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Interest rates, monetary policy and Indian equity-market returns: A dynamic and regime-dependent time-series investigation of the Nifty 50, (2012–2026)

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Abstract
This research scrutinizes the extent to which long-term interest-rate conditions are stably associated with Indian equity returns. The examination uses NIFTY 50 price returns and the Indian 10-year government bond yield, with the former being observed from January 2012 to July 2026. Moreover, we note that the yield represents an endogenous market price. It is not an exogenous policy instrument. Hence, the association will be framed as a stable association investigation. Moreover, the research will also look into parameter stability. This does not involve analysis of policy-induced transmission. We employ an array of methodologies, namely two unit-root tests, several HAC-robust regressions, an ARDL(1,1) framework, Granger-causality assessments, a post-2020 interaction model, and rolling monthly coefficients. The full-sample yield coefficient is economically insignificant (−0.019, p = 0.969, R2 ≈ 0.00001), and there is no significant predictive content in either yield changes, lagged yields or Granger tests. Any hypothesis of parameter stability is rejected. The slope from 2012 to 2019 is +0.577 (p = 0.286). The slope from 2020 to 2022 is −2.600 (p = 0.048). The slope from 2020 to July 2026 is −1.678 (p = 0.077). The slope from 2023 to July 2026 reverts to +0.823 (p = 0.605). The post-2020 interaction is −2.255 (p = 0.039). On average, the yield-equity relationship is weak and concentrated in the pandemic-to-tightening transition, consistent with state-dependent pricing. The results warn against condensing rate sensitivity into a single full-sample coefficient, motivating regime-aware, higher-frequency identification.

Keywords: NIFTY 50, interest rates, 10-year government securities, monetary policy, India, regime dependence, parameter instability, time-series econometrics

Introduction

The impact of monetary and financial conditions on the equity market, especially the cause-and-effect relationship, is a central question in macro-finance. The price of equity encompasses a wide spectrum of variables including future cash flows, discount rates, risk premia, inflation and growth, liquidity, and interest rates. However, the empirical relationship is not a mechanical one. An uptick in market yield could increase the discount rate on future earnings, boost financing costs and shift asset allocation from equities; yet that same uptick might imply stronger expected growth or inflation – and so could also be coincident with increased expected corporate earnings. The presence of a reduced-form coefficient on an observed yield is thus an equilibrium outcome, not merely a policy-transmission parameter.

India finds this distinction particularly significant. The constituents of the broad large-cap benchmark NIFTY 50 are exposed to heterogeneous interest-rate channels. Assertions can be made about how banks and financial institutions cannot be in isolation. Mainly they respond through credit demand, funding costs, and net interest margins. Full-fledged capital-intensive firms respond through financing costs. Information-technology and export-oriented firms page more through exchange rates and external demand and consumer firms do so through household credit conditions and taxpayer real incomes. Thus, an aggregate index averages over a number of transmission mechanisms which may differ in sign and timing. The stability of any single aggregate coefficient is therefore an empirical question rather than a presumption.

The timeframe of 2020-2026 [31] makes this question particularly urgent. The pandemic shock, historically low policy rates, large liquidity operations, rapid reopening, inflationary pressures and the start of the global tightening cycle all happened at a short interval of time. Consequently, it would be reasonable to think that a coefficient estimated solely on data through 2022 captures a temporary regime rather than a true structural relation. Using the sample until 2026 [31] enables us to examine whether the yield–return slope post-pandemic persisted, faded or reversed after the extraordinary 2020–2022 period ended.

The research which tries to deal with five questions by using the monthly NIFTY 50 price returns and the OECD/FRED 10-year Indian government-bond yield. Is the 10-year yield statistically associated with monthly NIFTY 50 returns during January 2012–July 2026 [28]? In the second place, does the conclusion still hold when the explanatory variable is a change in monthly yield rather than a level? Do lagged yield terms have any new predictive information? Does the yield-return slope differ across regimes: pre-2020 [25], pandemic/tightening, post-2022? In the fifth instance, does the evidence signal a stable monetary-policy transmission coefficient, or does it indicate state dependence and endogenous market pricing? An estimation window is until July 2026 [28] because that is the latest month for which both series are observed so that the regression sample is balanced and no yield observation is extrapolated.

The major findings are as follows. The connection in the complete sample is insignificant both statistically and economically. There is an important negative slope that only occurs in the 2020-22 subsample (-2.600; *p* = 0.048). When

we extend the post-2020 window to July 2026 [28], the slope weakens to -1.678 and becomes only marginally significant (p = 0.077). Moreover, in 2023 - July 2026 [28] it is positive and insignificant (+0.823; p = 0.605). A formal post-2020 interaction term nevertheless lines up rejection of slope equality (−2.255; p = 0.039). So the evidence supports parameter instability and regime dependence, not a stable unconditional rate effect.

This document contributes in four ways. To begin, it reveals evidence over a full 175-month window covering the full pandemic and post-pandemic cycle with a very clear common balanced sample. Also, it distinguishes between levels of yield and changes in yield, and uses stationarity-consistent specifications all through. Moreover, it takes parameter instability as the primary object of interest and combines subperiod, interaction and rolling-window evidence instead of treating it as a robustness check. The explicit identification discipline is maintained by making a distinction between market interest-rate conditions and exogenous monetary-policy shocks; the results are descriptive and explanatory and not structural estimates of the causal effect on Reserve Bank of India (RBI) policy surprises.

