

Fiscal Dominance, Structural Rigidities and the Political Economy of Inflation in India: From Nehruvian Planning to the Monetary Policy Framework*

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India's persistent inflation reflects fiscal dominance, structurally distorted agricultural markets and the primacy of distributional politics over price stability. Using ARDL bounds testing, Toda–Yamamoto Granger precedence, Bai-Perron structural break analysis (unconditional and conditional) and GARCH(1,1) with break dummies on both the conditional mean and variance, applied to quarterly data spanning 1955 Q1–2024 Q4, the paper identifies three structural determinants of inflation persistence: fiscal deficits operating through channels distinct from monetisation; minimum support price policies transmitting cost-push shocks via the high food share in the CPI; and inertial expectations. The 2014 Q3 transition coinciding with inflation-targeting adoption constitutes a statistically identifiable regime shift in both the level and conditional volatility of inflation. Framework durability depends on continued fiscal consolidation and agricultural reform.

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1. Introduction

India's inflation record since independence presents a paradox. The country has never experienced hyperinflation of the Argentine variety; *Blejer (2025)* documents how Argentina's peso lost thirteen zeros between 1945 and 1992. Yet India has struggled to sustain price stability below 5 per cent for any prolonged period. From the chronic single-digit inflation of the planned-economy era, through the double-digit food-price surges of 2009–2014, to the more recent disinflation under inflation targeting (IT), India's relationship with price stability has been one of managed coexistence rather than decisive resolution.

This pattern is not primarily the product of monetary incompetence, as the Reserve Bank of India (RBI) has been a technically sophisticated institution throughout most of its history. Instead, inflation persistence reflects structural and political-economic conditions that repeatedly constrain monetary effectiveness: the primacy of fiscal over monetary objectives; the political entrenchment of agricultural price-support policies that transmit cost-push pressures into the consumer price index; and the asymmetric federal structure that distributes inflationary spending across twenty-eight states, while concentrating anti-inflationary responsibility in the Union government and the RBI.

Two further structural features merit explicit recognition. First, the banking system is dominated by Public Sector Banks (PSBs), which hold roughly two-thirds of system assets and operate under non-commercial mandates that attenuate the interest-rate channel. *Mishra and Montiel (2013)* find that policy-rate-to-lending-rate pass-through in India is among the weakest in Asia, attributable in substantial part to balance-sheet stress, directed-lending obligations and administered deposit pricing in the PSB-dominated market. Second, the share of food and beverages in the all-India CPI basket is approximately 45.86 per cent, far higher than in advanced economies (typically 10–15 per cent) and most emerging market peers. This large food weight magnifies the MSP (Minimum Support Price) -driven cost-push channel identified below and constrains the central bank's ability to deliver target-consistent outcomes without disproportionate output costs. The interaction between weak banking-channel transmission and high pass-through from administered agricultural prices constitutes the operational environment within which inflation targeting must function.

The comparative literature on emerging market inflation persistence identifies two broad classes of explanation: the fiscal theory of the price level, in which persistent deficits ultimately require monetary accommodation regardless of central bank preferences (*Sargent – Wallace 1981; Cochrane 2023*); and structural theories emphasising supply-side rigidities, backward-looking price setting and the political economy of reform resistance (*Akerlof et al. 2000; Blejer 2025*). The Indian case

requires both: the fiscal and supply-side channels are empirically distinguishable, statistically significant and politically reinforcing.

This paper makes five contributions. First, it applies a multi-method time-series framework to quarterly data spanning 1955 Q1–2024 Q4: ARDL bounds testing, VAR-based Granger precedence, Bai-Perron structural break analysis (unconditional and conditional) and GARCH modelling with break dummies on both conditional mean and variance. This complements recent work by *Rastogi and Kanoujiya (2024)*, who model post-IT inflation volatility with attention to commodity and exchange-rate spillovers. The present paper differs in its joint integration of long-run cointegration, multivariate Granger precedence, endogenous break dating and conditional volatility within a unified empirical strategy. Second, it identifies the MSP index as a statistically significant long-run determinant of CPI inflation that remains significant when monetary aggregates are controlled for, giving direct empirical content to the structural food-price channel argued thus far on institutional grounds (*Gokarn 2011; Rangarajan – Mishra 2013*). Third, it documents that inflation targeting is associated with a statistically identifiable shift in both the level and conditional volatility of inflation, consistent with *Rastogi and Kanoujiya (2024)* and *Bhalla et al. (2026)*, while interpreting this cautiously in light of *Eichengreen and Gupta (2024)*, who find the post-2016 reaction function not demonstrably more hawkish. Fourth, it situates these findings within a comparative political-economy framework drawing on the Argentine experience (*Blejer 2025*) to illuminate the shared structural logic of inflation persistence where distributional politics constrain reform. Fifth, it draws concise institutional lessons from the Indian transition for other emerging market inflation-targeting frameworks, including those of Central and Eastern Europe.

A methodological caveat applies throughout. Following central bank practitioners with extensive operational experience, we treat statistical precedence relationships (whether from Granger-type tests or cointegration) as evidence compatible with hypothesised economic mechanisms rather than proof of fundamental causality, since in-sample co-movements need not correspond to exploitable causal mechanisms. Where the paper reports that one variable ‘Granger-causes’ another, the claim is one of statistical predictive content, not fundamental causal exploitability for policy.

The paper is organised as follows. *Section 2* reviews the historical record. *Section 3* presents data and methodology. *Section 4* reports the econometric results, including a robustness subsection (4.6) on unit-root robustness, intra-quarter endogeneity and regime-specific estimation. *Sections 5* through *7* provide the structural and political-economy interpretation, with *Section 7* incorporating the March 2026 framework renewal. *Section 8* draws concise comparative institutional lessons. *Section 9* concludes.

2. Inflation in India: the historical record

In the Indian context, inflation is not merely a monetary phenomenon; it is also influenced by agricultural conditions, Centre–State fiscal relations, and political factors. The evolution of Indian inflation since 1950 divides into four phases (*Table 1*). The first (1950s–1960s) saw modest inflation below 3 per cent on the WPI basis, suppressed by administered prices but structurally fragile.

The second (1970s–1980s) saw acceleration driven by oil shocks, droughts and fiscal expansion, with M3 growth above sixteen per cent per annum. The third (1991–2016) encompasses both the 1991–1993 crisis-stabilisation inflection and the resurgence of elevated food-price inflation under the United Progressive Alliance. The fourth (2017–2024) is the inflation-targeting era: lower average inflation but persistent structural pressures.

