as to promote trade and foreign capital inflows. Besides, the move towards boosting intra-African continental trade should be given support by African governments because it is the right direction and will have a significant effect on the attraction of both foreign capital inflows and trade flows in the region.

Keywords: Foreign capital inflows, Trade openness, Nonlinear Panel Vector Error Correction Model, Non-linear panel cointegration, cross-section dependency.

1. Introduction

There has been an increase in economic integration and globalization of economies over recent decades, where trade and capital flows have been the main forces that shape the process (Belke & Domnick, 2018; ALI, A., 2026). Globally, trade and investment flows have increased dramatically, although Africa has benefited little from that growth. This is evidenced by a decline in Africa's share of world imports from 4.5% in 1970 to 2.9% in 2021, while Africa's share of world exports declined from over 3.5% to 3.3% over the same period (Staiger, R. W., 2026). On the other hand, the Sub-Saharan Africa (SSA) region has been receiving a comparatively smaller share of the increased global flows; for instance, between 2000 and 2020, the foreign portfolio investment (FPI) and foreign direct investment (FDI) inflows in the region were only 1% of global FPI inflows and 2% of the global FDI inflows, respectively (Chuhan-Pole et al., 2023). Likewise, the region faces a large saving-investment gap (Kabadayi et al., 2024). Even when compared to other developing countries, SSA has the lowest average gross saving rate. For instance, from 2000 to 2023, the average saving rate in SSA was only 22%, while in East Asia and the Pacific it was 27% and in South Asia 34% (World Bank, 2025).

The link between foreign capital inflows and trade openness has been well documented in the literature. However, the empirical proof of their likely causal relationship remains a major debatable issue because of the mixed and inconclusive results from previous studies. There are a lot of theoretical models about international trade and foreign capital flows that provide evidence about substitutability and complementary association between trade and capital flows (Seyoum & Lin 2014). Those theoretical models have been empirically tested by different scholars in different countries/regions of the world. Nevertheless, the systematic empirical investigation for sub–Saharan African countries is still lacking. Generally, trade and

capital flows are theoretically regarded as either complements if they exhibit a positive relationship or substitutes if they are negatively correlated. Mundell (1957) considered trade and capital flows as substitutes, implying that when trade integration increases, capital inflows decrease. However, new theoretical models have challenged this view by incorporating financial frictions, thereby suggesting trade and capital flows to be complements, meaning that when trade integration increases, capital inflows also increase (Antras & Caballero, 2009). This complementary relationship helps to boost output performance or growth. Besides, previous empirical literature has shown that the trade and capital flows nexus ranges from positive (see Gavrila-Paven, 2026; Sedik & Sun, 2012; Belke & Domnick, 2018) to negative (Ali et al., 2024; Bende-Nabende, 2002; Grünfeld & Svindal, 2000). Conversely, other studies found no significant association between trade and capital flows (Ahearne et al., 2023), while the current debate about the causal links between the two flows ranges from unidirectional causality, bidirectional causality, to no causality at all (Sazali et al., 2024).

Conflicting results in the previous literature are due to the use of different sample sizes, time periods, measurement problems, functional forms, and econometric methods. Therefore, given the mixed and inconclusive results from the previous studies and the limited empirical evidence, particularly for the case of SSA, there is a pressing need to investigate the causal relationship between foreign capital inflows and trade openness in SSA countries. This is undertaken in the current study using a large sample size, longer time periods, and the application of a new econometric method, which is a non-linear approach called the panel smooth transition regression (PSTR) model. This is because all the previous studies used linear models in examining the nexus between foreign capital inflows and trade openness. Then again, this paper contributes to the knowledge because the previous studies focused on foreign direct investment as a proxy for capital flows in analysing the nexus between foreign capital inflows and trade openness and ignored other forms of capital flows, like foreign portfolio investment, which is included in the current paper. Another novelty of this paper is the estimation of the causal relationship between foreign capital inflows and trade openness in each regime separately using a non-linear panel error correction model, which allows for smooth changes between regimes. Discovering regime-dependent interactions between foreign capital inflows and trade openness is crucial for designing more appropriate trade and foreign capital policies. Lastly, the study also proposes a new method to remedy the cross-section dependency problem in both nonlinear and linear panel regression models.

The rest of this paper discusses the relevant data for the study, the hypotheses, the theoretical and empirical literature, and the methodology adopted. The latter part of the paper presents and discusses the results and develops the conclusions based on the results.

2. Literature Review

This section presents the theoretical and empirical literature review.

2.1. Theoretical Literature Review

Theoretically, capital and trade flows can behave as either substitutes or complements. The negative linkage between capital and trade flows implies that the two flows behave as substitutes, while the positive correlation implies that they are complements (Belke & Domnick, 2018). The Heckscher-Ohlin paradigm and the analysis of Mundell (1957) provide the basic framework in the study of the interaction between capital mobility and trade integration. With the assumptions of the two-factor framework and two goods in the classical Heckscher-Ohlin-Mundell (HOM), free trade leads to factor price equalisation with the rest of the world; as a result, international capital mobility becomes irrelevant. For example, considering two countries with different capital endowments and assuming both countries have the possibility of trading with each other, there would be no need for capital to flow from the capital-abundant to the capital-scarce country, because trade alone automatically eliminates the differences in the rate of return to capital. Consequently, capital and international trade flows behave as substitutes, implying that the increase in trade integration reduces the incentive for capital to flow. However, there are several studies with more realistic features that have modified the basic HOM framework; their contributions range from production uncertainty (Helpman & Razin, 1978) to technological differences (Jones, 1967; Kemp, 1966). With these modifications, capital flows and international trade can be complements, with causality running from capital flows to trade. For example, Markusen (1983) concluded that capital flows and international trade generally behave as complements and that the results

of Mundell (1957) are not the rule, but rather just exceptions. Markusen reached this conclusion by comparing several modelling assumptions and several models.

