Exchange Rate and Inflation in Nigeria: A Cointegrated VAR Approach

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DOI: https://doi.org/10.51244/IJRSI.2026.1309000241

Received: 05 October 2026; Accepted: 10 October 2026; Published: 22 October 2026

ABSTRACT

This study examines the long-run equilibrium relationship and dynamic interactions between the exchange rate and inflation in Nigeria using a Cointegrated Vector Autoregression (CVAR) estimated in its Vector Error Correction (VECM) representation. Monthly data on the exchange rate (constructed from National Bureau of Statistics annual rate ranges), inflation rate, monetary policy rate, broad money supply and Brent crude oil price for January 2015 to December 2025 (132 observations) are analysed. Augmented Dickey–Fuller and Phillips–Perron tests classify inflation, the policy rate, money supply and the oil price as integrated of order one, but the log exchange rate is more persistent than an I(1) series. The Johansen test with an unrestricted constant does not identify a cointegrating vector on the full sample, although one is identified with a restricted constant and in the pre-unification sample, and Gregory–Hansen and likelihood-ratio tests show that the long-run relation is not stable across the mid-2023 foreign exchange unification. Conditional on one cointegrating vector, the exchange rate bears the burden of adjustment to long-run disequilibrium, inflation is weakly exogenous, and Granger causality runs from inflation to the exchange rate rather than the reverse. Impulse responses and variance decompositions, checked under all 120 Cholesky orderings and with generalised responses, attribute well under one percent of inflation forecast error variance to exchange rate shocks, and a time-varying-parameter error-correction model does not improve on the VECM out of sample. The evidence therefore does not support strong exchange rate pass-through to inflation in this dataset, and the findings are provisional until monthly exchange rate, money supply and policy rate series replace the interpolated ones.

Keywords: Exchange Rate; Inflation; Exchange Rate Pass-Through; Cointegrated VAR; VECM; Nigeria

INTRODUCTION

Exchange rate movements and their pass-through to domestic prices are among the most closely watched macroeconomic relationships in any economy, and have remained a central concern for policymakers and researchers alike (Ebenezer, Kanayo & Andrew, 2022; Obstfeld & Rogoff, 1995; Taylor, 2000). Emerging market economies (EMEs) are engines of global growth, yet they remain uniquely vulnerable to external and internal economic shocks. Among the most persistent and challenging issues they face is the interplay between exchange rate volatility and domestic inflation. Exchange rate volatility, the degree of fluctuation in a currency's value relative to others, often measured by standard deviations or GARCH models, has intensified in many EMEs because of commodity price swings, volatile capital flows ("sudden stops"), shifts in global risk appetite and domestic political uncertainty. Unlike advanced economies with deep financial markets and credible monetary policy frameworks, EMEs often grapple with thin foreign exchange markets, limited policy credibility, structural rigidities and high levels of dollarization. These factors amplify the transmission of exchange rate fluctuations into the domestic price level, a phenomenon known as exchange rate pass-through (ERPT).

Nigeria, like many EMEs, has continued to grapple with the twin challenges of exchange rate volatility and rising inflation. Fluctuations in the value of the naira, driven by volatile global oil prices, unstable foreign capital inflows, poor fiscal discipline, an unreliable tax base and persistent structural weaknesses (Oluwayemisi et al., 2024), have had far-reaching consequences for price levels, production costs, trade dynamics and overall macroeconomic performance (Adeniyi, Omotosho & Akanbi, 2020). In response, the Central Bank of Nigeria (CBN) has pursued a flexible exchange rate regime, a loan-to-deposit ratio policy to boost lending, and an

ISSN No. 2321-2705 | DOI: 10.51244/IJRSI |Volume XIII Issue IX September 2026

infrastructural development fund (Ogundipe & Akinbobola, 2020), yet inflation remains a pressing concern: CBN data recorded inflation of 33.88 percent in October 2024 and 34.60 percent in December 2024, well above the single-digit target and eroding household purchasing power.

The empirical relationship between exchange rate movements and domestic inflation in Nigeria remains theoretically intuitive but statistically contested. While extensive literature has confirmed a positive long-run association between the naira–dollar exchange rate and the Consumer Price Index (Choudhri & Hakura, 2006; Ehikioya, 2019), the existing evidence is characterised by three unresolved statistical gaps that undermine policy reliability.

First, prior studies on Nigeria (e.g., Okoli et al. 2016, using GARCH, and earlier OLS-based studies) predominantly rely on single-equation models that treat the exchange rate as exogenously determined. GARCH-family models remain the appropriate tool for characterising the conditional volatility of the exchange rate itself; however, because GARCH is a single-equation, univariate technique, it cannot recover the long-run equilibrium relationship or the bidirectional short-run feedback between exchange rate and inflation that a system-wide CVAR/VECM framework is designed to capture. The two approaches answer different questions and are complementary rather than competing. Treating the exchange rate as exogenous nevertheless ignores the feedback between exchange rates, interest rates, money supply and oil prices in an open commodity-exporting economy, so the resulting estimates are likely to be biased and inconsistent. The paper that most directly motivates the present study, Okoli et al., models the Nigerian exchange rate–inflation relationship with a single-equation GARCH framework; its principal weakness is that it does not test for cointegration, structural breaks or the direction of causal adjustment, leaving the long-run equilibrium and the speed of adjustment after a shock unexamined.

Second, the Nigerian economy has undergone profound structural regime shifts, most notably the unification of the foreign exchange windows and the transition to a managed float in mid-2023. The extant literature largely employs linear ARDL or basic VECM frameworks that do not account for, or test, deterministic structural breaks, and whether the long-run equilibrium vector has shifted as a result remains unexplored.

Third, while cointegration is frequently tested, prior research rarely proceeds to test weak exogeneity. This is a serious oversight because identifying the long-run relationship (β) is insufficient for policy: policymakers need to know which variable, the exchange rate or inflation, actively adjusts to restore equilibrium after a shock. Without examining the speed of adjustment (the α coefficients), monetary authorities cannot distinguish whether intervening in the foreign exchange market reduces inflation, or whether tightening monetary policy to curb inflation stabilises the exchange rate.

This study therefore addresses these deficiencies by employing a full-system Cointegrated Vector Autoregression (CVAR) framework, with the Johansen (1991) procedure used to establish long-run equilibrium among the exchange rate, inflation, interest rate, money supply and oil price. Specifically, it integrates a structural break dummy, coded 0 for observations before the mid-2023 foreign exchange unification and 1 thereafter, into the deterministic component of the VECM to capture the regime shift, tests the cointegrating rank via the Johansen procedure, and applies weak exogeneity tests to determine the causal adjustment mechanism. The aim is to examine the long-run equilibrium relationship and dynamic interactions between the exchange rate and inflation in Nigeria. The objectives are:

i. to test the stationarity properties of the exchange rate, inflation, interest rate, money supply and oil price series in Nigeria;

ii. to examine the long-run equilibrium (cointegrating) relationship and the short-run dynamic interactions between exchange rate changes and inflation, the short-run dimension being essential because it captures the immediate transmission of exchange rate shocks to prices that is most relevant for the timing of policy response, and to identify the direction of causality (does the exchange rate drive inflation, or does inflation drive the exchange rate?);

iii. to model the exchange rate–inflation nexus using a CVAR framework; and

ISSN No. 2321-2705 | DOI: 10.51244/IJRSI |Volume XIII Issue IX September 2026

iv. to forecast the exchange rate and inflation in Nigeria.

The significance of the study lies first in its capacity to confirm or reject a genuine inflationary link running from exchange rate movements and to clarify the direction of causality. Inflation can raise risk premiums, import costs through precautionary pricing, wage demands and inflation expectations, feeding back into currency weakness; conversely, rising or unstable exchange rates erode currency value, trigger capital flight and force central banks into reactive policy tightening. A bidirectional, mutually reinforcing relationship is highly plausible in emerging markets such as Nigeria, which underscores the need for a framework that captures feedback in both directions. Second, the study contributes evidence on the magnitude and speed of pass-through, that is, whether the inflationary impact of exchange rate shocks is complete or partial and how quickly it materialises, with direct implications for the timing of policy response. Finally, because findings on pass-through vary considerably across emerging markets depending on policy credibility, trade openness and financial depth, a country-specific study of Nigeria adds to the comparative literature. The study covers Nigeria over the period January 2015 to December 2025 and concentrates on the long- and short-run impact of exchange rate movements on inflation.

LITERATURE REVIEW

Conceptual Review

Exchange rate. An exchange rate is the price at which one currency can be exchanged for another, and is a key variable in international trade, investment and capital flows (Krugman & Obstfeld, 2018). It can be expressed in nominal terms, as the direct quotation of one currency against another, or in real terms, which adjust the nominal rate for relative price levels between trading partners (Mankiw, 2020). For emerging markets, exchange rate movements significantly influence macroeconomic stability because of their integration into global capital markets.

Inflation. Inflation is the sustained rise in the general price level of goods and services in an economy over time, typically measured by the Consumer Price Index (CPI) or the Producer Price Index (PPI) (Blanchard & Johnson, 2017). It is commonly classified as demand-pull, driven by excess aggregate demand; cost-push, resulting from rising production costs; or structural, arising from supply-side bottlenecks and inefficiencies (Dornbusch et al., 2014). In emerging markets inflationary pressures tend to be more volatile because of reliance on imported goods, less diversified production structures and vulnerability to commodity price fluctuations.

