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MARKET PERCEPTIONS, MONETARY POLICY, AND CREDIBILITY

by Vincenzo Cuciniello*

Abstract

This paper presents novel time-varying estimates of the perceived monetary policy rule by financial markets, focusing on days of heightened inflation-linked swap rates' volatility corresponding to preliminary inflation release dates in the euro area. Findings reveal significant fluctuations in the perceived responsiveness of monetary policy to inflation, reflecting shifts in the ECB's concerns regarding price stability risks. Moreover, the sensitivity of this perceived responsiveness to monetary shocks varies based on prevailing inflation expectations, with tighter policy having a greater impact in high inflation environments. Lastly, a stronger perceived monetary policy response to inflation enhances policy credibility by dampening the sensitivity of long-term inflation expectations to short-term fluctuations.

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1 Introduction[1]

The inflation surge, linked to the post-Covid-19 recovery and intensified by the energy crisis, triggered a synchronized monetary policy tightening. Globally, almost 95% of central banks increased policy rates from early 2021 to mid-2023 (Bank for International Settlements 2023). In this heightened monetary policy period, understanding how adjustments influence the broader economic landscape is imperative. Two factors impact policy transmission effectiveness. First, the greater the public perception of the central bank's actions, the larger the impact on future expectations of interest rates, asset prices, spending decisions, and ultimately, inflation (Woodford 2005). Second, a more credible commitment to a long-run inflation target enhances macroeconomic stability (Orphanides and Williams 2005).

As Rajan (2023) noted: "Unfortunately, the kind of credibility needed to escape a regime of overly low inflation, which we had until recently, is different from the kind needed to curb high inflation, which we have now." The tension between these two types of credibility is even more acute if the inflation regime switch stems from a global shock to energy prices, which, by its nature, may require prudence in the response. For instance, the orthodox monetary policy prescription is to "look through" supply shocks, such as commodity price shocks, that are not assessed to leave a lasting imprint on potential output (Bodenstein et al. 2008). Attempting to offset such shocks with a policy rate increase would cause more inflation volatility rather than less, making it more challenging to meet the inflation target in the medium term.

This paper focuses on key questions arising from inflation regime switches: How have financial market perceptions of the European Central Bank's (ECB) interest rate reaction function evolved over time? How does the perceived responsiveness of monetary policy to inflation affect central bank credibility? Can monetary policy decisions influence financial

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market perception of the central bank's stance on inflation?

To address these questions, a forward-looking perspective is essential. Financial markets shape perceptions of the monetary policy reaction function based on central bank communications and actions, and these perceptions evolve over time in response to changing economic conditions and market expectations. The credibility of monetary policy hinges on the alignment between perceived and actual policy strategies, as well as the central bank's track record in achieving stated objectives.

This study employs a high-frequency approach utilizing publicly available data to gauge market participants' perceptions of the ECB's responsiveness to inflationary signals. The estimation focuses on discerning the time-varying sensitivity of short-term overnight index swap (OIS) rates to inflation-linked swap (ILS) rates, specifically on days coinciding with the release of the euro area's harmonised index of consumer prices (HICP) inflation flash data. Notably, these days exhibit elevated variance in ILS rates, indicating a propensity for fundamental shocks to drive heteroskedasticity within the data.[2] This methodological choice enables an examination of market sentiment regarding monetary policy reactions to inflation dynamics, offering insights into the underlying mechanisms influencing financial market behavior.

The findings indicate that the responsiveness of monetary policy to inflation, denoted by $\hat{\phi}_t$, varies over time, potentially reflecting changes in the ECB's focus on price stability. Notably, the rise in $\hat{\phi}_t$ post-2022 aligns with speeches by ECB officials emphasizing efforts to address inflationary pressures:[3]

- In August 2022, at the annual Jackson Hole central banking symposium, Isabel Schnabel emphasized the ECB's need for "determination" in curbing price rises, "even at the risk of lower growth and higher unemployment."

[2]Bahaj et al. (2023) also note a pronounced increase in the volatility of ILS rates in the UK during data release dates. While liquidity shocks could contribute to this volatility, the authors contend that the primary driver is the revelation of actual economic fundamentals on these specific dates.

[3]The recent normalization of monetary policy, starting in December 2021 with the cessation of net asset purchases within the PEPP and continuing with the discontinuation of net asset purchases under the APP in June 2022, followed by policy rate hikes in July 2022, has been unprecedented in both speed and magnitude.

- In August 2022, Philip R. Lane, at the Annual Meeting of the Central Bank Research Association, highlighted that "when inflation is both high and expected to remain above target for an extended period, the central bank must provide reassurance that it has the capability and determination to deliver its price stability mandate."

- In September 2022, Christine Lagarde pointed out that "[the] imperative to anchor inflation expectations helps explain why, over the last two policy meetings of the ECB's Governing Council, we raised our key interest rates by 125 basis points in total. This is the fastest change in rates in our history and it has sent a strong signal of our determination to return inflation to our medium-term target in a timely manner."

The analysis also examines $\hat{\rho}_t$, representing the perceived inertia in the ECB's interest rate rule, indicative of its tendency to spread policy responses. Findings suggest $\hat{\rho}_t$ rose after the ECB's forward guidance in 2013 and balance sheet expansions, but declined post-2022 due to the Council's data-driven approach to monetary policy tightening.