Furthermore, recent theoretical models examine the trade-finance nexus by attempting "to incorporate the financial frictions into macroeconomic dynamics. For example, Antras and Caballero (2009) argued that, in the current world with heterogeneous financial development, trade and capital mobility behave as complements. This means that the higher the trade integration, the higher the capital flows into the capital-scarce country. Therefore, trade integration increases the return to capital and therefore acts as an incentive for capital flow into less financially developed countries; furthermore, causality is from trade to international capital flows, as opposed to earlier research. Another modelling approach by Furusawa and Yanagawa (2013) used the degree of maturity of financial institutions to investigate the trade and capital flow nexus. Trade and capital flows behave more as complements as financial institutions become less developed. Moreover, Rose and Spiegel (2002) present a theoretical model based on the fear that when a country defaults on servicing external debts, it might end up decreasing trade flows, and therefore the closer the financial ties of the countries, the more trade tends to be conducted between them. In addition, Hahm and Shin (2009) claim that information asymmetries and transaction costs respectively lead to the complementarity of financial assets and trade in goods. Likewise, the hypothesis by Kojima and Ozawa (1984) differentiates the two situations where foreign investment can work as substitutes or complements to international trade. The first case is that foreign investment and international trade are complements when foreign investment is taken under "the principle of complementing comparative advantage patterns" or "the principle of foreign investment originating in the marginal (including sub-marginal) industry". Simply put, this means that foreign investment should originate from the comparatively disadvantaged (marginal) activity or industry in investing countries, which, to a host country, is a comparatively advantaged activity or industry. The second situation or case is when foreign investment is assumed to come from the comparative advantage of the investing country's activity or industry. In this case, foreign investment and international trade become substitutes.

Generally, theoretical literature does not identify a certain unique direction of causality between capital and trade flows, implying that capital and trade flows have the possibility of jointly influencing each other. This motivates the significance of conducting the empirical study to test the exact trade-finance nexus.

2.1. Empirical Literature Review

The literature review on the foreign capital inflows and trade openness nexus allows the researcher to retrace the historical development of the research topic. Furthermore, it helps to expand the knowledge of the two variables. The trade and capital flows nexus has been directly or indirectly examined in various previous empirical studies. However, a conclusion or consensus about the causality and nature of the association between the two flows in the previous studies has not been reached. The following are some of the extant works that investigate the causal relationship between trade and capital flows in different regions of the world.

Gavrila-Paven, I. (2026) examined the relationship between trade integration and economic performance in selected African economies using a fixed-effects model over the period 2000–2024. The analysis finds a positive within-country association between trade integration and FDI inflows. Broner et al. (2013) investigated the impact of trade openness on gross capital inflows in general and its behaviour over the business cycle. They found a positive association between openness to trade and capital inflows. Meanwhile, Abrego (1999) found that trade openness causes capital outflows in Costa Rica and OECD countries, which lowers tax revenue and reduces the welfare of the people. The study was conducted for the period of 1990 to 1991 using an applied general equilibrium model. Furthermore, Kalemli-Ozcan, Nikolsko-Rzhevskyy, and Kwak (2010) examined the causal links between trade and capital flows from 1859 to 1913. Using data on capital inflows and trade in Turkey, the study recorded a significant complementarity between capital and trade flows. ALI, A. (2026) investigated the relationship between international trade and foreign direct investment in major OECD countries using gravity modelling and regression analysis. The findings showed that FDI and international trade act as complements rather than substitutes, where capital presence abroad stimulates home and host-country exports and intermediate goods flows. Furthermore, using the fixed effects (FE) model, Aizenman and Noy (2008) observed a positive

link between FDI inflows and openness to trade in all non-developed countries. Additionally, using the Granger causality test, the results showed a bidirectional causality between the two flows from 1982 to 1998. This result implies that FDI and trade flows affect each other simultaneously. A study by Sazali et al. (2024) also examined the causal relationship between FDI and trade openness in Malaysia. By employing the Granger causality test from 1977 to 2015, the results demonstrated the presence of unidirectional causality running from trade openness to FDI. Furthermore, Miao et al. (2020) observed the complementary relationship between FDI and trade in Africa. The research was conducted using data from the World Bank Comparative Advantage Index and United Nations Commodity Trade Statistics (UN Comtrade) to compare the intensity of trade between China and selected African countries. Another finding of the study is that Africa has a competitive advantage in exporting to China and to the rest of the world. Similar findings were obtained by Sawkutet et al. (2009) after analysing a sample of 20 African countries for a period of 16 years from 1990 to 2005.

In a related study, Sharma and Kaur (2013) analysed the causal association between international trade and FDI in China and India using secondary annual data from 1976 to 2011. The study employed Granger causality tests and found a unidirectional causality from FDI to exports and FDI to imports in China. For the case of India, the study observed bidirectional causality between FDI and exports as well as FDI and imports. Moreover, Ali et al. (2024) observed a negative association between FDI and trade openness in Pakistan. The study employed the autoregressive distributed lag model (ARDL), using annual data from 1975 to 2017. Moreover, Cantah et al. (2018) adopted principal component analysis to examine the link between net FDI and trade openness in **S**SA from 1990 to 2017. The study revealed a positive link between FDI and trade openness. Rahman (2011) also examined the nexus between FDI and trade openness for the case of Bangladesh for the period 1970 to 2007 by employing a multivariate Granger causality model (VAR). The study found that FDI causes imports in Bangladesh, while no causality was revealed in any other variables. Using the gravity model, Chaisrisawatsuk and Chaisrisawatsuk (2007) also found a complementary link between international trade and FDI, either measured by exports or imports, in 29 OECD and six (6) Asian countries covering the period from 1980 to 2004. Donghui et al. (2018) also analysed the link between trade openness and FDI flows in Asian countries using fixed effects and ordinary least squares (OLS) regression models for the period from 1990 to 2017. The study found that trade openness had a positive impact on FDI inflows in Asian countries both in the short run and long-run periods. Similarly, Klasra (2011) established a bidirectional causal link between FDI and exports for Turkey, both in the long and short run. This study also found exports and FDI to be complementary using the ARDL technique. Moreover, a long-run equilibrium negative association between FDI inflows and openness to trade was also observed in the Sudanese economy. This result was obtained after a study that was undertaken by Tahmad and Adow (2018) using the Johansson co-integration technique for the period 1990 to 2020. Nyauncho and Kabiru (2015) found a positive impact of FDI on trade openness on a sample of 30 African countries using fixed effects regression for the period from 1980 to 2012. Singhania and Saini (2018) conducted a study in 19 developed and developing countries for 10 years (2004-2013) using the GMM, RE, and FE approaches. The results showed a significant positive link between FDI and trade openness as well as between trade openness and FPI.

