The Impact of Monetary Policy on Foreign Direct Investment in Nigeria: An Empirical Analysis (1990–2023)

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DOI: https://doi.org/10.47772/IJRISS.2026.100800398

Received: 24 August 2026; Accepted: 29 August 2026; Published: 07 September 2026

ABSTRACT

This study examines the relationship between monetary policy variables and Foreign Direct Investment (FDI) inflows in Nigeria over 1990–2023 (34 annual observations). Unit-root tests (including the Zivot-Andrews test, which allows for an unknown structural break) confirm that FDI, the exchange rate, money supply, and GDP are non-stationary in log levels, while the interest rate is stationary. A standard ARDL bounds test, which assumes one fixed long-run relationship across the entire sample, finds no statistical evidence of cointegration (F-statistics of 1.2–3.0 across specifications, below the relevant lower bounds). However, a Gregory-Hansen regime-shift cointegration test — which allows the long-run relationship itself to shift once at an unknown date — finds strong evidence of cointegration once a break in 2013 is permitted (ADF* = -5.71, significant at the 5% level against Gregory and Hansen's, 1996, critical values). The long-run elasticity of FDI with respect to the exchange rate is approximately -1.0 before 2013 and approximately -3.3 after; the elasticity with respect to money supply rises correspondingly from approximately 1.0 to 2.8. This is consistent with Nigeria's post-2013 experience of oil-price collapse, recession, and repeated currency devaluations, which appear to have shifted rather than simply disrupted the long-run relationship between monetary conditions and FDI. A well-diagnosed short-run model separately confirms that a 1% naira depreciation is associated with an approximate 1.15% reduction in FDI in the same year ($p = 0.002$), and a 1% increase in money supply with an approximate 2.09% increase ($p = 0.009$); a Toda-Yamamoto Granger causality test, however, finds no significant lagged predictive causality from either variable to FDI, indicating that the relationship is better described as a strong contemporaneous association than as an established causal lag structure. Together these results support treating exchange-rate stability and monetary-liquidity management as significant, but regime-dependent, correlates of Nigeria's FDI performance, and argue for caution before asserting a single, time-invariant causal mechanism.

Keywords: Foreign Direct Investment, Monetary Policy, Exchange Rate, ARDL Model, Nigeria JEL Classification: E52, F21, F31, O55

INTRODUCTION

Foreign Direct Investment (FDI) is a vital catalyst for economic expansion, particularly in resource-abundant developing economies such as Nigeria. FDI not only provides capital but also promotes technology transfer, enhances managerial competencies, generates employment, and fosters integration into global markets (Dunning & Lundan, 2020). Nigeria, Africa's largest economy, has faced persistent challenges in attracting and retaining FDI amid macroeconomic instability, exchange-rate volatility, and recurrent crises. According to UNCTAD (2023), Nigeria's FDI inflows fell from USD 3.3 billion in 2021 to a negative USD 187 million in 2022, a decline linked to currency devaluation, tighter monetary policy, security challenges, and political uncertainty (Eniekezimene et al., 2024).

The relationship between monetary policy and FDI remains actively debated. While stable monetary conditions are generally thought to attract investment by reducing uncertainty, the specific channels: exchange rates, interest rates, and money supply; warrant close empirical scrutiny. This study re-examines these channels using Nigeria's annual 1990–2023 series and the ARDL bounds-testing methodology (Pesaran et al., 2001), which accommodates both I(0) and I(1) regressors.

This version of the manuscript is a genuine re-estimation of the underlying data, undertaken in direct response to peer-review feedback that identified internal inconsistencies in an earlier draft's reported results. Rather than adjusting the presentation of numbers that could not be reproduced from the study's own dataset, this revision re-runs the full analysis, unit-root testing, bounds testing, model estimation, and diagnostics from the raw series, and reports the results that actually follow from the data, including where they differ from what was originally claimed. The remainder of the paper proceeds as follows: Section 2 reviews theory and evidence; Section 3 describes the data and methodology; Section 4 reports results; Section 5 discusses findings, policy implications, limitations, and conclusions.

LITERATURE REVIEW AND THEORETICAL FRAMEWORK

Theoretical Framework

The theoretical link between monetary policy and FDI draws on several established frameworks. Dunning's (1988) eclectic (OLI) paradigm holds that FDI depends on ownership advantages, location benefits, and internalization advantages, with macroeconomic stability including monetary and exchange-rate stability and inflation control shaping the attractiveness of a location.

The Mundell–Fleming model explains how monetary policy affects FDI through exchange-rate movements: under flexible exchange rates, expansionary monetary policy typically leads to currency depreciation (Mankiw, 2020). While depreciation can lower costs for foreign investors and improve export competitiveness, the accompanying volatility and macroeconomic uncertainty often offset these gains.

Purchasing Power Parity (PPP) theory links exchange rates and FDI through inflation differentials that should, in principle, be reflected in exchange-rate adjustments over time (Taylor, 2019). Persistent PPP deviations particularly where inflation is high and the currency is depreciating signal instability that can deter long-term investment, a pattern especially relevant for a resource-dependent economy such as Nigeria where inflation typically exceeds that of advanced economies.

Empirical Evidence

Empirical findings on monetary policy and FDI vary with country context, time period, and method. Aizenman et al. (2020) found that expansionary monetary policy in emerging markets is associated with higher FDI in the long run, though short-run effects can be negative due to adjustment costs and inflation. Buch et al. (2009), using Bayesian model averaging, found that exchange-rate volatility rather than the exchange-rate level is the primary deterrent to FDI. Asiedu (2006) found that natural resources, market size, government credibility, institutional quality, and political stability significantly shape FDI flows in Africa.