Exchange rate pass-through. ERPT is the extent to which changes in the exchange rate translate into domestic prices (Goldberg & Knetter, 1997). In emerging markets, depreciation of the domestic currency often raises import prices, which may cascade into general inflation through higher input costs and consumer prices. The magnitude of pass-through depends on trade openness, market competition, inflation expectations and the credibility of monetary policy (Taylor, 2000), and is generally higher in emerging economies because of weaker monetary policy anchors and greater import dependence.

Theoretical link between exchange rate and inflation. Three strands of theory explain the nexus. (i) Purchasing Power Parity (PPP) posits that exchange rate changes offset inflation differentials between countries, implying that sustained domestic inflation leads to currency depreciation (Cassel, 1918). (ii) The monetary approach to the exchange rate suggests that excessive money supply growth induces inflationary pressure and hence currency depreciation (Frenkel, 1976). (iii) The pass-through channel highlights how exchange rate fluctuations influence domestic inflation through import prices and production costs (Campbell & Krugman, 1993). In emerging markets, where exchange rate instability is often more severe, these channels operate with greater intensity, making currency stability a crucial component of inflation control.

Emerging market context. Emerging markets are economies experiencing rapid growth and industrialisation but with relatively underdeveloped financial systems (International Monetary Fund, 2022). They are typically more exposed to external shocks, including commodity price swings, global interest rate changes and exchange rate volatility (Calvo & Reinhart, 2002), and their higher import dependence and structural vulnerabilities make the interaction between exchange rate volatility and inflation stronger and more persistent than in advanced countries.

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Theoretical Review

Foundational models by Obstfeld and Rogoff (1995, 2000) and Taylor (2000) place the exchange rate at the centre of the monetary transmission mechanism and emphasise ERPT to domestic prices. Cross-country evidence indicates that pass-through is substantially more pronounced in emerging and developing economies than in advanced ones (Choudhri & Hakura, 2006), and Calvo and Reinhart (2002) introduced the notion of "fear of floating", whereby central banks in emerging markets avoid exchange rate flexibility because of concerns about inflationary repercussions. Devereux and Engel (2003) show how nominal rigidities and imperfect credibility can amplify the inflationary effects of exchange rate fluctuations in open economies, while Aghion et al. (2009) link exchange rate volatility to macroeconomic instability and reduced growth in financially underdeveloped economies. At the country level, Kiptui et al. (2005) for Kenya and Adu et al. (2015) for Ghana used structural VAR frameworks to document significant inflationary effects of exchange rate shocks.

Work on exchange rate determination has increasingly stressed instability in the underlying fundamentals. Caputo and Pedersen (2020), combining a DSGE framework with SVAR analysis, show that real exchange rate fluctuations are increasingly driven by country risk premium shocks, particularly under inflation targeting, as central banks prioritise inflation over exchange rate control. Model-averaging and meta-modelling techniques allow the weight on different fundamentals to change over time according to real-time forecasting performance. The European Central Bank (2021) found that PPP- and BEER-based equilibrium frameworks show gradual in-sample convergence of real exchange rates to equilibrium, whereas models anchored in current account fundamentals are less stable and less reliable for prediction. A 2022 post-Keynesian study of Brazil, using Thirlwall's growth constraint, found the real to have been overvalued between 2005 and 2014, with nominal appreciation helping to dampen inflation Nassif, A., Feijó, C., & Araújo, E. (2015). Overvaluation trend of the Brazilian currency in the 2000s: empirical estimation. *Revista de Economia Política, 35*(1). On forecasting, Baffour, Gyimah and Owiredu (2021) argue for hybrid approaches (technical, stochastic and neural-network models) over simple random-walk assumptions while reaffirming the role of inflation, interest rates and reserves; Pirayesh Neghab et al. (2023) find a dominant, time-varying role for crude oil prices in an interpretable machine-learning model of the CAD/USD rate; and a 2024 study of deep-learning models for the RMB/USD rate finds a transformer-based model (TSMixer) most effective when combined with macroeconomic indicators.

Oil and exchange rate linkages have also been examined in a structural VAR setting. Basnet and Upadhyaya (2015) study five Asian economies and find that oil price shocks to output, inflation and the real exchange rate are short-lived, fading within five to six quarters and explaining only a small share of variability, which implies that policy should emphasise short-term stabilisation. Alimi and Olaniran (2019), using a panel ARDL for Nigeria, South Africa and Egypt over 2005–2017, find that stock prices and exchange rates move in the same direction in South Africa but in opposite directions in Nigeria and Egypt, and that depreciation lowers stock prices in the full sample and the post-crisis period. Interest rates and inflation also shape this interaction, with tighter monetary policy generally weakening equity performance (Moales, 2008; Phylaktis & Ravazzolo, 2005).

The broader literature on exchange rate volatility and macroeconomic performance is mixed. Some authors find that exchange rate instability supports growth (Ozcelebi, 2018) while others find that it hampers it (Barguellil et al., 2018); differences are usually attributed to data, methodology, sample period and country characteristics (Phiri, 2018). For South Africa, Chamunorwa and Choga (2015) and Ngondo and Khobai (2018), using GARCH and ARDL respectively, find a negative effect of exchange rate instability on output, whereas Katusiime et al. (2016) report a positive effect of exchange rate shocks on GDP in Uganda. Balcilar et al. (2019) find prices to be stickier in South Africa than in Nigeria; Munthali et al. (2010) report a negative relationship between real effective exchange rate shocks and GDP in Malawi; Mahonye and Zengeni (2019) conclude that real devaluation in Zimbabwe is inflationary but may encourage exports; and Serenis and Tsounis (2014) show that volatility reduced trade in Malawi, Morocco and South Africa, with policy conclusions that were sensitive to the volatility measure used. Iyke and Odhiambo (2014) find that the growth effect of volatility depends on the exchange rate regime, depressing growth under fixed rates in 1960–1990 and supporting moderate-to-rapid growth under managed floating in 1990–2010. Panel evidence points in a similar direction: low exchange rate volatility appears necessary for growth in Central and Eastern Europe (Morina et al., 2020; Ioan et al., 2020), volatility lowers growth under flexible regimes in 45 developing and emerging countries (Barguellil et al., 2018), and the same sign is found for OECD economies (Janus & Riera-Crichton, 2015) and for Turkey (Alper, 2017). Lean and

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Ehigiamusoe (2019) show that exchange rate variability weakens the growth benefits of financial deepening in West Africa, and Hoang et al. (2020) find that a rising exchange rate affected growth in Vietnam through inflation.

The inflation–growth literature is equally unsettled. Early studies found that only high inflation harms growth (Fischer, 1993; Bruno & Easterly, 1995; Huybens & Smith, 1999), and the Mundell–Tobin theory holds that moderate inflation can stimulate capital accumulation. Estimated thresholds differ: 7.8 percent for five ASEAN countries (Thanh, 2015) and 17.9 and 14.5 percent for ECOWAS and SADC respectively (Mahawiya et al., 2020). Bittencourt et al. (2015) and Moyo and Roux (2019) find inflation detrimental to growth in the SADC region, whereas Phiri et al. (2018) and Anidiobu et al. (2018) report positive effects for Ghana, South Africa and Nigeria, and Taderera et al. (2021) find a positive effect in the Southern African Customs Union. On pass-through itself, Jasova et al. (2016) find that declining pass-through in 33 emerging and advanced economies reflects lower inflation, and Ha et al. (2019) find that the co-movement of exchange rate instability and inflation across 55 emerging and developing economies depends on each country's inflation threshold and period. In sum, pass-through is low and stable in advanced economies but higher and more damaging in emerging ones, so monetary authorities in African economies must weigh the effect of exchange rate policy on both price stability and capital market performance.

Empirical Review on Nigeria

Evidence for Nigeria is mixed in sign, significance and direction of causality. **GARCH and Granger causality (1970–2014).** Okoli et al. 2016 applied a GARCH(1,1) model and Granger causality tests to quarterly data and found one-way causality from inflation to real exchange rate volatility, suggesting that inflationary dynamics may drive currency instability; they recommend inflation targeting to reduce exchange rate fluctuations. **GARCH and VECM (1986–2022).** Another study *Exchange Rate Volatility on Inflation Dynamics in Nigeria: An ARCH-GARCH Approach*). Bello, Olayungbo and Folorunso use non-linear GARCH on 1986Q1–2019Q4 data (*Exchange Rate Volatility and Macroeconomic Performance in Nigeria*, IntechOpen). found that both nominal exchange rate changes and money supply growth significantly raise inflation in Nigeria, identifying currency volatility as a key inflation driver. **Structural VAR (2000–2023).** A recent study Ndume, M. M., & Akanegbu, B. N. (2025). Exchange rate volatility and inflation dynamics in Nigeria: A structural VAR approach. *Journal of Information Systems Engineering and Management, 10*(49s). using monthly data found that exchange rate shocks have an immediate and persistent inflationary impact, but Granger causality tests showed no predictive power from exchange rate volatility to inflation, suggesting that other forces also drive inflation. **FGLS (1981–2021).** A further study Impact of Exchange Rate Volatility on Selected ..."(the title was cut off in the result), *Nigerian Journal of Social Development* (AJOL, hosted under "ngjsd") reported a positive and statistically significant effect of the exchange rate on inflation. **OLS and Granger test (1970–2011).** In contrast, another analysis Inam, U. S., & Umobong, E. C. (2015). An empirical analysis of the relationship between exchange rate movements and economic growth in Nigeria. *European Journal of Business and Management*, *7*(30), 191–199. found an insignificant negative effect of the exchange rate on inflation and no causality between the two series, suggesting a weaker linkage.