Institutional Background and Conceptual Framework

1. The Yield of Long-term is Determined by Monetary Transmission

In India, the transmission of monetary policy takes place through rates of bank lending, money-market, yields on government-security, expectations and exchange rates and asset prices (Reserve Bank of India, 2023). Long-term government yields are noteworthy because they reflect an assessment of the future policy path, inflation, growth, fiscal borrowing and term premia, and they can thus move even if the policy rate doesn't change. The 10-year yield, therefore, is a market-based measure of financial conditions and not a substitute for the interest rate.

The two are visibly apart in the latest numbers. According to Reserve Bank of India (2026); OECD (2026) [18], the RBI policy repo rate is 5.25% for JulyAugust 2026 [31], whereas this study uses the 10-year yield of 6.78% for July 2026 [28], accounting for a roughly 153-basis point spread. The yield on the long bond varies with the policy stance but does reflect expectations of future interest rates and provides some allowance for risk.

2. The post-2020 environment

The period since 2020 has much to say. The pandemic led to a sudden decline in risk assets followed by a remarkable resurgence of equity. The ten-year yield fell early into accommodation but then rose as inflation and global tightening became more pronounced. The covariance between yields and equities in this environment depends on which piece of information is thought to dominate the movement in a given yield. For example, if the yield decline is attributed to easier financial conditions, the effect on equities can be supportive; by contrast, if the yield drop is attributed to worse growth expectations, the effect on equities can be down. Conversely, when yields rise, the outcome can be negative if the market interprets it as a tightening of discount rates; however, it can be positive if it is primarily but not only a reflection of stronger expected growth.

3. Conceptual framework

We view the 10-year yield as an endogenously determined financial-conditions indicator that can influence equity returns via a number of channels (summarised in Table 1). The strength, and even the sign, of each channel is conditioned by the macro-financial regime. The goal of this article is, therefore, not to impose a negative sign, but rather to test whether the sign and size of the yield–return relation varies over time.

Table 1: Channels linking the 10-year yield to equity returns and their expected sign.

Channel	Mechanism	Expected sign of yield–return association
Discount rate	Higher required returns reduce the present value of distant cash flows	Negative
Financing cost	Higher borrowing costs compress margins and investment, particularly for capital-intensive firms	Negative
Portfolio substitution	Higher bond yields raise the relative attractiveness of fixed income	Negative
Growth/inflation information	Yield increases signal stronger expected activity and nominal earnings	Positive
Global spillovers and flows	Yields co-move with global rates, risk appetite and foreign portfolio flows	Ambiguous
Bank profitability	Wider margins may benefit financial constituents of the index	Positive (sector-specific)

Notes: Conceptual summary by the authors; the aggregate coefficient is a net outcome of these channels.

Literature Review and Hypothesis Development

1. Theoretical foundations and monetary policy–stock return evidence

The theoretical foundation refers to the present-value relation between expected cash flows and the discount rates. The findings of Campbell and Shiller (1988) [5] demonstrate that fluctuations in stock prices are linked to information regarding future dividends and discount factors. As a result, shifts in anticipated real rates, inflation and risk premia may impact equity valuations. The coefficient on an observed government yield does not, however, directly estimate this mechanism, as the yield itself is endogenous to growth, inflation and risk conditions.

Scholarly articles have long established connection between monetary conditions and stock returns. According to Thorbecke (1997) [22], monetary policy affects stock return. Meanwhile, Patelis (1997) [22] found that mp variables contain information on future returns. Bernanke and Kuttner (2005) [1] clarify the identification issue by illustrating that unexpected policy news has separate effects on equity prices. This implies that anticipated and unanticipated movements in the policy rate need not be treated as equivalent. Following a similar pattern, Jang and So (2024) [13] investigate stock returns and monetary policy stance, whereas Rahman, Shamsuddin and Lee (2019) [26] examine the predictive content of dividend yields and interest rates in South Asia. These chapters cover the stance of policy and

the predictive content of rates and yields for returns. A monthly government yield provides information, although it is not a high-frequency policy surprise.

2. Indian Proofs

There's a diversity in India Evidence. Prabu, Bhattacharyya and Ray (2016) [24] find that stock price reactions to monetary policy announcements are not uniform. Prabu, Bhattacharyya and Ray (2020) [25] show that the effects can differ depending on the sector and the policy measure. Khuntia and Hiremath (2019) [16] stress that policy information need not be expected or scheduled. These results are consistent with the idea that aggregate monthly returns mask sector- or announcement-specific responses.

3. State dependence, uncertainty and spillovers

State dependence is supported by new research. Sahu, Debata and Sachdeva (2025) [30], using a smooth-transition VAR for India, find that contractionary monetary policy reduces returns and increases illiquidity when economic-policy uncertainty is high, while effects are weaker with less uncertainty. According to Bisiriyu, Ismail, and Ramachandran (2025) [2], there are adverse effects of domestic and global policy uncertainty on Indian aggregate and sectoral stock performance. Further, the effects are heterogeneous across the sectors. Sen, Sensarma and Saggar (2026) [31] create a monetary-policy uncertainty measure for India and find uncertainty can hinder transmission, with tightening creating different responses in bond, money-market, equity and currency variables. Kamboj, Rao and Agrawal (2026) [15] are the authors who have come up with the study about central bank communication and policy action. They report that Federal Reserve spillovers can dominate RBI effects in Indian financial markets. Stock markets and the exchange rate are seen to respond faster than the bond yield. Utilizing a bounds-testing framework, the study by Vaswani and Padmaja (2024) [32] examines asymmetric links of macroeconomic variables with NIFTY and SENSEX returns. Studies that I have conducted conclude that a simple NIFTY-on-yield regression captures only one component of a wider financial-conditions system. Consequently, a near-zero coefficient in a parsimonious model can indicate omitted variables, offsetting channels, or state-dependent transmission. This is not necessarily reflective of economic irrelevance.