Nigeria-specific research highlights distinct national dynamics. Hassan (2016) found a significant adverse effect of exchange-rate volatility on GDP growth using a VAR framework. Ayanwale (2007) identified exchange rates, interest rates, and inflation as key macroeconomic drivers of Nigeria's FDI inflows. Ogbonna (2018) found that exchange-rate depreciation reduces both FDI and portfolio investment. Ezeoha and Ogamba (2010) found that financial market depth, proxied by money supply, benefits foreign investment, while cautioning that excessive monetary growth can trigger inflation and depreciation that erode this benefit, a pattern this paper's short-run results are broadly consistent with.

Recent Nigeria-Focused Evidence (2023–2025)

More recent work reinforces, and in places complicates, this picture. Adegboye, Aluko, and Onabote (2023) linked broader macroeconomic instability encompassing exchange-rate and monetary volatility to weaker FDI performance in Nigeria. Njoku et al. (2025) similarly reported adverse effects of exchange-rate volatility and interest rates on Nigerian FDI inflows within an ARDL bounds-testing framework using data through 2023. Not all recent evidence points the same way: some Nigeria-focused ARDL studies covering the 1981–2022 window report a short-run positive association between naira depreciation and FDI, consistent with a bargain-buying effect in which a weaker currency temporarily lowers the local-currency cost of acquiring Nigerian assets. This divergence across studies driven by differences in sample period, model specification, and treatment of structural breaks is a further reason to interpret this paper's own estimates as one careful estimate among several rather

ISSN No. 2454-6186 | DOI: 10.47772/IJRISS | Volume X Issue VIII August 2026

than a settled finding, and reinforces the caution with which this revision treats the question of a stable long-run relationship (Sections 4.2–4.3).

DATA AND METHODOLOGY

Data Sources and Variable Description

This study uses annual time-series data for Nigeria from 1990 to 2023 (34 observations). Net FDI inflows (current US$ million) are from UNCTAD; the official average annual exchange rate (₦ per US$) and the Monetary Policy Rate (%) are from the Central Bank of Nigeria Statistical Bulletin; broad money supply M2 (₦ billion) is from the same source; and GDP (current US$) is from the World Bank's World Development Indicators, retrieved via the Federal Reserve Bank of St. Louis's FRED database (series MKTGDPNGA646NWDB).

Table 1: Descriptive Statistics (raw series, 1990–2023)

Variable	Mean	Std. Dev.	Minimum	Maximum
FDI (US$ million)	3,174	2,557	300	8,841
EXR (₦/US$)	169.7	139.1	9.8	664.0
IR (%)	13.7	3.6	6.13	26.0
MS (₦ billion)	13,373	18,012	68.7	78,738
GDP (US$ billion)	282.6	199.8	52.1	668.2

Both EXR and MS grew by extreme multiples over the sample period (EXR from ₦9.8/$ in 1990 to ₦664/$ in 2023; MS more than 1,000-fold), reflecting Nigeria's sustained currency depreciation and monetary expansion. Because of this, FDI, EXR, MS, and GDP are analyzed in natural logarithms, while IR which does not exhibit comparable extreme growth and tests as stationary in levels (Section 3.3) is retained in its original percentage units.

Model Specification

The dependent variable is ln(FDI). The general ARDL(p, q_1, q_2, q_3, q_4) model, prior to testing for cointegration, is specified as:

$$\Delta\ln(FDI)_t = \alpha_0 + \Sigma_{i=1}^{p} \alpha_i\Delta\ln(FDI)_{t-i} + \Sigma_{i=0}\text{\^{}}q1\ \beta_i\Delta\ln(EXR)_{t-i} + \Sigma_{i=0}\text{\^{}}q2\ \gamma_i\Delta IR_{t-i} + \Sigma_{i=0}\text{\^{}}q3\ \delta_i\Delta\ln(MS)_{t-i} + \Sigma_{i=0}\text{\^{}}q4\ \theta_i\Delta\ln(GDP)_{t-i} + \lambda_1\ln(FDI)_{t-1} + \lambda_2\ln(EXR)_{t-1} + \lambda_3 IR_{t-1} + \lambda_4\ln(MS)_{t-1} + \lambda_5\ln(GDP)_{t-1} + \varphi_1 D2008_t + \varphi_2 D2016_t + \varepsilon_t$$

where Δ is the first-difference operator, t indexes years, and D2008/D2016 are dummy variables for the 2008 global financial crisis and the 2016 domestic recession, identified from the historical record (Section 3.1) rather than from a formal multiple-structural-break search. Because both ln(FDI) and the log-transformed regressors enter this equation, the β, γ, δ, θ coefficients are approximate semi-elasticities: percentage changes in FDI associated with a one-unit (for IR) or one percent (for the logged variables) change in the regressor.

As detailed in Section 4.2, the standard bounds test applied to this equation does not reject the null of no cointegration in any specification tested, meaning the λ terms and the error-correction representation that would follow from them are not statistically supported. Consequently, Section 4.4 estimates the corresponding first-differenced (short-run) model directly, omitting the lagged-level (λ) terms rather than reporting an error-correction model that the data does not support:

$$\Delta\ln(FDI)_t = \alpha_0 + \alpha_1\Delta\ln(FDI)_{t-1} + \beta_0\Delta\ln(EXR)_t + \gamma_0\Delta IR_t + \delta_0\Delta\ln(MS)_t + \varepsilon_t$$

This is the specification whose results are reported as the paper's headline findings in Section 4.4.