Table 1 Inflation, monetary growth and fiscal balances by decade (quarterly averages) in India					
Period	Avg. WPI (%)	Avg. CPI (%)	Avg. M3 Growth (%)	Fiscal deficit (% GDP)	Key structural context
1951–1960	1.7	2.1	6.8	~2.0	Nehruvian ISI; FYP allocation
1961–1970	6.4	7.0	9.2	~3.0	Green Revolution transition; droughts
1971–1980	9.6	8.3	16.4	~4.5	Oil shocks; bank nationalisation
1981–1990	7.1	8.9	17.3	~8.0	Fiscal expansion; SLR peaks at 38.5 per cent
1991–2000	7.6	9.4	15.1	~5.8	BOP crisis; partial liberalisation
2001–2010	5.5	5.9	17.0	~4.8	High-growth decade; MSP acceleration
2011–2016	6.2	9.5	13.1	~4.9	Elevated food inflation; UPA II
2017–2024	4.1	5.4	10.2	~5.8	IT era; COVID shock; 2024–2025 disinflation
<i>Note: WPI = Wholesale Price Index; CPI = Consumer Price Index; M3 growth refers to broad money growth. Fiscal deficit figures are approximate central government deficits as a percentage of GDP. ISI: Import Substituting Industrialisation; FYP: Five-Year Plan, SLR; Statutory Liquidity Ratio; BOP: Balance of Payments; MSP: Minimum Support Price; UPA: United Progressive Alliance.</i>					
<i>Source: RBI Handbook of Statistics (2024), Ministry of Finance Annual Reports, World Bank World Development Indicators</i>					

As *Table 1* shows, a persistent positive association between fiscal deficits and inflation is visible across most sub-periods, interrupted only by the post-1991 reform episode and partially by the early IT period. The M3-inflation association is similarly visible in the pre-IT era, consistent with (though not itself proof of) the quantity theory at longer horizons. Such bivariate co-movements may reflect common drivers or accommodative monetary responses rather than monetary growth as an independent causal force; we therefore treat the association as compatible with monetary-theoretic interpretations without claiming an exploitable mechanism. The distinctiveness of 2011–2016 (high CPI inflation despite declining WPI and moderate money growth) points to the structural food-price channel that *Section 4* confirms formally.

3. Data and methodology

3.1. Data

The analysis uses quarterly data spanning 1955 Q1–2024 Q4, yielding 280 observations. The primary variable is all-India CPI inflation, the annualised quarterly log-difference of the price level. For the pre-2011 period, when the new all-India CPI was unavailable, we use the CPI (Industrial Workers) series spliced to the new series using standard index-linking consistent with RBI practice.¹

The explanatory variables span five categories: (i) *fiscal*: the central government fiscal deficit as a share of GDP (*fd*) and its components; (ii) *monetary*: broad money growth M3 (*m*) and the real interest rate (*r*); (iii) *agricultural price policy*: the all-commodity MSP index (*m**sp*), a weighted average of MSP levels for the twenty-three CACP commodities;² (iv) *demand and external*: the output gap (*gap*) from Hodrick-Prescott filtering, the nominal effective exchange rate (*er*), and the rupee crude-oil price index (*oil*);³ and (v) *expectations*: the RBI's quarterly household inflation-expectations survey (*pe*), available from 2006 Q3 and extended backwards using model-based fitted values. Consistent quarterly unemployment data are unavailable for the full sample; the output gap serves as the demand-side proxy. *Kapur (2013)* provides a complementary unemployment-based Phillips-curve analysis for the post-2010 period.

¹ CPI data for the pre-2011 period are based on the CPI-IW (Industrial Workers) series, re-based and spliced to the new all-India CPI series at 2011 Q4 using standard index-linking methodology consistent with RBI practice. Sensitivity analysis using alternative splicing approaches (CPI-AL series for the agricultural-labour weight; chain-linked weighted average across CPI-IW, CPI-AL and CPI-RL) yields qualitatively similar coefficients on the main variables and identical structural break dates; results are available from the author on request.

² The MSP index is constructed as a geometric weighted average of the CACP-announced MSP levels for twenty-three commodities, using procurement volume weights updated at five-year intervals. Full construction details are available from the author.

³ The output gap is estimated via one-sided Hodrick-Prescott filtering with smoothing parameter $\lambda = 1,600$ for quarterly data, applied to real GDP at constant 2011–2012 prices. The filter boundary problem is addressed using the RBI endpoint-constraint method.

All series are drawn from the *RBI Handbook of Statistics (2024)*, the DGCIS (Directorate General of Commercial Intelligence and Statistics), the CACP (Commission for Agricultural Costs and Prices) Annual Reports and the IMF International Financial Statistics (IFS). Models including expectations are estimated on 2006 Q3–2024 Q4 (73 observations, 68 after lags); for comparability, all four ARDL models in *Table 5* use this sample, and full-sample robustness checks (excluding *pe*) yield qualitatively similar coefficients on *fd*, M3, MSP and the output gap. Monetary and fiscal variables are in log-differences where required by the stationarity analysis.

One caveat concerns consistency across methodological revisions. The CPI basket and the IFS series underwent definitional changes in the period 1955–2024 (the 2011 CPI transition, periodic fiscal-deficit redefinitions, and IMF methodological updates) that may bias break estimates toward transition dates. Recomputing the principal Bai-Perron procedure on the RBI Handbook series alone and on the IFS series independently yields break dates differing by no more than two quarters, with the principal dates (1991 Q3, 2007 Q4, 2014 Q3) coinciding in both. A residual concern is that breakpoint tests may attribute significance to dates reflecting statistical revisions; we therefore treat the smaller breaks (notably 1997 Q2, near the 1999 fiscal-accounting reform) with caution and assign them no causal weight in the interpretation.

3.2. Econometric framework

The strategy proceeds in five stages, following time-series precedents on long-horizon regime-shift identification, cointegration under structural change and conditional-volatility modelling (*Hamilton 1994; Stock – Watson 2003; Lütkepohl 2005*).

In the first stage, we test for unit roots using the Augmented Dickey-Fuller (ADF) test with lag length selected by AIC:

$$\Delta y_t = \alpha + \beta t + \gamma y_{t-1} + \sum_{j=1}^p \delta_j \Delta y_{t-j} + \varepsilon_t \quad (1)$$

where $H_0: \gamma = 0$ (unit root) is tested against $H_1: \gamma < 0$ (stationarity), for specifications with intercept only and with intercept and trend.

In the second stage, we employ the ARDL bounds test of *Pesaran et al. (2001)*. The unrestricted error-correction model is:

$$\Delta \pi_t = \alpha_0 + \sum_{i=1}^p \alpha_i \Delta \pi_{t-i} + \sum_{j=0}^q \beta_j' \Delta X_{t-j} + \lambda_1 \pi_{t-1} + \lambda_2' X_{t-1} + \varepsilon_t \quad (2)$$

where X is the vector of explanatory variables. Lag orders p and q are selected by AIC with a maximum of four quarters; the optimal structure is ARDL(2,1,1,1,0,1,1,0). The bounds test evaluates $H_0: \lambda_1 = \lambda_2 = 0$ against the F-statistic critical values of *Pesaran et al. (2001)*. Long-run coefficients (*Table 5*) are recovered by the normalisation $\vartheta = -\lambda_2/\lambda_1$.

In the third stage, we estimate a reduced-form VAR with lag order selected by the Schwarz criterion and apply the *Toda – Yamamoto (1995)* procedure for Granger precedence tests valid regardless of integration order, avoiding pre-testing distortions.⁴ Precedence is interpreted as statistical predictive content, not exploitable causality. The framework establishes lag-based content, but cannot rule out intra-quarter endogeneity among inflation, fiscal and monetary variables within a quarter; *Section 4.6.* addresses this by re-estimating the ARDL model with lagged or instrumented contemporaneous regressors.

In the fourth stage, we apply the *Bai and Perron (1998, 2003)* multiple-break test in two specifications: an unconditional one regressing inflation on a constant only, and a conditional one that additionally includes the contemporaneous fiscal deficit, M3 growth and the MSP index. The conditional specification tests whether these structural channels jointly account for the observed regime shifts. The test allows for m breaks in:

$$\pi_t = x_t' \beta_j + \varepsilon_{jt} \text{ for } t = T_{j-1} + 1, \dots, T_j, j = 1, \dots, m+1 \quad (3)$$

with break dates $\{T_1, \dots, T_m\}$ estimated jointly with regime-specific parameters by minimising the global sum of squared residuals. The supF and double-maximum tests (UDmax, WDmax) evaluate the presence and number of breaks.

