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Government Budget Deficits and Price Stability in Nigeria: An ARDL Bounds Testing Approach, 1981–2024

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Abstract

Persistent fiscal deficits in Nigeria have coincided with sustained inflationary instability, yet rigorous, methodologically transparent empirical evidence on the deficit–inflation nexus over an extended sample that incorporates the structural reforms of 2023–2024 remains scarce. This study addresses that gap by examining the effect of government budget deficits on consumer price inflation in Nigeria from 1981 to 2024 using 44 annual observations drawn from the Central Bank of Nigeria Statistical Bulletin and the World Bank World Development Indicators. Following a comprehensive battery of unit-root tests (ADF, Phillips–Perron, KPSS) supplemented by the Zivot–Andrews structural-break procedure that classified all variables as integrated of order one, the Autoregressive Distributed Lag (ARDL) bounds testing approach of Pesaran, Shin, and Smith (2001) is applied, with optimal lag selection undertaken via the Akaike Information Criterion and corroborated by the Schwarz Bayesian and Hannan–Quinn criteria. The selected ARDL(1, 1, 0, 1) specification generates an ARDL bounds F-statistic of 7.351, decisively rejecting the null of no cointegration at the 1% level against both Pesaran et al. (2001) and Narayan (2005) finite-sample critical values. In the long run, a one-percentage-point increase in the deficit-to-GDP ratio is associated with a 3.55-percentage-point increase in the Consumer Price Index inflation rate ($p = .042$, 95% CI excludes zero), the study's most robust and policy-relevant finding. The error-correction term of −0.689 ($p < .001$) indicates that approximately 68.9% of any short-run disequilibrium is corrected within one year, while Granger causality tests confirm unidirectional temporal precedence from deficits to inflation ($F = 17.645$, $p = .0001$). The model satisfies the Breusch–Godfrey serial-correlation test, the Breusch–Pagan–Godfrey heteroskedasticity test, the Jarque–Bera normality test, the Ramsey RESET specification test, and the CUSUM and CUSUMSQ stability tests, confirming parameter constancy across successive macroeconomic regimes. The evidence is consistent with the Neoclassical and monetarist predictions of deficit-induced inflation and supports the Fiscal Theory of the Price Level interpretation whereby unbacked nominal liabilities force the price level to adjust. Policy responses should prioritise statutory fiscal rules, entrenched central-bank independence against deficit monetisation, and non-oil revenue diversification.

Keywords: *Budget Deficit; Inflation; Price Stability; ARDL Bounds Test; Cointegration; Nigeria.*

1. Introduction

Three macroeconomic objectives; sustainable growth, productive employment, and price stability, anchor the policy architecture of every modern state. Price stability, the focus of the present paper, preserves the real value of income and contracts, anchors inflation expectations, and supports the efficient transmission of monetary policy. In Nigeria, the maintenance of price stability has proved persistently elusive. Headline inflation averaged 19.81% over 1981–2024, breached 70% during the 1995 hyperinflationary episode, and reached 34.80% by December 2024, a level that systematically exceeds the Central Bank of Nigeria's implicit 9% upper target (Central Bank of Nigeria [CBN], 2024; National Bureau of Statistics [NBS], 2024).

Coincident with this inflationary record, government budget deficits have widened dramatically, from approximately 3.30% of GDP on average across the sample to 3.7% of GDP in 2024, breaching the three-per-cent statutory ceiling stipulated by the Fiscal Responsibility Act (2007). Total Federal Government revenue of ₦21.0 trillion in 2024 was insufficient to cover expenditure of ₦35.02 trillion, while debt servicing absorbed approximately 47% of total outlays (Budget Office of the Federation, 2024; Debt Management Office [DMO], 2024). The sustained deficit–inflation co-movement observed in

the aggregate data motivates the core research question addressed in this paper: does the Nigerian fiscal deficit exert a statistically significant long-run effect on consumer price inflation once the conduct of monetary policy and the trajectory of broad money are controlled for?

Theoretical debate on this question remains unresolved. The Keynesian framework permits deficit spending to stimulate aggregate demand without necessarily generating inflation when output is below potential (Keynes, 1936). The Neoclassical school anticipates that persistent deficits generate inflationary pressure via either resource crowding-out or the seigniorage channel under monetisation (Bernheim, 1989; Elmendorf & Mankiw, 1999). The Ricardian Equivalence Hypothesis treats deficits as macroeconomically neutral (Barro, 1974). The Fiscal Theory of the Price Level (FTPL) advanced by Leeper (1991) and developed by Sims (1994), Woodford (1995), and Cochrane (1998, 2023) provides a fourth and complementary perspective in which the price level adjusts endogenously to equilibrate the intertemporal budget constraint whenever debt issuance is not credibly backed by future surpluses. Sargent and Wallace (1981) demonstrated that even non-monetised domestic deficits ultimately generate inflation once treasury–central bank coordination forces monetary expansion, a trilemma that retains empirical relevance to Nigeria's reliance on statutory "Ways and Means" advances.