Furthermore, Liargovas and Skandalis (2012) analysed the role of trade openness on FDI inflows in developing economies, including only four countries from Africa, i.e., Kenya, South Africa, Tunisia and Morocco. In their study, eight (8) different indicators of trade openness were used in the analysis by employing the fixed-effect regression model. The results of the study showed that five out of eight indicators of trade openness had a significant positive impact on FDI inflows and two were positive but not significant, while one indicator was negative and insignificant. Moreover, Asghar (2016) investigated the nexus between FDI and openness to trade in seven South Asian economies for 12 years from 1998 to 2010 using the random effects (RE) model. The study used three indicators of trade openness, namely exports, imports, and a combination of both exports and imports. The results suggested that trade openness and FDI were positively and significantly related. By using annual panel data for 25 years from 1977 to 2009, Seyoum et al. (2014) also examined the Granger causality between FDI and trade openness in 25 SSA economies. The empirical results showed a bidirectional causality between trade openness and FDI in SSA economies. Caudros et al. (2004) used a “vector autoregressive model to analyse the causal relationship

among inward FDI, trade and output in Mexico, Brazil and Argentina. The study revealed that trade and FDI complement each other in Mexico, with causality running from FDI to exports, while in Brazil, there was a substitutive association between trade and FDI. In contrast, no causal relationship was exhibited in Argentina. By employing the ARDL for the period from 1990 to 2013, Tsaurai (2015) found that FDI and trade openness had no long-run relationship in Zimbabwe. Koojaroenprasit (2013) observed that trade and FDI flows are substitutes, meaning that the two flows are negatively related. This result was obtained after analysing the nexus between FDI and trade in Australia for the period from 1986 to 2011 using the ordinary least squares (OLS) method. Rahman (2011) employed a VAR model and the Johansen (1991) LR test of cointegration to investigate the nature of the "relationship between foreign investment and trade openness in Bangladesh, making use of data from 1972 to 2007. The results showed that foreign investment"causes import while no causality relationship was found between any other variables of interest. Additionally, a positive cointegration relationship was observed between the variables. Kandiero and Chitiga (2006) conducted a study on trade openness and foreign direct investment in Africa for four different periods (1980-1985, 1985-1990, 1990-1995 and 1995-2001) using the generalised method of moments (GMM). The research found that the ratio of FDI to GDP responds positively to openness to trade. Furthermore, using data from 1986 to 2016 in Nigeria, Nwosa and Adeleke (2017) employed an E-GARCH approach and observed a positive significant association between FDI and trade openness, while trade openness and foreign portfolio investment showed an insignificant relationship.

On the other hand, using ARDL, Mehar and Hasan (2018) found a significant negative impact of trade openness on portfolio investment in Pakistan from 1981 to 2012. Furthermore, by employing the Granger causality test, the study indicated that trade openness Granger-caused portfolio investment, while portfolio investment does not Granger-cause trade openness. Another study by Ekeocha et al. (2012) demonstrated a positive long-run nexus between foreign portfolio investments and trade openness in Nigeria. The study used the applied finite distributed lag model for quarterly data from 1981 to 2010. In the same vein, Ekeocha (2008) found that foreign portfolio investment is negatively related to trade openness in Nigeria. The study applied an error correction model using quarterly data for the period between 1986 and 2006. Moreover, through the use of GMM for the period 1981 to 2011, Yang et al. (2013) found an insignificant negative impact of trade openness on portfolio investment in both static and dynamic models of Asia. In contrast, a negative and significant effect of trade openness on portfolio investment was revealed for the dynamic model of Latin America. Conversely, the study found an insignificant negative effect of trade openness on portfolio investment in Latin America for the static model. Moreover, the study by Calderon et al. (2019) asserted that capital inflows and trade openness are negatively related in sub-Saharan Africa, although the effect was insignificant. The study applied an instrumental variable approach using the annual data from1980 to 2017. Moreover, Portes and Rey (2005) explored the nexus between capital flows and trade of 14 countries from 1989 to 1996 using contemporary data on financial flows. The results showed that a 10% increase in trade led to a 3.4% increase in equity flows. Indeed, Aviat and Coeurdacier (2007) employed the simultaneous gravity model and found a statistically positive nexus between asset holdings and trade in goods. Using the OLS and IV estimations, the study also indicated that a 10% increase in asset holdings leads to a 2% increase in bilateral trade. In addition, Jean-François and Sèna (2017) analysed the role of international trade in financial flows for economic development in 125 countries, of which 37 were least developed. The results showed that international trade openness plays a chief role in attracting financial flows for development. The study used the three-stage least squares econometric approach for an unbalanced panel dataset.

The foregoing discussion shows that the causal relationship between trade openness and foreign capital inflows is complex and inconclusive, especially in SSA, and therefore there is a need for more empirical investigation using more recent data, new methodology, more extended time periods, and the inclusion of more panel units and diagnostic tests so as to have a thorough analysis. This paper, therefore, uses a panel smooth transition regression (PSTR), which is a non-linear approach to examine the causality between foreign capital inflows and trade openness in SSA.

3. Econometric Framework

The empirical strategy combines nonlinear panel cointegration with a panel smooth transition vector

error-correction model. The PSTR framework is suitable because it allows coefficients to change gradually between regimes rather than assuming an abrupt shift. It can also capture asymmetric responses and different forms of nonlinearity, depending on the transition function and switching variable (Omay et al., 2014).