Despite this body of work, studies of the Nigerian case rarely combine a full-system cointegration framework with tests of weak exogeneity or an explicit treatment of the post-2023 regime change. This is the gap the present study is designed to address.

METHODOLOGY

Data Description

The study uses monthly data on five variables for Nigeria from January 2015 to December 2025 (132 observations): the exchange rate (EXR), inflation (INF), the monetary policy rate (INT), broad money supply (M2) and the Brent crude oil price (OIL), as described in Table 1. The exchange rate, money supply and oil price enter the model in natural logarithms (LEXR, LM2 and LOIL). All computations were carried out in Python.

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[Author check: the exchange rate is derived from NBS annual rate ranges by linear interpolation within each year (2015 and 2025 are fitted to the NBS annual average). Money supply rises by an identical increment each month for 2016–2024 and the policy rate is constant at 14 percent from September 2016 to December 2024, which suggests these series are also interpolated or carried forward. Replace them with monthly CBN series before submission and re-run all tables.]

Table 1: Variable Description and Sources

Variable	Symbol	Description	Measurement	Source
Exchange Rate	EXR	Naira per US Dollar (Official Rate)	Monthly path derived from NBS annual rate ranges (linear within year)	NBS
Inflation Rate	INF	Consumer Price Index (CPI) - All Items	Year-on-year percentage change	NBS
Interest Rate	INT	Monetary Policy Rate (MPR)	Percentage	CBN
Money Supply	M2	Broad Money Supply	Billions of Naira	CBN
Oil Price	OIL	Brent Crude Oil Price	US Dollars per barrel	Macrotrends/FRED

Analysis Workflow

1. **Descriptive analysis:** summary statistics, correlation matrix and time plots of the five series.
2. **Stationarity:** Augmented Dickey–Fuller (ADF) and Phillips–Perron (PP) unit root tests, supplemented by the Zivot–Andrews test.
3. **Lag selection and cointegration:** information criteria for the VAR lag order and the Johansen trace and maximum-eigenvalue tests.
4. **Estimation and diagnostics:** VECM estimation, residual and stability diagnostics, Granger causality and weak exogeneity tests.
5. **Dynamic analysis:** orthogonalised impulse response functions (IRF) and forecast error variance decomposition (FEVD).
6. **Forecasting:** 24-month dynamic forecasts and evaluation of out-of-sample accuracy.

Model Specification

The joint dynamics of the exchange rate, inflation, interest rate, broad money supply and oil price are modelled with a CVAR, estimated in its VECM representation once the variables are confirmed to be integrated of the same order and cointegrated. Letting Y_t = (LEXR, INF, INT, LM2, LOIL)′, the underlying VAR(p) in levels is

$$Y_t = A_1 Y_{t-1} + A_2 Y_{t-2} + \ldots + A_p Y_{t-p} + \varepsilon_t \quad (1)$$

which, following the Johansen (1991) procedure, is re-parameterised into VECM form:

$$\Delta Y_t = \Pi Y_{t-1} + \Gamma_1 \Delta Y_{t-1} + \Gamma_2 \Delta Y_{t-2} + \ldots + \Gamma_{p-1} \Delta Y_{t-p+1} + \varepsilon_t \quad (2)$$

Here $\Pi = \alpha\beta'$ is decomposed into the r cointegrating vectors β, which define the long-run equilibrium relationship among the five variables, and the adjustment (weak exogeneity) coefficients α, which measure the speed at which

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each variable returns to equilibrium after a shock. The Γ_i matrices capture the short-run dynamics, that is, the effect of past changes in the system on current changes. The rank of Π is tested using two likelihood-ratio statistics. The trace statistic is

$$\lambda_{trace}(r) = -T \sum_{i=r+1}^{k} \ln(1 - \hat{\lambda}_i) \quad (3)$$

which tests the null that the number of cointegrating vectors is at most r against the alternative that it is greater than r, and the maximum-eigenvalue statistic is

$$\lambda_{max}(r, r+1) = -T \ln(1 - \hat{\lambda}_{r+1}) \quad (4)$$

which tests the null of r cointegrating vectors against the alternative of r + 1. In both, T is the sample size and $\hat{\lambda}_i$ are the estimated eigenvalues of the Π matrix, ordered from largest to smallest. If the variables are cointegrated, a VECM is the appropriate specification; otherwise a VAR in first differences is used. The Johansen test is preferred to the Engle–Granger two-step method because it is suited to multivariate systems.

Following the Pantula (1989) principle, Case 2 (restricted constant, no trend) is compared with Case 3 (unrestricted constant, no trend); Case 3 is the baseline because the log exchange rate and log money supply drift. Critical values are simulated for the exact sample size, lag order and deterministic design. Because the sample spans the mid-2023 foreign exchange unification, stability of the long-run relation is tested formally rather than through a deterministic dummy alone: (i) the Gregory and Hansen (1996) residual-based cointegration test with a level or regime shift at an unknown date, (ii) likelihood-ratio tests, with the break fixed at June 2023, of a step dummy in the cointegrating space and of regime-specific cointegrating vectors (Hansen & Johansen, 1999; Johansen, Mosconi & Nielsen, 2000), and (iii) pre- and post-unification sub-sample Johansen estimates. The January 2025 discontinuity in the inflation (CPI rebasing), policy-rate and money-supply series is handled with a restricted step dummy in the cointegrating space and an unrestricted impulse dummy. Granger causality is tested with the Toda and Yamamoto (1995) procedure, which remains valid for integrated variables.

Stationarity Tests

A non-stationary series can lead to spurious regression, so the stationarity of each variable is examined formally. The primary test is the Augmented Dickey–Fuller (ADF) test, which estimates

$$\Delta Y_t = \alpha + \beta t + \gamma Y_{t-1} + \sum_{i=1}^{p} \delta_i \Delta Y_{t-i} + \varepsilon_t \quad (5)$$

The null hypothesis (H_0: $\gamma = 0$) is that the series contains a unit root, against the alternative (H_1: $\gamma < 0$) that it is stationary. The lag length p is selected using the Akaike (AIC) or Schwarz (SIC) information criterion. As a robustness check, the non-parametric Phillips–Perron (PP) test, which corrects for serial correlation and heteroscedasticity in the error terms, is also applied. Because the sample includes the pronounced 2023 foreign exchange unification and the 2025 devaluation, unit root testing is supplemented with the Zivot and Andrews (1992) test, which allows for one endogenous structural break in the level, trend or both. Unlike the standard ADF test, which assumes no break and can be biased towards non-rejection of the unit root null when a genuine break is present, the Zivot–Andrews test selects the break date endogenously by minimising the t-statistic on the unit root coefficient across all candidate break points, and compares it with Zivot–Andrews rather than Dickey–Fuller critical values. This guards against concluding non-stationarity, and hence obtaining a biased Johansen rank, because of a structural break rather than a genuine unit root.

Diagnostic Tests

After estimation, the adequacy of the VAR/VEC model is checked with a series of residual tests. Serial correlation is tested with the Lagrange Multiplier (LM) test,

$$LM = s(\tilde{\theta})'[I(\tilde{\theta})]^{-1} s(\tilde{\theta}) \sim \chi^2(q) \quad (6)$$

where $s(\tilde{\theta})$ is the score vector and $I(\tilde{\theta})$ the information matrix, both evaluated at the restricted estimate, and q is the number of restrictions. Normality of the residuals is tested with the Jarque–Bera test,

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$$JB = n[S^2/6 + (K - 3)^2/24] \sim \chi^2(2) \quad (7)$$

where n is the sample size, S the sample skewness and K the sample kurtosis. Normality is not strictly required for consistency but is desirable for the validity of standard inference. Constant residual variance is examined with the Autoregressive Conditional Heteroscedasticity (ARCH) LM test. In an ARCH(q) model the conditional variance depends on past squared errors,

$$\sigma_t^2 = \alpha_0 + \alpha_1\varepsilon_{t-1}^2 + \alpha_2\varepsilon_{t-2}^2 + \ldots + \alpha_q\varepsilon_{t-q}^2 \quad (8)$$

and if $\alpha_1, \ldots, \alpha_q$ are jointly zero the variance is constant. Finally, stability is assessed from the inverse roots of the characteristic AR polynomial, all of which should lie inside the unit circle, and parameter stability can additionally be examined with the CUSUM test of Brown, Durbin and Evans (1975).