Empirical evidence specific to Nigeria is fragmented and methodologically uneven. Earlier studies (Onwioduokit, 1999; Chinweobo, 2013; Ifeakachukwu & Adebiyi, 2017) reported a positive deficit–inflation relationship, frequently employing small samples and a single cointegration technique. More recent contributions (Nwosa et al., 2024; Tule et al., 2019) reaffirm the deficit–inflation transmission while noting the role of money-supply expansion and exchange-rate depreciation as proximate mechanisms. Three gaps persist: (i) the temporal coverage of most published studies terminates before the structural reforms of 2023–2025; (ii) the integration properties of the variables are rarely tested with complementary ADF, Phillips–Perron, and KPSS procedures supplemented by structural-break testing; and (iii) the diagnostic and stability adequacy of estimates is seldom reported with full transparency.

This paper addresses these gaps by applying the ARDL bounds testing approach of Pesaran et al. (2001) to the deficit–inflation relationship in Nigeria over 1981–2024, the most temporally comprehensive sample assembled for the Nigerian case to date. It integrates a multi-test unit-root battery (Said & Dickey, 1984; Phillips & Perron, 1988; Kwiatkowski, Phillips, Schmidt, & Shin, 1992; Zivot & Andrews, 1992), provides the linked error-correction representation, conducts Granger causality tests, and reports a comprehensive diagnostic and stability battery. The remainder of the paper is organised as follows: Section 2 reviews the theoretical and empirical literature; Section 3 presents the data, variables, and methodological framework; Section 4 reports the empirical results; Section 5 discusses the findings in relation to theory and prior evidence; and Section 6 concludes with policy implications.

2. Literature Review

The fiscal–inflation nexus has been the subject of extensive theoretical inquiry. The Keynesian transmission mechanism (Keynes, 1936; Hicks, 1937; Mundell, 1962) presumes short-run price stickiness and renders inflation an outcome of demand pressures that arise only at high levels of resource utilisation. The Neoclassical synthesis of Elmendorf and Mankiw (1999) treats persistent deficits as a source of upward pressure on real interest rates, with subsequent crowding-out of investment and, where deficit monetisation is operative, of the price level. Bernheim (1989) demonstrates that under monetisation the seigniorage channel delivers a quantitatively significant inflationary consequence, the formal condition for which is the absence of credible central-bank independence from the fiscal authority.

The FTPL formalises a complementary channel: when fiscal policy is "active" (does not adjust primary surpluses to stabilise the real value of outstanding debt) and monetary policy is "passive," the price level becomes the residual variable that equilibrates the intertemporal government budget constraint (Leeper, 1991; Sims, 1994; Woodford, 1995; Cochrane, 1998, 2023). Under the FTPL, monetary expansion is not a precondition: the mere issuance of nominal liabilities without a credible future-surplus backing forces the price level to rise. Bajo-Rubio et al. (2009) provide an applied implementation relevant to developing economies. Sargent and Wallace (1981)'s "unpleasant monetarist arithmetic" demonstrates an equivalent result under a rational-expectations monetary regime in which the central bank delays inflation but cannot ultimately prevent it.

Empirical investigations of the Nigerian deficit–inflation nexus have produced convergent findings. Onwioduokit (1999) documented bidirectional short-run but unidirectional long-run causality from deficits to inflation. Chinweobo (2013) and Ifeakachukwu and Adebiyi (2017) confirmed a positive long-run relationship with strong transmission through the monetary channel. Nwosa et al. (2024) reported a significant positive deficit–inflation coefficient in cointegration analyses spanning four decades, while Tule et al. (2019) highlighted money-supply expansion as the principal proximate channel. These studies motivate the present re-examination within a rigorously specified ARDL bounds-testing framework over the most temporally comprehensive sample assembled for the Nigerian case to date, thereby extending both the currency and the methodological transparency of the existing evidence base.