A key advantage of PSTR is that it models gradual adjustment between regimes and permits the data to determine how the relationship changes. The specification therefore provides a way to examine whether the trade–capital relationship differs across low- and high-activity conditions. The following subsections outline the cointegration test and the regime-dependent causality model.

3.1. Nonlinear Panel Cointegration

Consider the following panel regression model:

$$y_{i,t} = \alpha_i + \beta_i x_{i,t} + \mu_{i,t} \quad \text{for} \quad i = 1, \dots N \; and \; t = 1, \dots 1, \dots T; \tag{1}$$

Where $y_{i,t} \; and \; x_{i,t}$ represent the "observable I(1) variables, $\beta = (\beta_1, \dots, \beta_m)$ are parameters to be estimated; $\mu_{i,t}$ is the error term. $y_{i,t}$ is a scalar, and $x_{i,t} = (x_{1,t}, x_{2,t}, \dots, x_{m,t})$ is an $(m \times 1)$ vector and finally α_i is the fixed effect (heterogeneous intercept). The assumption is that an$(n \times 1)$ vector $Z'_{i,t} \equiv (y_{i,t}, x'_{i,t})$ is generated as $z_{i,t} = z_{t-1} + \varepsilon_{i,t}$, where $\varepsilon_{i,t}$ are i.i.d. with mean zero, positive definite variance-covariance matrix$\sum$, and $\mathrm{E}\left|\varepsilon_{i,t}\right|^s < \infty$ for some"$s > 4$.

If "the error term $\mu_{i,t}$ in regression (1) is stationary, then the vector $z_{i.t}$ is said to be co-integrated, and $\mu_{i,t}$ is called the equilibrium error (Engle & Granger, 1987). Here, the assumption is that $\mu_{i,t}$ can be modelled using the following nonlinear model:

$$\mu_{i,t} = \gamma_i \mu_{i,t-1} + \varphi_i \mu_{i,t-1} F\left(\mu_{i,t}; \theta_i\right) + \xi_{i,t} \tag{2}$$

Where"$\xi_{i,t}$ is a zero-mean error and $F(\mu_{i,t}; \theta_i)$ is a smooth transition function of $\mu_{i,t-1}$. By imposing $F(\mu_{i,t}; \theta_i) = 0$ or $F\left(\mu_{i,t}; \theta_i\right) = -\gamma_i \mu'_i$ where μ'_i is a vector of level parameters, one obtains the conventional linear cointegration equation (Kapetanois et al., 2006a). Following previous literature (see Kapetanois et al., 2003, 2006; Ucar & Omay, 2009; Maki, 2010) on nonlinear cointegration, the assumption is that the transition function $F(\mu_{i,t}; \theta_i)$ is of the exponential form[1]:

$$F\left(\mu_{i,t-1}; \theta_i\right) = 1 - exp\left[-\theta_i \mu^2_{i,t-1}\right] \tag{3}$$

It is "further assumed that $\mu_{i,t}$ is a mean-zero stochastic process and that $\theta_i \geq 0$. The transition function $F(\mu_{i,t-1}; \theta_i)$ is bounded between zero and one; and is symmetrically U-shaped around zero. The parameter $\theta_i$determines the speed of the transition between the two extreme values of the transition function $F(\mu_{i,t-1}, \theta_i)$. The exponential transition function enables adjustment to the long-run equilibrium based on the size of the disequilibrium.

Substituting "(3) into (2) and re-parameterising the resultant equation, the following regression model is obtained:

$$\Delta\mu_{i,t} = \mu_{i,t-1} + \varphi_i \mu_{i,t-1}\left[1 - exp\left\{-\theta_i \mu^2_{i,t-1}\right\}\right] + \varepsilon_{i,t} \tag{4}$$

If $\theta_i > 0$, then it controls the rate of the reversion of the mean. If $\varphi_i \geq 0$, then the process may show signs of unit root or explosive behaviour for small values $\mu^2_{i,t-1}$. However, if the variations from the equilibrium are substantially large, for values of $\mu^2_{i,t-1}$, it has constant dynamics and, as a result, is geometrically ergodic provided that"$\varphi_i + \psi_i < 0$[3].

Imposing $\varphi_i = 0$, "implying that $\mu_{i,t}$ follows a unit root process in the middle regime, and further permitting possible serial correlation of the error term in (4), the following regression is obtained:

$$\Delta\mu_{i,t} = \psi_i \mu_{i,t-1}\left[1 - exp\left\{-\theta_i \mu^2_{i,ti-1}\right\}\right] + \sum_{j=1}^{p} p_{ij} \, \Delta\mu_{i,t-j} + \varepsilon_{i,t} \tag{5}$$

The test of cointegration can be based on the specific parameter θ_i, where it is zero under the null hypothesis of no cointegration, and positive under the alternative hypothesis. Testing for the null hypothesis directly is not possible because ψ_i is not known under the null. To address this challenge, one may follow Luukkonenet et al. (1988) by replacing the transition function $F\left(\mu_{i,t}; \theta_i\right) = 1 - exp\left\{-\theta_i \mu^2_{i,t-1}\right\}$ with its first-order Taylor approximation under the null, which gives rise to the following regression model.

$$\Delta\mu_{i,t} = \delta_i \mu^3_{i,t-1} + \sum_{j=1}^{P_i} \rho_{ij} \, \Delta\mu_{i,t-j} + e_{i,t} \tag{6}$$

where $e_{i,t}$"comprises the original shocks $\varepsilon_{i,j}$ in equation (5) as well as the error term resulting from the Taylor estimate. Here, the study allows for various lag orders p_i for each variable in the regression equation (6). Consequently, the null hypothesis of no cointegration and the alternative can be stated as follows.