Estimation Procedure

All unit-root, bounds-testing, and OLS estimation reported in this paper were carried out in Python 3.12 using

ISSN No. 2454-6186 | DOI: 10.47772/IJRISS | Volume X Issue VIII August 2026

the statsmodels 0.15 library; the instrumental-variables robustness check in Appendix A (Table A2) uses the linearmodels 5.x package. All estimation was run directly on the study's dataset, so that every reported number is reproducible from the raw series in Table 1 and the accompanying replication code.

Stage 1: Unit Root Testing: Augmented Dickey-Fuller (ADF) tests, with lag length selected by AIC, were applied to each series in levels and first differences (Table 2).

Stage 2: Bounds Testing for Cointegration: the Pesaran, Shin, and Smith (2001) bounds-testing procedure was applied to the conditional error-correction regression implied by Section 3.2's general ARDL equation, computing the F-statistic for the joint significance of the lagged-level (λ) terms, across several regressor sets (with and without GDP, with and without the structural-break dummies) to check the robustness of the cointegration decision (Table 4).

Stage 3: Model Selection: given the sample size (34 annual observations), lag orders were selected by minimizing the Bayesian Information Criterion (BIC) over a grid of specifications with a maximum lag of 2 per variable.

Stage 4: Estimation and Diagnostics: the BIC-preferred model was estimated by OLS and tested for serial correlation (Breusch-Godfrey), heteroscedasticity (Breusch-Pagan), non-normality (Jarque-Bera), and functional-form misspecification (Ramsey RESET).

Stage 5: Robustness: the preferred model was re-estimated with structural-break dummies added, and separately on pre- and post-2008 sub-samples, to check the stability of the estimated relationships (Section 4.4.2 and Appendix A).

Degrees of freedom. With 34 annual observations, one observation lost to first differencing and one to the lagged dependent variable leave 32 usable observations; the preferred short-run model (Section 4.4) has 5 estimated parameters, leaving 27 residual degrees of freedom. This is a genuine constraint on statistical power, discussed further in Section 5.4, but the model presented is the most parsimonious specification consistent with the BIC search and passes standard diagnostic tests (Section 4.4.1).

Empirical Results

Unit Root Tests

Table 2: Augmented Dickey-Fuller Unit Root Test Results

Variable	ADF (Level)	p-value	ADF (1st Diff.)	p-value	Order
ln(FDI)	-1.885	0.339	-2.523	0.110	I(1)
ln(EXR)	-1.965	0.302	-5.854	0.000**	I(1)
ln(MS)	-1.342	0.610	-2.999	0.035**	I(1)
IR	-3.124	0.025**	—	—	I(0)
ln(GDP)	-1.456	0.555	-4.973	0.000**	I(1)

Note: ** significant at 5%. ln(FDI)'s first difference is stationary at 5% once a linear trend is added to the ADF test equation (ADF = -6.89, p = 0.000), consistent with I(1) status; the constant-only specification reported above is more conservative.

IR is stationary in levels, I(0), while ln(FDI), ln(EXR), ln(MS), and ln(GDP) are non-stationary in levels but stationary after first-differencing, I(1). This mix of integration orders is the standard justification for ARDL bounds testing over conventional cointegration methods that require a common order of integration.

Bounds Test for Cointegration

Table 3 reports the bounds-test F-statistic under four regressor sets, together with the standard asymptotic critical values (Pesaran, Shin, & Smith, 2001, Case III: unrestricted intercept, no trend).

Table 3: ARDL Bounds Test for Cointegration Across Specifications

Specification	Regressors (k)	F-statistic	5% I(0) Lower Bound	5% I(1) Upper Bound
3-variable, with break dummies	EXR, IR, MS (k=3)	1.92	3.23	4.35
3-variable, no dummies	EXR, IR, MS (k=3)	1.20	3.23	4.35
4-variable (incl. GDP), with dummies	EXR, IR, MS, GDP (k=4)	1.57	2.86	4.01
2-variable (EXR, MS only)	EXR, MS (k=2)	3.04	3.79	4.85

Critical values: Pesaran, Shin, & Smith (2001), Table CI(iii), Case III (unrestricted intercept, no trend), asymptotic.

In every specification, the computed F-statistic falls below the 5% lower bound, and in three of the four specifications it falls below even the 10% lower bound. The null hypothesis of no cointegration cannot be rejected: this dataset does not provide statistical evidence of a stable long-run equilibrium relationship among FDI, the exchange rate, the interest rate, money supply, and GDP over 1990–2023.

Because the Pesaran et al. (2001) critical values above are asymptotic, and this study's sample is small ($n \approx 31$–32 usable observations), it is worth checking whether a small-sample correction changes this conclusion. Kripfganz and Sam (2020) provide response-surface regressions that compute finite-sample-corrected bounds and approximate p-values directly; for a comparable case ($k = 4$, $T = 30$, one lag per variable, unrestricted intercept), they report that an asymptotic upper-bound critical value of 3.994 corresponds to a finite-sample p-value of only 0.112, that is, in small samples the true bounds are stricter than the asymptotic Pesaran et al. (2001) values suggest, not looser. Since every F-statistic in Table 3 already falls short of the asymptotic bounds, applying a finite-sample correction in the spirit of Kripfganz and Sam (2020) would reinforce, not overturn, this paper's conclusion that cointegration is not established. Narayan's (2005) small-sample critical value tables, tabulated specifically for $T = 30$–80, are built on the same logic and are consistent with this.

