INTERNATIONAL NETWORK FOR ECONOMIC RESEARCH

INFER WORKING PAPER SERIES

INTERNATIONAL NETWORK FOR ECONOMIC RESEARCH

No. 17 | 2025

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GINN William *(Labcorp, Coburg University of Applied Sciences)*
SAADAOUI Jamel *(University Paris 8)*
SALACHAS Evangelos *(University of the Aegean)*

Website:
https://infer-research.eu/

Contact:
publications@infer.info

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William Ginn[a,b], Jamel Saadaoui[d], Evangelos Salachas[c]

[a] *Labcorp, Sr. Economist / Data Scientist, Artificial Intelligence, Durham, NC, USA.*
[b] *Coburg University of Applied Sciences, Germany.*
[c] *University Paris 8, IEE, LED, Saint-Denis, France.*
[d] *University of the Aegean, Athens, Greece.*

Abstract

This paper examines how U.S. monetary policy shocks influence asset price bubbles under different inflation regimes. Using data from 1998–2023, we show that the transmission of policy is neither constant nor time-invariant. Standard local projection (LP) estimates suggest modest average effects, but including the COVID-19 period reveals that these relationships weaken. Employing time-varying local projections (TVP-LP), we document sharp shifts in transmission during the Global Financial Crisis and the pandemic, motivating a nonlinear approach. Nonlinear VAR-LP estimates uncover clear asymmetries: in high-inflation states, monetary tightening deflates bubbles by raising financing costs and constraining risk-taking; in low-inflation states, the same shocks amplify bubbles by raising expected inflation, narrowing credit spreads, and boosting equity returns. We interpret this inversion as evidence of a *state-contingent speculative signaling channel*, whereby tightening is perceived as a signal of stronger demand or implicit accommodation, encouraging further speculation. This mechanism highlights that safeguarding financial stability requires more than interest rate adjustments alone—it demands explicit attention to the inflationary environment and investor perceptions.

Keywords: Financial stability, Asset Prices, Booms and Busts, Local Projections.
JEL: E21, E32, D12, C32

Email addresses: William.Ginn.OBA@said.oxford.edu (William Ginn), jamelsaadaoui@gmail.com (Jamel Saadaoui), esalachas@aegean.gr (Evangelos Salachas)

Highlights

- Monetary policy shocks affect bubbles differently across inflation regimes.
- Constant-parameter LP estimates obscure shifts in transmission after COVID.
- TVP-LP and non-linear models reveal regime-dependent monetary effects.
- In high-inflation states, tightening reduces bubbles via conventional channels.
- In low-inflation states, tightening amplifies bubbles through a *state-contingent speculative signaling channel*.

1. Introduction

The interaction between inflation, monetary policy, and asset price bubbles has been a persistent topic of debate in macroeconomics and finance. A growing body of literature suggests that low-inflation environments may be particularly conducive to the emergence of speculative market behavior. According to Kaliva and Koskinen (2008), "... inflation is negatively related to the log dividend–price ratio, and that a low inflation rate has a tendency to create optimistic expectations, or outright bubble behavior." In contrast, high inflation acts as a corrective force, reducing the incentive for speculative risk-taking and deflating these bubbles. Kaehler and Weber (2023) further show that financial crises often lead to higher inflation, reinforcing the view that inflation may act as a natural constraint on speculative dynamics.

This asymmetry is especially important when evaluating the effectiveness of monetary policy across states. During the Global Financial Crisis (GFC), inflation spiked as a result of systemic financial stress and commodity volatility. In this high-inflation setting, monetary policy surprises were more likely to suppress speculative market behavior, aligning with the view that inflation serves as a check on asset bubbles. Indeed, Evgenidis and Malliaris (2020) argue that the GFC was not preceded by a bubble, consistent with our finding that rising inflation during that time discouraged excessive risk-taking.

Recent studies, however, have questioned traditional views on monetary policy's role in managing asset price bubbles. The conventional *risk-taking channel of monetary policy* (RTC)—formalized by Borio and Zhu (2012) and further developed in Gambacorta (2009), Bruno and Shin (2015), and Jiménez et al. (2014)—emphasizes how accommodative policy compresses risk premia, lowers credit spreads, and encourages leverage and risk-taking. While this literature has focused primarily on credit markets and risk appetite, rather than explicitly on bubbles, its mechanism provides a natural foundation for understanding how monetary conditions may influence speculative dynamics in asset prices. In this framework, tightening is expected to discipline markets and deflate bubbles.

By contrast, Galí and Gambetti (2015) present theoretical and empirical evidence that monetary tightening may paradoxically inflate the bubble component of asset prices. Adding to this debate, Evgenidis and Malliaris (2020) employ a time-varying VAR and find that contractionary policy reduces both fundamentals and the bubble component, lending support to the conventional RTC view and standing in opposition to the paradox by Galí and Gambetti (2015).

Other recent contributions also highlight the limits of monetary policy in low-inflation regimes. Ahmed et al. (2024) show that monetary transmission to economic activity weakens when interest rates remain low for prolonged periods. Borio and Zhu (2012) emphasize that under conditions of low and stable inflation, financial imbalances often emerge not via consumer prices but through credit booms and asset price surges. They caution that modest tightening in such environments may even be misinterpreted as a signal of macroeconomic strength, thereby encouraging further risk-taking. Consistent with this view, Christiano et al. (2010) document that nearly every major stock market bubble in the past two centuries occurred in periods of low inflation.

Attention to inflation is also state-dependent. Korenok et al. (2023) find that public concern with inflation—as measured by Google search activity—rises sharply when inflation exceeds a 4% threshold. Similarly, Weber et al. (2025) show that higher inflation induces households and firms to become more attentive and informed, suggesting that informational frictions are weaker in high-inflation states. The COVID-19 pandemic provides a natural experiment in this regard: while inflation initially remained subdued, retail investor activity surged, amplifying speculative behavior. Ülkü et al. (2023) and Chiah et al. (2022) document a sharp rise in retail participation during the pandemic, while Ozik et al. (2021) show that fintech platforms enabled retail investors to supply liquidity during stress episodes.

This paper builds on and extends this literature by investigating how monetary policy surprises affect asset price bubbles, with particular attention to the COVID-19 period and the role of inflation states. Our analysis builds on the empirical framework of Gao and Martin (2021), who use sentiment-based methods to measure the speculative component of asset prices. More recently, Blot et al. (2024) examine how monetary policy surprises influence bubble formation, concluding that expansionary surprises tend to reduce speculative behavior. However, their analysis ends before 2020 and therefore misses the critical episode of inflationary dynamics during and after COVID-19, shaped by supply chain disruptions (Benigno et al., 2022; Diaz et al., 2023; Asadollah et al., 2024; Ascari et al., 2024; Ginn, 2024; Ginn and Saadaoui, 2025*a*,*b*).

To fill this gap, we make three contributions. First, we extend the analysis to cover the pandemic and post-pandemic period, a context where the interaction of inflation, speculation, and monetary policy is particularly salient. Second, we employ a time-varying (TV) local projection model to capture shifts in transmission during turbulent episodes such as the GFC and COVID-19. Third, we embed speculative bubbles into both linear and nonlinear VAR-LP systems to jointly model expectations, credit spreads, and earnings. This framework allows us to show that the transmission of monetary policy to bubbles is state-contingent. In high-inflation regimes, monetary tightening curbs speculation, consistent with the conventional RTC. In low-inflation regimes, however, tightening amplifies bubbles—a novel *state-contingent speculative signaling channel*.

Accordingly, our results reconcile seemingly contradictory findings in the literature. The effects documented by Borio and Zhu (2012) and Evgenidis and Malliaris (2020) apply in high-inflation states, while the paradoxical amplification highlighted by Galí and Gambetti (2015) emerges in low-inflation states. Our evidence suggests that the effectiveness of monetary policy in restraining bubbles depends critically on the inflationary environment. This implies that financial stability cannot be managed through interest rates alone. Instead, policymakers must account for broader macro-financial conditions, including inflation expectations, credit dynamics, and investor sentiment, when evaluating the consequences of monetary interventions.

The remainder of this paper is structured as follows. Section 2 describes the data. Section 3 outlines the modeling methodology. Section 4 presents new empirical evidence on monetary policy and speculative sentiment. Section 5 concludes.

2. Data

We use monthly data from September 1997 to December 2023. The dataset includes stock price, dividends data and earnings data,[1] which is adjusted for real terms using the U.S. Consumer Price Index (CPI). Additionally, inflation expectations[2] and bond spread data[3] are considered. Monetary policy surprise data are sourced from Bauer and Swanson (2023), capturing unexpected components of U.S. monetary announcements. Following Blot et al. (2024), we also include Chicago Board Options Exchange (CBOE) Volatility Index (VIX) and 5-year interest rate as a proxy for the time-varying risk premia and the role of unconventional monetary policies, respectively. We also control for the pandemic period and its aftermath by including the Global Supply Chain Pressure Index (GSCPI) from Benigno et al. (2022) as a proxy for supply chain disruptions.[4] Table 1 summarizes the data. Figure 1 plots the economic data.