3. Methodology

3.1 Data and Variables

Annual data covering fiscal years 1981–2024 (N = 44) are drawn from the Central Bank of Nigeria Statistical Bulletin (2024), the National Bureau of Statistics (2024), and the World Bank World Development Indicators (2024). The dependent variable is PRSP, the annual percentage change in the Consumer Price Index, operationalised as the inverse measure of price stability. The principal regressor of theoretical interest is BD, the consolidated federal budget deficit expressed as a percentage of GDP. The control variables are BR, the end-of-period Central Bank of Nigeria policy (bank) rate, and MS, broad money (M2) expressed as a percentage of GDP.

Table 1. Variable Definitions and Measurement

Variable	Symbol	Definition	Unit / Measurement
Price stability	PRSP	Annual percentage change in consumer price index	Percent; annual
Budget deficit	BD	(Government expenditure − government revenue) / GDP	Percent of GDP; annual
Bank rate	BR	Central Bank of Nigeria monetary policy rate	Percent p.a.; annual
Money supply	MS	Broad money (M2) / GDP	Percent of GDP; annual

Source. Author's compilation based on CBN (2024), NBS (2024), and World Bank WDI (2024).

3.2 Descriptive Statistics

Table 2 reports the descriptive statistics for the four variables. Inflation exhibits the greatest dispersion (SD = 16.54; skewness = 1.653), reflecting hyper-inflationary episodes of the early 1990s and the post-2020 re-acceleration. The budget-deficit distribution is symmetric around a mean of 3.30% of GDP with a maximum of 9.10%.

Table 2. Descriptive Statistics

Variable	N	M	Mdn	Max	Min	SD	Skew.	Kurt.
PRSP (%)	44	19.805	13.100	72.800	4.700	16.544	1.653	4.510
BD (% GDP)	44	3.302	3.000	9.100	0.100	2.258	0.489	2.480
BR (%)	44	13.365	13.250	27.250	6.000	4.328	1.044	3.890
MS (% GDP)	44	18.609	18.150	31.200	10.800	4.381	0.700	2.940

Note. N = number of observations; M = mean; Mdn = median; SD = standard deviation. Source. Author's computation using EViews 13.

3.3 Estimation Framework

Following a rigorous preliminary unit-root testing protocol that confirmed all four variables are integrated of order one, employing the Augmented Dickey–Fuller (Said & Dickey, 1984), Phillips–Perron (Phillips & Perron, 1988), and Kwiatkowski–Phillips–Schmidt–Shin (Kwiatkowski et al., 1992) tests, supplemented by the Zivot–Andrews (1992) procedure to allow for a single endogenous structural break, the ARDL bounds testing approach of Pesaran, Shin, and Smith (2001) is applied. The ARDL framework is particularly appropriate for the present sample size (N = 44) because it remains valid irrespective of whether the regressors are purely I(0), purely I(1), or mutually cointegrated, and it yields consistent estimates in small samples when the lag order is correctly selected.

The conditional error-correction representation associated with the long-run cointegrating relationship is specified as:

$$\Delta PRSP_t = \alpha_0 + \Sigma_i \alpha_i \Delta PRSP_{t-i} + \beta_0 BD_t + \Sigma_j \beta_j \Delta BD_{t-j} + \gamma_0 BR_t + \Sigma_n \gamma_n \Delta BR_{t-n} + \delta_0 MS_t + \Sigma_p \delta_p \Delta MS_{t-p} + \lambda ECT_{t-1} + \varepsilon_t \quad \ldots\ldots\ldots(3.1)$$

where λ denotes the speed-of-adjustment coefficient (expected to lie within (−1, 0)); ECT_{t-1} is the lagged residual of the long-run cointegrating regression; and ε_t is a white-noise disturbance. The

theoretical coefficient of interest is β_0, the long-run multiplier on the deficit-to-GDP ratio, which theoretical priors and the Nigerian empirical literature (Nwosa et al., 2024; Tule et al., 2019) lead us to expect to be positive and statistically significant under the Neoclassical and monetarist frameworks.

Optimal lag-order selection is undertaken by minimising the Akaike Information Criterion (AIC) within a candidate ARDL(p, q_1, q_2, q_3) specification space reported in Table 3, with the Schwarz Bayesian Criterion (SBC) and the Hannan–Quinn Criterion (HQC) reported for cross-validation.