Weak Exogeneity Test

The adjustment coefficients α indicate the direction in which each variable responds to the error-correction term but do not by themselves establish which variables are weakly exogenous with respect to the long-run relationship. This is tested with the system likelihood-ratio test of Johansen (1991), which restricts $\alpha = H\psi$, where H is an n × s selection matrix, and compares the restricted eigenvalues λi^* with the unrestricted eigenvalues λi,

$$LR = T \Sigma i=1r \ln[(1 - \lambda i^*)/(1 - \lambda i)] \sim \chi 2(r(n - s)) \quad (9)$$

A variable for which H0: $\alpha i = 0$ cannot be rejected is weakly exogenous: it does not adjust to restore long-run equilibrium, and the burden of adjustment falls on the other variables in the system. Block restrictions (for example, that only the exchange rate and inflation adjust) are tested in the same way, t-ratios on α are computed from the equations conditional on the estimated cointegrating vector, and results are reported with and without the 2025 discontinuity dummy.

Impulse Response and Variance Decomposition

To trace the effect of a one-standard-deviation shock to one variable on the current and future values of all variables, orthogonalised impulse response functions are generated using the Cholesky decomposition. The ordering of the variables, here exchange rate, inflation, interest rate, money supply and oil price, imposes a contemporaneous causal ordering and should be kept in mind when interpreting the results. The FEVD complements the IRF by quantifying the proportion of the h-step-ahead forecast error variance of each variable that is attributable to its own shocks and to shocks to the other variables. Writing the VAR in moving-average form,

$$y_t = \Sigma_{s=0}^{\infty} \Psi_s u_{t-s}, \quad \Psi_0 = I \quad (10)$$

and orthogonalising the innovations through $u_t = P\varepsilon_t$ with $Var(\varepsilon_t) = I$ and $PP' = \Sigma_u$, so that $\Theta_s = \Psi_s P$, the share of the h-step forecast error variance of variable i explained by shocks to variable j is

$$FEVD_{ij}(h) = \Sigma_{s=0}^{h-1}(e_i'\Theta_s e_j)^2 / \Sigma_{s=0}^{h-1}\Sigma_{k=1}^{n}(e_i'\Theta_s e_k)^2 \quad (11)$$

where ei is the i-th selection vector. The generalised (Pesaran & Shin, 1998) impulse responses and decomposition, which do not depend on the variable ordering, are reported alongside, and the Cholesky results are re-estimated under all 120 orderings of the five variables to report the range of responses and variance shares across orderings.

Forecast Evaluation

Forecasts from the estimated VECM and from a time-varying-parameter error-correction VAR (TVP-ECM; forgetting-factor Kalman filter with $\lambda = 0.99$ and $\kappa = 0.96$, following Koop & Korobilis, 2013) are evaluated against a random walk and a random walk with drift using three standard error measures, the Mean Absolute

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Percentage Error (MAPE), the Mean Percentage Error (MPE) and the Mean Absolute Deviation (MAD), together with the RMSE:

$$MAPE = (100/n) \Sigma_{t=1}^{n} |A_t - F_t| / |A_t| \quad (12)$$

$$MPE = (100/n) \Sigma_{t=1}^{n} (A_t - F_t) / A_t \quad (13)$$

$$MAD = (1/n) \Sigma_{t=1}^{n} |A_t - F_t| \quad (14)$$

where At is the actual value, Ft the forecast and n the number of forecast origins in the evaluation window. MAPE is scale-free; MPE retains the sign of the error; and MAD reports the average absolute error in the original units. Forecasts are produced on an expanding window (rolling origin, first origin January 2023) at horizons of 1, 3, 6 and 12 months, the cointegrating vector is re-estimated at every origin, iterated increments are truncated to the historical range of increments, and the January 2025 discontinuity month is excluded as a target for all models. Forecasts are compared with the random walk with drift using the Diebold and Mariano (1995) test on the squared-error loss differential with the Harvey, Leybourne and Newbold (1997) small-sample correction; a negative statistic indicates that the model outperforms the benchmark, and the test is reported only when the number of origins exceeds 2h + 2.

RESULTS

Descriptive Statistics

Table 2 presents summary statistics of the variables in levels. The exchange rate averaged ₦558.13/US$ over the sample and ranged from ₦196.76 to ₦1,600.00; because it is built from NBS annual ranges, its movements within a year are linear, and the large depreciations of 2023 and 2024 appear as steps between years. Inflation averaged 15.65 percent, ranging from 6.79 to 34.80 percent. All series are positively skewed. The interest rate (kurtosis 8.45) and money supply (6.24) are markedly leptokurtic, whereas the exchange rate (3.87), inflation (3.64) and oil price (3.37) are close to the normal benchmark. The Jarque–Bera statistics reject normality for all variables except the oil price at the 5 percent level, a common feature of macroeconomic time series that motivates a dynamic multivariate framework such as the CVAR/VECM rather than a static OLS regression.

Table 2: Descriptive Statistics of the Variables (Levels)

Statistic	**EXR**	**INF**	**INT**	**M2**	**OIL**
Mean	558.13	15.65	14.90	46535.60	66.17
Median	370.00	13.45	14.00	41290.11	66.59
Maximum	1600.00	34.80	27.50	124400.00	116.80
Minimum	196.76	6.79	11.00	21288.24	21.04
Std. Dev.	418.70	7.08	4.00	25222.96	17.17
Skewness	1.54	1.23	2.61	1.94	0.16
Kurtosis	3.87	3.64	8.45	6.24	3.37
Jarque-Bera	55.99	35.76	313.28	140.43	1.35
Probability	0.0000	0.0000	0.0000	0.0000	0.5104
Observations	132	132	132	132	132

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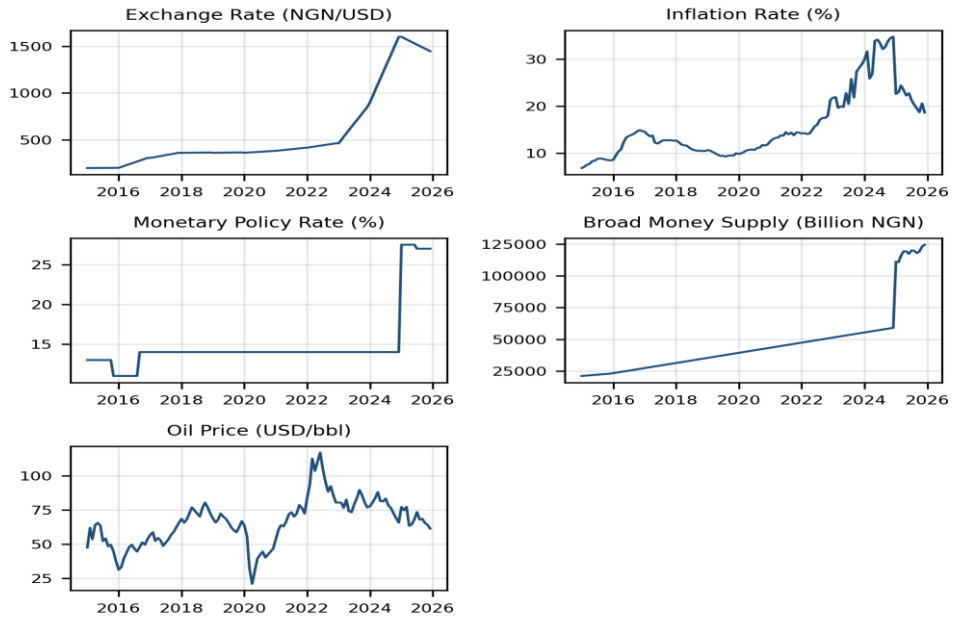


Figure 1: Trend of Exchange Rate, Inflation, Interest Rate, Money Supply and Oil Price, 2015M1–2025M12

Figure 1 shows the five series. The exchange rate rises in annual linear segments with steps between years, a feature of its construction, with the sharpest increases in 2023 and 2024. Inflation rises from 2022 and reaches 34.8 percent in December 2024 before falling to 22.6 percent in January 2025, a discontinuity that coincides with the NBS CPI rebasing rather than an economic shock. The monetary policy rate and broad money supply also change level abruptly in January 2025, and the oil price shows the more cyclical, mean-reverting pattern typical of a global commodity price. These discontinuities are handled with a step and impulse dummy in the VECM (see Model Specification).

Correlation Analysis

Table 3 reports pairwise correlation coefficients among the variables in levels. The exchange rate is strongly correlated with money supply (0.884), inflation (0.804) and the interest rate (0.761), and weakly correlated with the oil price (0.309). These are simple contemporaneous associations and, since the series are non-stationary, should not be interpreted causally; they nonetheless motivate a multivariate cointegration framework capable of separating short-run co-movement from long-run equilibrium relationships.