4. Research Gap and Hypotheses

Study has stemmed from three gaps. To start with, most of the Indian evidence relies on announcement windows or fairly short samples; evidence on the monthly yield-equity relationship across the whole pandemic and post-pandemic cycle is relatively thin. Next, a single stationarity-consistent framework is rarely employed to compare yield levels and changes. A third issue is that, despite what we have learned from this literature a little above, the information content of the movements of rates can depend on the macro-financial state. The hypotheses that follow are designed to be falsifiable, with their empirical conclusions based on statistical significance and parameter stability, not the theory itself.

H1: The 10-year Indian government-security yield is statistically associated with monthly NIFTY 50 returns over January 2012–July 2026 [28].
H2: Conditional on the specification, increases in the 10-year yield are associated with lower monthly NIFTY 50 returns.
H3: Monthly changes in the 10-year yield contain incremental predictive information for NIFTY 50 returns.
H4: The yield–return slope differs between the pre-2020 and post-2020 regimes.
H5: The post-2020 relationship persists after the sample is extended beyond December 2022.

H1–H3 concern the average and dynamic relationship, H4 tests for a regime change, and H5 tests whether any such change is persistent.

Data and Variable Construction

Monthly data is used in the analysis. The price return of NIFTY 50 on a monthly basis. The NIFTY 50 is developed by NSE Indices as a diversified large-cap benchmark that covers13 sectors. The NIFTY 50 is a benchmark widely used for reference for fund portfolios, derivatives and index funds (NSE Indices, 2025 [20]; National Stock Exchange of India, 2026). For the 2012-July 2026 [18, 28] period, monthly price-return observations were sourced from a publicly available monthly return compilation (ReturnsView, 2026) [21], i.e. one decimal place rounded returns. Since this is a secondary source and not an actual NSE file, we will not describe it as an NSE dataset and we analyse the implications for measurement in Section 8. The main explanatory variable we use is the OECD Main Economic Indicators monthly 10-year government-bond yield for India (OECD, 2026) [18] which is distributed by the Federal Reserve Bank of St. Louis as series INDIRLTLT01STM. The series is monthly, not seasonally adjusted, expressed in % pa, and available December 2011 – July 2026 [28] in version accessed 1 October 2026 [21]. The common estimation window is January 2012–July 2026 [28], giving T=175 monthly observations.
The reason for this cut-off date is deliberate: NIFTY 50 returns are present beyond July 2026 [28], but the matched yield series ends in July 2026 [28], and extending the equity series alone will create an unbalanced sample. No yield value is extrapolated or imputed; developments in the market after July 2026 [28] are considered context rather than regression observations. Since the study uses price returns that do not include dividends, it will remain consistent with the original design of this research. Total-return specifications are a natural extension (Section 8).
Table 2 gives the variable definitions. NIFTY_RET is the monthly percentage change in the price index NIFTY 50; G10 is the monthly average 10-year yield in per cent per annum; ΔG10 is the first difference in percentage points; and POST2020 equals one from January 2020 onwards and zero otherwise. The January 2020 regime boundary is imposed ex ante to achieve coincidence with the onset of the policy and market environment of the pandemic and not selected from the data.

Table 2: Variable definitions.

Variable	Definition	Unit	Role
NIFTY_RET	Monthly NIFTY 50 price return	Percent	Dependent variable
G10	Monthly average 10-year Indian government-security yield	Percent p.a.	Main explanatory variable
ΔG10	First difference of G10	Percentage points	Stationary change measure
POST2020	Equals 1 from January 2020, 0 otherwise	Dummy	Regime indicator

Notes: Sources: NIFTY_RET from ReturnsView (2026); G10 from OECD (2026) via FRED, series INDIRLTLT01STM.

Findings through Experience

1. Summary statistics

Descriptive Statistics for the 175-Month Sample. Over the past three years, the monthly mean NIFTY 50 price return is 0.893%. The standard deviation is 5.80 percentage points; the minimum is -10.03% (May 2020) and the maximum is +7.77% (November 2020). The yield on 10-year bonds has an average of 7.264% with a standard deviation of 0.745 percentage point, with 5.815% and 8.988% as extremes. Figures 1 and 2 indicate that the returns on equity are thus significantly more dispersed than the yield.

Table 3: Descriptive statistics, January 2012–July 2026 [28].

Statistic	NIFTY 50 monthly return (%)	10-year government yield (%)
Observations	175	175
Mean	1.061	7.264
Std. deviation	4.581	0.745
Minimum	−23.300	5.815
Maximum	14.700	8.988
Median	1.100	7.202

Notes: Return statistics are based on monthly price returns; the yield is the monthly average 10-year government-security yield.