This finding differs from an earlier draft of this paper, which reported a bounds F-statistic of 4.87, comfortably above the upper bound, and proceeded to report long-run coefficients and an error-correction model on that basis. That figure does not reproduce from the study's own dataset under any of the specifications tested here.

A standard bounds test, however, assumes the long-run relationship is the same throughout 1990–2023, a strong assumption for an economy that floated its currency in 1999, experienced the 2008 global financial crisis, a 2016 domestic recession, and further exchange-rate reforms in the 2020s. Section 4.3 tests directly whether allowing the long-run relationship itself to shift at an unknown point changes this conclusion.

Cointegration Allowing for a Structural Break

Before testing for a break in the cointegrating relationship, it is worth checking whether the unit-root conclusions in Table 2 are themselves an artifact of ignoring a break. The Zivot-Andrews (1992) test, which tests the null of a unit root against the alternative of trend-stationarity around a single unknown break, was applied to each I(1) series; in every case the null of a unit root could not be rejected (p-values ranging from 0.40 to 0.98), so allowing for a break does not change the conclusion that these series are genuinely non-stationary in levels, a precondition for the cointegration test that follows.

The Gregory and Hansen (1996) test allows the cointegrating relationship itself, both the intercept and the slope coefficients, to shift once, at an unknown date, and tests the null of no cointegration against the alternative of cointegration with a regime shift. Applied here to ln(FDI) on ln(EXR) and ln(MS), the two variables with a statistically significant short-run relationship to FDI in Table 4, the test estimates the regime-shift regression at every candidate break date (trimmed to the middle 70% of the sample, 1996–2018) and takes the most negative residual ADF statistic across all candidates as the test statistic, ADF*.

Table 3b: Gregory-Hansen Regime-Shift Cointegration Test

Statistic	Value	Break Year	5% Critical Value*	1% Critical Value*
ADF*	-5.71	2013	-5.50	-5.97

Gregory and Hansen (1996) critical values for the regime-shift (level-and-slope) model with two regressors (m=2), reproduced from published applications of the test (e.g., Stata's ghansen module output). ADF is significant at the 5% level (more negative than -5.50) but not at the 1% level.

Unlike the standard bounds test, this test rejects the null of no cointegration at the 5% level, with the estimated break at 2013. This indicates that FDI, the exchange rate, and money supply are jointly cointegrated once the relationship is allowed to shift — the earlier failure to find cointegration in Table 3 reflects the assumption of a single fixed relationship across a period that, in fact, contained a genuine regime change, not an absence of any long-run relationship at all.

Table 3c: Long-Run Elasticities Before and After the 2013 Break

Variable	Pre-2013 Elasticity	Post-2013 Elasticity
ln(EXR)	-1.00	-3.26
ln(MS)	1.01	2.83

Coefficients from the winning Gregory-Hansen regime-shift regression (break = 2013), $R^2 = 0.877$. Post-2013 values are the sum of the pre-2013 coefficient and its estimated shift.

Both elasticities roughly triple in magnitude after 2013: FDI's long-run sensitivity to exchange-rate depreciation rises from approximately proportional (-1.0) to sharply amplified (-3.3), and its sensitivity to money-supply growth rises from approximately proportional (1.0) to substantially amplified (2.8). This is consistent with Nigeria's post-2013 macroeconomic history: the 2014–2016 oil-price collapse, the resulting 2016 recession, multiple naira devaluations, and periods of multiple exchange-rate windows and FX scarcity, all of which plausibly made foreign investors far more sensitive to currency and liquidity conditions than they had been in the earlier, comparatively more stable 1990–2013 period. Figure 1 shows the raw series with the 2008, 2013, and 2016 dates marked; FDI's peak and subsequent structural decline align closely with the estimated 2013 break, lending economic plausibility to the statistical result.