In the fifth stage, we model both the conditional mean and variance of inflation using GARCH(1,1) (*Bollerslev 1986*) augmented with structural-break dummies in both equations:

$$\pi_t = \mu + \varphi D_{IT} + \rho \pi_{t-1} + \varepsilon_{jt} \quad \varepsilon_t \mid \Omega_{t-1} \sim N(0, h_t) \quad (4)$$

$$h_t = \omega + \alpha \varepsilon_{t-1}^2 + \beta h_{t-1} + \gamma D_{IT} \quad (5)$$

⁴ We follow the convention adopted in *Toda and Yamamoto (1995)* of estimating the VAR in levels with $k + d_{max}$ lags, where k is the optimal lag length and d_{max} is the maximum order of integration of the variables, then applying the modified Wald test on the first k lag coefficients only. This approach yields χ^2 test statistics with standard asymptotic distributions regardless of the integration order of the variables.

where D_{IT} equals one from 2016 Q1, φ captures the IT-associated mean shift, and γ the variance shift. Including φ is consistent with the Bai-Perron evidence (Section 4.4.) that IT adoption was associated with a level shift, not merely a volatility shift. The sum $\alpha + \beta$ measures volatility persistence. Equations (4)–(5) are estimated jointly by quasi-maximum likelihood with Bollerslev-Wooldridge robust standard errors.

4. Empirical results

4.1. Unit root tests

Table 2 reports ADF results for all variables in levels and first differences. With the exception of the output gap, which is stationary in levels by construction, all variables are integrated of order one, $I(1)$, failing to reject the unit-root null in levels but strongly rejecting it in first differences (MacKinnon 1996), consistently across both specifications. This motivates the ARDL bounds approach, valid for any mixture of $I(0)$ and $I(1)$ regressors. As shown in Section 4.6., re-estimating with procedures allowing endogenous breaks (Zivot – Andrews 1992; Lee – Strazicich 2003) preserves the integration-order classification.

Table 2
Augmented Dickey-Fuller unit root test results

Variable	Levels– Intercept	Levels– Trend	Lags	Δ –Intercept	Δ –Trend	Lags	Integration order
CPI inflation (π)	–1.42	–1.61	3	–6.83***	–6.79***	2	$I(1)$
WPI inflation (π_w)	–1.67	–1.88	2	–7.12***	–7.08***	1	$I(1)$
Fiscal deficit (fd)	–1.29	–1.47	4	–5.94***	–5.90***	3	$I(1)$
M3 money growth (m)	–2.01	–2.14	2	–6.41***	–6.38***	1	$I(1)$
MSP index (m_{sp})	–1.55	–1.72	3	–5.78***	–5.74***	2	$I(1)$
Output gap (gap)	–3.12**	–3.34**	2	–7.91***	–7.85***	1	$I(0)$
Exchange rate (er)	–1.83	–1.99	3	–6.02***	–5.97***	2	$I(1)$
Global oil price (oil)	–1.47	–1.63	2	–6.25***	–6.21***	1	$I(1)$
Real interest rate (r)	–2.88*	–3.05**	3	–6.78***	–6.74***	2	$I(1)$
Inflation expectations (pe)	–1.96	–2.12	4	–5.55***	–5.51***	3	$I(1)$

Note: ADF t-statistics for the null of a unit root. Lag length selected by AIC with maximum lag of 8. *** $p < 0.01$; ** $p < 0.05$; * $p < 0.10$ relative to MacKinnon (1996) critical values. Sample: 1955 Q1–2024 Q4 ($N = 280$).

As a preliminary multicollinearity check, the pairwise correlation matrix (*Appendix Table 10*) shows all absolute pairwise correlations below 0.50 and variance inflation factors well below the conventional threshold of five, indicating that multicollinearity does not substantively distort the ARDL estimates.

4.2. ARDL bounds testing

Table 3 presents ARDL bounds test F-statistics for seven specifications. The full model yields an F-statistic of 6.84, well above the upper $I(1)$ bound of 4.35 at 5 per cent, confirming cointegration; five of seven specifications exceed the upper bound. The isolated fiscal-dominance model ($\pi = f(fd)$ alone) is inconclusive, suggesting that fiscal deficits operate through multiple channels rather than purely through monetisation (*Pesaran et al. 2001*).

Model specification	F-statistic	Lower bound $I(0)$ (5%)	Upper bound $I(1)$ (5%)	Cointegration decision
$\pi = f(fd, m, msp, gap, er, oil)$	6.84***	3.23	4.35	Cointegrated
$\pi = f(fd, m, msp)$	5.91***	3.23	4.35	Cointegrated
$\pi = f(m, gap, er, oil, pe)$	7.12***	3.23	4.35	Cointegrated
$\pi = f(msp, fd, r, gap)$	5.47***	3.23	4.35	Cointegrated
$\pi w = f(fd, m, msp, er, oil)$	4.88**	3.23	4.35	Cointegrated
$\pi = f(fd)$ – fiscal dominance baseline	3.01	3.23	4.35	Inconclusive
$\pi = f(m)$ – monetary baseline	5.22***	3.23	4.35	Cointegrated

Note: F-statistics test H_0 : no long-run relationship. Critical values from Pesaran et al. (2001), Case III (unrestricted intercept, no trend). *** $p < 0.01$; ** $p < 0.05$. All models include a structural break dummy for 1991 Q3. Lag length selected by AIC (maximum four quarters).

Long-run coefficients are reported in *Table 4*: *Model 1* (full), *Model 2* (fiscal and MSP), *Model 3* (monetary and demand) and *Model 4* (food-price emphasis). All four use the common 2006 Q3–2024 Q4 sample (68 observations after lags) for comparability with the expectations variable; full-sample checks excluding *pe* yield qualitatively similar coefficients (available on request). Across all specifications the coefficients are signed consistently with theory and significant at conventional levels.

Table 4**ARDL long-run coefficient estimates: dependent variable is CPI inflation (π)**

Regressor	Model 1 (full)	Model 2 (fiscal)	Model 3 (monetary)	Model 4 (food-price)	Expected sign
<i>Fiscal deficit (fd)</i>	0.312*** (0.081)	0.387*** (0.094)	—	0.241** (0.097)	+
<i>M3 money growth (m)</i>	0.418*** (0.073)	—	0.511*** (0.088)	0.329*** (0.076)	+
<i>MSP index (msp)</i>	0.287*** (0.068)	0.341*** (0.079)	—	0.348*** (0.072)	+
<i>Output gap (gap)</i>	0.196** (0.084)	—	0.241*** (0.079)	—	+
<i>Exchange rate (er)</i>	0.143** (0.062)	—	0.167** (0.071)	—	+
<i>Global oil price (oil)</i>	0.098* (0.053)	—	0.112** (0.057)	—	+
<i>Inflation expectations (pe)</i>	0.221*** (0.066)	—	0.198*** (0.071)	0.244*** (0.069)	+
<i>Constant</i>	-1.204** (0.491)	-0.874* (0.508)	-1.567*** (0.472)	-0.991** (0.467)	—
R^2	0.847	0.721	0.803	0.776	
Adjusted R^2	0.821	0.701	0.783	0.754	
AIC	-2.341	-1.978	-2.219	-2.108	
Durbin-Watson	1.943	1.871	1.988	1.924	
Breusch-Godfrey LM p	0.312	0.271	0.394	0.341	
RESET test p	0.481	0.398	0.512	0.443	
Observations	68	68	68	68	

Note: Long-run coefficients recovered from the UECM normalisation $\vartheta = -\lambda_2/\lambda_1$. Newey-West HAC-robust standard errors in parentheses. ‘—’ indicates variable excluded in columns of models. *** $p < 0.01$; ** $p < 0.05$; * $p < 0.10$. All four models estimated on the common 2006 Q3–2024 Q4 sample ($N = 73$, reducing to 68 after lags).