Table 3: Correlation Matrix of the Variables (Levels)

Variable	**EXR**	**INF**	**INT**	**M2**	**OIL**
EXR	1.000	0.804	0.761	0.884	0.309
INF	0.804	1.000	0.309	0.557	0.455
INT	0.761	0.309	1.000	0.928	0.139
M2	0.884	0.557	0.928	1.000	0.308
OIL	0.309	0.455	0.139	0.308	1.000

Unit Root Tests

Table 4 reports the ADF and PP results. At levels, no statistic is below the 5 percent critical value of −2.891, so a unit root cannot be rejected for any variable. After first differencing, inflation, the interest rate, money supply and the oil price reject a unit root in both tests and are therefore I(1). The log exchange rate is the exception: its

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first difference remains non-stationary under both ADF (-2.34) and PP (-2.49), so it is more persistent than an I(1) process. This is consistent with its construction as a piecewise-linear path, whose growth rate is itself a slowly moving series, and it means the I(1) pre-condition of the Johansen test is not met for this series (Objective i).

Table 4: ADF and Phillips–Perron Unit Root Test Results

Variable	ADF level	ADF 1st diff	PP level	PP 1st diff	5% CV	Order of Integration
Exchange Rate (log)	-1.291	-2.339	0.562	-2.491	-2.891	Not I(1): first difference non-stationary
Inflation Rate	-1.607	-3.299	-1.512	-13.399	-2.891	I(1)
Interest Rate	-0.583	-10.894	-0.489	-11.403	-2.891	I(1)
Money Supply (log)	0.525	-11.112	0.720	-11.659	-2.891	I(1)
Oil Price (log)	-2.856	-8.578	-2.412	-9.079	-2.891	I(1)

Lag Order Selection

Table 5 reports the lag order selection criteria. All four criteria (AIC, BIC, HQIC and FPE) are minimised at lag 2, which is therefore adopted for the VAR in levels, implying one lag of the first-differenced terms in the corresponding VECM.

Table 5: VAR Lag Order Selection Criteria

Lag	AIC	BIC	HQIC	FPE	nobs
1	-20.6947	-20.0363	-20.4272	1.0293e-09	131
2	-21.3394	-20.1262	-20.8465	5.4111e-10	130
3	-21.1242	-19.3507	-20.4036	6.7398e-10	129
4	-21.0173	-18.6777	-20.0667	7.5650e-10	128
5	-20.8210	-17.9096	-19.6382	9.3389e-10	127
6	-20.6215	-17.1324	-19.2040	1.1656e-09	126
7	-20.5568	-16.4841	-18.9023	1.2844e-09	125
8	-20.4620	-15.7994	-18.5679	1.4775e-09	124

Cointegration Test

Table 6 reports the Johansen rank test with an unrestricted constant, using critical values simulated for the sample size, lag order and design. The trace statistic (67.52) and the maximum-eigenvalue statistic (33.11) fall below their 5 percent critical values (75.44 and 36.41), so the null of no cointegration (r = 0) is not rejected on the full sample, and the result is unchanged when the 2025 discontinuity dummy is added (trace 82.49 against 88.56). The rank evidence is specification-dependent (Table 16): with a restricted constant r = 0 is rejected (trace 104.76 against 81.70) and the maximum-eigenvalue test selects r = 1, and in the pre-unification sample a cointegrating

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relation is clearly identified (trace 127.67 against 78.18). Cointegration is therefore not established over the full 2015–2025 sample with a drifting constant. Because the pre-unification evidence and economic theory point to a single long-run relation, the VECM with r = 1 is retained as a maintained hypothesis, and all VECM-based results (Tables 7–14) should be read as conditional on that assumption (Objective ii).

Table 6: Johansen Cointegration Rank Test (Unrestricted Constant, Simulated Critical Values)

r <=	Eigenvalue	Trace Stat	Trace CV (95%)	Max-Eig Stat	Max-Eig CV (95%)	Trace Reject	Max-Eig Reject
0	0.2248	67.52	75.44	33.11	36.41	No	No
1	0.1103	34.41	52.28	15.20	29.25	No	No
2	0.0832	19.21	32.42	11.29	21.96	No	No
3	0.0585	7.93	18.72	7.84	15.33	No	No
4	0.0007	0.09	7.89	0.09	7.89	No	No

Long-Run Relationship

Table 7 reports the normalised cointegrating vector, conditional on r = 1. Setting the cointegrating relationship to zero and solving for the log exchange rate yields the estimated long-run equation

Table 7: Normalised Cointegrating Vector (Long-Run Equation)

Variable	Coefficient
LEXR	1.0000
INF	-0.0830
INT	0.1610
LM2	-1.0600
LOIL	0.2012
D(2025M1)	-1.1795

LEXR = 0.0830 INF − 0.1610 INT + 1.0600 LM2 − 0.2012 LOIL + 1.1795 D(2025M1)(15)

The long-run coefficient on inflation is positive (0.083: a one percentage point rise in inflation is associated with an equilibrium naira depreciation of about 8 percent), and the money-supply elasticity (1.06) is close to unity, as the monetary approach to the exchange rate predicts. The interest rate and oil price enter with negative signs, consistent with higher yields and higher oil receipts supporting the naira. The dummy coefficient absorbs the January 2025 level discontinuity. Because rank is not confirmed on the full sample and the vector is unstable across the 2023 regime (Table 15), these magnitudes are indicative.

Short-Run Dynamics and Error Correction

The short-run dynamics and speed-of-adjustment coefficients are reported in Table 8, with t-ratios in parentheses from the equations conditional on the estimated cointegrating vector. The error-correction term enters the exchange rate equation with the expected negative sign (−0.0182, t = −5.60), so the exchange rate closes about

1.8 percent of the disequilibrium each month. It is positive and significant in the money supply equation (0.0065, t = 2.74) and insignificant in the inflation, interest rate and oil price equations. The January 2025 impulse dummy is significant in the exchange rate, inflation, interest rate and money supply equations, confirming the discontinuity in those series.

Table 8: VECM Short-Run Coefficients and Adjustment Coefficients

Regressor	**ΔLEXR**	**ΔINF**	**ΔINT**	**ΔLM2**	**ΔLOIL**
ECT(-1)	-0.0182 (-5.60)	-0.4597 (-0.88)	-0.0546 (-0.42)	0.0065 (2.74)	-0.0171 (-0.41)
DLEXR(-1)	0.6668 (11.94)	5.1981 (0.58)	1.1910 (0.53)	0.0534 (1.31)	-0.0236 (-0.03)
DINF(-1)	-0.0011 (-1.81)	-0.3247 (-3.49)	-0.0029 (-0.12)	0.0009 (2.13)	-0.0015 (-0.20)
DINT(-1)	0.0017 (0.82)	0.2036 (0.60)	-0.0481 (-0.56)	-0.0001 (-0.05)	0.0420 (1.57)
DLM2(-1)	-0.0606 (-1.26)	-9.9028 (-1.28)	1.1751 (0.60)	-0.0044 (-0.13)	-1.0332 (-1.69)
DLOIL(-1)	-0.0012 (-0.18)	-0.2784 (-0.26)	-0.1541 (-0.56)	0.0005 (0.11)	0.2293 (2.66)
Constant	-0.0546 (-5.39)	-1.2799 (-0.79)	-0.2062 (-0.50)	0.0295 (4.00)	-0.0490 (-0.38)
Impulse 2025M1	-0.0674 (-7.04)	-13.4414 (-8.76)	13.3584 (34.52)	0.6330 (90.74)	0.1381 (1.14)

Residual Diagnostics and Stability

Table 9 reports the residual diagnostics. The LM test does not reject the null of no residual serial correlation at the 5 percent level for the exchange rate (p = 0.633) and interest rate (p = 0.812) equations, but rejects it for inflation (p = 0.016), money supply (p < 0.001) and the oil price (p = 0.029); ARCH effects are present in the inflation, money supply and oil-price equations. The Jarque–Bera test rejects residual normality in every equation, which, as noted in the Methodology, does not affect consistency but qualifies the reliability of standard inference. The companion matrix has four unit roots, equal to n − r by construction, and the remaining eigenvalue moduli (0.966, 0.691, 0.280, 0.208, 0.055 and 0.011) lie inside the unit circle, so the system is stable.

Table 9: VECM Residual Diagnostic Tests

Equation	**LM(4) stat**	**LM p-value**	**JB stat**	**JB p-value**	**ARCH-LM(4)**	**ARCH p-value**
LEXR	2.565	0.633	2977.0	0.000	1.336	0.855
INF	12.195	0.016	635.8	0.000	37.048	0.000
INT	1.583	0.812	24095.9	0.000	0.106	0.999
LM2	32.843	0.000	1347.3	0.000	41.318	0.000
LOIL	10.822	0.029	339.8	0.000	39.907	0.000

Granger Causality

Table 10 reports Toda–Yamamoto Granger causality tests for all 20 ordered pairs (VAR(2) augmented with one extra lag and the 2025 dummies). At the 5 percent level the rejections are: INF → EXR (F = 3.72, p = 0.027). Inflation Granger-causes the exchange rate (F = 3.72, p = 0.027), whereas the exchange rate does not Granger-cause inflation (F = 2.50, p = 0.087). Tests that combine lagged differences with the error-correction term give

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the same direction: the exchange rate equation responds to inflation and the error-correction term ($p < 0.001$), but inflation does not respond to the exchange rate and the error-correction term ($p = 0.075$) (Objective ii).