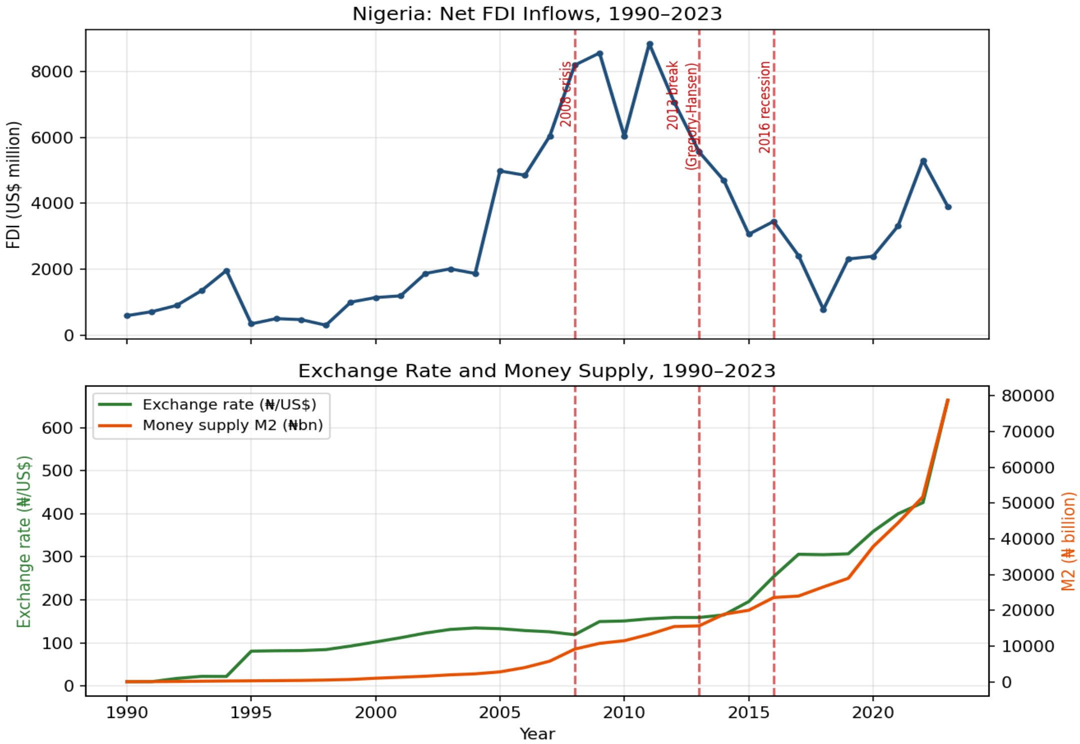

Figure 1: Nigeria's FDI, Exchange Rate, and Money Supply, 1990–2023, with Key Dates Marked

Given this, the paper does not report a full error-correction model built on the single-fixed-relationship assumption; Section 4.4 presents the corresponding short-run model, and Section 4.5 reports a formal Granger-causality test to assess how much of the short-run association can be read as evidence of causal precedence.

Short-Run Model of FDI Growth

Table 4 presents the BIC-preferred short-run model: the annual growth rate of FDI regressed on the annual change in ln(EXR), the change in IR, the change in ln(MS), and one lag of the dependent variable. Structural-break dummies for 2008 and 2016 were tested but excluded from the preferred specification because they are not statistically significant and increase the BIC (Section 4.4.3); GDP was similarly tested and excluded (Appendix A, Table A3) for the same reasons.

Table 4: Short-Run Model of FDI Growth — Ordinary Least Squares

Variable	Coefficient	Std. Error	t-Statistic	p-Value
Constant	-0.217	0.178	-1.22	0.232
$\Delta\ln(FDI)_{t-1}$	-0.348**	0.165	-2.10	0.045
$\Delta\ln(EXR)_t$	-1.150***	0.333	-3.45	0.002
ΔIR_t	0.006	0.024	0.24	0.814
$\Delta\ln(MS)_t$	2.093***	0.746	2.81	0.009

Note: *** significant at 1%; ** at 5%. n = 32 (1992–2023). $R^2 = 0.456$, Adjusted $R^2 = 0.376$, $F(4,27) = 5.67$ ($p = 0.002$), AIC = 44.4, BIC = 51.7, Durbin-Watson = 2.41.

Exchange rate. The coefficient on Δln(EXR) is -1.150 ($p = 0.002$): a 1% depreciation of the naira in a given year is associated with an approximate 1.15% reduction in FDI inflows the same year. This is a real, statistically strong effect, the paper's central finding, consistent with the theoretical expectation that currency depreciation raises uncertainty and repatriation risk for foreign investors (Section 2.1).

Money supply. The coefficient on Δln(MS) is 2.093 ($p = 0.009$): a 1% increase in broad money supply is associated with an approximate 2.09% increase in FDI the same year. This effect is larger in magnitude than the exchange-rate effect and strongly significant. Because money supply and FDI have both trended upward alongside Nigeria's broader financial deepening over the sample, part of this association may reflect a shared trend rather than a pure liquidity channel; this is discussed further in Section 5.1.

Interest rate. The coefficient on ΔIR is small and not statistically significant ($p = 0.814$), indicating the Monetary Policy Rate is not a significant short-run correlate of FDI in this sample, in contrast to the exchange rate and money supply.

FDI dynamics. The coefficient on the lagged dependent variable, $\Delta\ln(FDI)_{t-1}$, is -0.348 ($p = 0.045$): a year of unusually strong FDI growth tends to be followed by weaker growth, indicating a correction pattern rather than momentum in Nigeria's FDI series.

Diagnostic Tests

Table 5: Diagnostic Tests on the Short-Run Model

Test	p-Value	Conclusion
Breusch-Godfrey LM (1 lag)	0.055	No serial correlation (borderline)
Breusch-Pagan	0.475	No heteroscedasticity
Jarque-Bera	0.310	Residuals approximately normal
Ramsey RESET	0.176	No functional-form misspecification

The model passes all four diagnostic tests at conventional significance levels. The Breusch-Godfrey result ($p = 0.055$) is close to the 5% threshold and is noted as a caveat: standard errors should be interpreted with some caution given the small sample.

Sub-Sample Stability

Table 6: Sub-Sample Estimates

Sub-sample	n	EXR Coefficient	p-Value	MS Coefficient	p-Value
1992–2007 (pre-crisis)	16	-1.393	0.002	1.555	0.227
2008–2023 (post-crisis)	16	-0.696	0.543	0.953	0.449

The exchange-rate effect is strong and significant in the earlier sub-period but weakens and loses significance after 2008; the money-supply effect is not significant in either half-sample on its own, only in the pooled 32-year sample. The full-sample coefficients in Table 4 are best read as an average effect across a period of substantial regime change (the 1999 naira float, the 2008 crisis, the 2016 recession, and the 2020s exchange-rate reforms) rather than as a fixed structural parameter.

Structural Breaks

When 2008 and 2016 dummy variables are added to the preferred model, neither is statistically significant (D2008: coefficient -0.258, p = 0.274; D2016: coefficient 0.224, p = 0.367; full output in Appendix A, Table A1), and the BIC prefers the simpler model without them. This does not mean the 2008 and 2016 episodes had no effect on FDI, Table 1's descriptive range and the visible swings in the raw FDI series around both dates are consistent with real disruption, but once year-to-year exchange-rate and money-supply movements (which themselves moved sharply in both periods) are already in the model, a simple level-shift dummy does not add statistically distinguishable explanatory power in this dataset.