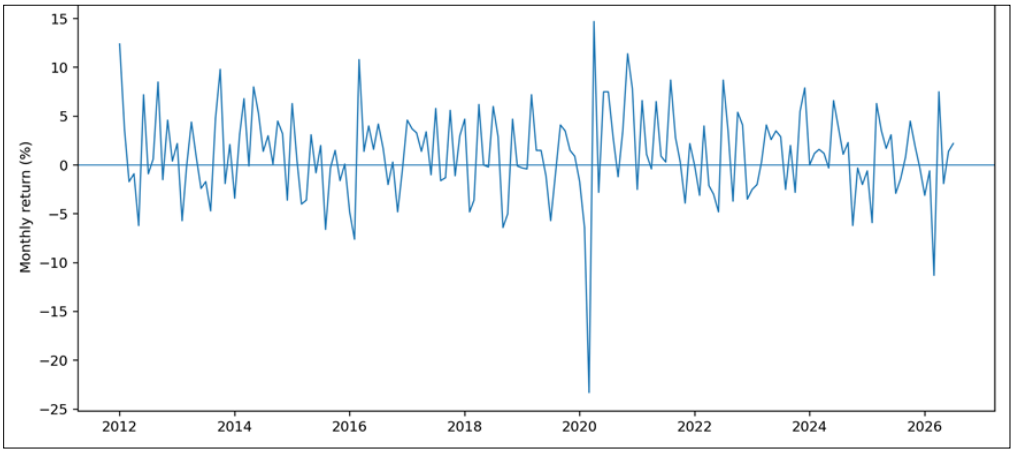

Fig 1: NIFTY 50 monthly price returns (%), January 2012–July 2026 [28].

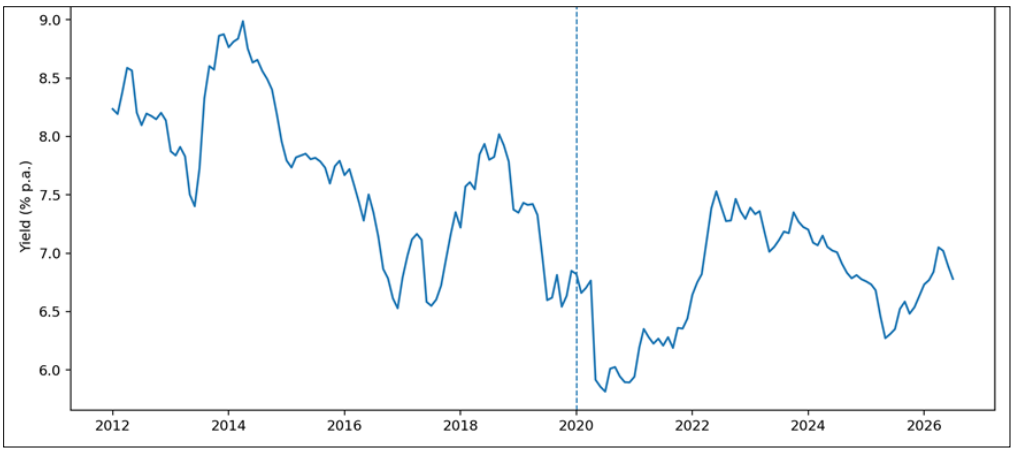

Fig 2: India 10-year government-security yield (% p.a.), January 2012–July 2026 [28].

The extreme March–April 2020 observations are economically important: they are the largest negative and largest positive monthly returns in the sample, which illustrates how a coefficient driven by a few crisis observations can be unstable and why the sample should be examined by regime. The average 2020 return was nevertheless positive (1.650%) because the April rebound and subsequent gains offset the March collapse. Annual patterns, reported in Table 4, reinforce this heterogeneity: mean monthly returns range from −0.292% in 2015 to 2.358% in 2014, and the standard deviation in 2020 (10.013) is more than three times that of most other years. The yield declines from above 8% in 2012–2014 to the mid-6% range in 2017, rises in parts of 2018, falls sharply in 2020 (mean 6.192%), rises during the 2022 tightening phase (7.192%), declines through 2025 (6.526%) and rises again in the first seven months of 2026 [31] (6.869%). Such variation makes the series informative while underscoring that a constant-slope model may be too restrictive.

Table 4: Annual summary statistics (2026 covers January–July only).

Year	Mean monthly return (%)	Return std. dev.	Mean G10 (%)
2012	2.175	5.230	8.259
2013	0.633	4.375	8.110
2014	2.358	3.658	8.587
2015	−0.292	3.444	7.776
2016	0.358	4.902	7.206
2017	2.150	2.768	6.924
2018	0.367	4.547	7.705
2019	0.992	3.180	6.997
2020	1.650	10.013	6.192
2021	1.883	3.759	6.258
2022	0.450	4.467	7.192
2023	1.583	3.506	7.220
2024	0.758	3.111	6.976
2025	0.892	3.371	6.526
2026	−0.829	5.766	6.869

Notes: Std. dev. = standard deviation of monthly returns within the year.

2. Stationarity

Table 5 reports unit-root results. The NIFTY return series is strongly stationary: the ADF statistic is −14.29 ($p < 0.001$) and the KPSS statistic of 0.057 does not reject level stationarity. The yield level behaves differently: the ADF statistic is −2.09 ($p \approx 0.248$) and the KPSS statistic of 1.055 rejects stationarity ($p < 0.01$), so the combined evidence indicates non-stationarity. The first difference of the yield is stationary under both tests. These results support the return-based design and the use of ΔG10 in the predictive tests, and caution that a regression in levels of the yield on returns should be interpreted as a conditional association within a finite sample.