Table 10: Toda–Yamamoto Granger Causality Test Results

Null Hypothesis	**F-stat**	**df1**	**df2**	**p-value**	**Decision**
INF does not Granger-cause LEXR	3.718	2	111	0.0274	Reject H0
INT does not Granger-cause LEXR	0.165	2	111	0.8482	Fail to reject H0
LM2 does not Granger-cause LEXR	1.079	2	111	0.3434	Fail to reject H0
LOIL does not Granger-cause LEXR	0.989	2	111	0.3751	Fail to reject H0
LEXR does not Granger-cause INF	2.496	2	111	0.0870	Fail to reject H0
INT does not Granger-cause INF	0.298	2	111	0.7427	Fail to reject H0
LM2 does not Granger-cause INF	1.558	2	111	0.2152	Fail to reject H0
LOIL does not Granger-cause INF	0.166	2	111	0.8470	Fail to reject H0
LEXR does not Granger-cause INT	0.307	2	111	0.7364	Fail to reject H0
INF does not Granger-cause INT	0.147	2	111	0.8636	Fail to reject H0
LM2 does not Granger-cause INT	0.005	2	111	0.9948	Fail to reject H0
LOIL does not Granger-cause INT	0.149	2	111	0.8619	Fail to reject H0
LEXR does not Granger-cause LM2	0.264	2	111	0.7683	Fail to reject H0
INF does not Granger-cause LM2	0.859	2	111	0.4265	Fail to reject H0
INT does not Granger-cause LM2	2.234	2	111	0.1119	Fail to reject H0
LOIL does not Granger-cause LM2	0.055	2	111	0.9462	Fail to reject H0
LEXR does not Granger-cause LOIL	0.579	2	111	0.5620	Fail to reject H0
INF does not Granger-cause LOIL	0.207	2	111	0.8134	Fail to reject H0
INT does not Granger-cause LOIL	1.152	2	111	0.3197	Fail to reject H0
LM2 does not Granger-cause LOIL	1.137	2	111	0.3246	Fail to reject H0

Note: Table 10 has been regenerated programmatically from the data. The two rows that previously showed identical statistics now show $F = 0.147$ ($p = 0.8636$) for INF → INT and $F = 0.166$ ($p = 0.8470$) for LOIL → INF. The Toda–Yamamoto procedure is valid for integrated variables.

Weak Exogeneity

Table 11 reports the Johansen likelihood ratio tests of the restriction $\alpha i = 0$ for each variable, estimated on the full sample (2015M1–2025M12, $n = 132$) in a VECM(2) with one cointegrating relation and the 2025 discontinuity dummy. The exchange rate adjusts strongly ($\alpha = -0.0182$, $t = -5.60$; $LR = 18.77$, $p < 0.001$) and

money supply adjusts weakly (α = 0.0065; LR = 6.67, p = 0.010), whereas inflation (p = 0.463), the interest rate (p = 0.622) and the oil price (p = 0.864) are weakly exogenous. The block tests in the note sharpen the result: the joint restriction that only the exchange rate, inflation and money supply adjust is not rejected (p = 0.855), the restriction that only inflation adjusts is rejected decisively (p < 0.001), and the restriction that only the exchange rate adjusts is not rejected at conventional levels (p = 0.126). The burden of adjustment toward long-run equilibrium therefore falls on the exchange rate, not on inflation. The result for the exchange rate is robust to the dummy: without it α = −0.0147 (LR = 15.35, p < 0.001) and no other variable rejects weak exogeneity.

Table 11: Weak Exogeneity (α-Restriction) Likelihood Ratio Test

Variable	Restriction	α (unrestricted)	LR statistic	df	p-value	Decision
EXR	αEXR = 0	-0.0182	18.77	1	< 0.001	Reject H0 (adjusts)
INF	αINF = 0	-0.4597	0.54	1	0.463	Do not reject H0
INT	αINT = 0	-0.0546	0.24	1	0.622	Do not reject H0
M2	αM2 = 0	0.0065	6.67	1	0.010	Reject H0 (adjusts)
OIL	αOIL = 0	-0.0171	0.03	1	0.864	Do not reject H0

Note: Johansen (1991) system LR test of α = Hψ, distributed χ2(r(n − s)); χ2(1) for the single restrictions above. Block restrictions: only EXR and INF adjust, χ2(3) = 7.19 (p = 0.066); only EXR adjusts, χ2(4) = 7.19 (p = 0.126); only INF adjusts, χ2(4) = 28.69 (p < 0.001); only EXR, INF and INT adjust, χ2(2) = 6.84 (p = 0.033); only EXR, INF and M2 adjust, χ2(2) = 0.31 (p = 0.855).

Impulse Response Functions

Orthogonalised IRFs were generated using a Cholesky decomposition with the ordering exchange rate, inflation, interest rate, money supply and oil price over a 24-month horizon, and were re-estimated under all 120 orderings and as generalised IRFs; Figure 2 shows the baseline, the range across orderings and the generalised response. A one-standard-deviation exchange rate shock changes inflation by only 0.03 to 0.08 percentage points after six months across orderings (generalised: 0.08), and the response is negligible and turns negative by the 24-month horizon (-0.11 to -0.03). By contrast, a one-standard-deviation inflation shock raises the log exchange rate by about 4 percent after 12 months (3.9 to 4.1 percent across orderings; generalised 4.1), and this response is stable across orderings. In this dataset the dynamic link therefore runs from inflation to the exchange rate rather than from the exchange rate to prices.

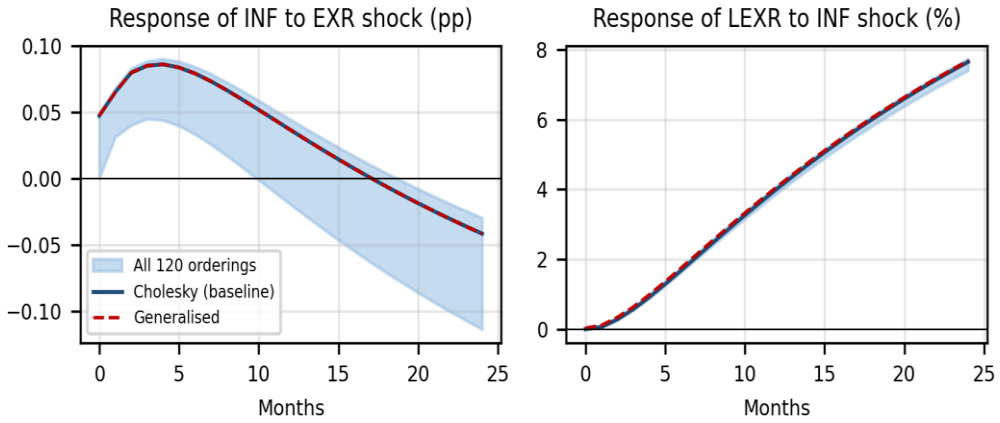


Figure 2: Response of Inflation to an Exchange Rate Shock, and Response of Exchange Rate to an Inflation Shock

ISSN No. 2321-2705 | DOI: 10.51244/IJRSI |Volume XIII Issue IX September 2026

Figure 3 extends the analysis to the full system. Inflation responds mainly to its own shock (peak response 1.86 percentage points). Interest rate and oil price shocks lower inflation by up to about 0.5 and 0.3 percentage points over 24 months, while the response to the exchange rate shock never exceeds 0.09. The generalised responses (dashed) are close to the Cholesky responses for the exchange rate, inflation and interest rate shocks but differ for money supply and oil price shocks, so those two panels are sensitive to the identification scheme.

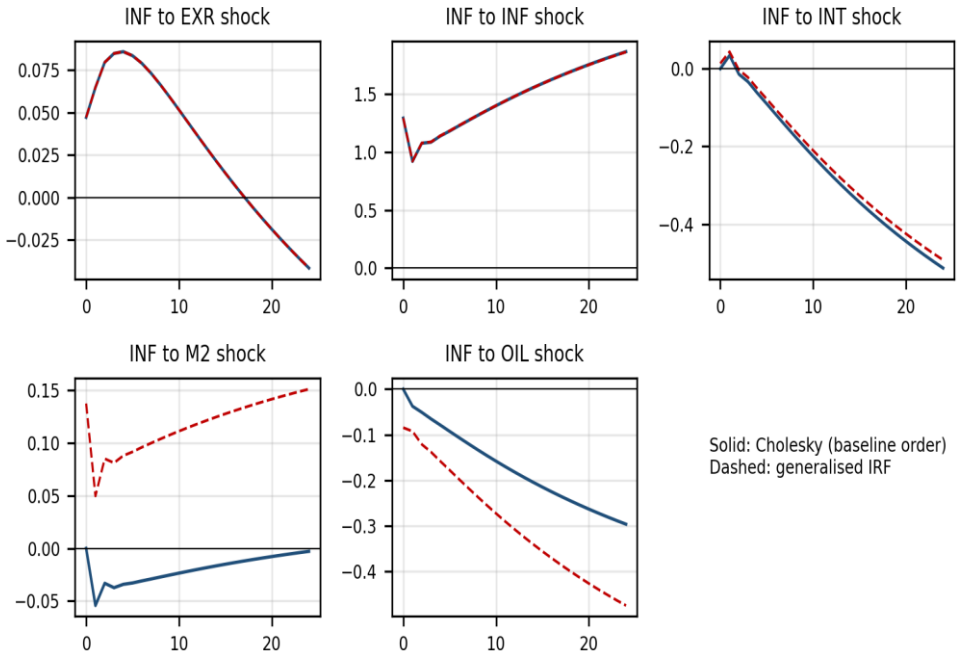


Figure 3: Response of Inflation to Shocks in All Five System Variables

Forecast Error Variance Decomposition

Figure 4 shows the FEVD of inflation. Inflation's own shock accounts for 99.9 percent of its forecast error variance at one month, 99.0 percent at six months, 97.6 percent at 12 months and 94.4 percent at 24 months. Exchange rate shocks account for 0.13, 0.45, 0.31 and 0.12 percent at the same horizons, interest rate shocks for 3.9 percent at 24 months, and the generalised decomposition gives similar exchange rate shares (0.13, 0.31 and 0.11 percent at 1, 12 and 24 months). Across all 120 Cholesky orderings the exchange rate's share of inflation variance at 12 months ranges only from 0.06 to 0.35 percent, so the result does not depend on the ordering. The reverse is not true: inflation shocks explain 41–47 percent of exchange rate forecast error variance at 12 months across orderings (generalised: 43 percent).