The *Model 1* fiscal-deficit coefficient of 0.312 ($p < 0.01$) implies that a one-point rise in the deficit-to-GDP ratio is associated with a 0.31-point rise in steady-state CPI inflation, holding other conditions constant, robust across *Models 2* and *4*. The significance of *fd* conditional on *m* indicates that the fiscal effect operates beyond simple monetisation, consistent with the fiscal theory of the price level.

The *MSP* coefficient of 0.287 ($p < 0.01$, largest in *Model 4* at 0.348) is the paper’s most policy-consequential finding, confirming that agricultural price-support policies are an independent structural driver of CPI inflation through a cost-push channel orthogonal to monetary conditions. A 10-per cent MSP increase (roughly the annual hike in 2011–2014) is associated with a 2.87-point rise in steady-state inflation, an effect amplified by the high food share in the CPI. The expectations

coefficient (0.221 in *Model 1*) captures inertia: a one-point rise in prior-period household expectations is associated with a 0.22-point rise in realised inflation, consistent with backward-looking price setting that sustains inflation independently of contemporaneous demand or monetary conditions.

4.3. VAR Granger precedence analysis

Table 5 reports *Toda – Yamamoto (1995)* precedence tests from a VAR(4) in levels; we use ‘precedence’ rather than ‘causality’ to underscore that these evaluate predictive content, not fundamental causality. The pattern is clear and asymmetric: fiscal deficits, money growth, MSP prices, exchange-rate movements and global oil prices all statistically precede CPI inflation, while the reverse is generally not supported. The strongest signal comes from M3 ($\chi^2 = 21.54$, $p = 0.000$), followed by MSP (18.92, $p = 0.001$), inflation expectations (16.74, $p = 0.002$) and the fiscal deficit (14.32, $p = 0.006$). The bidirectional link between expectations and realised inflation is consistent with a self-reinforcing mechanism. The magnitude ranking does not by itself establish a hierarchy of causation: a stronger signal for money growth than for the deficit may reflect either greater predictive content or lower measurement noise in monetary aggregates.

Table 5				
Toda–Yamamoto Granger precedence test results (VAR system, 4 lags)				
Null hypothesis	χ^2 statistic	df	p-value	Decision
$fd \rightarrow \pi$	14.32***	4	0.006	Reject
$\pi \rightarrow fd$	3.41	4	0.492	Fail to reject
$m \rightarrow \pi$	21.54***	4	0.000	Reject
$\pi \rightarrow m$	5.87	4	0.209	Fail to reject
$mSP \rightarrow \pi$	18.92***	4	0.001	Reject
$\pi \rightarrow mSP$	2.14	4	0.710	Fail to reject
$er \rightarrow \pi$	9.67**	4	0.046	Reject
$\pi \rightarrow er$	6.43	4	0.169	Fail to reject
$oil \rightarrow \pi$	11.28**	4	0.024	Reject
$\pi \rightarrow oil$	1.98	4	0.740	Fail to reject
$pe \rightarrow \pi$	16.74***	4	0.002	Reject
$\pi \rightarrow pe$	8.41*	4	0.077	Reject at 10%
$gap \rightarrow \pi$	7.82*	4	0.098	Reject at 10%
$\pi \rightarrow gap$	4.11	4	0.391	Fail to reject

Note: $X \rightarrow Y$: X does not statistically precede Y . *Toda – Yamamoto (1995)* modified Wald test statistics, valid for $I(1)$ or cointegrated variables. VAR lag order = 4 selected by Schwarz BIC. *** $p < 0.01$; ** $p < 0.05$; * $p < 0.10$. Sample: 1955Q1–2024Q4 ($N = 280$ for base variables; 2006 Q3–2024 Q4 for pe). The tests evaluate statistical predictive precedence rather than fundamental economic causality.

4.4. Bai-Perron structural break analysis

We implement the Bai-Perron procedure in two specifications. The unconditional one regresses inflation on a constant only, identifying breaks in the unconditional mean; the conditional one adds the fiscal deficit, M3 growth and the MSP index.

Table 6 reports the unconditional results: seven significant breaks over the sample. The supF test strongly rejects no breaks ($p < 0.001$), and the double-maximum tests (UDmax = 31.4, WDmax = 38.2, both $p < 0.001$) confirm multiple breaks.

Table 6 Bai-Perron multiple structural break test results: unconditional specification, CPI inflation series, 1955 Q1–2024 Q4				
Break no.	Break date	95% CI	Regime avg. CPI (%)	Economic narrative
1	1973 Q3	[1972 Q4 – 1974 Q2]	9.81	First oil shock; drought confluence
2	1984 Q1	[1983 Q2 – 1984 Q4]	9.14	Peak fiscal expansion; pre-crisis buildup
3	1991 Q3	[1991 Q1 – 1992 Q1]	13.24	BOP crisis; stabilisation inflection
4	1997 Q2	[1996 Q4 – 1997 Q4]	5.22	Post-reform disinflation consolidates
5	2007 Q4	[2007 Q2 – 2008 Q2]	9.87	Global food and fuel shock; UPA II expansion
6	2014 Q3	[2014 Q1 – 2015 Q1]	4.41	IT credibility gains; supply-side favourable
7	2020 Q2	[2019 Q4 – 2020 Q4]	6.93	COVID-19 pandemic supply disruption
<i>Note: Bai and Perron (1998, 2003) sequential procedure with maximum of 8 breaks and trimming parameter $\epsilon = 0.10$. Break dates estimated by minimising the global sum of squared residuals. 95% confidence intervals from the asymptotic distribution. Regime average CPI is the mean annualised quarterly inflation within each regime.</i>				

Table 7 reports the conditional results, replacing the constant with a regression on lagged inflation, the fiscal deficit, M3 growth and the MSP index. Only three breaks now survive: 1991 Q3 (BOP crisis), 2007 Q4 (global food and fuel shock), and 2014 Q3 (IT credibility transition). The four unconditional breaks absorbed by the structural regressors provide direct evidence that the fiscal, monetary and food-price channels jointly account for the medium-frequency level shifts in Indian inflation. The three surviving breaks correspond to events operating through channels orthogonal to the regressors: external crisis, global commodity shock and institutional regime change.

Table 7**Bai-Perron multiple structural break test results: conditional specification (regressors: π_{-1} , fd, m, msp), CPI inflation series, 1955 Q1–2024 Q4**

Break no.	Break date	95% CI	Sup F-stat	Interpretation
1	1991 Q3	[1991 Q1 – 1992 Q2]	27.4***	BOP crisis: external break orthogonal to regressors
2	2007 Q4	[2007 Q2 – 2008 Q2]	19.8***	Global commodity shock: cost-push break
3	2014 Q3	[2014 Q1 – 2015 Q1]	23.1***	IT credibility transition: institutional break

Note: Conditional Bai-Perron procedure with maximum of 8 breaks and trimming parameter $\epsilon = 0.10$. Regressors: one lag of CPI inflation, fiscal deficit, M3 growth and MSP index. UDmax = 24.7, WDmax = 31.3 (both $p < 0.001$).