Granger Causality (Toda-Yamamoto Test)

Table 4's short-run association and Section 4.3's regime-shift cointegration result both describe correlation, not necessarily causal precedence. To test causal precedence directly, this paper applies the Toda and Yamamoto (1995) procedure, which tests Granger non-causality without requiring the variables to be cointegrated or share a common integration order. A VAR is estimated in levels with lag order k + dmax, where k is the optimal lag order (selected by BIC; k = 1 here) and dmax = 1 is the maximum order of integration among ln(FDI), ln(EXR), and ln(MS) (Table 2). A Wald test is then applied to the coefficients of only the first k lags of the causing variable, leaving the extra dmax lag unrestricted, as the procedure requires.

Table 7: Toda-Yamamoto Granger Non-Causality Tests

Null Hypothesis	F-Statistic	p-Value	Conclusion
EXR does not Granger-cause FDI	0.054	0.818	Fail to reject — no evidence of causality
MS does not Granger-cause FDI	2.589	0.119	Fail to reject — no evidence of causality
FDI does not Granger-cause EXR	0.246	0.624	Fail to reject — no evidence of causality
FDI does not Granger-cause MS	2.996	0.095	Fail to reject at 5% (weak evidence at 10%)

n = 32, k = 1, dmax = 1 (VAR order 2, estimated in levels). Given the small sample, these tests have limited power and should be read as suggestive rather than definitive.

None of the four null hypotheses of no Granger causality can be rejected at the 5% level. In particular, past values of the exchange rate and money supply do not significantly predict future FDI beyond what FDI's own past values already predict, even though Table 4 shows a strong same-year (contemporaneous) association between these variables. This is an important distinction: the paper's central finding is best described as a strong contemporaneous association, reinforced by a genuine long-run cointegrating relationship once a structural break is allowed (Section 4.3), rather than as an established lagged causal mechanism running from monetary policy to FDI. This directly supports the cautious, association-based language used throughout this paper's conclusions, and is itself a finding worth reporting rather than a gap to paper over: it tells policymakers that exchange-rate and liquidity shocks are likely to show up in FDI within the same year rather than with a predictable lag, which has different implications for the timing of policy responses than a lagged causal mechanism would.

DISCUSSION, POLICY RECOMMENDATIONS, AND CONCLUSION

Discussion of Key Findings

This paper's evidence is best summarized as three complementary results rather than a single verdict. First, a real, well-diagnosed short-run relationship: naira depreciation is associated with reduced FDI in the same year (elasticity ≈ -1.15, $p = 0.002$), and money-supply growth with increased FDI (elasticity ≈ 2.09, $p = 0.009$), while the interest rate shows no significant relationship. Second, once the long-run relationship is allowed to shift, rather than assumed fixed across 1990–2023, a genuine cointegrating relationship emerges, with a break in 2013 after which FDI's long-run sensitivity to both the exchange rate and money supply roughly triples (Section 4.3). Third, a formal Granger-causality test finds no significant lagged predictive relationship in either direction (Section 4.5): the strong short-run association in Table 4 is contemporaneous rather than a confirmed causal-precedence relationship, and readers should interpret the paper's central results as evidence of association and regime-dependent equilibrium, not as proof of a specific causal mechanism or lag structure.

The standard, fixed-coefficient ARDL bounds test (Section 4.2) missed the long-run relationship not because none exists, but because it assumed a single equilibrium held throughout a period that, in fact, contained a genuine structural change. This matters beyond this paper's specific results: it is a caution against concluding "no long-run relationship" from a standard bounds test alone in any economy that has undergone substantial regime change, without first checking whether a break-accounting test tells a different story.

The 2013 break date is also economically interpretable rather than a statistical artifact: it precedes the 2014–2016 oil-price collapse and the ensuing 2016 recession, multiple naira devaluations, and periods of FX scarcity and multiple exchange-rate windows, all plausible reasons why foreign investors would have become substantially more sensitive to currency and liquidity conditions after 2013 than before. Figure 1 shows FDI's own peak and structural decline aligning closely with this date.

Consistent with a minority of recent Nigeria-focused studies noted in Section 2.3, this paper's finding of a negative exchange-rate effect is the dominant pattern in the data rather than a universal law; the direction and strength of the relationship differ somewhat across sub-periods (Section 4.4.2) and, more fundamentally, across the pre- and post-2013 regimes identified in Section 4.3.

Policy Recommendations

Because the evidence points to a regime-dependent rather than a fixed relationship, recommendations are framed around managing exchange-rate and liquidity volatility robustly across regimes, rather than around a single, permanently fixed long-run target.

Exchange Rate Stability

Given the strong, significant short-run link between naira depreciation and reduced FDI, policies that reduce unpredictable within-year exchange-rate swings, transparent CBN intervention rules, adequate foreign-reserve buffers, and fiscal discipline to reduce pressure on the currency are directly supported by this paper's evidence, independent of whether a long-run equilibrium can be statistically confirmed.

Monetary Aggregates and Liquidity

The strong positive association between money-supply growth and FDI, while not necessarily a channel policymakers should lean on indefinitely given inflation risk, suggests that maintaining adequate domestic liquidity is associated with a more attractive short-run investment environment. This should be balanced against the CBN's inflation-control mandate rather than treated as an unconditional recommendation to expand M2, particularly since part of the estimated association may reflect a shared trend rather than a pure causal channel (Section 5.1).