Following the log-linearization of the price-dividend ratio as proposed by Campbell and Shiller (1988), we express the relationship between prices, dividends, expected returns, and dividend growth as:

$$(p_t - d_t) \approx \kappa + \rho(p_{t+1} - d_{t+1}) - (r_{t+1} - g_{t+1}), \tag{1}$$

where p_t denotes the logarithm of prices, d_t the logarithm of dividends, r_{t+1} the return, g_{t+1} the dividend growth rate, κ is a constant, and ρ is the persistence parameter ($0 < \rho < 1$). Equation

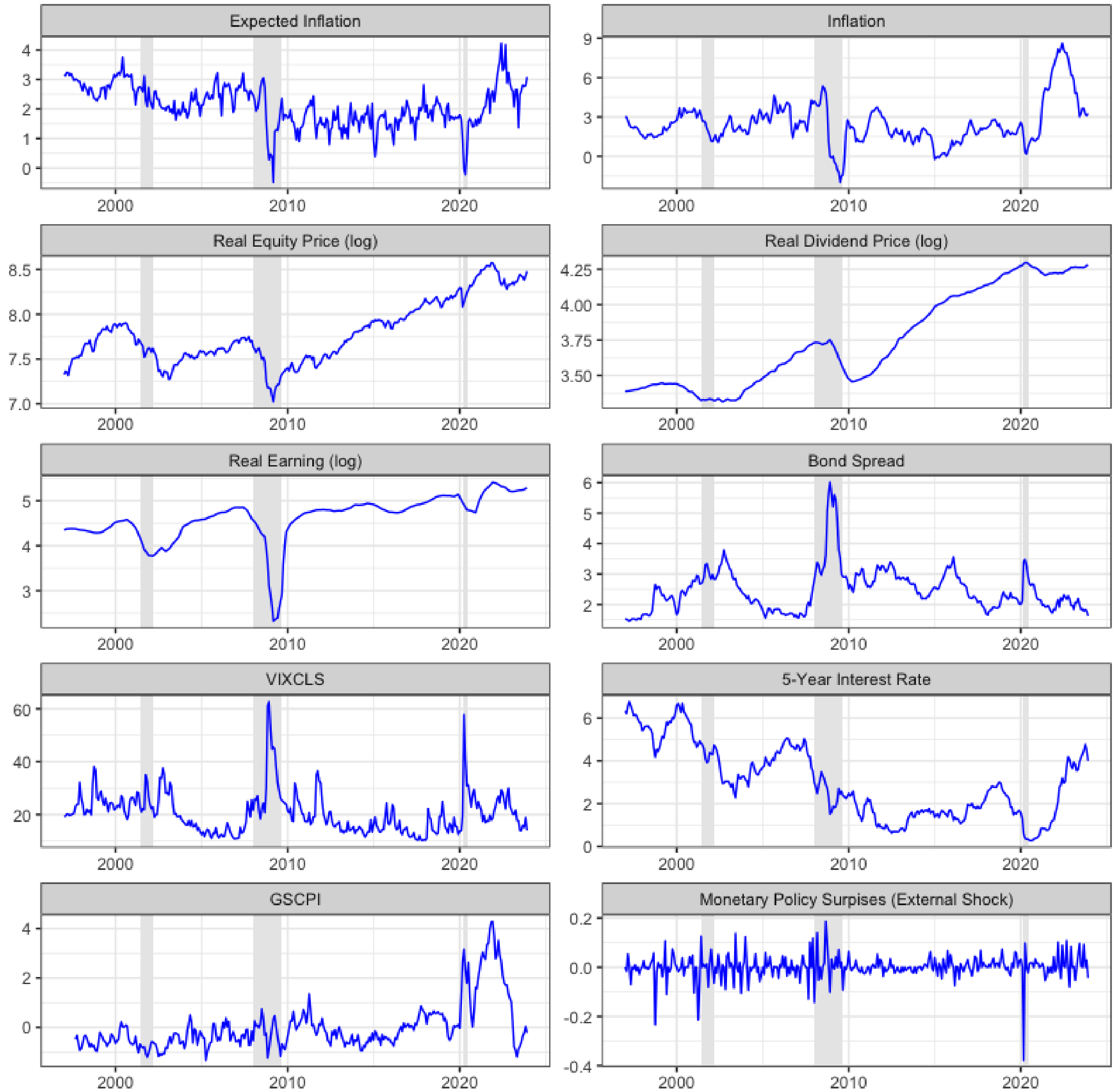
[1]See Shiller (2015), the data is available via `http://www.econ.yale.edu/~shiller/data.htm`. Real earnings is an indicator of corporate profitability, adjusted for inflation. In the context of monetary policy, earnings can be sensitive to unexpected policy shifts, as these policies influence not only the cost of capital but also aggregate demand. Monetary policy surprises can affect corporate earnings by influencing borrowing costs, investment decisions, and consumer spending.

[2]Inflation expectations have become particularly relevant in the aftermath of the COVID-19 pandemic, as economic uncertainty has been exacerbated by supply chain disruptions. By focusing on inflation expectations, this study aims to capture how these forward-looking variables respond to unexpected monetary policy changes.

[3]The BAA interest rate spread, defined as the difference between the yield on BAA-rated corporate bonds and the yield on risk-free Treasury securities, serves as a measure of credit market conditions and risk perceptions. A widening spread signals increased risk aversion among investors, whereas a narrowing spread indicates greater investor confidence and lower perceived credit risk.

[4]GSCPI is calculated using a combination of various indicators that reflect the state of global supply chain conditions. It incorporates data on shipping costs, freight rates, inventory levels and delays (Benigno et al., 2022). High (low) values of the GSCPI indicate significant disruptions (stable conditions) in global supply chains.

Figure 1: Economic Data



Shaded areas indicate NBER recession dates.

(1) captures the intertemporal trade-off between expected returns and dividend growth. To link dividend growth expectations with risk-free and excess returns, we further adopt the following specification from Gao and Martin (2021):

$$E_t(g_{t+1}) = r_{f,t+1} + E_t(r_{t+1} - r_{f,t+1}) - E_t(r_{t+1} - g_{t+1}), \tag{2}$$

Table 1: Data Description

Item	Symbol	Source	Description
Aggregate CPI	$\ln P_t$	FRED	CPI: All items (mnemonic CPIAUCSL)
Inflation Expectations	$\Delta \ln P_t^E$	FRED	1-Year Inflation Expectations (mnemonic EXPINF1YR)
Bond spread	S_t	FRED	Measure of credit market conditions (mnemonic BAA10YM)
Earnings	E_t	Shiller	Real earnings (see [a])
Equity price	P_t^S	Shiller	Real equity price (see [a])
Dividend price	D_t	Shiller	Real dividend price (see [a])
VIX	VIX_t	FRED	CBOE Volatility Index (mnemonic VIXCLS)
5-Year Rate	R_t^{5-year}	FRED	Yield on 5-Year U.S. Treasury Securities (mnemonic GS5)
GSCPI	$GSCPI_t$	NY Fed	Global Supply Chain Pressure Index (see Benigno et al., 2022)

[a]: http://www.econ.yale.edu/~shiller/data.htm.

where $r_{f,t+1}$ is the risk-free rate and $E_t(\cdot)$ denotes expectations conditional on information at time t. Equation (2) highlights how expected dividend growth is related to the risk-free component, expected equity premium, and the gap between returns and dividend growth. Consistent with Blot et al. (2024), we exclude the contribution of the risk-free rate and equity risk premium components in this sentiment indicator, as we aim to investigate the impact of monetary policy beyond its effect on the risk-free rate and the risk premium.

These two equations provide the theoretical foundation for our subsequent empirical analysis of price-dividend dynamics. The dividend yield "bubble" component is derived following Gao and Martin (2021). We follow Gao and Martin (2021) and decompose the log price-dividend ratio using a Campbell-Shiller log-linear approximation. The approach isolates the speculative component of equity prices beyond expected dividend growth and returns. Specifically, expected return-growth differentials are estimated via:

$$\mathbb{E}_t(g_{t+1} - r_{t+1}) = c + \frac{1 - \rho' \phi}{1 - \rho'} y_t, \tag{3}$$

where $y_t = \log\left(1 + \frac{D_t}{P_t^S}\right)$, $\rho' = e^{-\bar{y}}$, and ϕ is estimated from an AR(1) process on y_t. This forward-looking measure captures sentiment-driven deviations in stock prices. We plot the dividend yield component by Gao and Martin (2021) based on the Campbell–Shiller approximation in Figure 2.