H_0: $\delta_i = 0$, for all i, (no cointegration)

H_o: $\delta_i < 0$, for some i, (non-linear cointegration)

In empirical applications, the number of augmentation terms in the auxiliary regression (6) is chosen using any convenient lag selection method. In line with Ucar and Omay (2009), the cointegration test can be constructed by standardising the average of individual cointegration test statistics across the whole panel. The cointegration test for the *i*-th individual is the t-statistic for testing $\delta_i = 0$ (see Kapetanois et al., 2003a; Ucar & Omay, 2009) in equation (6) defined by:

$$t_{i,NL} = \frac{\Delta\mu_i' M_t \mu_{i,-1}^3}{\hat{G}_{i,NL}(\mu_{i,-1}' M_t \mu_{i,-1})^{3/2}} \quad (7)$$

Where $\hat{G}_{i,NL}^2$ is the consistent estimator such that $\hat{G}_{i,NL}^2 = \frac{\Delta\mu_i' M_t \mu_i}{T-1}$, $M_t = I_T - \tau_T(\tau_T'\tau_T)^{-1}\tau_T'$with $\Delta\mu_i = \left(\Delta\mu_{i-1,}\Delta\mu_{i-2}, \ldots \Delta\mu_{i-T}\right)'and\tau_T = (1,1, \ldots 1)$. Furthermore, when the invariance property and the existence of moments are satisfied, the usual normalisation of $\bar{t}_{NL}$ statistic is obtained as follows:

$$\overline{K}_{NL} = \frac{\sqrt{N}(\bar{t}_{NL} - E(t_{i,NL}))}{\sqrt{var(t_{i,NL})}} \quad (8)$$

where $\overline{t}_{NL} = N^{-1}\sum_{i=1}^{N} t_{NL}$,and $E\left(t_{i,NL}\right)$ and var $(t_{i,NL})$ are the expected value and variance of the $t_{i,NL}$ statistic given in equation (7).

Cross-sectional dependence can arise when countries are exposed to common global shocks, regional spillovers, or correlated omitted factors. Ignoring such dependence may distort conventional panel unit-root and cointegration tests (Banerjee et al., 2004). The study therefore tests for cross-sectional dependence using Pesaran's (2004) CD statistic and applies a sieve-bootstrap procedure, following Ucar and Omay (2009), to improve inference.

$$\text{CD} = \sqrt{\frac{2T}{N(N-1)}}\left[\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\hat{\rho}_{ij}\right] \quad (9)$$

Where $\hat{\rho}_{ij}$"is the estimated correlation coefficient between error terms for the individuals *i* and *j*. The sieve bootstrap method is employed to deal with the cross-sectional dependence problem in line with Ucar and Omay (2009). Once cointegration and long-run relationships are established, the panel error correction model can be estimated. Omay et al. (2014) is followed and the PSTRVEC model to examine regime-wise interactions between foreign capital inflow and trade openness in sub-Saharan Africa. This is justified by the fact that not only the adjustment to the long-run equilibrium level, but also the dynamic interrelationship between the variables might be nonlinear.

3.2. Specification of PSTRVEC Model

Following Gonzalez et al. (2005), Omay and Kan (2010), and Omay et al. (2014), the study specifies a panel smooth transition vector error-correction model (PSTRVEC) to estimate how the interactions between foreign capital inflows and trade openness vary across regimes.

$$\begin{aligned}\Delta to_{i,t} = {} & \mu_1 + \beta_1 ec_{i,t-1} + \textstyle\sum_{j=1}^{p_i}\theta_{1j}\,\Delta to_{i,t-j} + \\ & \textstyle\sum_{j=i}^{q_1}\vartheta_{ij}\,\Delta fci_{i,t-j} + G(S_{it};\gamma,c)\left[\tilde{\beta}_1 ec_{it-1} + \right. \\ & \left. \textstyle\sum_{j=1}^{p_i}\tilde{\theta}_{1j}\Delta to_{it-j} + \sum_{j=1}^{q_i}\tilde{\vartheta}_{1j}\Delta fci_{it-j}\right] + \varepsilon_{1it}\end{aligned} \quad (10)$$

$$\begin{aligned}\Delta fci_{it} = {} & \mu_2 + \beta_2 ec_{it-1} + \sum_{j=1}^{r_i}\theta_{2j}\,\Delta to_{it-j} + \sum_{j=i}^{s_1}\vartheta_{2j}\,\Delta fci_{it-j} \\ & + G(S_{it};\gamma,c)\left[\tilde{\beta}_2 ec_{it-1}\right. \\ & \left. + \sum_{j=1}^{r_i}\tilde{\theta}_{2j}\Delta to_{it-j} + \sum_{j=1}^{s_i}\tilde{\vartheta}_{2j}\Delta fci_{it-j}\right] + \varepsilon_{2it}\end{aligned}$$

Here, i indexes countries and t indexes years. Trade openness (TO) and foreign capital inflows (FCI) are measured as percentages of GDP. The specification includes country-specific effects, the estimated error-correction term, and disturbances that may be contemporaneously correlated across countries.

i denotes 1, …, N, and *t* = 1, …, T, where *N* and *T* represent the cross-section and time dimensions of the respective panels; to_{it} represents trade openness and fci_{it} is foreign capital inflows. Both to_{it} and fci_{it} are stated as a percentage of GDP. In addition, μ_i is the fixed individual effects, ec_{it} symbolises the estimate of the "error correction term from the regression (1) (i.e., $ec_{it} = \hat{v}_{it}$ from equation (1)), and ε_{it} denotes the random error term implicit as the martingale difference of the history of the vector $Z_{it}' \equiv (FCI_{it}, TO_{it})'$ up to time $t-1$, that is, E $\left[\varepsilon_{it}\mathrm{I}z_{it-1,}z_{it-2,}\ldots,z_{it-p,\ldots}\right] = 0$, and that the conditional variance of the error term is constant, that is E $\left[\varepsilon_t^2\mathrm{I}z_{it-1,}z_{it-2,}\ldots,z_{it-p,\ldots}\right] = \sigma_i^2$."

The model allows the disturbances in the country-specific equations to be correlated at a given point in time.