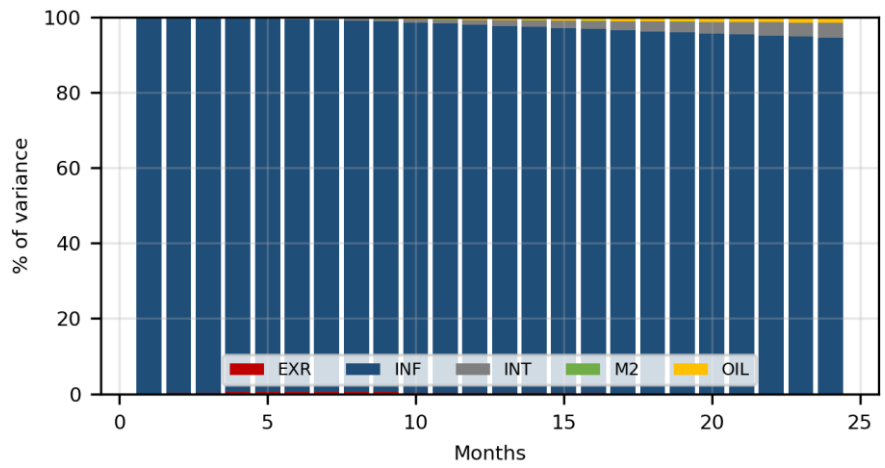


Figure 4: Forecast Error Variance Decomposition of Inflation

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Forecasts

Using the estimated VECM, the exchange rate and inflation (together with the interest rate, money supply and oil price) were forecast dynamically for the 24 months after the end of the sample, January 2026 to December 2027 (Objective iv), conditional on r = 1 and on the 2025 level shift remaining in place. Selected values are presented in Table 12 and Figure 5. Absent further shocks, the naira is projected to strengthen gently from ₦1,431/US$ in January 2026 to about ₦1,287 in March 2027 and then drift back to ₦1,337 by December 2027, while inflation eases from 19.1 percent to about 18.1 percent in early 2027 before edging up to 18.7 percent.

Table 12: Selected VECM Forecasts, 2026–2027

Month	EXR	INF	INT
Jan-2026	1431.3	19.08	26.96
Mar-2026	1396.8	18.70	26.86
May-2026	1363.9	18.46	26.76
Jul-2026	1336.1	18.28	26.67
Sep-2026	1314.6	18.16	26.59
Nov-2026	1299.4	18.08	26.51
Jan-2027	1290.4	18.06	26.45
Mar-2027	1287.2	18.09	26.39
May-2027	1289.5	18.16	26.34
Jul-2027	1297.0	18.27	26.30
Sep-2027	1309.4	18.42	26.27
Nov-2027	1326.6	18.61	26.25

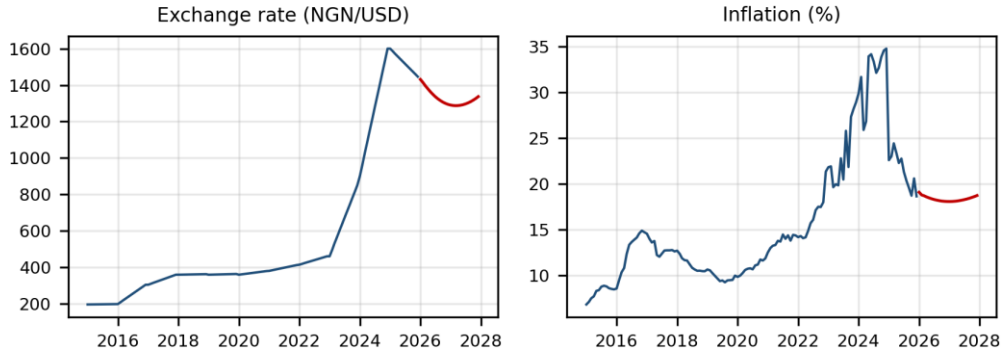


Figure 5: Actual and Forecast Exchange Rate and Inflation, 2015M1–2027M12

Out-of-Sample Forecast Accuracy

To judge whether the VECM adds value relative to naive benchmarks, rolling-origin forecasts (expanding window, first origin January 2023) were compared with a random walk with drift using the Diebold–Mariano

test with the Harvey–Leybourne–Newbold correction (Table 13: one-month horizon, targets February 2023 to December 2024; the January 2025 discontinuity is excluded for all models). For the exchange rate the VECM's RMSE is ₦11.8 against ₦36.9 for the benchmark (DM = -8.35, $p < 0.001$); for inflation the VECM's RMSE of 3.03 percentage points is slightly larger than the benchmark's 2.85 and the difference is insignificant (DM = 0.70, $p = 0.493$). The exchange rate result should not be over-read: because the series is a within-year linear path its changes are highly predictable, and any model that tracks the recent slope will beat a random walk.

Table 13: Out-of-Sample Forecasting Accuracy Against a Random Walk with Drift (One-Month Horizon, 2023M2–2024M12)

Variable	RMSE (VECM)	RMSE (RW-drift)	DM statistic	P-value
Exchange rate (EXR, ₦)	11.82	36.94	-8.35	< 0.001
Inflation (%)	3.03	2.85	0.70	0.493

Table 14 reports MAPE, MPE and MAD for the VECM at the one-month horizon: MAPE is 1.20 percent for the exchange rate and 8.87 percent for inflation, and the MPE of +0.23 and -1.45 percent indicates a slight under-forecast of the exchange rate and over-forecast of inflation. Accuracy deteriorates with the horizon (Table 18): at 12 months the VECM's MAPE is 15.1 percent for the exchange rate and 15.6 percent for inflation.

Table 14: Out-of-Sample Forecast Accuracy of the VECM (MAPE, MPE, MAD; One-Month Horizon)

Variable	MAPE (%)	MPE (%)	MAD
Exchange rate (EXR, ₦)	1.20	+0.23	9.64
Inflation (%)	8.87	-1.45	2.34

Structural Break Tests

Table 15 reports formal tests for a shift in the long-run relation at the mid-2023 foreign exchange unification. The Gregory–Hansen test (regressand LEXR, four regressors) does not reject the null of no cointegration against either a level shift (ADF* = -5.27, estimated break March 2021; 5 percent critical value −5.57) or a regime shift (ADF* = -5.96, break September 2018; 5 percent critical value −6.32). The likelihood-ratio tests with the break fixed at June 2023 reject constancy of the cointegrating vector across regimes ($\chi2(5) = 43.72$, $p < 0.001$) and of the whole relation including the level ($\chi2(6) = 48.29$, $p < 0.001$), and reject the absence of a level shift at the 5 percent level ($\chi2(1) = 4.57$, $p = 0.033$). Table 16 shows that a cointegrating relation is clearly identified before the unification but not on the full sample with an unrestricted constant. The post-unification sample has only 31 observations, too few for the five-variable system, so only a three-variable system (LEXR, INF, LM2) is estimated there; the likelihood-ratio p-values rely on asymptotic $\chi2$ approximations and should be confirmed by bootstrap. Taken together, the tests indicate that the long-run relation is not stable across the 2023 regime change, so neither a single full-sample vector nor a deterministic dummy alone is an adequate description.

Table 15: Structural Break Tests for the Long-Run Relation

Test	Statistic	Critical value / df	Result
Gregory–Hansen, level shift (C)	-5.27	5% CV: −5.57	Do not reject no cointegration
Gregory–Hansen, regime shift (C/S)	-5.96	5% CV: −6.32	Do not reject no cointegration
LR: no level shift (step dummy in β), break 2023M6	4.57	$\chi2(1)$	$p = 0.033$

LR: β constant across regimes, level shift allowed	43.72	$\chi2(5)$	$p < 0.001$
LR: whole relation constant (no break)	48.29	$\chi2(6)$	$p < 0.001$

Note: Gregory–Hansen (1996) critical values for four regressors. LR tests are asymptotic $\chi2$ tests with a known break date; the post-break sample is short, so bootstrap p-values are recommended.