Figure 2: Dividend Yield Component (GM21)

3. Methodology

We employ the LP framework, developed by Jordà (2005), to estimate the dynamic responses of dividend yield to a monetary policy shock. As a starting point, we extend the baseline LP model of Blot et al. (2024) in two dimensions: first, by lengthening the sample to include the pandemic and post-pandemic period, and second, by incorporating a time-varying (TV) LP specification. More broadly, we adopt multiple LP-based frameworks to ensure robustness and to capture different dimensions of monetary policy transmission. The constant-parameter LP provides a benchmark consistent with Blot et al. (2024), but may mask important temporal instabilities. The TVP-LP addresses this by allowing for potential changes in policy effects over time. The VAR-LP combines the flexibility of LP with the structural identification advantages of VARs, enabling joint dynamics between macro-financial variables. Finally, the non-linear VAR-LP captures state dependence, recognizing that the effects of monetary shocks can vary systematically across inflationary states. Each model is discussed in turn.

3.1. Time-Series Local Projections

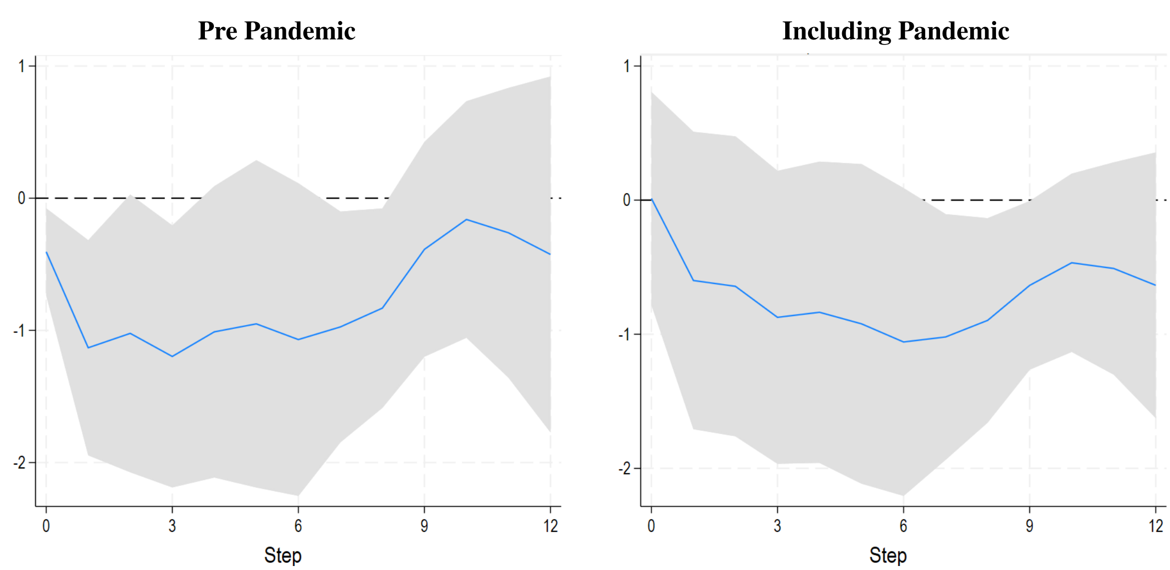
We estimate the LP model as follows:

$$y_{t+h} = \alpha_h + \beta_h MPS_t + \sum_{j=1}^{p} \alpha_h^{'} \mathbf{x}_{t-j} + v_{t+h}, h = 0, 1, ..., T \quad (4)$$

where y_t is the dividend yield bubble component (Gao and Martin, 2021). α is an intercept term. MPS_t is a one-standard deviation shock to a transitory increase in monetary tightening (Bauer and Swanson, 2023). The vector $\mathbf{x}_t = \{VIX_t, \Delta R_t^{5-year}, GSCPI_t\}$ includes the VIX index, change in the 5-year interest rates and GSCPI. These variables are used to control for macroeconomic fundamentals. The inclusion of VIX_t and ΔR_t^{5-year} follows directly from Blot et al. (2024), where we additionally include GSCPI, since we extend the dataset[5] to capture the COVID-19 pandemic and its aftermath, which fueled supply chain disruptions. v_t relates to the error term. The coefficient β_h in Equation (4) traces out the effect of the shock at time t on the dividend yield at time $t + h$.

Figure 3 plots the (time-invariant) LP IRFs, based on two sample periods: pre-pandemic sample period (left-pane, covering 268 months from 1997:09 to 2019:12) and the full sample (right-pane, covering 316 months from 1997:09 to 2023:12). The IRFs during the pre-pandemic period indicate a monetary surprise reduces the bubble component at the onset of the shock and is statistically significant for multiple periods. These findings align with Blot et al. (2024) relating to the sample period prior to the COVID-19 pandemic. However, when extending the time domain to include the pandemic period and its aftermath, the statistical significance is largely diminished.[6]

Figure 3: Response of Monetary Shock to Price-Dividend



[5]The time sample in the paper by Blot et al. (2024) ends in December 2019, hence prior to the global pandemic.

[6]While the sign of the impulse response for the full sample (including the pandemic period) is similar, much of the statistical significance is diminished (periods 8 and 9 are statistically significant).

3.2. Time-varying Local Projections

The constant-parameter LP model provides initial insight into the relationship between monetary policy surprises and the bubble component of asset prices. However, once the sample is extended to include the COVID-19 period and its aftermath, the estimated impulse responses become statistically insignificant, suggesting that the relationship may no longer hold uniformly across time. In particular, Figure 3 shows that while the pre-pandemic IRFs are negative and significant for several horizons, the inclusion of the pandemic period weakens the results considerably. These differences in the IRFs across sample periods warrant further analysis.

Accordingly, the instability in the response function raises an important question: has the effect of monetary policy shocks changed over time? A model with constant coefficients is not necessarily appropriate to accommodate such structural variation. To address this, we turn to the TVP-LP framework. Unlike standard LPs or VARs, the TVP-LP allows for time-varying changes in the impulse response coefficients over time without imposing a specific form of state switching. This approach is particularly well-suited to capturing evolving policy transmission mechanisms—especially across turbulent periods such as the GFC and COVID-19.

We extend the analysis by estimating a TVP-LP method, introduced by Inoue et al. (2024), to investigate the impact of a monetary policy shock on the bubble component. Olea et al. (2024) recently provided formal proof that conventional LP confidence intervals for impulse responses remain robust despite misspecification. In the context of our study, the TVP-LP method facilitates time-varying dynamic insights into the causal impact of monetary policy shocks. The TVP-LP has been applied to the recent econometric applications. Saadaoui et al. (2025) analyze the responsiveness of critical minerals prices to geopolitical risk. Ginn and Saadaoui (2025*c*) estimate the impact of geopolitical risks on monetary policy. This approach is particularly suited to our research.

Unlike other VAR/LP-based models that incorporate time variation (like TVP-(B)VAR or state-dependent LPs), the TVP-LP does not require a parametric specification of the instability process.

The TVP-LP generates a path estimator for each time horizon.[7] This allows for flexibility in time variation without imposing uniformity in the response across horizons.

We extend the LP model in Equation (4) to incorporate the time-varying effects as follows:

$$y_{t+h} = \alpha_{t+h} + \beta_{h,t+h} MPS_t + \sum_{j=1}^{12} \alpha'_{j,t+h} \mathbf{x}_{t-j} + v_{t+h}, h = 0, 1, ..., T \quad (5)$$