The most policy-relevant break is the 2014 Q3 conditional break, coinciding with the disinflationary transition preceding formal IT adoption. The shift from the elevated UPA-era regime (9.87 per cent unconditional average, *Table 6*) to the IT-era regime (4.41 per cent) is the largest single reduction across all transitions, suggesting that the RBI credibility commitment from 2013, the Urjit Patel Committee report of January 2014 (*RBI 2014*) and the February 2015 framework constituted a genuine regime change rather than cyclical moderation. That this break survives conditioning (when the BOP-crisis and commodity-shock breaks would not, absent the regressors) is direct evidence that the IT transition operates through a credibility channel empirically distinct from the structural drivers.

4.5. GARCH conditional mean and volatility analysis

Table 8 presents GARCH(1,1) estimates with break dummies on both the conditional mean (ϕ) and variance (γ). Across all periods, the sum $\alpha + \beta$ is near unity (0.910–0.942), indicating near-integrated, highly persistent inflation volatility (*Bollerslev 1986*). In the full sample, the IT dummy enters significantly negative in both the mean ($\phi = -1.823$, $p < 0.01$) and variance ($\gamma = -0.641$, $p < 0.01$) equations, confirming that IT adoption is associated with a statistically significant downward shift in both the conditional level and volatility of inflation, consistent with *Section 4.4.* and with international evidence (*Bernanke et al. 1999; Lin – Ye 2009; Bhalla et al. 2026*).

Table 8
GARCH(1,1) conditional mean and volatility estimates: CPI inflation, 1955 Q1–2024 Q4

Parameter	Full (1955–2024)	Pre-IT (1955–2015)	IT era (2016–2024)	Interpretation
<i>Mean equation</i>				
Constant (μ)	6.523*** (0.467)	7.214*** (0.538)	4.882*** (0.321)	Long-run mean inflation
IT mean dummy (φ)	–1.823*** (0.541)	–	–	IT lowers conditional mean
AR(1) persistence (ρ)	0.628*** (0.069)	0.712*** (0.088)	0.488*** (0.064)	Inertial component
<i>Variance equation</i>				
Constant (ω)	0.412*** (0.118)	0.487*** (0.137)	0.318** (0.142)	Baseline variance
ARCH term (α)	0.287*** (0.074)	0.341*** (0.091)	0.198** (0.082)	Shock impact on volatility
GARCH term (β)	0.654*** (0.088)	0.601*** (0.104)	0.712*** (0.097)	Volatility persistence
$\alpha + \beta$ (persistence sum)	0.941	0.942	0.910	Near unit-root volatility
IT variance dummy (γ)	–0.641*** (0.198)	–	–	IT reduces conditional variance
Log-likelihood	–139.4	–118.4	–31.2	
AIC	3.221	3.494	2.817	
Ljung-Box Q(10) p	0.408	0.388	0.521	
ARCH-LM test p	0.341	0.291	0.477	
Observations	276	243	33	

Note: GARCH(1,1) estimates by quasi-maximum likelihood with Bollerslev-Wooldridge robust standard errors (in parentheses). $D_{IT} = 1$ from 2016Q1; both mean (φ) and variance (γ) dummies are included in the full-sample specification. The pre-IT and IT-era columns are sub-sample estimates without dummies, reported for comparison. *** $p < 0.01$; ** $p < 0.05$. Diagnostics: Ljung-Box Q(10) on standardised residuals; ARCH-LM on squared standardised residuals. The IT-era sub-sample ($N = 33$) is illustrative; the full-sample dummy specification provides the primary inference.

Comparing sub-samples, the ARCH coefficient (α) falls from 0.341 (pre-IT) to 0.198 (IT era), indicating a smaller immediate impact of shocks on conditional variance. The GARCH coefficient (β) remains high, so shocks remain persistent but strike with lower initial force, consistent with the credibility gains of a rules-based framework. These findings align with *Rastogi and Kanoujiya (2024)*, who document a post-IT volatility reduction; our framework additionally identifies a statistically significant downward level shift through the mean-equation dummy, which their specification does not estimate. Their documented volatility spillovers from crude oil, gold and exchange rates correspond to the level effects captured by our exchange rate and oil coefficients in *Table 4*.

4.6. Robustness: break-robust unit roots, intra-quarter endogeneity and regime-specific estimation

Three additional exercises test the robustness of the principal inferences.⁵

Break-robust unit-root inference. The ADF test loses power under structural breaks, potentially yielding spurious non-rejection for series stationary around a broken trend. Re-estimating with *Zivot – Andrews (1992)*, permitting one endogenous break, and *Lee – Strazicich (2003)*, allowing two, yields t-statistics for CPI inflation, the fiscal deficit, M3 growth and the MSP index of -4.18 , -4.02 , -4.51 and -3.98 , respectively, all failing to reject the unit-root null at 5 per cent (critical value -5.08). Endogenous break dates cluster at 1991 Q3 and 2014 Q3. The $I(1)$ classification is therefore robust; the ADF non-rejection in *Table 2* does not reflect break-induced distortion.

Intra-quarter endogeneity. To assess whether the contemporaneous fiscal and monetary coefficients reflect within-quarter simultaneity, we re-estimate *Model 1* (i) replacing contemporaneous regressors with first lags and (ii) instrumenting them with second and third lags in an IV-ARDL framework. The long-run coefficients on the fiscal deficit (0.291 and 0.305, both $p < 0.01$) and the MSP index (0.262 and 0.279, both $p < 0.01$) remain significant and close to the baseline estimates of 0.312 and 0.287, while the M3 coefficient declines slightly (to 0.382 and 0.395), but stays highly significant. The principal long-run associations are therefore not driven by intra-quarter simultaneity, though the precise contemporaneous magnitudes cannot be pinned down without external instruments that are unavailable in this setting.

Regime-specific estimation. Re-estimating the full ARDL model separately on the four regimes defined by the conditional breaks (pre-liberalisation, 1955 Q1–1991 Q2; post-reform pre-2007, 1991 Q3–2007 Q3; UPA-era food inflation, 2007 Q4–2014 Q2; and IT era, 2014 Q3–2024 Q4) shows the MSP coefficient positive and significant in all four, rising from 0.21 to 0.34, consistent with the growing political-economic salience of MSP. The fiscal coefficient is significant in the pre-liberalisation and IT regimes but weaker in the intervening periods, consistent with the fiscal channel operating differently under financial repression and credible commitment than under partial liberalisation. The expectations coefficient is significant only in the IT regime, consistent with expectations becoming anchored to the official target only after the 2015 framework. These estimates, though based on shorter samples, confirm that the structural channels operate across configurations with interpretable variation. Full tables are available on request.

⁵ The structural-break-allowing unit-root and cointegration procedures in *Section 4.6*. (*Zivot – Andrews 1992*; *Lee – Strazicich 2003*; *Gregory – Hansen 1996*) are estimated using the standard implementations in R packages ‘urca’ and ‘tseries’. Full output is available from the author on request.

5. The Nehruvian legacy and the architecture of fiscal dominance

The institutional roots of India's fiscal dominance (evidenced empirically in *Tables 4* through *7*) lie in the model of economic governance established under Jawaharlal Nehru. The nationalisation of fourteen commercial banks in 1969 deepened financial repression through SLR and CRR requirements. As *McKinnon (1973)* showed, financial repression of this type is inherently inflationary: it suppresses saving and necessitates higher monetary growth to finance investment. Import-substitution industrialisation created a high-cost industrial sector generating cost-push pressures; as *Bhagwati and Desai (1970)* documented, the licensing system ('licence raj') created rents for incumbents, discouraged productivity improvement, and generated persistent inflationary pressure.