Table 16: Cointegration Rank Evidence Across Specifications and Sub-Samples

Specification	**Trace (r = 0)**	**5% CV**	**Conclusion**
Full sample, unrestricted constant (Case 3)	67.52	75.44	r = 0 not rejected
Full sample, Case 3 with 2025 dummy	82.49	88.56	r = 0 not rejected
Full sample, restricted constant (Case 2)	104.76	81.70	r = 0 rejected; max-eigenvalue selects r = 1
Pre-unification (2015M1–2023M5), Case 3	127.67	78.18	r = 0 rejected
Post-unification (2023M6–2025M12), 3 variables	46.79	n/a	T = 31: no reliable critical values

Ordering Robustness

Table 17 summarises the responses and variance shares across all 120 Cholesky orderings and the generalised counterparts. The response of inflation to an exchange rate shock ranges from 0.033 to 0.084 percentage points at six months, −0.020 to 0.044 at 12 months and -0.113 to -0.030 at 24 months; none is economically large, and the generalised responses lie inside these ranges. The exchange rate's share of inflation variance never exceeds 0.35 percent at 12 months under any ordering (generalised: 0.31 percent), so the finding that exchange rate shocks explain almost none of inflation variance is not an artefact of the ordering. The response of the exchange rate to an inflation shock and the share of exchange rate variance explained by inflation are likewise insensitive to ordering.

Table 17: Sensitivity of IRF and FEVD Results to Variable Ordering

Statistic	**Min**	**Median**	**Max**	**Generalised**
Response of INF to EXR shock, 6 months (pp)	0.033	0.054	0.084	0.079
Response of INF to EXR shock, 12 months (pp)	-0.020	0.006	0.044	0.036
Response of INF to EXR shock, 24 months (pp)	-0.113	-0.080	-0.030	-0.042
Share of INF variance from EXR, 1 month (%)	0.00	0.03	0.15	0.13
Share of INF variance from EXR, 12 months (%)	0.06	0.15	0.35	0.31
Share of INF variance from EXR, 24 months (%)	0.10	0.11	0.13	0.11
Response of LEXR to INF shock, 12 months (%)	3.89	4.00	4.10	4.07
Share of EXR variance from INF, 12 months (%)	41.0	43.6	46.6	43.0

ISSN No. 2321-2705 | DOI: 10.51244/IJRSI |Volume XIII Issue IX September 2026

Time-Varying-Parameter Forecasts

Table 18 compares the VECM with the TVP-ECM, a random walk with drift and a driftless random walk on rolling origins from January 2023 with targets up to December 2024 (before the 2025 discontinuity). For the exchange rate the VECM has the lowest RMSE and MAPE at every horizon (MAPE of 1.20, 4.02, 8.37 and 15.12 percent at 1, 3, 6 and 12 months, against 2.44, 5.00, 11.28 and 27.86 percent for the TVP-ECM), and both beat the random walk with drift. For inflation the differences are small and statistically insignificant at horizons up to six months (DM p-values above 0.49): the TVP-ECM has a slightly lower RMSE than the VECM at 1, 3 and 6 months (2.92 against 3.03; 3.87 against 4.10; 4.20 against 4.66) but a higher one at 12 months (6.26 against 5.99). The time-varying-parameter specification therefore does not deliver a consistent improvement in accuracy over the VECM in this dataset. Only 12 forecast origins are available at the 12-month horizon, so Diebold–Mariano tests are not reported there.

Table 18: Rolling-Origin Forecast Accuracy by Model and Horizon (2023M1–2024M12)

Variable	h	RMSE VECM	RMSE TVP-ECM	RMSE RW-drift	MAPE VECM	MAPE TVP-ECM	MAPE RW-drift
EXR (₦)	1	11.82	23.03	36.94	1.20	2.44	3.95
EXR (₦)	3	46.61	54.20	113.82	4.02	5.00	11.41
EXR (₦)	6	121.01	142.43	236.22	8.37	11.28	21.73
EXR (₦)	12	206.75	412.58	498.13	15.12	27.86	39.80
INF (pp)	1	3.03	2.92	2.85	8.87	8.49	7.37
INF (pp)	3	4.10	3.87	3.76	13.00	12.00	11.05
INF (pp)	6	4.66	4.20	4.51	14.03	12.27	13.06
INF (pp)	12	5.99	6.26	7.49	15.64	14.59	21.39

Note: RMSE and MAD in original units (₦ per US$ for EXR; percentage points for INF); MAPE in percent. The January 2025 discontinuity month is excluded as a target for all models.

DISCUSSION

The results do not support a strong, stable exchange rate-to-inflation pass-through in Nigeria over 2015–2025 on the data used here. The Johansen test with an unrestricted constant does not confirm a cointegrating vector on the full sample, although one is clearly identified before the 2023 unification, and the likelihood-ratio tests show that the long-run relation shifted afterwards. Conditional on one vector, the exchange rate does the adjusting to long-run disequilibrium while inflation is weakly exogenous, Granger causality runs from inflation to the exchange rate, exchange rate shocks explain well under one percent of inflation forecast error variance in every Cholesky ordering and in the generalised decomposition, and inflation shocks explain 41 to 47 percent of exchange rate variance. This is closer to the Nigerian studies that find causality running from inflation to exchange rate dynamics (Okoli et al., 2016) than to the literature reporting strong and persistent pass-through in emerging and import-dependent economies (Obstfeld & Rogoff, 1995; Calvo & Reinhart, 2002; Choudhri & Hakura, 2006), and it means the "pass-through magnitude and direction of causality" question posed in the Introduction is answered, for this dataset, with a small pass-through and a direction from inflation to the exchange rate.

Three qualifications are essential. First, the exchange rate is derived from NBS annual ranges with linear interpolation within years, so month-to-month exchange rate dynamics, and therefore the impulse responses, Granger tests and short-run forecasts that involve it, partly reflect the construction of the series rather than market

behaviour; the log exchange rate also fails the I(1) pre-condition. Second, the money supply and policy rate series show long stretches of constant increments or values and abrupt level changes in January 2025, and inflation is discontinuous at the CPI rebasing; these are handled with a dummy but not resolved. Third, the long-run relation is unstable across the 2023 regime, and the 31-month post-unification sample is too short for reliable inference. The findings should therefore be treated as provisional until monthly exchange rate, money supply and policy rate series are used. The forecasts are conditional scenarios, and the time-varying-parameter error-correction model does not improve on the VECM.

CONCLUSION

This study examined the long-run equilibrium relationship and dynamic interactions between the exchange rate and inflation in Nigeria using monthly data from January 2015 to December 2025 and a CVAR/VECM framework. Inflation, the policy rate, money supply and the oil price are I(1), while the log exchange rate is more persistent. The Johansen test with an unrestricted constant does not identify a cointegrating relation on the full sample, and Gregory–Hansen and likelihood-ratio tests indicate that the long-run relation shifted with the 2023 foreign exchange unification. Conditional on one relation, the exchange rate adjusts to disequilibrium, inflation is weakly exogenous, causality runs from inflation to the exchange rate, and exchange rate shocks explain almost none of the variance of inflation under any Cholesky ordering or in the generalised decomposition. A time-varying-parameter model does not improve forecast accuracy, and VECM forecasts for 2026–2027 are best read as conditional scenarios. These conclusions are provisional until monthly series replace the interpolated exchange rate, money supply and policy rate data.

RECOMMENDATIONS

i. Because the evidence does not show exchange rate shocks to be a dominant driver of inflation variance in this sample, policy should not rest on a single pass-through estimate; the Central Bank of Nigeria should monitor influence in both directions, including the response of the exchange rate to inflation and monetary conditions.

ii. Given that the exchange rate, not inflation, adjusts to long-run disequilibrium, policies that anchor inflation and money growth are consistent with greater exchange rate stability, and monetary responses to inflation pressure should be timely rather than reactive.

iii. Policies that reduce reliance on imported food, fuel and industrial inputs, for example investment in domestic refining capacity and agricultural productivity, would be expected to weaken the exchange rate–inflation link over time, since import dependence is a plausible transmission channel.

iv. The tests show that the long-run relation shifted around the 2023 unification, so future monitoring and policy analysis should distinguish between the pre- and post-liberalisation exchange rate regimes rather than relying solely on relationships estimated over the pooled sample.

v. The forecasts should be used by policymakers as one conditional scenario among several and updated regularly as new data become available, rather than treated as a fixed prediction.

Limitations and suggestions for further research

The exchange rate, money supply and policy rate series used here are derived or interpolated, and replacing them with observed monthly series is the most important extension. Second, the post-unification sample has only 31 observations, so the break tests rely on asymptotic approximations and bootstrap-based critical values should be used as a robustness check; the Gregory–Hansen critical values are tabulated values for four regressors. Third, the five-variable system is estimated with a short lag structure dictated by 132 monthly observations, and a longer or higher-frequency series would permit richer dynamics. Finally, ordering robustness has been assessed through all Cholesky permutations and generalised responses, but identification with theory-based restrictions in a structural VECM was not attempted and remains a useful extension.

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