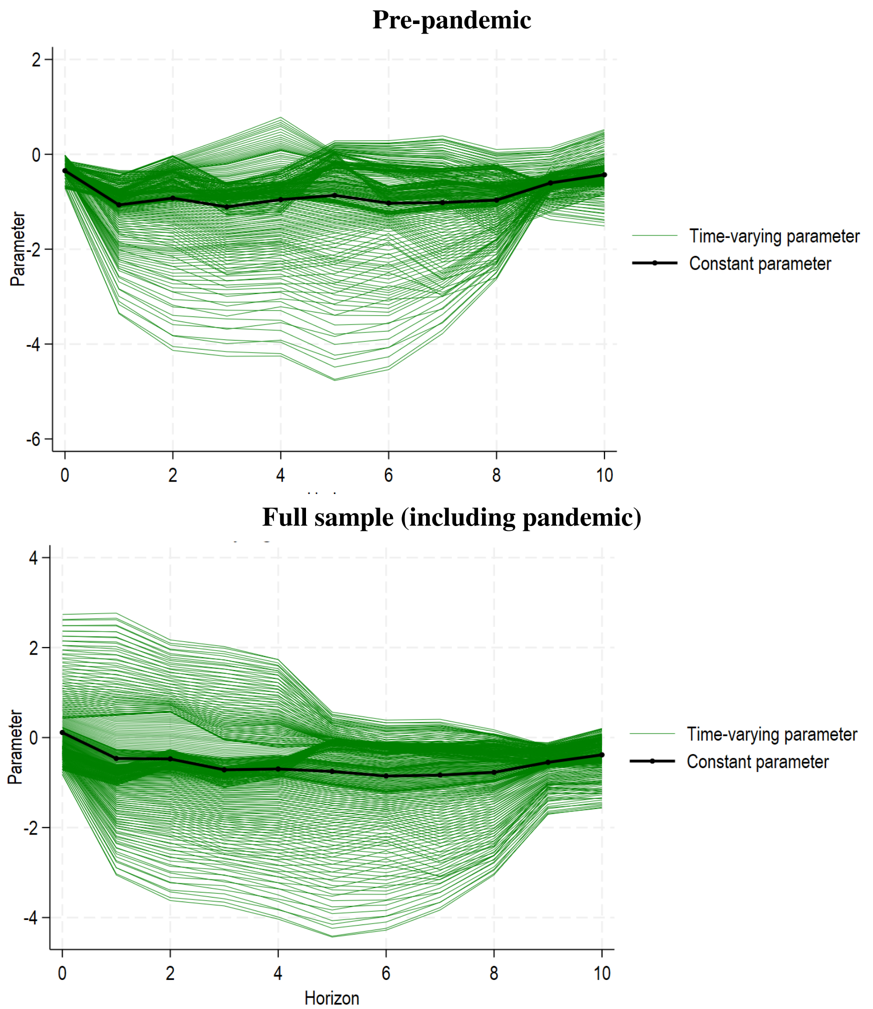
Figure 4 plots the TVP-LP IRFs of price-dividend to a monetary policy shock based on two sample periods: pre-pandemic sample period (bottom-pane, covering 1997:09 to 2019:12) and the full sample (top-pane, covering 1997:09 to 2023:12). The IRFs during the pre-pandemic period indicate the a monetary surprise reduces the bubble component at the onset of the shock and is statistically significant for multiple periods. While the sign of the impulse response for the full sample (including the pandemic period) is similar, much of the statistical significance is diminished (periods 8 and 9 are statistically significant).

Figure 5 plots the TVP-LP IRFs, illustrating that the effect of monetary policy on bubbles is not stable over time. During the GFC, when inflation was elevated, monetary tightening clearly reduced speculative pressures. By contrast, in the years preceding and during the COVID-19 pandemic—characterized by subdued inflation and unprecedented liquidity—the same shocks were associated with an increase in bubbles. This inversion underscores the importance of inflationary context for the transmission of monetary policy. In low-inflation environments, tightening may signal continued accommodation or fail to discipline risk-taking, thereby fueling rather than restraining speculative behavior.

Figures 9 and 10 plots the IRFs from a TVP-LP model, estimating the impact of a monetary policy shock on price-dividend yield across different horizons (i.e., 1 to 12 months). The solid line represents the time-varying response of dividend yield to a monetary policy shock, while the dashed lines indicate the 95% confidence interval.

[7]For instance, at horizon 1, our sample spans T = 312 months (from January 1998 to December 2023) and the horizon h = 1 month, resulting in 311 impulse response functions at horizon 1.

Figure 4: Bubble Price Reaction to Monetary Policy Shock (TVP-LP)

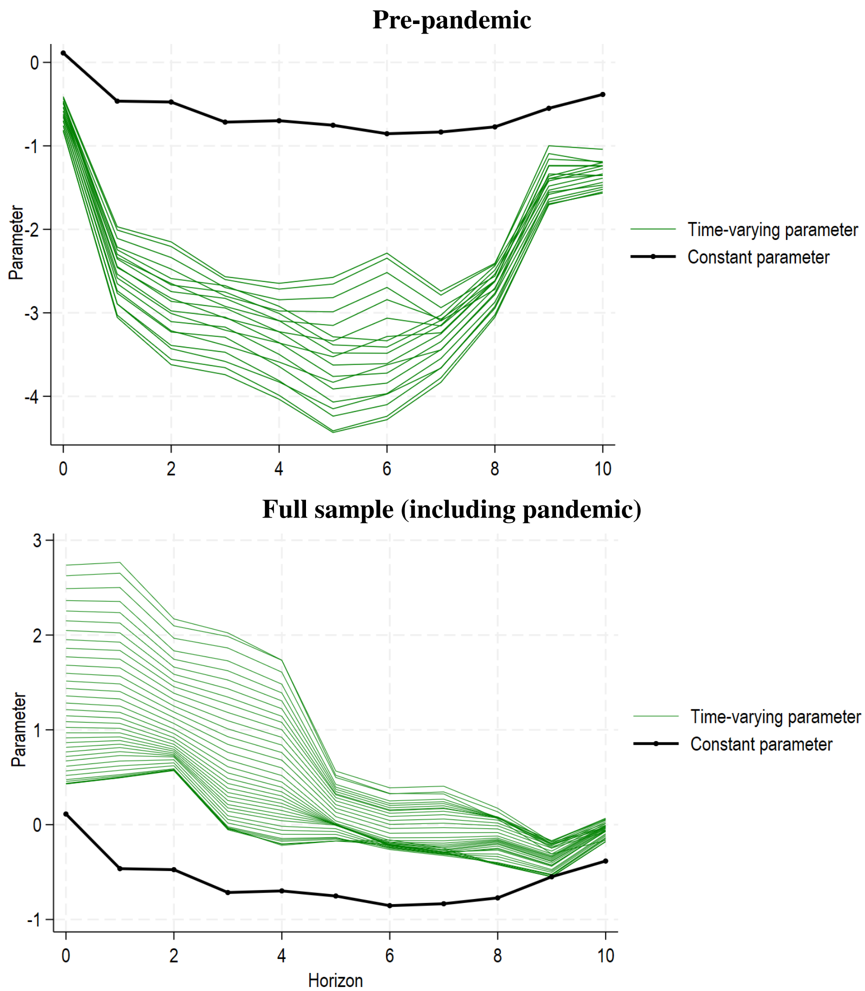


Note: The time-varying parameters (green lines) are separated based on the pre-pandemic period; GFC period; and pandemic period. The black line indicates the constant parameter LP model.

The response of the dividend yield appears to fluctuate around zero prior to GFC period, yet becomes significantly negative for a number of periods thereafter. Interestingly, the response turns positive at the beginning of 2020, coinciding with the onset of the pandemic period.

Our findings resonate with both theoretical and policy perspectives in the literature. Mishkin (2010) emphasizes that not all bubbles are alike: credit-driven bubbles, such as those preceding the GFC, are deeply intertwined with lending and financial stability, making them more responsive to monetary tightening. By contrast, equity-only bubbles like the late-1990s tech boom have weaker credit linkages and limited macroeconomic spillovers. Consistent with this distinction, we find that

Figure 5: Bubble Price Reaction to Monetary Policy Shock (TVP-LP)



the bubble component responds strongly to monetary surprises during the GFC but more weakly around the tech bubble. Likewise, Christiano et al. (2010) show in a structural model that when inflation is low, a central bank that narrowly targets inflation may destabilize asset markets and the macroeconomy. Our empirical results confirm this prediction: during the onset of the low-inflation COVID period, monetary tightening amplified rather than restrained bubbles. Together, these insights suggest that the interaction between monetary policy and bubbles depends critically on both the nature of the bubble and the inflationary environment, highlighting the limits of rules that focus narrowly on inflation stabilization.

A comparison between the time series LP and TVP-LP models reveals significant differences in the responses. In the time series LP model, the IRFs (Figure 3) assume uniform effects of monetary policy surprise shocks, which can "mask" important differences with respect to state dependence. By contrast, the TVP-LP estimates in Figure 5 uncover substantial heterogeneity across periods.

During the Global Financial Crisis, monetary tightening shocks were followed by a pronounced and persistent decline in the bubble component, consistent with the disciplining effect of higher rates in an environment of elevated inflation and financial stress. However, at the onset of the COVID-19 pandemic, the response pattern shifted: monetary surprises were associated with a positive response of the bubble component, suggesting that in a low-inflation, highly accommodative environment, conventional tightening shocks may have fueled rather than contained speculative dynamics.

Taken together, the TVP-LP results indicate that the transmission of monetary policy to bubbles is likely state-dependent, and that constant-parameter models understate this instability. These findings motivate a more explicit modeling of state dependence, which we address in the VAR-LP framework below.

3.3. Linear VAR-LP

A common representation of a VAR is as follows:[8]

$$\boldsymbol{y_t} = \boldsymbol{\alpha} + B(L)\boldsymbol{y_t} + \boldsymbol{\epsilon_t} \tag{6}$$

where $\boldsymbol{\alpha}$ is a vector of intercept term; B(L) is an autoregressive lag polynomial; and $\boldsymbol{\epsilon}_t$ is a vector of white noise error terms. Multiplying equation (6) by $\boldsymbol{A}_0$ yields a VAR represented in structural form:

$$\boldsymbol{A}_0\boldsymbol{y_t} = \boldsymbol{\zeta} + \boldsymbol{A_0}B(L)\boldsymbol{y_t} + \boldsymbol{u_t} \tag{7}$$

for $\boldsymbol{\zeta} = \boldsymbol{A}_0\boldsymbol{\alpha}$ and $\boldsymbol{u}_t = \boldsymbol{A}_0\boldsymbol{\epsilon}_t$. Matrix $\boldsymbol{A}_0$ captures the contemporaneous relationship between the variables. To identify $\boldsymbol{A}_0$, a standard Cholseky decomposition is employed imposing a lower triangular matrix.