5.1. The political economy of fiscal dominance

The analogy with Peronist Argentina is instructive but imperfect. *Blejer (2025)* identifies three core inflationary mechanisms of Peronist policy: structural wage-price dynamics sustained by politicised union bargaining; fiscal populism financed by monetary expansion; and backward-looking indexation embedding expectations into contracts. India's Nehruvian framework exhibits all three, with the wage-price mechanism operating through the MSP-agricultural labour nexus rather than urban trade union bargaining, and indexation through the dearness allowance system for government employees, a formal mechanism transmitting past inflation into current wage costs.

What India shares with Argentina is the 'inflationary trap': a political-institutional configuration in which the structural sources of inflation are sufficiently politically embedded that no government has been able or willing to address them comprehensively. Monetary policy can moderate inflationary episodes but cannot resolve the underlying distortions without complementary fiscal and supply-side reforms that the political economy constrains.

5.2. Labour market structure and the flat Phillips curve

Although wage-price dynamics operate primarily through the agricultural-MSP nexus, the broader labour market warrants discussion. *Behera et al. (2018)* documents a substantial flattening of the unemployment-inflation trade-off over the post-2010 period, with the unemployment-gap coefficient rendered insignificant in most specifications once supply-side controls enter. This is consistent with our finding that the *Model 1* output-gap coefficient (0.196, *Table 4*), while positive and significant, is substantially smaller than those on the fiscal deficit (0.312),

M3 growth (0.418) and the MSP index (0.287). In an economy where cost-push channels operate through administered agricultural prices and the formal-sector labour market is a small minority of employment, the demand-side Phillips curve operates with reduced potency relative to the supply-side and fiscal channels. Consistent full-sample quarterly unemployment data are unavailable; we view this as a limitation for future research building on the post-2010 sample.

6. The political economy of food inflation: India's MSP trap

The ARDL result that the MSP index carries a long-run coefficient of 0.287–0.348 on CPI inflation (significant and robust across specifications) is the paper's most policy-consequential finding. It confirms empirically what institutional analysis suggests: the Minimum Support Price system is a structural generator of inflation qualitatively distinct from, and empirically separable from, the fiscal and monetary channels. The effect is amplified by the high food-and-beverages share (≈ 45.86 per cent) in the CPI, so that even moderate MSP increases translate into substantial headline movements.

The MSP system was established in 1966 as a response to food security vulnerabilities. Between 2007 and 2014, paddy MSP increases averaged 14.3 per cent per annum, far above general inflation. As *Rangarajan and Mishra (2013)* documented, these increases substantially exceeded non-food inflation, creating a secular upward trend in food prices that monetary policy could not address without excessive output and employment losses.

The precedence result (*Table 5*) that *msp* precedes π ($\chi^2 = 18.92$, $p = 0.001$), while π does not precede *msp* ($\chi^2 = 2.14$, $p = 0.710$) establishes MSP's asymmetric predictive content for inflation, consistent with a political-economy interpretation: MSP levels are set by electoral considerations rather than inflationary conditions, generating an exogenous driver of headline inflation unresponsive to monetary instruments. We re-emphasise that statistical precedence is not fundamental causal exploitability; nonetheless the institutional evidence on MSP determination is consistent with the asymmetric pattern in the data.

The Farm Laws episode of 2020–2021 provides a natural experiment for the political-economy constraint. The three laws were withdrawn in November 2021 following sustained protests, the most explicit recent illustration of the trap: a reform with strong long-run macroeconomic benefits overridden by concentrated political costs, structurally analogous to *Blejer's (2025)* account of successive Argentine governments unable to dismantle the legacies of Peronist labour policy.

7. The inflation-targeting framework: credibility, recent disinflation and the 2026 renewal

The 2014 Q3 break in both Bai-Perron specifications (*Tables 6 and 7*) and the significant negative IT dummies in the GARCH mean ($\phi = -1.823$, $p < 0.01$) and variance ($\gamma = -0.641$, $p < 0.01$) equations together confirm that flexible inflation targeting represents a genuine institutional regime change. *Table 9* contextualises India's IT performance in comparative perspective.

Table 9 Inflation targeting performance: India in comparative perspective, 2017–2024					
Country	IT adopted	CPI target (%)	Avg. inflation	Fiscal deficit avg. (% GDP)	Assessment
India	2016	4±2	5.4	~5.8	Partial; structural constraints persist
Indonesia	2005	3±1	3.4	~2.6	Successful; fiscal complement
South Korea	1998	2±0.5	1.5	~1.4	Strong; institutional credibility
Brazil	1999	3.5±1.5	5.1	~7.1	Mixed; fiscal-monetary tensions
Philippines	2002	3±1	3.9	~3.1	Successful; commodity exposure
Turkey	2006	5±2	42.7	~3.5	Failed; political interference
Argentina	2016*	N/A	51.4	~5.4	Abandoned; institutional collapse

*Note: Averages for 2017–2024. * Argentina adopted IT in 2016, but effectively abandoned the framework by 2019. Fiscal deficit figures are approximate general-government balances. The seven economies span three Asian emerging market peers (Indonesia, South Korea, the Philippines), one Latin American comparator with a comparable fiscal challenge (Brazil), and two cases of IT failure or abandonment (Turkey, Argentina). Advanced economy adopters are excluded as less informative for evaluating the institutional preconditions of sustained target adherence in an emerging market context.*

Sources: IMF (2024), national central bank publications, World Bank World Development Indicators (<https://databank.worldbank.org>)

India's average CPI inflation under IT (5.4 per cent) is well below its pre-IT average and below comparable economies with weaker frameworks, but exceeds the 4-per cent target midpoint by 1.4 points and remains above Asian peers such as Indonesia (3.4 per cent) and the Philippines (3.9 per cent). The gap reflects precisely the structural constraints identified econometrically: MSP-driven food inflation and persistent fiscal deficits.

An important qualification comes from *Eichengreen and Gupta (2024)*, who find the RBI's post-2016 reaction function not demonstrably more hawkish than before, raising the question of how much observed disinflation reflects a genuine policy-stance shift versus favourable supply-side and global conditions. Our findings are not in conflict with theirs: the break and volatility-reduction results (*Tables 6–8*) capture outcomes (the conditional level and variance of inflation), while their analysis examines policy inputs. The two reconcile if the framework operated primarily through the expectations channel and institutional commitment rather than systematically more aggressive rate movements, consistent with the post-IT decline in expectations captured by our ARDL coefficient on *pe*.

7.1. Recent disinflation and the March 2026 framework renewal

Headline CPI inflation moderated through 2024 and fell within the 4 ± 2 per cent band by early 2026, despite a central-government deficit above 5 per cent of GDP and an unreformed MSP regime. In March 2026, the Government and the RBI renewed the framework (retaining the 4-per cent target and ± 2 -point band) for a further 5-year cycle, a decision *Gupta (2026)* reports commanded wide professional support.

This disinflation is consistent with the structural framework rather than evidence against it. The structural channels operate in the long-run cointegrating sense: cyclical disinflation from favourable food supply conditions, base effects and global commodity moderation can coexist with structural pressures that re-emerge under less favourable circumstances, as in the 1997–2003 disinflation followed by the 2007–2014 food price acceleration. The GARCH evidence that conditional volatility fell substantially in the IT era is itself a partial answer to the structural critique: even where the conditional mean remains susceptible to structural drivers, a tighter conditional distribution implies fewer extreme episodes, and the institutional commitment provides an anchor absent in the pre-IT regime.

On the appropriateness of the 4-per cent target, our estimates imply that a 3-per cent deficit (the FRBM⁶ target) and MSP indexation aligned with general inflation would together reduce steady-state inflation by roughly 1.4–1.9 points relative to the IT-era average of 5.4 per cent, bringing it to about 3.5–4.0 per cent. A lower target (e.g. 3 per cent, as in Indonesia and the Philippines) would be consistent with the evidence only if accompanied by these structural reforms; it cannot be achieved through monetary policy alone without excessive output and employment losses given weak transmission. The retention of the 4-per cent target in March 2026 thus represents a reasonable institutional choice, given the unresolved structural constraints; any reduction should be coordinated with, and conditional on, measurable fiscal and agricultural market reform.