Table 3. Lag-Order Selection Criteria

Specification	AIC	SBC	HQC	Selection
ARDL(1, 1, 0, 1)	6.754*	6.968	6.842	Selected
ARDL(1, 1, 1, 1)	6.812	7.012	6.893	
ARDL(2, 1, 1, 1)	6.911	7.214	7.041	
ARDL(2, 2, 1, 1)	7.002	7.315	7.142	

*Note. * denotes the AIC-minimising candidate. The selected specification, ARDL(1, 1, 0, 1), enters PRSP, BD, and MS with one lag each and BR at the contemporaneous level (zero lags). Source. Author's computation using EViews 13.*

Cointegration is tested via the ARDL bounds procedure against the Case III critical values of Pesaran et al. (2001), corroborated by the Narayan (2005) finite-sample refinement for samples of 30–45 observations. Diagnostic adequacy is evaluated through the Breusch–Godfrey LM serial-correlation test, the Breusch–Pagan–Godfrey heteroskedasticity test, the Jarque–Bera normality test, the Ramsey RESET specification test, and the CUSUM and CUSUMSQ structural-stability tests (Brown, Durbin, & Evans, 1975). All estimations are performed using EViews 13.

4. Results

4.1 Cointegration: ARDL Bounds Test

Table 4 reports the ARDL bounds test for cointegration in Model 3 (PRSP = f(BD, BR, MS)). The computed F-statistic of 7.351 exceeds the upper I(1) critical bound of 4.66 at the 1% level of significance, decisively rejecting the null hypothesis of no cointegration at every conventional significance level. The same conclusion is reached when the Narayan (2005) finite-sample critical values appropriate for N ≈ 40–45 are employed, reinforcing the robustness of the cointegration finding.

Table 4. ARDL Bounds Test for Cointegration

Test statistic	Value	Significance	I(0) bound	I(1) bound
F-statistic (Wald)	7.351**	10%	2.37	3.20
k (regressors)	3	5%	2.79	3.67
Specification	ARDL(1,1,0,1)	2.5%	3.15	4.08
Adjusted obs.	43	1%	3.65	4.66
Decision	Reject null at 1%			

*Note. ** denotes rejection of the no-cointegration null at the 1% level. Critical values follow Pesaran et al. (2001) Case III; same decision confirmed by Narayan (2005) finite-sample critical values. Source. Author's computation using EViews 13.*

4.2 Long-Run Coefficients

Table 5 reports the estimated long-run equilibrium coefficients. The model explains approximately 58% of the variation in long-run inflation (R^2 = .5843; adjusted R^2 = .5150; $F(3, 39) = 8.432$, $p < .001$).

Table 5. Long-Run Coefficients

Variable	Coefficient	SE	t-statistic	p-value
BD	3.5497	1.6850	2.1067	.0422**
BR	1.3557	0.7327	1.8504	.0725*
MS	−0.0280	0.7346	−0.0381	.9698
Constant	−7.5747	16.0690	−0.4714	.6402

*Note. ** p < .05; * p < .10. R^2 / Adj. R^2 / $F(3,39)$ = .5843 / .5150 / 8.432 (p < .001). Source. Author's computation using EViews 13.*

The coefficient on BD is positive and statistically significant at the 5% level (β = 3.5497, p = .0422, 95% CI approximately [.014, 7.085]), providing the principal empirical finding of this paper: in the long-run equilibrium, a one-percentage-point rise in the deficit-to-GDP ratio is associated with a 3.55-percentage-point increase in the CPI inflation rate. The coefficient on BR is positive and marginally significant at the 10% level (β = 1.3557, p = .0725), consistent with cost-push transmission through the

monetary-policy rate. The coefficient on MS is statistically indistinguishable from zero (β = −0.0280, p = .9698), providing evidence that once the budget deficit and the policy rate are controlled for, broad money has no independent long-run effect on inflation—a result consistent with the theoretical proposition that monetary aggregates in Nigeria are endogenous to fiscal conditions.

4.3 Short-Run Error Correction Representation

Table 6 reports the associated error-correction model (ECM), which characterises the short-run dynamics and the speed of adjustment toward long-run equilibrium.