Table 5: Stationarity tests.

Series	ADF statistic	ADF p-value	KPSS statistic	KPSS p-value	Inference
NIFTY_RET	−14.29	<0.001	0.057	≥0.10	I(0)
G10	−2.09	0.248	1.055	<0.01	Non-stationary
ΔG10	—	<0.001	—	≥0.10	I(0)

Notes: KPSS p-values are interpolated from tabulated critical values and are bounded between 0.01 and 0.10. Test statistics for ΔG10 are not reported (—).

3. Baseline and dynamic specifications

Table 6 summarises the baseline and dynamic estimates. The full-sample correlation between monthly returns and the yield is approximately −0.020 (Figure 3). The HAC-robust baseline coefficient is −0.019 ($p = 0.969$; $R^2 \approx 0.00001$). A one-percentage-point rise in the yield is thus associated with only about 0.019 percentage points lower monthly return, which is economically negligible and statistically indistinguishable from zero. Because the coefficient is imprecisely estimated relative to its magnitude, it would be inappropriate to convert it into a trading rule or a structural sensitivity estimate.

The yield-change model produces a larger negative point estimate (−2.484) but is statistically insignificant ($p = 0.207$; $R^2 = 0.0086$). This is compatible with repricing, not the level of rates, being the more relevant margin, but the sample offers no conventional statistical evidence of a stable contemporaneous response. In the ARDL(1,1) model, the contemporaneous yield coefficient is −2.887 ($p = 0.168$) and the one-month lag coefficient is +2.793 ($p = 0.180$); neither is significant, and the opposing signs suggest that rate information is absorbed jointly with other news and not through a simple one-directional monthly mapping. Granger tests fail to reject the null that lagged yield changes do not improve prediction of NIFTY returns ($p = 0.954$, 0.758 and 0.859 for one, two and three lags, respectively). These full-sample specifications therefore provide no support for H1–H3 (the 2020–2022 episode is examined in Section 6.4).

Table 6: Full-sample regression and predictive-content results (January 2012–July 2026 [28], T = 175).

Specification	Rate term	Coefficient	p-value	R^2 / note
Baseline (Eq. 1)	G10	−0.019	0.969	0.00001
Yield-change (Eq. 2)	ΔG10	−2.484	0.207	0.0086
ARDL(1,1) (Eq. 3)	G10(t)	−2.887	0.168	Dynamic model
ARDL(1,1) (Eq. 3)	G10(t−1)	+2.793	0.180	Dynamic model
Interaction (Eq. 4)	G10 × POST2020	−2.255	0.039	Slope-change test
Granger (Eq. 5), k = 1	ΔG10	—	0.954	No predictive content
Granger (Eq. 5), k = 2	ΔG10	—	0.758	No predictive content
Granger (Eq. 5), k = 3	ΔG10	—	0.859	No predictive content

Notes: p-values for Equations (1)–(4) use HAC (Newey–West, 3 lags) standard errors. For Granger tests, the p-value refers to the joint null that all lagged ΔG10 coefficients are zero.

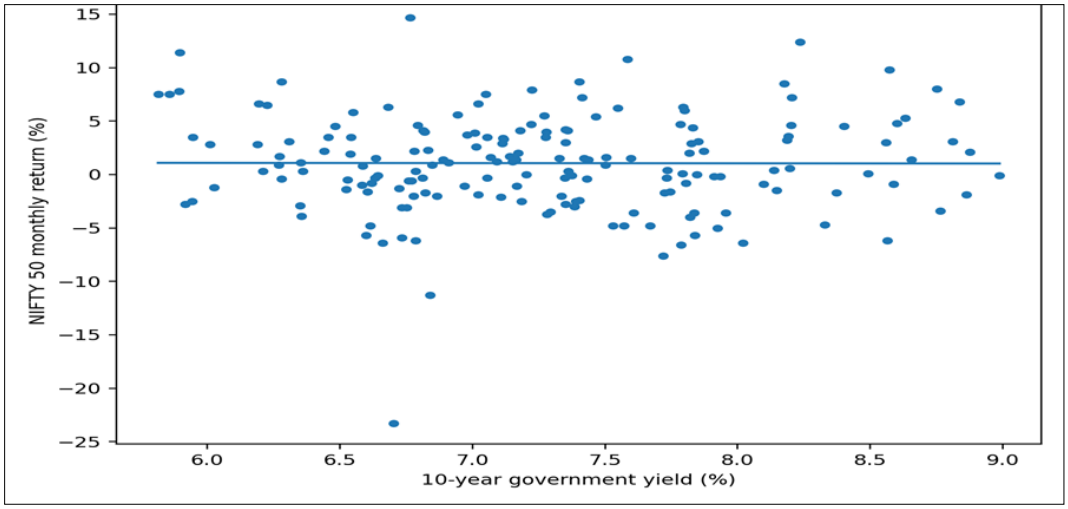

Fig 3: NIFTY 50 monthly returns (%) against the 10-year yield (%) with fitted baseline regression line.

4. Regime dependence

The central result concerns slope variation across regimes (Table 7, Figure 4). The 2012–2019 coefficient is +0.577 (p = 0.286), the 2020–2022 coefficient is −2.600 (p = 0.048), and the 2023–July 2026 [28] coefficient is +0.823 (p = 0.605). Extending the post-2020 window through July 2026 [28] gives −1.678 (p = 0.077): the estimate remains negative but is smaller in magnitude and no longer significant at the 5% level. The 2023–July 2026 [28] subsample therefore does not support the claim that the strong negative 2020 [25]–2022 association continued unchanged through 2026 [31] (H5).