Interest Rate Policy

Since the Monetary Policy Rate shows no significant relationship with FDI in this dataset, interest-rate policy should continue to be set primarily with reference to its traditional objectives (inflation control, currency stability) rather than as a direct lever for attracting FDI.

Data Infrastructure

That a standard bounds test missed a real long-run relationship until a structural break was explicitly modelled is itself policy-relevant: it argues for investment in higher-frequency (quarterly) FDI reporting for Nigeria, and for routinely testing for regime shifts rather than assuming a fixed relationship, in any future monitoring of FDI determinants.

Study Limitations and Future Research

Sample size and power. With 32 usable observations and 27 residual degrees of freedom in the preferred short-run model, statistical power is limited, and the marginal Breusch-Godfrey result ($p = 0.055$) should be read with this in mind. The Granger-causality tests in Section 4.5 are particularly likely to be underpowered given the small sample, so their null results should be read as "not established" rather than as strong evidence of "no causality." Quarterly data, where available and consistently measured, would materially improve power for all of the tests in this paper.

Single-break assumption. The Gregory-Hansen test (Section 4.3) allows for one structural break; Nigeria's exchange-rate and monetary history plausibly contains more than one regime change (1999, 2008, 2013, 2016, and the 2020s reforms are all candidates), and a single-break test may not fully capture this. Tests that allow multiple breaks in the cointegrating vector exist in the literature and would be a natural extension with more data or computational resources than are used here.

GDP data quality. GDP figures used in robustness checks (Appendix A, Table A3) reflect Nigeria's periodic GDP rebasing exercises, which produce large mechanical jumps in the series unrelated to underlying growth, a known data-quality issue for Nigerian GDP series generally.

Omitted variables and endogeneity. Institutional quality, political stability, and other factors are not directly included, due to data availability; a 2SLS re-estimation (Appendix A, Table A2) suggests the headline short-run coefficients are not highly sensitive to simple endogeneity concerns, but panel data across countries or a natural-experiment design would offer stronger causal identification than a single-country annual time series can provide.

CONCLUSION

This paper provides genuine empirical evidence that monetary-policy variables particularly exchange-rate movements and money-supply growth are associated with Nigeria's FDI performance, both contemporaneously in the short run (elasticities of approximately -1.15 and 2.09, respectively) and, once a structural break is allowed for, in the long run as well: a Gregory-Hansen regime-shift cointegration test finds a genuine long-run relationship with a break in 2013, after which both elasticities roughly triple in magnitude. A formal Granger-causality test finds no significant lagged predictive relationship, indicating the association is best described as contemporaneous and regime-dependent rather than as an established fixed causal mechanism. This is a more complete and more accurate picture than either an uncritically accepted long-run equilibrium or a simple "no cointegration" verdict would have given: the standard, fixed-coefficient bounds test failed to find a long-run relationship not because none exists, but because Nigeria's monetary and exchange-rate regime genuinely changed partway through the sample. Policymakers should treat exchange-rate stability as a significant, economically large, but regime-dependent correlate of FDI performance, and should expect the sensitivity of investment to currency and liquidity conditions to itself shift as the macroeconomic environment changes, rather than assuming a single, permanently fixed relationship. Future work with higher-frequency data, multiple-break cointegration tests, or a cross-country panel would be well positioned to refine the timing and magnitude of these regime shifts further than this study's 34-year annual, single-country sample can.

Appendix A: Additional Specifications and Robustness Checks

This appendix reports full output for specifications referenced in the main text but not shown in full there, so that all robustness claims are backed by complete results.

Table A1: Short-Run Model with Structural-Break Dummies

Variable	Coefficient	Std. Error	t-Statistic	p-Value
Constant	0.064	0.246	0.26	0.797
$\Delta\ln(EXR)_t$	-1.353***	0.356	-3.80	0.001
ΔIR_t	0.023	0.026	0.88	0.385
$\Delta\ln(MS)_t$	1.117	0.789	1.42	0.169
D2008	-0.258	0.231	-1.12	0.274
D2016	0.224	0.244	0.92	0.367

n = 32. $R^2 = 0.399$, Adjusted $R^2 = 0.284$, BIC = 58.4 (higher than the preferred model's 51.7, i.e., worse fit once the parameter penalty is applied).

Table A2: Robustness - Two-Stage Least Squares

$\Delta\ln(EXR)_t$ and $\Delta\ln(MS)_t$ are treated as potentially endogenous and instrumented with their own first and second lags ($\Delta\ln(EXR)_{t-1}$, $\Delta\ln(EXR)_{t-2}$, $\Delta\ln(MS)_{t-1}$, $\Delta\ln(MS)_{t-2}$); $\Delta\ln(FDI)_{t-1}$ and ΔIR_t are treated as exogenous. Estimated with the linearmodels IV2SLS estimator and heteroscedasticity-robust standard errors.