⁶ Fiscal Responsibility and Budget Management.

7.2. Residual vulnerabilities of the Indian IT framework

The fiscal-monetary coordination problem of *Sargent and Wallace (1981)* remains the central vulnerability. The central government's budget deficit has averaged approximately 5.8 per cent of GDP during the IT period, against the FRBM target of 3.0 per cent. At 2.8 points above target, the implied inflationary pressure from the fiscal stance alone is approximately 0.87 points of steady-state inflation on the *Model 1* coefficient of 0.312, a substantial fraction of the 1.4-point gap between average IT-era inflation and the 4-per cent midpoint. The durability of the framework thus remains contingent on fiscal consolidation that political-economy constraints have so far prevented.

8. Institutional lessons for other inflation-targeting economies

The Indian experience (a multi-decade transition from fiscal dominance under Nehruvian institutions to a credible inflation-targeting framework adopted in 2016 and renewed in 2026) offers institutional lessons of comparative interest, including for the Central and Eastern European (CEE) economies whose own disinflationary trajectories were anchored by the European Union accession process and the Maastricht criteria. We draw three high-level structural takeaways, while emphasising that the macroeconomic structures, openness and policy instruments of India and the CEE economies differ fundamentally, so that the comparison is illustrative rather than a template for direct policy transfer.

First, the source of disinflationary credibility differs. The CEE transition to inflation targeting in the late 1990s and early 2000s was underwritten by an external commitment device (the prospective EU accession process and the fiscal discipline embedded in the Maastricht criteria) and proceeded comparatively rapidly. India, lacking an equivalent regional anchor, built credibility entirely through domestic institutional innovation over a substantially longer horizon. The high-level lesson is that domestically-anchored credibility-building takes longer and is more exposed to political-economic shocks during the transition, but can prove durable once established, as the 2014 Q3 break and the framework's continued survival suggest. Economies whose external anchors are weakening may find the Indian case of domestic institutional commitment instructive.

Second, structural cost-push channels operating independently of the monetary stance can drive headline inflation through periods of disinflation, and their political economy differs across contexts. In India, the channel is agricultural and domestic-political in origin, the MSP system administered by an institution responsive to electoral constituencies. In CEE economies, the analogous channel is principally

energy-related and largely externally determined, as the 2022–2023 episode illustrated. The shared high-level lesson is that a framework operating only through the demand channel, without addressing cost-push channels in an institutionally-appropriate way, will exhibit persistent target deviations during cost-push stress; the appropriate institutional response must be calibrated to the political economy of the relevant channel rather than a single template.

Third, fiscal-monetary coordination conditions durability everywhere. The Indian evidence that a 5.8-per cent deficit contributes roughly 0.87 points of inflationary pressure relative to a 3-per cent benchmark generalises directly: frameworks operating without fiscal consolidation should expect persistent target overshoots even where expectations are nominally anchored. The broader implication is that the institutional architecture of inflation targeting cannot be specified independently of the structural-political-economy configuration in which it operates.

9. Conclusions and policy implications

This paper has applied a comprehensive time-series framework to establish four empirical propositions about India's inflation persistence. First, CPI inflation, the fiscal deficit, M3 growth, the MSP index and most external variables are $I(1)$ and cointegrated, a classification robust to break-allowing unit-root procedures (*Section 4.6.*). Second, the ARDL estimates indicate significant positive long-run effects of fiscal deficits ($\beta \approx 0.31\text{--}0.39$), the MSP index ($\beta \approx 0.29\text{--}0.35$) and M3 growth ($\beta \approx 0.33\text{--}0.51$) on CPI inflation, robust to replacing contemporaneous regressors with lagged or instrumented counterparts and operating with interpretable variation across regimes. Third, Granger precedence runs from fiscal, monetary and agricultural price variables to inflation but not in reverse, with the strongest signal from M3 growth and the MSP index. Fourth, IT adoption constitutes a statistically identifiable break in both the level ($\phi = -1.823$) and conditional volatility ($\gamma = -0.641$) of inflation, with the 2014 Q3 break surviving the inclusion of structural regressors, consistent with a credibility interpretation.

These results support the central argument: India's inflation persistence reflects a structural political-economy configuration in which monetary policy operates in an environment shaped by fiscal dominance and agricultural price politics. The analogy with *Blejer's* (2025) characterisation of Argentine inflation as institutional rather than purely monetary is empirically grounded, though the magnitudes differ by orders of magnitude: both exhibit an 'inflationary trap' in which the structural drivers are sufficiently embedded that successive governments have been unable or unwilling to address them. The 2024–2025 disinflation and the March 2026 renewal demonstrate

the partial success of reform since 2014; the structural channels nonetheless remain operative and would re-emerge under less favourable conditions.

The comparative discussion in *Section 8* places the Indian transition in the broader context of emerging market disinflationary reform. The principal lesson is that the institutional architecture of inflation targeting cannot be specified independently of the structural-political-economy configuration in which it operates: the Indian MSP-driven and CEE energy-driven cost-push channels are different manifestations of the same underlying problem, each requiring responses calibrated to its own political economy.

The policy implications are demanding. A durable solution requires three complementary reforms. First, genuine fiscal consolidation to a deficit of 3 per cent or below, which on the *Model 1* coefficient could alone reduce steady-state inflation by approximately 0.87 points, bringing average IT-era inflation to about 4.5 per cent and within striking distance of the target midpoint. Second, a politically sequenced reform of the MSP–FCI⁷ complex: on the coefficient of 0.287–0.348, moderating annual MSP increases to align with general inflation could contribute a further 0.5–1.0-point reduction in steady-state food-price inflation. Third, a strengthening of Monetary Policy Committee (MPC) institutional independence to protect the GARCH-confirmed volatility reduction from reversal, and continued work on the transmission channel to address the PSB-related rigidities documented by *Mishra and Montiel (2013)*.

Whether India can achieve this combination is the defining question for its macroeconomic future. The evidence documents both the genuine progress of the IT era and the structural constraints that remain. The institutional architecture of price stability has been erected; the foundations in the real and political economy remain contested. The interpretation advanced here (that statistical evidence is compatible with the structural-political-economy framework rather than proof of fundamental causality) is consistent with the empirical record and with the operational experience of central banks in similar contexts.

⁷ Food Corporation of India (FCI).

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Author contributions

The sole author was responsible for all aspects of this manuscript. The author has read and agreed to the published version of the manuscript.

Data Availability

The data and code that support the findings of this study are available from the author upon reasonable request.

Ethical Approval

Not applicable. This study did not involve human participants or sensitive personal data requiring institutional ethical approval.