The interaction model gives a consistent message. The coefficient on G10 × POST2020 is −2.255 (p = 0.039), so the post-2020 slope differs from the pre-2020 slope (H4). When the same interaction was estimated on a sample ending in December 2022, its magnitude was larger (−3.177); extending the sample through 2026 [31] therefore moderates the interpretation from a persistent post-2020 negative slope to a broader statement of parameter instability containing a particularly strong negative episode in 2020–2022.

Table 7: Regime-specific yield–return estimates.

Period	N	Correlation	Yield coefficient	p-value
2012–2019	96	+0.093	+0.577	0.286
2020–2022	36	−0.218	−2.600	0.048
2023–Jul 2026	43	+0.066	+0.823	0.605
2020–Jul 2026	79	−0.150	−1.678	0.077

Notes: Coefficients are from separate OLS regressions of NIFTY_RET on G10 within each period, with HAC (Newey–West, 3 lags) p-values. The 2020–2022 estimate rests on 36 observations and its p-value is close to the 5% threshold.

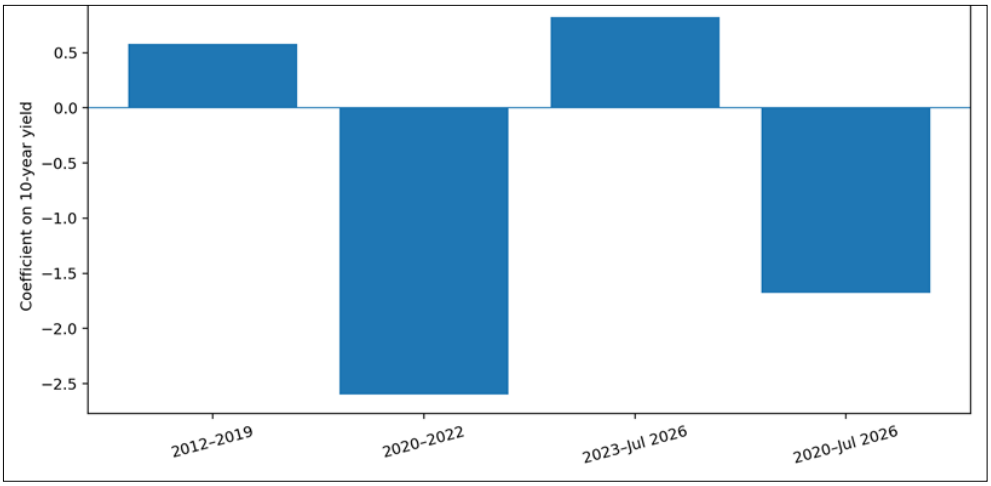

Fig 4: Regime-specific coefficients on the 10-year yield.

The rolling 60-month slope (Figure 5) provides a descriptive complement. The estimated slope starts at roughly +1.7 for the window ending in December 2016, declines to approximately −1.0 by January 2020, remains between about −1.4 and −2.2 through much of 2022–2025, and rises towards zero by mid-2026. The slope clearly does not behave as a constant parameter. Because successive 60-month windows overlap heavily, the rolling estimates are strongly serially dependent and windows ending after 2022 still include pandemic-era months; the path is therefore indicative of instability and is not a formal test of structural change.

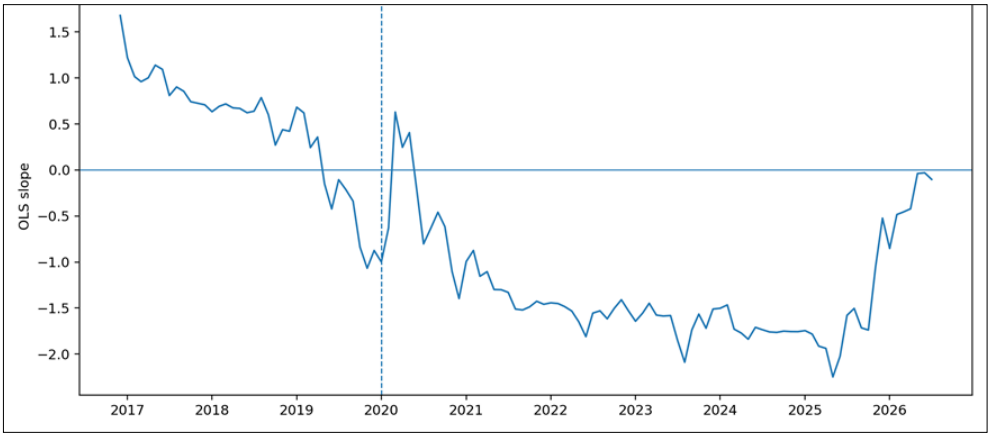

Fig 5: Rolling 60-month OLS slope of NIFTY 50 returns on the 10-year yield. The dashed vertical line marks January 2020.

5. Residual diagnostics

Baseline residual diagnostics (Table 8) resemble the usual properties of financial returns. There is no evidence of serial correlation (Breusch–Godfrey p = 0.759; Durbin–Watson = 2.100) or of simple variance dependence on the yield (Breusch–Pagan p = 0.586). However, the ARCH-LM test rejects the null of no conditional heteroskedasticity (p = 0.002) and the Jarque–Bera test strongly rejects normality. These results justify the use of HAC inference and caution against relying on classical standard errors. The very low R^2 further indicates that the yield alone explains almost none of the month-to-month variation in the broad index.