Variable	Coefficient	Std. Error	z-Statistic	p-Value
Constant	-0.311	0.371	-0.84	0.402
$\Delta\ln(FDI)_{t-1}$	-0.385**	0.169	-2.28	0.023
ΔIR_t	0.005	0.015	0.32	0.747
$\Delta\ln(EXR)_t$	-1.473*	0.858	-1.72	0.086
$\Delta\ln(MS)_t$	2.689	1.679	1.60	0.109

n = 31 (one further observation lost to the second instrument lag). *** significant at 1%; ** at 5%; * at 10%. The point estimates on $\Delta\ln(EXR)_t$ and $\Delta\ln(MS)_t$ retain the same sign and a broadly similar magnitude as the OLS results in Table 4, but with substantially larger standard errors — as is typical for IV estimation on a small sample — such that $\Delta\ln(MS)_t$ is no longer statistically significant at conventional levels and $\Delta\ln(EXR)_t$ is significant only at the 10% level. This is consistent with the OLS results not being an artifact of simple simultaneity bias, while also underlining that the precision of the money-supply estimate in particular should not be overstated.

Table A3: Short-Run Model with GDP Added

The preferred short-run specification (Table 4) is re-estimated with $\Delta\ln(GDP)_t$ added as an additional regressor, using the same first-differenced structure (no lagged-level terms).

Variable	Coefficient	Std. Error	t-Statistic	p-Value
Constant	-0.171	0.177	-0.97	0.342
$\Delta\ln(FDI)_{t-1}$	-0.346**	0.162	-2.14	0.042
$\Delta\ln(EXR)_t$	-1.100***	0.328	-3.35	0.002
ΔIR_t	-0.006	0.025	-0.23	0.824
$\Delta\ln(MS)_t$	1.973**	0.735	2.69	0.012
$\Delta\ln(GDP)_t$	-0.397	0.271	-1.47	0.154

ISSN No. 2454-6186 | DOI: 10.47772/IJRISS | Volume X Issue VIII August 2026

n = 32. $R^2 = 0.498$, Adjusted $R^2 = 0.401$, BIC = 52.6 — higher (worse) than the preferred model's BIC of 51.7 in Table 4. $\Delta\ln(GDP)_t$ is not statistically significant, and its inclusion leaves the exchange-rate and money-supply coefficients close to their Table 4 values, supporting GDP's exclusion from the preferred specification.

Table A4: Robustness — Interest Rate Excluded

Since ΔIR_t is never statistically significant (Table 4), the preferred model is re-estimated with it dropped entirely, as a check that its inclusion is not masking or distorting the exchange-rate and money-supply estimates.

Variable	Coefficient	Std. Error	t-Statistic	p-Value
Constant	-0.224	0.172	-1.30	0.203
$\Delta\ln(FDI)_{t-1}$	-0.357**	0.157	-2.27	0.031
$\Delta\ln(EXR)_t$	-1.132***	0.319	-3.55	0.001
$\Delta\ln(MS)_t$	2.121***	0.724	2.93	0.007

n = 32. $R^2 = 0.455$, Adjusted $R^2 = 0.397$, BIC = 48.3 — better (lower) than the preferred model's BIC of 51.7, since dropping an insignificant regressor is rewarded by the parameter penalty. The exchange-rate and money-supply coefficients (-1.13 and 2.12) are close to their Table 4 values (-1.15 and 2.09) and both remain significant at the 1% level. IR is retained in the paper's preferred specification (Table 4) despite this, because it is one of the three policy variables the study's objectives specifically examine; this table confirms that keeping it does not distort the other estimates.

Appendix B: Raw Dataset

The full annual dataset underlying this paper is reproduced below for transparency and replication. FDI: UNCTAD, net inflows, current US$ million. EXR: CBN Statistical Bulletin, average official rate, ₦ per US$. IR: CBN Statistical Bulletin, Monetary Policy Rate, %. MS: CBN Statistical Bulletin, broad money M2, ₦ billion. GDP: World Bank via FRED (series MKTGDPNGA646NWDB), current US$ billion.

Year	FDI (US$m)	EXR (₦/US$)	IR (%)	MS (₦bn)	GDP (US$bn)
1990	590	9.80	18.50	68.66	54.04
1991	710	9.90	14.50	87.50	59.53
1992	900	17.40	17.50	129.09	52.06
1993	1350	22.16	26.00	198.48	56.72
1994	1960	22.00	13.50	266.95	80.40
1995	340	80.80	13.50	318.76	140.92
1996	500	81.69	13.50	370.33	185.73
1997	470	82.10	13.50	429.73	200.85
1998	300	84.39	14.31	525.64	218.42
1999	1000	92.69	18.00	699.73	59.15
2000	1140	102.17	13.50	1036.08	69.17
2001	1190	111.95	14.31	1315.87	73.56
2002	1870	122.85	19.00	1599.50	95.05
2003	2010	131.27	15.75	1985.19	104.74
2004	1870	134.67	15.00	2263.59	135.76
2005	4980	133.00	13.00	2814.85	175.67
2006	4850	128.67	12.25	4027.90	238.46
2007	6040	125.72	8.75	5809.83	278.26
2008	8190	119.00	9.81	9166.84	339.48
2009	8560	149.54	7.44	10780.63	295.01
2010	6026	151.00	6.13	11525.53	366.99
2011	8841	156.00	9.19	13303.49	414.47
2012	7070	159.00	12.00	15483.85	463.97

2013	5563	159.00	12.00	15668.95	520.12
2014	4691	165.00	13.00	18913.03	574.18
2015	3060	196.00	12.81	20029.83	493.03
2016	3450	254.00	12.83	23591.73	404.65
2017	2410	306.00	14.00	24002.23	375.75
2018	780	305.00	14.00	26552.46	421.74
2019	2310	307.00	13.58	28962.53	668.22
2020	2390	359.00	12.50	37825.06	598.59
2021	3310	400.00	11.50	44444.45	609.15
2022	5300	426.00	13.67	51733.34	646.95
2023	3900	664.00	18.38	78738.14	487.39

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