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References

- Akerlof, G.A. – Dickens, W.T. – Perry, G.L. (2000): *Near-rational wage and price setting and the long-run Phillips curve*. Brookings Papers on Economic Activity, 2000(1): 1–44. <https://doi.org/10.1353/eca.2000.0001>
- Bai, J. – Perron, P. (1998): *Estimating and testing linear models with multiple structural changes*. Econometrica, 66(1): 47–78. <https://doi.org/10.2307/2998540>
- Bai, J. – Perron, P. (2003): *Computation and analysis of multiple structural change models*. Journal of Applied Econometrics, 18(1): 1–22. <https://doi.org/10.1002/jae.659>
- Behera, H. – Wahi, G. – Kapur, M. (2018): *Phillips curve relationship in an emerging economy: Evidence from India*. Economic Analysis and Policy, 59: 116–126. <https://doi.org/10.1016/j.eap.2018.06.001>
- Bernanke, B.S. – Laubach, T. – Mishkin, F.S. – Posen, A.S. (1999): *Inflation targeting: Lessons from the international experience*. Princeton University Press.
- Bhagwati, J. – Desai, P. (1970): *India: Planning for industrialization*. Oxford University Press.
- Bhalla, S.S. – Bhasin, K. – Loungani, P. (2026): *Inflation targeting and monetary policy in India*. Economic and Political Weekly, 61(3). <https://doi.org/10.71279/epw.v61i3.44715>
- Blejer, M.I. (2025): *Understanding the persistence of inflation in Argentina: From Peron to Milei*. Financial and Economic Review, 24(4): 5–21. <https://doi.org/10.33893/FER.24.4.5>
- Bollerslev, T. (1986): *Generalized autoregressive conditional heteroskedasticity*. Journal of Econometrics, 31(3): 307–327. [https://doi.org/10.1016/0304-4076\(86\)90063-1](https://doi.org/10.1016/0304-4076(86)90063-1)
- Cochrane, J.H. (2023): *The fiscal theory of the price level*. Princeton University Press. <https://doi.org/10.2307/j.ctv2sbm8kh>
- Eichengreen, B. – Gupta, P. (2024): *Inflation targeting in India: A further assessment*. NCAER Working Paper No. 174. National Council of Applied Economic Research. <https://www.ncaer.org/publication/inflation-targeting-in-india-a-further-assessment>
- Gokarn, S. (2011): *Food inflation: This time it's different*. Kale Memorial Lecture, Reserve Bank of India, 9 December. <https://www.bis.org/review/r111215e.pdf>. Downloaded: 27 March 2026.
- Gregory, A.W. – Hansen, B.E. (1996): *Residual-based tests for cointegration in models with regime shifts*. Journal of Econometrics, 70(1): 99–126. [https://doi.org/10.1016/0304-4076\(96\)041685-7](https://doi.org/10.1016/0304-4076(96)041685-7)

- Gupta, P. (2026): *Inflation targeting in India – the past, the present and the future*. Speech at Joint Seminar and Discussion ‘India’s inflation targeting framework’ and ‘regional economic outlook for Asia and Pacific’, National Council of Applied Economic Research (NCAER), New Delhi, 6 May. <https://www.bis.org/review/r260506b.htm>. Downloaded: 9 September 2026.
- Hamilton, J.D. (1994): *Time series analysis*. Princeton University Press. <https://doi.org/10.1515/9780691218632>
- IMF (2024): *World economic outlook: Policy pivots, rising threats*. International Monetary Fund, October. <https://www.imf.org/-/media/files/publications/weo/2024/october/english/text.pdf>. Downloaded: 27 March 2026.
- Kapur, M. (2013): *Revisiting the Phillips curve for India and inflation forecasting*. Journal of Asian Economics, 25: 17–27. <https://doi.org/10.1016/j.asieco.2012.12.002>
- Lee, J. – Strazichich, M.C. (2003): *Minimum Lagrange multiplier unit root test with two structural breaks*. Review of Economics and Statistics, 85(4): 1082–1089. <https://doi.org/10.1162/003465303772815961>
- Lin, S. – Ye, H. (2009): *Does inflation targeting make a difference in developing countries?*. Journal of Development Economics, 89(1): 118–123. <https://doi.org/10.1016/j.jdevco.2008.04.006>
- Lütkepohl, H. (2005): *New introduction to multiple time series analysis*. Springer.
- MacKinnon, J.G. (1996): *Numerical distribution functions for unit root and cointegration tests*. Journal of Applied Econometrics, 11(6): 601–618. [https://doi.org/10.1002/\(SICI\)1099-1255\(199611\)11:6%3C601::AID-JAE417%3E3.0.CO;2-T](https://doi.org/10.1002/(SICI)1099-1255(199611)11:6%3C601::AID-JAE417%3E3.0.CO;2-T)
- McKinnon, R.I. (1973): *Money and capital in economic development*. Brookings Institution.
- Mishra, P. – Montiel, P. (2013): *How effective is monetary transmission in low-income countries? A survey of the empirical evidence*. Economic Systems, 37(2): 187–216. <https://doi.org/10.1016/j.ecosys.2012.12.001>
- Pesaran, M.H. – Shin, Y. – Smith, R.J. (2001): *Bounds testing approaches to the analysis of level relationships*. Journal of Applied Econometrics, 16(3): 289–326. <https://doi.org/10.1002/jae.616>
- Rangarajan, C. – Mishra, P. (2013): *India’s external sector: Do we need to worry?*. Economic and Political Weekly, 48(7): 52–59. <https://www.jstor.org/stable/23391309>
- Rastogi, S. – Kanoujiya, J. (2024): *The volatility spillover effect of macroeconomic indicators and strategic commodities on inflation: evidence from India*. South Asian Journal of Business Studies, 13(2): 180–200. <https://doi.org/10.1108/SAJBS-10-2021-0387>

RBI (2014): *Report of the expert committee to revise and strengthen the monetary policy framework*. Urjit Patel Committee, Reserve Bank of India, January. <https://the1991project.com/sites/default/files/2024-12/2014%20Urjit%20Patel%20Committee%20Report.pdf>. Downloaded: 27 March 2026.

RBI (2024): *Handbook of statistics on the Indian economy*. Reserve Bank of India.

Sargent, T.J. – Wallace, N. (1981): *Some unpleasant monetarist arithmetic*. Federal Reserve Bank of Minneapolis Quarterly Review, 5(3). <https://doi.org/10.21034/qr.531>

Stock, J.H. – Watson, M.W. (2003): *Forecasting output and inflation: The role of asset prices*. Journal of Economic Literature, 41(3): 788–829. <https://doi.org/10.1257/jel.41.3.788>

Toda, H.Y. – Yamamoto, T. (1995): *Statistical inference in vector autoregressions with possibly integrated processes*. Journal of Econometrics, 66(1–2): 225–250. [https://doi.org/10.1016/0304-4076\(94\)01616-8](https://doi.org/10.1016/0304-4076(94)01616-8)

Zivot, E. – Andrews, D.W.K. (1992): *Further evidence on the great crash, the oil-price shock, and the unit-root hypothesis*. Journal of Business & Economic Statistics, 10(3): 251–270. <https://doi.org/10.2307/1391541>

Appendix

Table 10

Pairwise correlation matrix of explanatory variables, 1955 Q1–2024 Q4

	Δfd	Δm	Δmsp	gap	Δer	Δoil	Δpe
Δfd	1.00						
Δm	0.42	1.00					
Δmsp	0.38	0.31	1.00				
gap	0.15	0.18	0.12	1.00			
Δer	–0.21	–0.18	–0.14	–0.09	1.00		
Δoil	0.19	0.22	0.27	0.11	–0.32	1.00	
Δpe	0.34	0.41	0.47	0.18	–0.22	0.29	1.00

Note: Pearson correlation coefficients across the estimation sample, computed in first differences for I(1) variables and in levels for the I(0) output gap. All absolute pairwise correlations are below 0.50, and variance inflation factors (untabulated) range from 1.18 (gap) to 2.31 (Δpe), well below the conventional threshold of 5, indicating that multicollinearity does not substantively distort the ARDL coefficient estimates. The highest correlation is between Δmsp and Δpe (0.47), consistent with the role of food-price dynamics in shaping household expectations.