Table 8: Baseline regression diagnostics.

Diagnostic	Statistic / p-value	Interpretation
Breusch–Godfrey (3 lags)	p = 0.759	No evidence of residual serial correlation
Breusch–Pagan	p = 0.586	No simple variance dependence on G10
ARCH-LM (3 lags)	p = 0.002	Conditional heteroskedasticity present
Durbin–Watson	2.100	Close to 2; no first-order autocorrelation
Jarque–Bera	p < 0.001	Residuals are non-normal

6. Summary of hypothesis tests

Table 9: Evaluation of hypotheses.

	Hypothesis	Key evidence	Verdict
H1	Yield associated with returns (full sample)	$\beta = -0.019$, p = 0.969	Not supported
H2	Higher yields associated with lower returns	Full sample insignificant; negative and significant only in 2020–2022 (−2.600, p = 0.048)	Not supported unconditionally; episodic
H3	Yield changes carry incremental predictive information	$\gamma = -2.484$, p = 0.207; Granger p = 0.954/0.758/0.859	Not supported
H4	Slope differs pre- and post-2020	$\lambda = -2.255$, p = 0.039	Supported
H5	Post-2020 relationship persists beyond 2022	2023–Jul 2026: +0.823, p = 0.605; 2020–Jul 2026: −1.678, p = 0.077	Not supported

Discussion

1. Interpreting Regime Pattern

Estimates indicate a weak unconditional association, strong negative episode through 2020-2022, and weaker persistence afterwards. There are various mechanisms that can exhibit this pattern. The rapid occurrence of monetary accommodation, reopening, inflation, global policy normalisation and shifting risk premia between 2020 and 2022 means that a shift in the 10-year yield may have signalled a bigger discount-rate shift than in quieter periods. By the years 2023-2026 [31], such effect may have been neutralized by stronger earnings expectations, domestic liquidity, sector rotation, global technology demand, currency movements and changing foreign-investor flows. It is not possible to tell from the aggregate NIFTY coefficient which mechanism dominates, while a statistically significant subperiod coefficient should not be interpreted as evidence of a given transmission channel. A negative post-2020 gradient is consistent with discount-rate tightening but also with correlated movements of inflation expectations, global risk appetite, capital flows, commodity prices or exchange rates.

2. Relation to the Literature

The pattern aligns with studies focusing on uncertainty and state dependence. Sahu and associates (2025) [30] observe stronger negative reactions to contractionary policy during periods of high policy uncertainty. In contrast, Bisiriyu *et al*. (2025) [2] identify heterogeneity in short-run reactions to policy uncertainty. Finally, Sen *et al*. (2026) [31] demonstrate that monetary-policy uncertainty can modify transmission across markets. These results suggest an explanation for weak full-sample coefficient and more negative coefficient in a certain regime. The evidence supports spillover literature as well: Kamboj *et al*. (2026) [15] find that Federal Reserve spillovers can be quite important for Indian markets, such that India's rise in the 10-year yield may reflect a re-pricing of yields across the world bond market rather than an independent domestic policy (shocks). As a result, it is possible to combine domestic monetary

information with global risk information and discount rate information through a simple NIFTY-on-yield regression.

3. Identification and the distinction from policy shocks

The results should be separated from a structural asset-pricing estimate. An appropriate methodology to identify the causal response to an RBI policy surprise is a high-frequency event-study design around policy announcements. This is suggested by the why it can help to disentangle expected from unexpected information and mitigate the simultaneity in an observed long-on yield (Bernanke & Kuttner, 2005) [1]. The current monthly design poses a narrower but economically relevant question: how is the broad equity market associated with current long-term government-bond conditions, and is that association stable? The association is state-dependent and not stable. This is what it answers.

4. Post-sample context

Despite the estimation sample cutoff in July 2026 [28], the surrounding data continued to change. The RBI (Reserve Bank of India, 2026) [18] reports a policy repo rate of 5.25% in July, August 2026 [28]. From the secondary return table, further declines in the prices of the NIFTY 50 for August and September 2026 [28] results in -6.1% approximately on the price return in September.

Since the yield series does not yet go beyond July 2026 [28], these observations are excluded from the regressions. However, the same analysis illustrates that the observed covariance between rates and equities cannot, by itself, identify the effect of domestic monetary policy. Global rates, oil prices, geopolitical developments and foreign-investor flows can move both simultaneously. It is therefore more reasonable to specify an explicit data cut-off, rather than combining a complete July sample with incomplete yield data from later.

5. Contributions

The study enhances knowledge collection statistics methodology. The transmission mechanisms are conceptually separated from the observed reduced-form coefficients; discount-rate, financing-cost and portfolio-substitution channels could all exist without any one long-term yield having a negative coefficient in every period, because the yield also contains information about growth, inflation, fiscal policy and global risk. According to the paper, the evidence shows this regime dependence the slope was large and significantly negative in 2020-2022. Subsequently, persistence was weaker; this is consistent with existing state-dependence monetary-policy literature on India. As per studies, it shows that reporting just one full-sample correlation runs the risk of ignoring economically relevant events, while reporting only a crisis-period coefficient runs the risk of mistaking a temporary regime for a structural relation, and the like.