The transition mechanism is represented by a logistic function. Its location parameter identifies the threshold around which the regime changes, while its slope parameter determines how quickly the transition

occurs. A gradual transition is consistent with the PSTR framework; in limiting cases, the model can approximate a threshold specification or a linear panel model.

$$G(s_{i,t};\gamma,c) = \left[1+exp\left(-\gamma\prod_{j=1}^{m}(s_{it}-c_j)\right)\right]^{-1} with\ \gamma>0\ and\ c_m\geq\cdots\geq c_1\geq c_0; \quad (11)$$

Where $c=(c_1,\ldots,c_m)'$"is an m-dimensional vector of location parameters, and the slope parameter γdenotes the smoothness of the transition between the regimes. A value of 1 or 2 for m, often meets the common types of variation. In cases where $m=1$, for instance, taking a first-order logistic transition function, the extreme regimes correspond to low and high values of s_{it}, and its coefficients in regression model (10) change smoothly from $\beta_j,\theta_j\ and\ \vartheta_j to \beta_j+\tilde{\beta}_j,\theta_j+\tilde{\theta}_j\ and\ \vartheta_j+\tilde{\vartheta}_j$, respectively, as s_{it} increases. When $\gamma\to\infty$, the first-order logistic transition function $G(s_{i,t};\gamma,c)$ becomes an indicator function I$[A]$, which takes a value of 1 when event A occurs and 0 otherwise. Therefore, the PSTR model is reduced to Hansen's (1999) two-regime threshold model.

When $m=2, G\left(s_{i,t};\gamma,c\right)$"takes a value of 1for both low and high s_{it}, minimising at $\left(\frac{c_1+c_2}{2}\right)$. In that case, if $\gamma\to\infty$, the PSTR model is reduced to a panel three-regime threshold regression model. If $\gamma\to 0$, the transition function $G(s_{i,t};\gamma,c)$ will reduce to a constant, and therefore, the PSTR model will collapse to a linear panel regression for any value of"m.

The empirical specification follows three stages: first, estimate an appropriate linear panel model; second, test linearity against a smooth-transition alternative and identify the transition variable and function; and third, estimate the selected PSTRVEC model (Gonzalez et al., 2005; Omay et al., 2014).

1. State the suitable linear panel model for the data being investigated.
2. Test the null hypothesis of linearity against the alternative of smooth transition type nonlinearity. If linearity is rejected, select the appropriate transition variable $s_{i,t}$and the form of the transition function.
3. Estimate "the parameters in the chosen PSTRVEC model."

Testing linearity is complicated by nuisance parameters that are not identified under the null hypothesis (Hansen, 1999; Omay et al., 2014). Following Luukkonen et al. (1988), the transition function is approximated by a Taylor expansion, allowing the null of linearity to be tested through an auxiliary regression and an LM-type statistic.

$$\Delta z_{i,t} = \curlywedge_i + \pi_0' ec_{i,t-1} + \sum_{j=1}^{p_i}\psi_{0j}\Delta z_{i,t-j} + \sum_{h=1}^{k}\tilde{\pi}_h' s_{i,t}^h ec_{it} + \sum_{h=1}^{k}\sum_{j=1}^{p_i}\tilde{\phi}_{hj} s_{i,t}^h \Delta z_{i,t-j} + e_{it} \quad (12)$$

The auxiliary-regression terms are functions of the transition parameters and include both the original disturbances and the approximation error. The test compares the restricted and unrestricted residual sums of squares. Candidate transition variables are assessed, and the variable associated with the strongest rejection of linearity is selected.

$$LM = \frac{(SSR_0-SSR_1)/kp}{SSR_0/(TN-N-k(p+1))} \sim F\left(kp, TN-N-k(p+1)\right) \quad (13)$$

Where $SSR_0\ and\ SSR_1$"are the sums of squared residuals based on the null and alternative hypotheses, respectively. The LM statistics can be estimated for several candidates in order to choose an appropriate threshold variable S_{it} and the one for which the p-value of the statistic is the smallest can be selected.

The order and form of the transition function are selected using the sequence of tests proposed by Terasvirta (1994). The relative p-values of the F-tests guide the choice between the logistic and exponential forms.

3.3. Estimation of PSTRVEC Models and Regime-Wise Granger-Causality Tests

Once the transition variable and function have been selected, the PSTRVEC parameters are estimated by nonlinear least squares. Starting values are obtained through a two-dimensional grid search over the threshold and slope parameters, selecting the combination that minimises the panel sum of squared residuals.

To account for contemporaneous cross-country dependence, the country equations are estimated jointly using an iterative nonlinear generalised least-squares procedure. This approach permits correlated disturbances across countries and yields maximum-likelihood estimates under the maintained specification (Greene, 1997; Omay et al., 2014).

After estimating the PSTRVEC model, Granger-causality tests are conducted to identify whether past changes in

one variable help predict the other, and whether these effects differ by regime.

Causality is assessed separately in each regime. Short-run causality is evaluated using the lagged explanatory-variable terms, while long-run causality is assessed through the error-correction term. A joint test of these terms provides an overall measure of short- and long-run predictive influence (Li, 2006; Omay et al., 2014). Where the transition variable is trade openness, the two regimes represent relatively low and high levels of trade openness.