The impulse responses corresponding to the reduced form and structural form shocks are,

[8] Note that we abstract from deterministic terms with the exception of the intercept term for expositionary purposes.

respectively, generalized as follows (Kilian and Kim, 2011):

$$\Phi_h^{VAR} = \sum_{l=1}^{h} \Phi_{h-l}^{VAR} B_l, h \in \{1, 2, ..., H\} \tag{8}$$

and

$$\Theta_h^{VAR} = \Phi_h^{VAR} A_0^{-1}, h \in \{1, 2, ..., H\} \tag{9}$$

An alternative approach to estimate the reduced form impulse responses is to fit a linear projection as follows:

$$\boldsymbol{y}_{t+h} = \boldsymbol{\alpha}_h + \mathbf{B}_1^h \boldsymbol{y}_t + \mathbf{B}_2^h \boldsymbol{y}_{t-1} + \ldots + \mathbf{B}_p^h \boldsymbol{y}_t + \phi^h \mathbf{x}_{t-1} + \boldsymbol{\epsilon}_{t+h} \tag{10}$$

where $\boldsymbol{y}_t^T$ is the transpose of a column vector α_h is an intercept term. $\mathbf{B}_i^h$ are autoregressive coefficients for future horizon h = 1,...,H. ϵ_{t+h} is a disturbance. As there is serial correlation present in the error terms, the Newey-West correction is used for standard errors. The corresponding structural impulse responses are:

$$\Theta_h^{LP} = \Phi_h^{LP} A_0^{-1}, h \in \{1, 2, ..., H\} \tag{11}$$

where A_0^{-1} is recovered from the estimated model in equation (7) (see Kilian and Kim, 2011).

The vector of macroeconomic variables includes the following:

$$\boldsymbol{y}_t^T = [\Delta \ln P_t^E, S_t, \Delta \ln E_t, B_t] \tag{12}$$

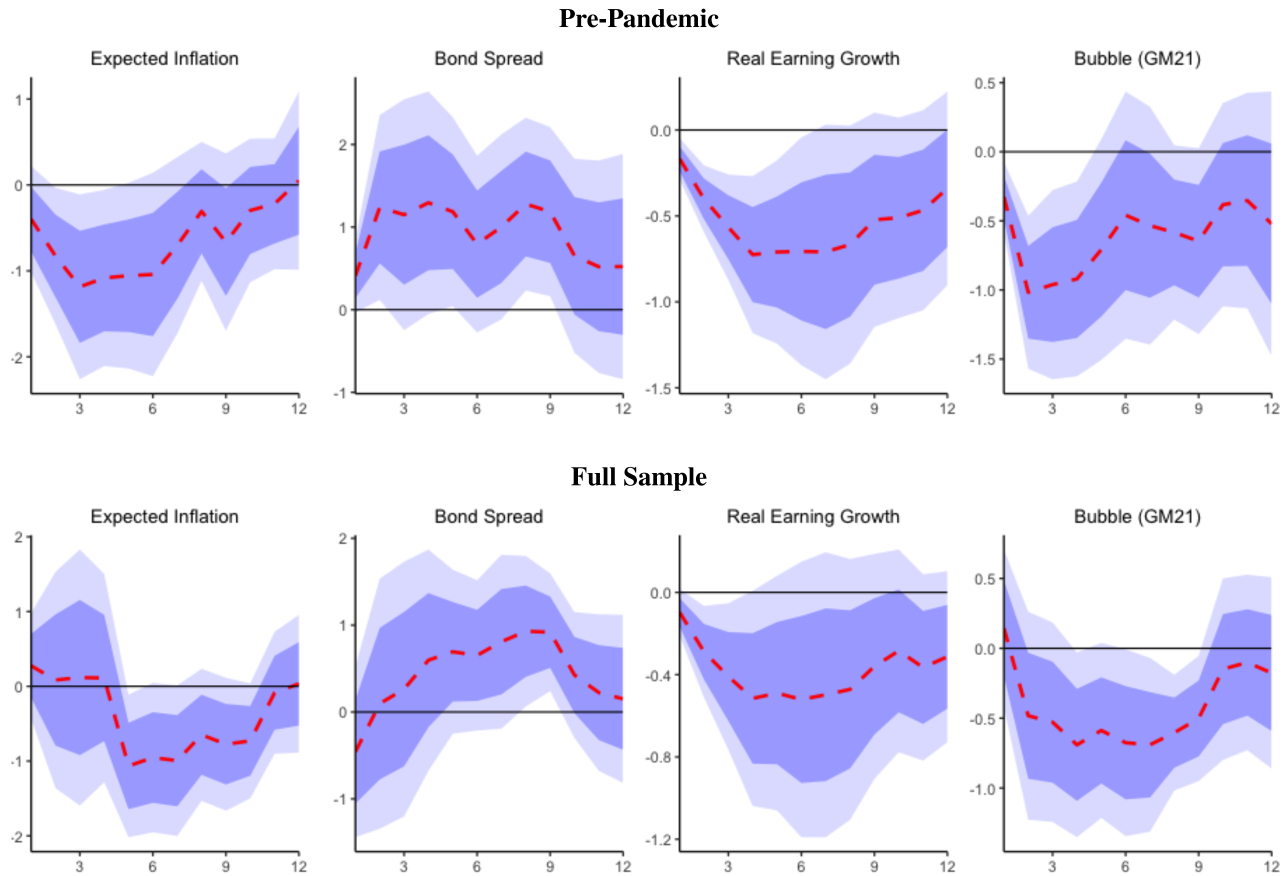
where $\boldsymbol{y}_t^T$ includes inflation expectations, bond spread, real earnings growth and the bubble component. The choice of variables reflects the macroeconomic and financial channels through which monetary surprises are likely to operate. Inflation expectations are central to the transmission of monetary policy and provide a forward-looking perspective on economic conditions. Given the crucial role of forward-looking behavior in financial economics, expected inflation has broad implications for understanding and modeling asset price bubbles, as expectations about future inflation influence investor behavior, interest rates, and ultimately the valuation of financial assets. As Coibion and Gorodnichenko (2012) show, there is gradual adjustment in expectations (due to rigidities). We posit that can lead to mispricing of assets for extended periods if agents do not revise their beliefs about earnings or interest rates, which may then continue to drive asset prices up, feeding a bubble. By including both inflation expectations and the BAA interest rate spread, the model captures the interplay between monetary policy and credit market conditions. The bond spread reflects financial conditions and perceived credit risk (Gilchrist and Zakrajšek, 2012), while real earnings growth is a proxy for the real side of financial markets. The bubble component captures the deviation of asset prices from fundamentals due to speculative behavior and market sentiment (Campbell and Shiller, 1988; Gao and Martin, 2021; Blot et al., 2024).

The IRFs for the Linear VAR-LP model to a monetary policy surprise are shown in Figure 6, decomposed for the pre-pandemic period (top panel) and the full sample (bottom panel). Across specifications, expected inflation declines on impact, indicating that monetary surprises have a contractionary effect in the short run. Bond spreads widen, consistent with tighter financial conditions, while real earnings growth falls, reflecting a slowdown in fundamentals.

The most striking difference arises in the response of the bubble component. In the pre-pandemic period (top panel), monetary policy surprises lead to a clear and statistically significant decline in the bubble measure, suggesting that tighter policy reliably deflated speculative dynamics when inflation was relatively higher. In contrast, once the pandemic period is included (bottom panel), the response of the bubble component becomes muted and statistically insignificant. This divergence highlights a structural shift: the sensitivity of bubbles to monetary policy shocks diminished in the post-pandemic environment, where inflation was initially subdued and policy accommodation unprecedented.

Taken together, the comparison suggests that the disciplining effect of monetary policy on bubbles held before the pandemic but weakened substantially thereafter. This change provides further evidence of state dependence and reinforces the need to move beyond linear models toward frameworks that explicitly account for state heterogeneity.

Figure 6: Linear Model IRFs



Note: IRFs from a one standard deviation monetary policy shock.