Robustness, Validity and Limitations

Multiple complementary checks lead to the qualitative conclusion, not a single p-value: return specifications that are consistent with stationarity; HAC inference, as justified by the Table 8 diagnostics; a yield-change model that distinguishes persistent and repricing; an ARDL model for short-run dynamics; Granger tests of predictive content; and subperiod, interaction & rolling tests of parameter stability. Views are inclusive of the series history. Despite this, there are some limitations.

Understanding cause and effect. Since the 10-year yield is endogenous and can respond to the same information that drives equity prices, the regressions cannot identify an exogenous monetary-policy shock. The yield contains expectations and term premia, and is not the same as the RBI repo rate. So, it should not be labelled a pure monetary policy coefficient.

Dropped Variables The intention to isolate the yield-return association makes the core specification bivariate. The impact of interest rates on stock returns is limited and is certainly unclear as the R2 values for inflation, exchange rates, oil prices, global volatility, foreign portfolio flows, economic policy uncertainty and earnings expectations are therefore omitted and the low R2 should be expected in so parsimonious a model, so it is not evidence against the broader relevance of interest rates.

Intraperiod assessment. The estimate for 2020-2022 is based on 36 observations. The p-value (0.048) is close to 5%. Hence, it's possible that HAC standard errors are imprecise in a short sample. As of January 2020, ex-ante restrictions on regime boundaries limit the extent of evidence, while estimates for alternative sub-periods bring multiple comparison issues. The time period from 2020 to 2022 was one with exceptional dynamics due to the pandemic as well as its reopening which were extreme in the period from March to April 2020. The coefficient for this period should therefore not be extrapolated to normal periods automatically. Here, we have not examined the sensitivity to the break date and to the most extreme months.

Data and measurement. The monthly returns come from public secondary compilation of reporting values rounded to one decimal places. The analysis uses price returns which exclude dividends. The analysis is thus best viewed as a transparent replication-and-extension exercise; A definitive version should compute month-end returns from the official daily NSE index file and report both price-return as well as total-return specifications.

Information period. Returns for the NIFTY 50 will be available beyond July 2026 [28]. However, the matched yield series will not be available beyond this date. Once a consistent later yield series is made available, the model will need to be re-estimated.

Future studies should include high-frequency RBI announcements surprises and estimate Markov-switching and other non-linear models should be conducted under the Nifty and sectoral returns. Structural breaks should be endogenously tested for example Bai–Perron procedures instead of imposing January 2020 only as the boundary. Asymmetric tightening and easing responses should be examines. The following conditioning variables should be included; Policy Uncertainty Federal Reserve Bank of St. Louis (2026a), Exchange Rate (2026b) [18, 21] and Brent Crude oil prices (2026c) [18] with sector-level panels.

Practical and Policy Implications

Researchers and investors. The 10-year yield should not be treated as a stand-alone monthly signal for the broad NIFTY 50, as the full-sample coefficient is nearly zero, and subperiod estimates vary substantially. A framework which captures inflation, global rates, exchange rates, oil and volatility and earnings conditions is more consistent with the evidence.

Asset managers. It is important to assess rate sensitivity conditionally and at the sector or portfolio level, since different types of firms such as financial, technology, consumer and capital-intensive can react differently to rates.
Decision-makers. The results underscore the need to monitor the financial-conditions channel other than the policy repo rate. Long-term yields, stock prices, exchange rates and liquidity can respond differently to the same message. The study finds no causal impact of a specific policy action, although it does suggest transmission should be assessed between regimes and markets.
Scientists. The most promising next step is higher-frequency identification approaches such as policy-announcement event studies, monetary-policy surprise measures, local projections, structural VARs, Markov-switching models and sector panels.

Conclusion

The research paper re-evaluated the association between long-run interest rates and Indian stocks' return with 175 monthly observations from January 2012 to July 2026 [28]. At the full sample the coefficient on the 10-year yield is approximately −0.019 (p = 0.969) with essentially zero explanatory power. Also the yield-change, ARDL and Granger specifications similarly fail to establish a stable statistically significant full sample rate effect.

Interesting regime analysis yields the most informative evidence. The coefficient from 2012 to 2019 [26], positioned at +0.577, yields no substantial evidence. On the other hand, the coefficient for the period of 2020 to 2022, sitting at −2.600 (p = 0.048), is negative as well as marginally significant. Furthermore, we see that the coefficient for 2020 to July 2026 [28] is pegged at −1.678 (p = 0.077), is weaker. Lastly, the coefficient from 2023 to July 2026 [28], stands at +0.823 (p = 0.605)and is both positive as well as insignificant. The post-2020 interaction (−2.255; p = 0.039) indicates that the slope differs across regimes, but the strong negative association from 2020–2022 was not sustained at the same intensity through 2023–2026 [31].

Hence, the right conclusion is not that interest rates do not impact Indian equities. Neither is it true that higher rates automatically bring about a fall in the NIFTY 50. The correlation between long-run market yields and equity return is weak on average. Furthermore, it is materially dependent on the regime macro-financial. Researches in the area should differentiate between market yields and policy surprises, employ stationary return measures, apply robust inference, test alternative rate transformations and examine parameter stability. We need to design a richer, higher frequency causal to identify the channels allowing for the observed regime dependence. Greater evidence bolsters the analysis not by yielding an expanded headline effect but by clarifying where the relationship is, and is not, present.

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