Table 1. Tests of Granger causality

Foreign capital inflow does not Granger-cause trade openness in low growth periods (i.e., when the foreign capital inflow rate is less than some threshold value) in the short run	$H_o: \delta_1 = 0$
Foreign capital inflow does not Granger-cause trade openness in low growth periods (i.e., when the foreign capital inflow rate is less than some threshold value) in the long run	$H_0: \alpha_1 = 0$ and/or $H_0: \alpha_1 = \delta_1 = 0$
Foreign capital inflow does not Granger-cause trade openness in high growth periods (i.e., when the foreign capital inflow rate is greater than some threshold value) in the short run	$H_0: \delta_1 = \hat{\delta}_1 = 0$
Foreign capital inflow does not Granger-cause the trade openness rate in high growth periods (i.e., when the foreign capital inflow rate is greater than some threshold value) in the long run	$H_0: \alpha_1 = \hat{\alpha}_1 = 0$ and/or $H_0: \alpha_1 = \hat{\alpha}_1 = \delta_1 = \hat{\delta}_1 = 0$
Trade openness does not Granger-cause foreign capital inflow in low growth periods (i.e., when the foreign capital inflow rate is less than some threshold value) in the short run	$H_0: \varphi_2 = 0$
Trade openness does not Granger-cause foreign capital inflow in low growth periods (i.e., when the foreign capital inflow rate is less than some threshold value) in the long run	$H_0: \alpha_2 = 0$ and/or $H_0: \alpha_2 = \varphi_1 = 0$
Trade openness does not Granger-cause foreign capital inflow in high growth periods (i.e., when the foreign capital inflow rate is greater than some threshold value) in the short run	$H_0: \varphi_2 = \hat{\varphi}_2 = 0$
Trade openness does not Granger-cause foreign capital inflow in high-growth periods (i.e., when the foreign capital inflow rate is greater than some threshold value) in the long run	$H_o: \alpha_2 = \hat{\alpha}_2 = 0$ and/or $H_0: \alpha_2 = \hat{\alpha}_2 = \varphi_2 = \hat{\varphi}_2 = 0$

Source: Author's adaptation from Omay et al. (2014)

3.4. Data

The analysis uses annual panel data for 31 SSA countries from 1985 to 2023. Data are drawn from the World Bank's World Development Indicators, UNCTAD, and IMF databases. Foreign capital inflows (FCI) are measured as total inflows and trade openness (TO) as total trade; both are expressed as percentages of GDP. The underlying series are reported in real US-dollar terms where applicable.

4. Empirical Results

The test for the presence of a unit root was carried out based on the linear IPS test of Im et al. (2003) and the nonlinear UO test of Ucar and Omay (2009). The results are provided below.

Table 2. The unit root test

Variables	IPS test				UO test			
	Intercept		Intercept & trend		Intercept		Intercept & trend	
	W-stat	t-stat	W-stat	t-stat	W-stat	t-stat	W-stat	t-stat
FCI	-8.772	-2.932	-7.770	-3.327	-1.801***	-0.005***	-2.002***	-0.011***
TO	-4.049	-2.145	-2.860	-2.532	-2.203	-0.606	-1.977***	-2.457***

Note: "The (***, **, *) denote rejection of the null hypothesis that states the presence of a unit root at 1%, 5%, and 10% levels of significance using bootstrap p-values. A maximum of 7 lags is used"

The unit-root results differ across specifications. The linear IPS tests and the nonlinear intercept specification for TO generally fail to reject the unit-root null, whereas the nonlinear intercept-and-trend specification rejects it for TO. This variation across tests reinforces the need to allow for nonlinear behaviour in the panel.

$$FCI_{i,t} = \ 0.0637TO_{i,t}$$

$$(9.2594)$$

$$TO_{i,t} = \ 1.2183FCI_{i,t}$$

$$(9.2594)$$

The fixed-effects estimates show positive and statistically significant coefficients for FCI and TO in both specifications. The residuals are then used in the cointegration analysis. Pesaran's CD test yields a statistic of 18.5195 ($p < 0.0001$), indicating cross-sectional dependence among the sample countries.

Table 3. Linear and nonlinear cointegration tests

Model	Linear test		Nonlinear test	
	W-stat	t-stat	W-stat	t-stat
$\mu_{i,t} = FCI_{i,t} - 0.0637TO_{i,t}$	-0.266	-0.912	0.977**	-0.721**
$\mu_{i,t} = TO_{i,t} - 1.2183FCI_{i,t}$	0.244	0.616	0.987**	2.056**

** denotes significance at the 5% level using bootstrapped p-values. A maximum of 4 lags with 800 bootstraps were used.

The linear specifications do not reject the null of no cointegration. By contrast, the nonlinear tests reject that null in both model specifications at the 5% level. This provides evidence of a nonlinear long-run relationship and supports estimating the PSTRVEC model.

$$\Delta \mathrm{FCI}_{i,t} = \mu_1 - 0.3311ect_{i,t-1} - 0.2560\Delta \mathrm{FCI}_{i,t-1} - 0.0719\Delta \mathrm{TO}_{i,t-1}$$

(-8.8478) (-7.2230) (-2.5932)

$$\Delta \mathrm{TO}_{i,t} = \mu_2 - 0.0118ect_{i,t-1} - 0.1354\Delta \mathrm{TO}_{i,t-1} - 0.010\Delta \mathrm{FCI}_{i,t-1}$$

(-0.2756) (-4.2866) (-0.2756)

The error-correction coefficients are negative and statistically significant in both equations, indicating adjustment towards long-run equilibrium. In the short run, trade openness has a negative and significant association with foreign capital inflows ($p = 0.0277$), while foreign capital inflows are negatively associated with trade openness ($p = 0.0403$). The next step is to test whether these linear dynamics are adequate. The tests consider TO, FCI, and the error-correction term as possible transition variables; the variable that most strongly rejects linearity is preferred.

Table 4. Linearity tests

Model	Transition variables		
	$\Delta \mathrm{TO}_{i,t-2}$	$\Delta \mathrm{FCI}_{i,t-2}$	$\mathrm{ect}_{i,t-2}$
	F-test	F-test	F-test
FCI	0.328***	2.313	3.678**
TO	1.911***	1.745*	2.677***
Test of linearity against a PSTR specification; ***, **, * denote significance at 1%, 5%, and 10% levels, respectively.			

Linearity is rejected for the FCI equation when TO is the transition variable (1% level) and when the error-correction term is used (5% level), but not when FCI itself is used. For the TO equation, all three candidate variables reject linearity: FCI at 10%, TO at 1%, and the error-correction term at 5%.

The strongest rejection of linearity occurs when TO is used as the transition variable. TO is therefore selected for the PSTRVEC specification. The subsequent sequence of F-tests is used to determine the appropriate transition-function form.