3.4. Non-Linear VAR-LP Model

Using a non-linear VAR-LP, we specify the prediction of y_{t+h} to differ according to an high inflationary ("HI") and a low inflationary ("LI") state of the economy (Ramey and Zubairy, 2018). That is, y_t is estimated that includes a smooth transition function $F(\zeta_t)$ that represents the state of the economy:

$$F(\zeta_{i,t}) = \frac{\exp(-\gamma z_{i,t})}{1 + \exp\left(\gamma z_{i,t}\right)}, \quad \gamma > 0, \quad |\zeta_t| < \infty \tag{13}$$

where ζ_t is a standardized transition variable and γ controls the degree of smoothness of the transition between states. The model has the following functional form:

$$\mathbf{y}_{t+h} = F(\zeta_{t-1})(\alpha^h_{HI}+\phi^h_{HI}\mathbf{x}_t+\sum_{i=1}^{p}\mathbf{B}^{h+1}_{HI}\mathbf{y}_{t-i})+(1-F(\zeta_{t-1}))(\alpha^h_{LI}+\phi^h_{NE}\mathbf{x}_{t-1}+\sum_{i=1}^{p}\mathbf{B}^{h+1}_{LI}\mathbf{y}_{t-i})+\boldsymbol{\epsilon}_{t+h} \quad (14)$$

The coefficient $\mathbf{B}^{h+1}_{HI}$ ($\mathbf{B}^{h+1}_{LI}$) corresponds with matrices of parameters during a high (low) inflationary state. Following Auerbach and Gorodnichenko (2012), the transition function is dated $t-1$ to avoid contemporaneous feedback from policy actions with regard to the state of the economy (i.e., $F(\zeta_{t-1})$). Furthermore, we follow Dufrénot et al. (2024) by estimating γ via Logistic Smooth Transition AutoRegressive (LSTAR) model.[9]

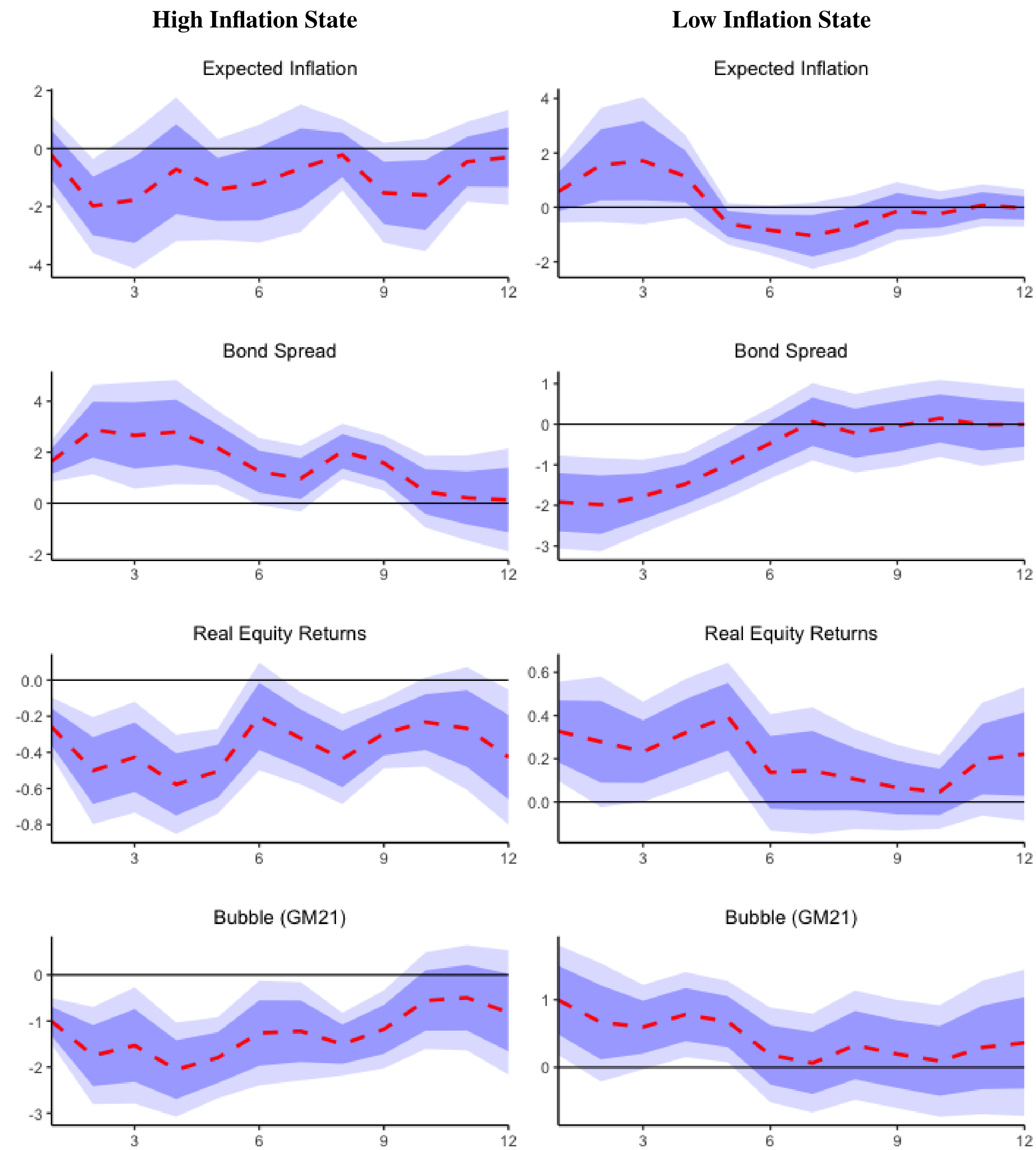
Figure 7 plots the IRFs from a one standard deviation monetary policy shock, estimated separately for high and low inflation states over the period 1998–2023. The transition function is presented in Figure 8, which plots the evolution of the transition variable $F(z_t)$ along with NBER recessions. By construction, a high value of the transition variable in Figure 8 corresponds to periods of low inflation.

The non-linear VAR-LP estimates in Figure 7 highlight pronounced asymmetries in monetary transmission across inflation regimes. In high-inflation states (left column), a monetary tightening shock reduces expected inflation, widens bond spreads, lowers equity returns, and deflates the bubble component. This response reflects the conventional credit and expectations channels of monetary policy: higher rates tighten financial conditions and discipline speculative behavior, thereby restraining asset price excesses.

By contrast, in low-inflation states (right column), the responses invert. Monetary tightening raises expected inflation, compresses bond spreads, lifts equity returns, and amplifies the bubble component. Rather than dampening speculation, policy surprises appear to stimulate it. We interpret this paradoxical response as evidence of a *state-contingent speculative signaling channel*. In environments where inflation is subdued, unexpected tightening may be read by investors as a

[9] Accordingly, we estimate $\gamma = 1.91$ via LSTAR model (see Appendix, Section 5.2).

Figure 7: Non-Linear Baseline Model IRFs (Full Sample)

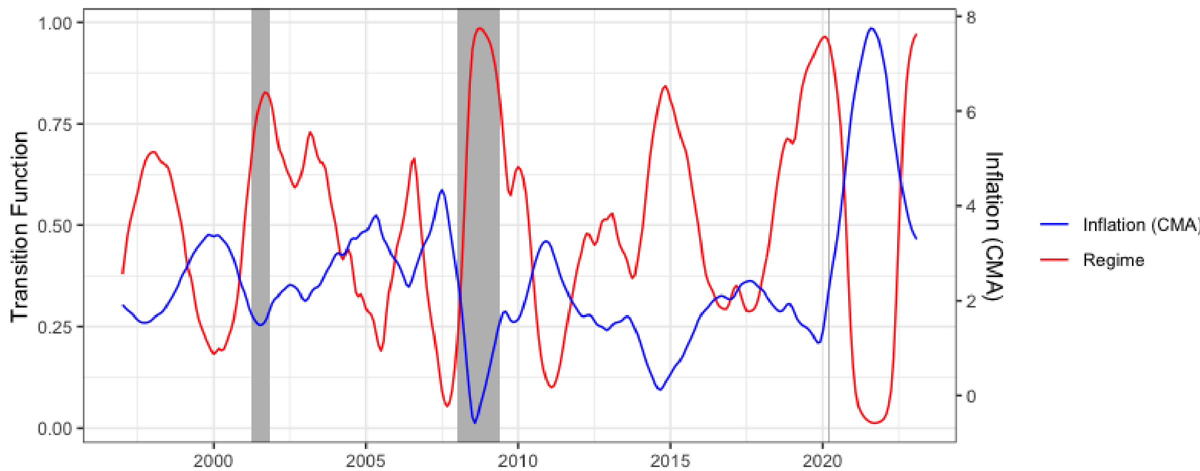


Note: IRFs from a one standard deviation monetary policy shock.

signal of stronger underlying demand or anticipated future accommodation, encouraging greater risk-taking and speculative activity.

This mechanism extends the established *risk-taking channel of monetary policy* (Borio and Zhu, 2012), which traditionally links speculation to persistently loose financial conditions. Our results

Figure 8: Non-Linear Baseline Model Transition (Full Sample)



The figures shows the weighted state (i.e., F(z)) based on the twelve month centered moving average of inflation. Shaded areas indicate NBER recession dates.

show that the risk-taking mechanism itself is regime-dependent: in high-inflation states, tightening curbs bubbles through conventional financial and expectations channels, whereas in low-inflation states, the same shock can paradoxically fuel bubbles through its signaling effects. This asymmetric transmission resonates with Christiano et al. (2010), who demonstrate that when inflation is low, a central bank narrowly focused on price stability may inadvertently destabilize asset markets.