Table 6. Error-Correction Model

Variable	Coefficient	SE	t-statistic	p-value
ΔBD	−0.7953	0.9136	−0.8706	.3898
ΔMS	−3.7594	1.8566	−2.0249	.0503*
ECM_{t-1}	−0.6891	0.1078	−6.3907	.0000***
Constant	1.2416	1.8124	0.6851	.4973

Note. ECM_{t-1} *= lagged error-correction term. ΔBR is absent from the short-run component owing to the ARDL(1, 1, 0, 1) specification assigning zero short-run lags to BR. *** p < .01; * p < .10.* R^2 */ Adj.* R^2 */ DW = .509 / .485 / 2.119. Source. Author's computation using EViews 13.*

The error-correction term is negative, large in absolute magnitude, and highly significant (λ = −0.6891, t = −6.391, p < .001), confirming both the existence of a stable long-run cointegrating relationship and dynamic stability of the model. The magnitude implies that approximately 68.91% of any deviation from the long-run price-stability equilibrium is corrected within one year. The short-run coefficient on ΔBD is statistically insignificant (β = −0.7953, p = .3898), confirming that contemporaneous changes in the budget deficit do not immediately affect the price level, the principal deficit effect operates through the long-run equilibrium relationship rather than through contemporaneous dynamics.

4.4 Granger Causality

Budget deficits unidirectionally Granger-cause inflation with extreme statistical significance (F = 17.645, p = .0001), establishing temporal precedence of fiscal imbalance on price deterioration. The reverse direction, inflation Granger-causing deficits is not supported (F = 1.187, p = .2824). Further, inflation Granger-causes money supply (F = 5.047, p = .0303), supporting the proposition advanced by Sargent and Wallace (1981) that monetary aggregates in fiscally dominated regimes are endogenous to inflationary dynamics rather than independently causal.

Table 7. Pairwise Granger Causality Tests

Null hypothesis	F-statistic	p-value	Decision
BD ⇢ PRSP	17.645	.0001***	Reject (BD → PRSP unidirectional)
PRSP ⇢ BD	1.187	.2824	Accept null
BR ⇢ PRSP	1.015	.3198	Accept null
MS ⇢ PRSP	0.029	.8658	Accept null
PRSP ⇢ MS	5.047	.0303**	Reject null (PRSP → MS)
BD ⇢ MS	3.347	.0748*	Reject at 10% (BD → MS)

*Note. ⇢ denotes "does not Granger-cause." Lag order = 1 (selected via VAR lag criteria). 43 observations (1982–2024). ***p < .01; **p < .05; *p < .10. Source. Author's computation using EViews 13.*

4.5 Diagnostic and Stability Tests

The model passes all five diagnostic tests at conventional significance levels, supporting the reliability of statistical inference. Parameter stability across all macroeconomic regimes of the 1981–2024 sample is particularly important, as it indicates that the long-run deficit–inflation relationship has remained structurally invariant through the Structural Adjustment Programme of the 1980s, the debt-relief period of the 2000s, the 2016 recession, the COVID-19 pandemic shock, and the 2023–2025 structural reforms.

Table 8. Diagnostic and Stability Tests

Diagnostic test	Statistic	Decision
Breusch–Godfrey LM (serial correlation)	F(2, 39) = 1.243, p = .303	No serial correlation
Breusch–Pagan–Godfrey (heteroskedasticity)	F(7, 39) = 0.876, p = .537	Homoskedastic
Jarque–Bera (residual normality)	χ^2 = 2.418, p = .299	Normal residuals

Diagnostic test	Statistic	Decision
Ramsey RESET (functional form)	F(1, 38) = 0.524, p = .475	Correctly specified
CUSUM and CUSUMSQ (parameter stability)	Within 5% bounds throughout	Stable parameters

Note. CUSUM and CUSUMSQ plots are retained in Appendix A. Source. Author's computation using EViews 13.

4.6 Formal Hypothesis Test

The null hypothesis H_{03} (the budget deficit exerts no statistically significant long-run effect on the Nigerian inflation rate) is decisively rejected across all four complementary tests: the ARDL bounds cointegration test (F = 7.351 > upper 1% bound of 4.66), the long-run coefficient test (β = 3.5497, p = .0422), the error-correction term test (λ = −0.6891, p < .001), and the Granger causality test (F = 17.645, p = .0001). The alternative hypothesis H_{13} is accepted.

5. Discussion

The positive long-run coefficient on the budget deficit (β = 3.5497, p = .0422) is the principal finding of this paper. A one-percentage-point increase in the deficit-to-GDP ratio corresponds to a 3.55-percentage-point increase in the long-run CPI inflation rate. Three interpretations are consistent with this finding.

First, the Neoclassical crowding-out and monetarist interpretation aligns with the present estimate. Persistent deficits must either be financed by domestic borrowing, which compresses the pool of loanable funds available to the private sector and lifts real interest rates, or be accommodated by monetary expansion through "Ways and Means" advances or other statutory instruments. Either path is inflationary. The structurally large budget deficits documented over 1981–2024, the elevated money-supply share of GDP, and the absence of independent inflation outcomes from MS once BD and BR are controlled for, jointly support the proposition that fiscal imbalance has been the proximate inflationary force in Nigeria. The insignificant coefficient on MS, in particular, indicates that monetary aggregates have not behaved exogenously with respect to inflation; rather, they appear to operate as a conduit through which the fiscal imbalance is propagated to the price level.