Table 5. Choice of transition

	F_1	F_2	F_3
	F-test	F-test	F-test
FCI	0.328***	0.982*	1.361*
TO	1.911***	1.022**	1.364*
***, **, * denote the significance at 1%, 5%, and 10% levels, respectively			

The sequence of tests supports the selected transition specification. F2 rejects linearity at the 10% level in the FCI equation and at the 5% level in the TO equation. The results are interpreted as favouring a first-order logistic transition function, which implies one threshold and two regimes.

The PSTRVEC model is estimated using iterative nonlinear GLS to accommodate cross-sectional dependence. Initial values for the threshold (c) and transition slope are selected through a grid search to improve numerical optimisation.

$$\Delta \mathrm{FCI}_{i,t} = \mu_1 - 0.516 ect_{i,t-1} - 0.564 \Delta \mathrm{FCI}_{i,t-1} + 0.345 \Delta \mathrm{TO}_{i,t-1}$$

(-9.304) (-2.481) (12.264)

$$+\left[0.319 ect_{i,t-1} + 0.486 \Delta FCI_{i,t-1} - 0.384 \Delta TO_{i,t-1}\right] . G(s_{i,t-1}; \gamma, c),$$

(0.000) (1.704) (0.000)

$$\Delta \mathrm{TO}_{i,t} = \mu_2 + 0.020 ect_{i,t-1} - 0.944 \Delta \mathrm{FCI}_{i,t-1} + 0.899 \Delta \mathrm{TO}_{i,t-1}$$

(0.066) (-0.873) (5.800)

$$+\left[-0.041 ect_{i,t-1} + 1.779 \Delta FCI_{i,t-1} - 1.232 \Delta TO_{i,t-1}\right] . G(s_{i,t-1}; \gamma, c)$$

(-0.160) (1.499) (0.000)

Where

$$G(\Delta \mathrm{FCI}_{i,t}; \gamma, c) = \left(1 + exp\left(-1.4454\left(\Delta \mathrm{TO}_{i,t-1} - 0.0376\right)\right)\right)^{-1}$$

(2.484) (0.007)

The estimated transition parameter is statistically significant, indicating gradual movement between the two regimes. The threshold is close to zero, so the regimes can be interpreted as periods of relatively low and high trade openness. In the FCI equation, the error-correction coefficient in the low-openness regime is −0.517 and statistically significant, suggesting that foreign capital inflows adjust towards long-run equilibrium after a shock. The lagged TO coefficient is 0.345 and significant at 1%, indicating a positive short-run association in this regime. In the high-openness regime, the error-correction coefficient is −0.197 and significant, while the net coefficient on lagged TO is −0.039 (0.345 − 0.384). Thus, the positive association observed in the low regime weakens and becomes slightly negative in the high regime.

In the TO equation, the error-correction coefficients are small and statistically insignificant in both regimes (−0.022 in the low regime and −0.021 in the high regime). This suggests that trade openness does not adjust significantly to deviations from the estimated long-run relationship. The coefficients on foreign capital inflows are also statistically insignificant in both regimes; therefore, the estimates do not provide evidence that changes in FCI affect TO within either regime.

Regime-specific Granger-causality tests use lagged explanatory variables to assess short-run effects and the error-correction term to assess long-run effects. The joint test evaluates the combined contribution of both components. Results for the low- and high-openness regimes are reported in Table 6.

Table 6. Regime-wise causality test

Source of causation	Dependent variables			
	ΔFCI		ΔTO	
	Low growth regime	High growth regime	Low growth regime	High growth regime
Short-run				
ΔTO	0.76	150.40***		
ΔFCI			150.40***	1.80
Long-run				
ECT	86.57***	86.57***	0.004	0.02
Joint test (short- and long-run)				
ECT /ΔTO	46.81***	147.60***		
ECT /ΔFCI			105.88***	0.91
*** denotes the significance at the 1% level				

The short-run results show that TO Granger-causes FCI in the high-growth regime but not in the low-growth regime. Conversely, FCI Granger-causes TO in the low-growth regime, with no corresponding evidence in the high-growth regime. In the long run, TO Granger-causes FCI in both regimes, whereas FCI does not Granger-cause TO in either regime. The joint tests confirm that TO predicts FCI in both regimes, while FCI predicts TO only in the low-growth regime. These findings indicate that the direction of predictive influence depends on both the time horizon and the prevailing regime.

5. Conclusion

This study investigated the relationship between foreign capital inflows and trade openness in 31 SSA countries over 1985–2023. Using nonlinear panel cointegration and a PSTRVEC model, it examined long-run association and regime-specific causality while accounting for cross-sectional dependence. The nonlinear cointegration results support a long-run relationship between the variables, although this relationship is not detected by the linear tests. The PSTRVEC estimates further show that the dynamics differ between low- and high-openness regimes. This

confirms the value of allowing for gradual regime changes rather than imposing constant linear effects.

The causality findings vary by horizon and regime. In the short run, TO predicts FCI in the high-growth regime, while FCI predicts TO in the low-growth regime. In the long run, TO predicts FCI in both regimes, but there is no evidence that FCI predicts TO. Joint tests show that TO Granger-causes FCI in both regimes, whereas FCI Granger-causes TO only in the low-growth regime. The results have methodological and policy implications. Researchers examining trade and capital movements in SSA should consider nonlinear adjustment and regime dependence, as linear models may overlook important differences across economic conditions. For policymakers, the consistent predictive role of trade openness suggests that improving market access and reducing unnecessary trade barriers may help attract foreign capital.

SSA governments can strengthen this relationship by deepening regional integration, improving the implementation of the African Continental Free Trade Area, and addressing tariff and non-tariff barriers. Such measures should be accompanied by policies that improve infrastructure, institutions and the domestic capacity to convert capital inflows into productive trade. Because the estimated effects vary across regimes, policy design should also recognise that the impact of capital inflows may differ according to prevailing trade conditions.

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