For robustness, the IRFs for the pre-pandemic period is provided in the Appendix (see Figure 11). Accordingly, the empirical evidence suggests that a monetary policy shock does not reduce an asset price bubble during a low inflationary period. Overall, these findings highlight the state-dependent nature of monetary policy transmission. Monetary surprises have stronger effects on bond spreads and inflation expectations during expansionary periods, while asset price responses exhibit greater volatility in non-expansionary states. These results suggest that policymakers should account for economic conditions when designing monetary interventions, as their effectiveness varies across different macroeconomic states.

4. Conclusion

This paper examines how monetary policy surprises affect the speculative component of equity prices across inflation states using an LP framework. Our results show that the transmission of monetary shocks to asset price bubbles is strongly state-dependent. In high-inflation environments, monetary tightening reliably deflates speculative excesses by raising financing costs and constraining risk-taking—consistent with the conventional credit and expectations channels.

By contrast, during low-inflation periods such as the COVID-19 era, monetary tightening produced the opposite response: expected inflation increased, bond spreads narrowed, equity returns rose, and bubbles expanded. A Time-Varying LP (TVP-LP) approach reveals that these shifts are not uniform across time but evolve markedly across episodes such as the Global Financial Crisis and the pandemic, underscoring the importance of accounting for structural change in the transmission mechanism.

Our results reconcile seemingly contradictory findings in the literature. In high-inflation states, we find that monetary tightening lowers expected inflation, widens credit spreads, reduces equity returns, and deflates the bubble component. This outcome is consistent with the risk-taking channel of Borio and Zhu (2012) and with the evidence of Evgenidis and Malliaris (2020), who show that contractionary policy reduces both fundamentals and speculative components. By contrast, in low-inflation states, the transmission inverts: monetary tightening raises expected inflation, narrows spreads, boosts equities, and amplifies bubbles. This dynamic aligns with the paradoxical amplification effect highlighted by Galí and Gambetti (2015), but only under conditions of subdued inflation. Taken together, our evidence points to a novel *state-dependent speculation channel*—an extension of the risk-taking channel—whereby the impact of monetary policy on bubbles depends critically on the prevailing inflation regime. Taken together, our findings imply that financial stability cannot be safeguarded through interest rate policy alone. Instead, the effectiveness of monetary policy depends critically on the inflationary context, liquidity conditions, and investor perceptions. Our identification of a *state-contingent speculative signaling channel* extends the established risk-taking view by showing that monetary policy can amplify speculation even when rates rise, provided inflation is low. This regime-dependent inversion underscores the need for central banks to consider the informational and behavioral dimensions of policy actions, not only their mechanical effects on financing costs.

From a policy perspective, these results raise important concerns. If monetary tightening can, under certain conditions, increase speculative behavior rather than curb it, this challenges the conventional toolkit used by central banks to manage financial stability. Policymakers must consider the broader macro-financial environment—including inflation expectations, liquidity conditions, and behavioral responses—when interpreting the likely effects of their actions. Future research should investigate the mechanisms behind these state-dependent responses, such as the role of forward guidance, investor sentiment, or the credibility of policy signals. Understanding the nonlinear and time-varying nature of monetary transmission is increasingly vital for formulating effective policy in a post-pandemic world characterized by uncertainty, structural change, and evolving market dynamics.

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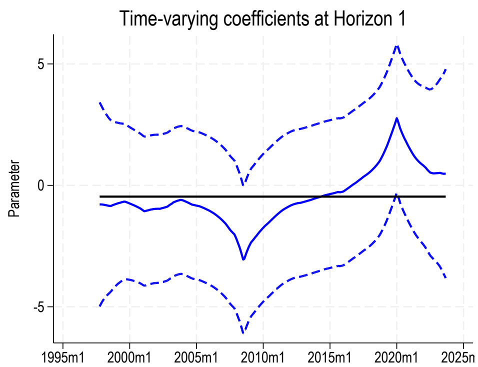
5. Appendix

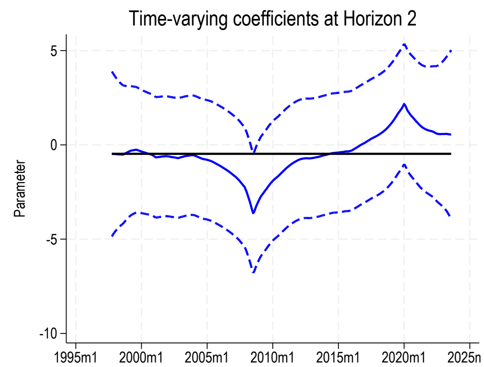
5.1. Impulse Responses to Monetary Shock (TVP-LP Model)

Figure 9: Bubble Price Reaction to Monetary Policy Shock at Different Horizons

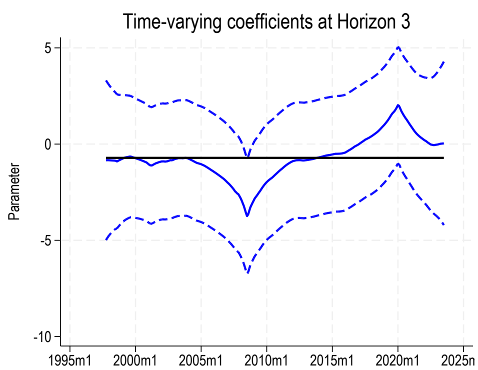
Horizon 1



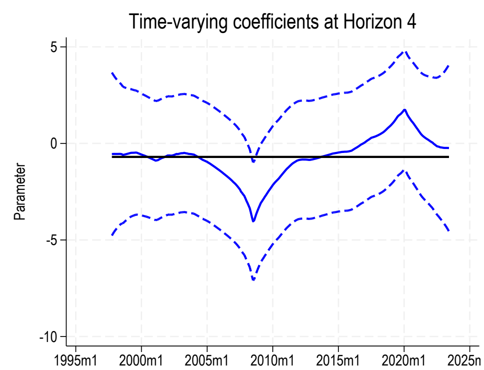
Horizon 2



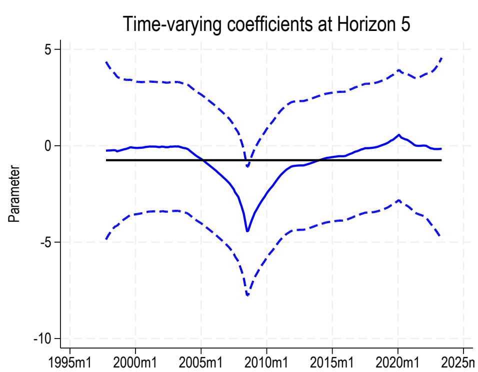
Horizon 3



Horizon 4



Horizon 5



Horizon 6

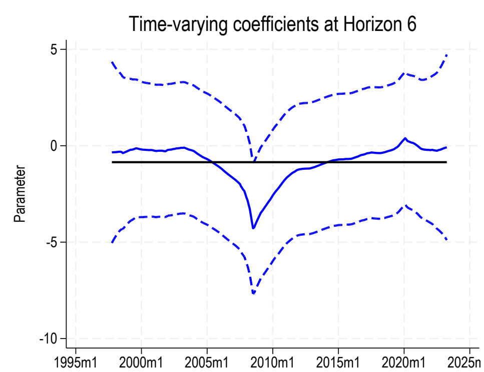
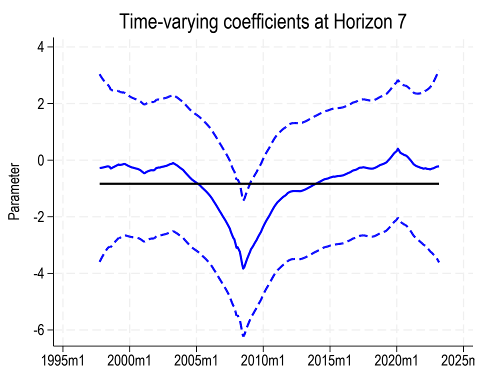
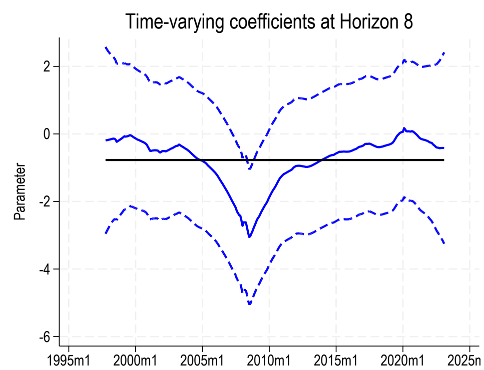


Figure 10: Bubble Price Reaction to Monetary Policy Shock at Different Horizons, continued