Second, the FTPL interpretation is consistent with the long-run coefficient and the unidirectional Granger causality from deficits to inflation. Under the FTPL, the price level is the residual that equilibrates the intertemporal government budget constraint (Leeper, 1991; Sims, 1994; Woodford, 1995). Nigeria's reliance on statutory Ways and Means advances, elevated debt-service ratios, and periodic exchange-rate crises are precisely the conditions in which FTPL channels are activated: where future primary surpluses do not credibly back outstanding nominal liabilities, the price level must rise to restore equilibrium. The decisive Granger-causal ordering (deficits → inflation) is consistent with the FTPL directional prediction.

Third, the magnitude of the deficit–inflation coefficient exceeds plausible values that could be generated purely through Keynesian demand pressure (which requires high capacity utilisation and operates only in the short run). The Keynesian framework predicts that deficit-financed expenditure stimulates output without generating sustained inflation when there is substantial economic slack. Nigeria's 44-year deficit–inflation relationship has operated in a sustained manner and is consistent with a long-run structural channel of transmission rather than with a short-run cyclical pressure. The insignificance of the short-run ΔBD coefficient is also consistent with the long-horizon mechanism implied by both Neoclassical crowding-out and FTPL channels: fiscal deficits affect the price level through cumulative adjustment rather than through instantaneous shocks.

Comparison with the prior Nigerian empirical literature is favourable. Onwioduokit (1999), Chinweobo (2013), Ifeakachukwu and Adebiyi (2017), Nwosa et al. (2024), and Tule et al. (2019) all document positive long-run deficit–inflation relationships with statistical significance in similar magnitudes to that recovered in the present study. The present estimate's 3.55-percentage-point long-run elasticity is broadly consistent with these earlier findings while extending the sample through the structural reforms of 2023–2024 and applying a more rigorous diagnostic and stability protocol. The result strengthens the cumulative empirical case for the primacy of fiscal imbalance in Nigeria's inflation dynamics.

6. Conclusion and Policy Implications

This paper has examined the long-run effect of government budget deficits on the Consumer Price Index inflation rate in Nigeria over the 1981–2024 period using the ARDL bounds testing approach. By

assembling the most temporally comprehensive sample to date and subjecting the estimates to a full battery of unit-root, cointegration, diagnostic, and stability tests, the study contributes both novel empirical evidence and enhanced methodological transparency to the Nigerian fiscal–inflation literature.

The principal empirical finding is that a one-percentage-point increase in the deficit-to-GDP ratio is associated with a 3.55-percentage-point increase in long-run CPI inflation ($\beta = 3.5497$, $p = .0422$). Cointegration is confirmed ($F = 7.351$, $p < .01$), the error-correction term indicates that approximately 68.9% of disequilibrium is corrected within one year ($\lambda = -0.6891$, $p < .001$), and Granger causality establishes that deficits unidirectional Granger-cause inflation ($F = 17.645$, $p = .0001$). The model passes all standard diagnostic and stability tests.

The evidence is consistent with the Neoclassical crowding-out and monetarist predictions, and with the Fiscal Theory of the Price Level, while inconsistent with the Keynesian multiplier hypothesis and the Ricardian Equivalence Hypothesis. Sustained deficit financing in Nigeria operates as a primary driver of inflation rather than as a neutral or output-stimulating instrument.

Three policy implications emerge. First, the Central Bank of Nigeria's statutory ceiling of three per cent of GDP on the federal budget deficit should be enforced through binding institutional mechanisms, including an independent fiscal council empowered to publish compliance assessments. Sustained breaches of this ceiling, as in the 2024 deficit of 3.7% of GDP, should trigger automatic expenditure-containment rules rather than political expediency. Second, the statutory prohibition on central-bank financing of fiscal deficits ("Ways and Means" advances) should be entrenched against short-term override to forestall the FTPL-type equilibrium in which the price level adjusts to clear the intertemporal budget constraint. Third, the deficit–inflation elasticity identified here implies that closing the non-oil revenue gap through tax-base broadening, formal-sector incorporation, and digital revenue infrastructure is a more inflation-neutral stabilisation strategy than recourse to additional expenditure compression in productive sectors.

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