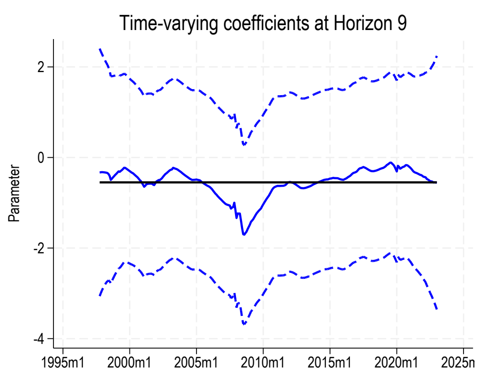
Horizon 7



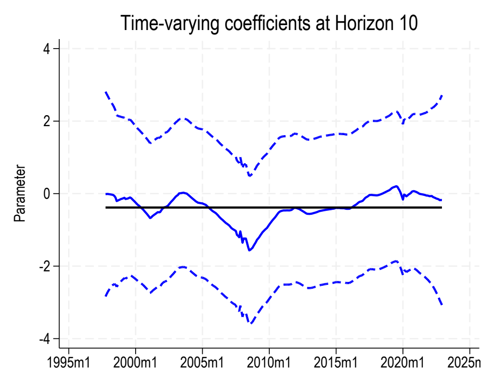
Horizon 8



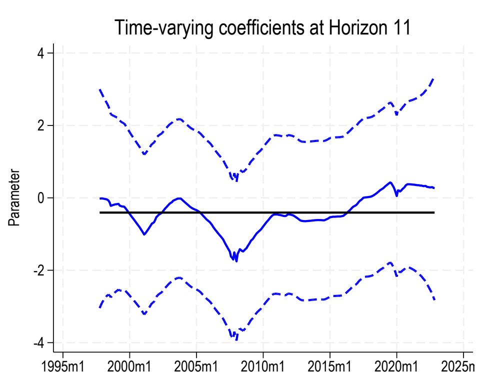
Horizon 9



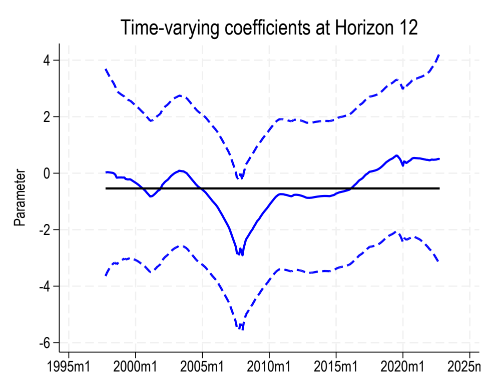
Horizon 10



Horizon 11



Horizon 12



5.2. Logistic STAR Model

The Logistic Smooth Transition Autoregressive (LSTAR) model is used to estimate the smoothness parameter γ, which governs the transition between low and high states in the local projections model (Equation 14). The estimated LSTAR dynamics are incorporated into the local projections (LP) specification as follows:

$$\begin{aligned} y_t = &\Big(\beta_{0,H} + \beta_{1,H}, \tau_t + \phi_{H,1} y_{t-1} + \phi_{H,2} y_{t-2} + \phi_{H,3} y_{t-3}\Big) F(\zeta_{t-1}; \gamma, \theta) + \\ &\Big(\beta_{0,L} + \beta_{1,L}, \tau_t + \phi_{L,1} y_{t-1} + \phi_{L,2} y_{t-2} + \phi_{L,3} y_{t-3}\Big) \big[1 - F(\zeta_{t-1}; \gamma, \theta)\big] + \varepsilon_t, \end{aligned} \tag{15}$$

where $F(\zeta_{t-1}; \gamma, \theta)$ is the logistic transition function, θ is the estimated threshold, and γ controls the speed of transition between states.

Table 2: LSTAR Model Estimation Results

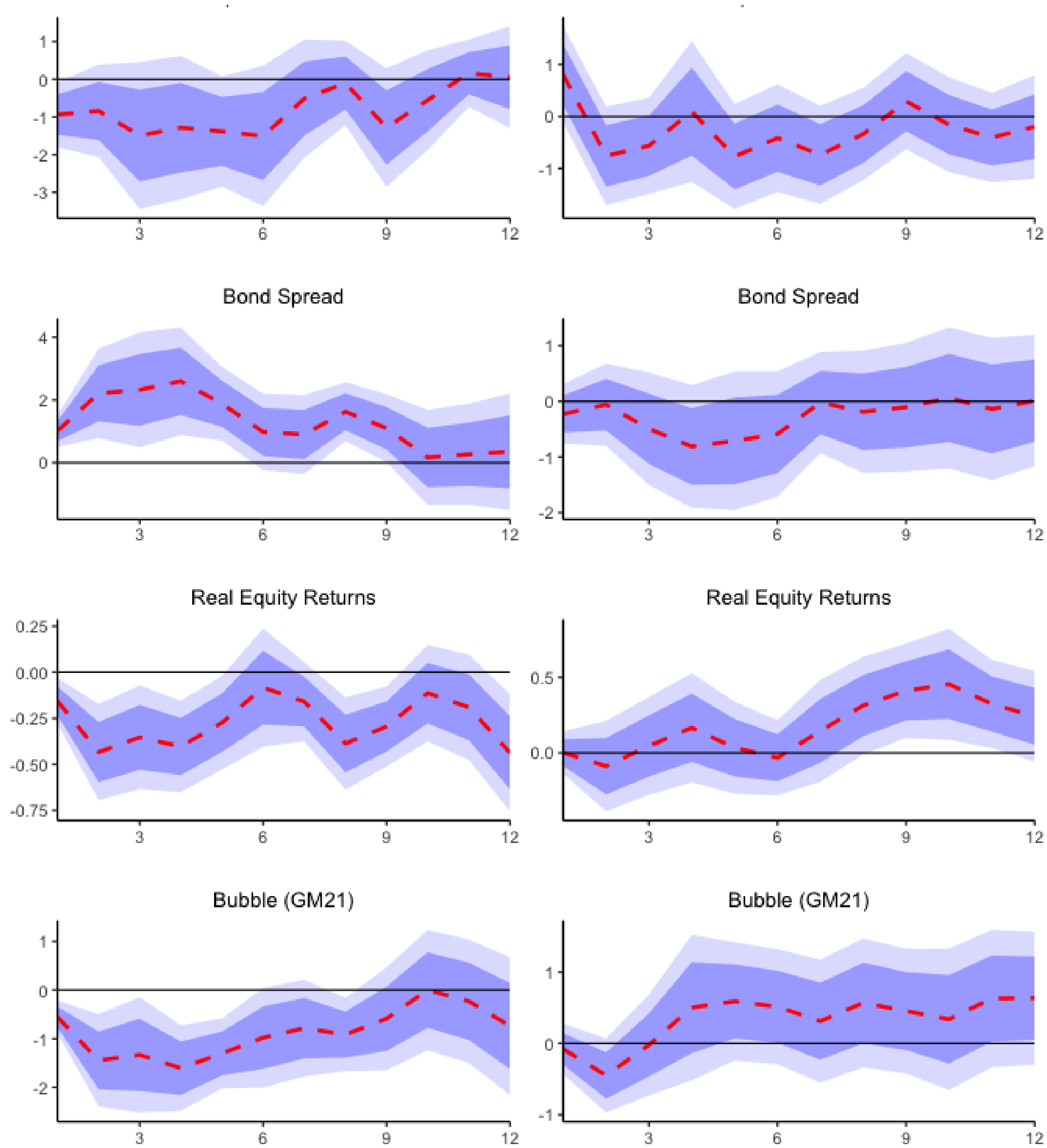
Variable	**Estimate**	**Std. Error**	**t-value**	**p-value**
Low state				
Constant ($\beta_{0,L}$)	0.0255	0.0081	3.142	0.0017**
$\phi_{L,1}$	1.9123	0.0680	28.117	0.000***
$\phi_{L,2}$	-0.9196	0.1195	-7.696	0.000***
$\phi_{L,3}$	-0.0043	0.0562	-0.077	0.938
High state				
Constant ($\beta_{0,H}$)	-0.0034	0.0188	-0.183	0.855
$\phi_{H,1}$	0.5183	0.2806	1.847	0.064*
$\phi_{H,2}$	-0.9918	0.5472	-1.813	0.070*
$\phi_{H,3}$	0.4793	0.2677	1.791	0.073*
Transition Parameters				
Smoothing (γ)	1.9108	1.0264	1.862	0.063*
Threshold (θ)	2.9976	0.7845	3.821	0.0001***

The non-linearity test of the full-order LSTAR model, compared against the full-order AR model, yields F-values of 3.103 with corresponding p-values of 0.062. These results indicate significant non-linearity in the data.

5.3. Pre-pandemic Impulse Responses to Monetary Shock (VAR-LP Model)

Figure 11: Non-Linear Baseline Model IRFs (1998-2019)

High Inflation State | **Low Inflation State**



Note: IRFs from a one standard deviation monetary policy shock.