Overall, the observed fluctuations in the perceived responsiveness of monetary policy to inflation over time align with the ECB's evolving concerns regarding price stability risks, exhibiting higher levels during tightening cycles with data-dependent policy and lower levels during easing cycles and economic uncertainty periods. This observation is consistent with the conclusions drawn by Bauer et al. (2022), who estimate the Federal Reserve's time-varying policy rule using surveys.

In the second part of the paper, I assess the credibility of central banks in anchoring long-term inflation expectations by estimating their sensitivity to short-term expectations, contingent on $\hat{\phi}_t$. The analysis reveals that a robust perceived monetary policy response to inflation signifies strong policy credibility, as it dampens the sensitivity of long-term inflation expectations to short-term fluctuations. For instance, with the average value of $\hat{\phi}_t$ during the recent tightening period, a one percentage point increase in 1-year ILS rates corresponds to only a 4 basis points increase in 5-year, 5-year ILS rates, suggesting a near-zero pass-through and increased credibility in anchoring long-term expectations.

Finally, the paper concludes by highlighting that the sensitivity of $\hat{\phi}_t$ to monetary shocks depends on prevailing inflation expectations among financial market participants. In scenarios anticipating high inflation, monetary tightening amplifies the sensitivity of future policy expectations to inflationary pressures. Conversely, in low inflation environments, monetary surprises exhibit reduced efficacy in altering the perceived policy reaction function. This asymmetry may stem from the extended zero bound on short-term nominal interest rates and the ECB's forward guidance since 2013. Additionally, information shocks do not significantly impact $\hat{\phi}_t$, suggesting its efficacy in capturing financial markets' expected monetary policy reaction function.

Related literature. A growing number of studies estimate monetary policy rules from financial and survey data. Hamilton et al. (2010), for instance, estimate the market-perceived rule using futures prices and lagged actual variables to instrument for market expectations. Similarly, Carvalho and Nechio (2014) utilize consumers' forecasts to estimate the Federal Reserve's forecaster-perceived policy rule. The closest paper is Bauer et al. (2022), who exploit variation across forecasters and forecast horizons to establish a time-varying relationship between Fed funds rate forecasts, inflation forecasts, and output gap forecasts in the United States. I go beyond these prior studies by employing an identification strategy using heteroscedasticity in daily-frequency data and focusing on monetary policy credibility in the euro area.

I further extend the macro-finance literature on the relationship between monetary policy and financial markets (e.g. Cochrane and Piazzesi 2002, Gürkaynak et al. 2005, Hanson and Stein 2015, and Nakamura and Steinsson 2018) and investigate the state-dependence of monetary policy (e.g. Tenreyro and Thwaites 2016, Jordà et al. 2020, Alpanda et al. 2021, Eichenbaum et al. 2022, and Santoro et al. 2014). My contribution lies in demonstrating that market expectations regarding the central bank's responsiveness to inflation can be influenced by monetary policy shocks in periods characterized by high inflation expectations, contrasting with observations in low inflation environments.

This paper also relates to the literature on long-run inflation expectation dynamics (e.g. Beechey et al. 2011, Gürkaynak et al. 2010, Gürkaynak et al. 2007, Stanisławska and Paloviita 2021, and Corsello et al. 2021). Beechey et al. (2011) demonstrate that long-run inflation expectations are more firmly anchored in the euro area compared to the United States. However, Corsello et al. (2021) find that from the second half of 2013, long-horizon inflation expectations in the euro area have become sensitive to short-horizon inflation forecasts and negative inflation surprises. Additionally, utilizing the ECB Consumer Expectation Survey, Stanisławska and Paloviita (2021) observe no heightened co-movement between medium-term inflation expectations among euro area consumers and short-term inflation expectations during the recent inflationary period. I contribute to this literature by providing evidence that a stronger perceived monetary policy response to inflation enhances policy credibility by dampening the sensitivity of long-term inflation expectations to short-term fluctuations.

A road map. Section 2 describes the data. Section 3 depicts the baseline specification for the monetary rule. Section 4 discusses the pattern of estimated inflation-gap and inertial coefficients over time. Section 5 examines the response of long-term inflation expectations to short-term ones, contingent on expectations regarding monetary policy responsiveness to inflation. Section 6 demonstrates that the central bank's perceived responsiveness to inflation is influenced by actual ECB decisions in a high-anticipated inflation environment. Section 7 concludes.

2 Data

I integrate three data sources at different frequencies to analyze the interplay between monetary policy, market perceptions, and credibility. Firstly, high-frequency intraday market reactions to monetary policy announcements are sourced from the Euro Area Monetary Policy Database, accessible at `https://www.ecb.europa.eu/pub/pdf/annex/Dataset_EA-`

`MPD.xlsx`. Second, daily ILS rates come from Bloomberg, while OIS rates, along with macro releases and forecasts, are obtained from Refinitiv. The third source, Consensus Economics, provides short- and long-term expectations for Euro-area GDP growth rates on a quarterly basis.

2.1 High-frequency intraday data

ECB monetary policy surprises. From November 2001 to December 2014, the ECB Governing Council made monthly policy decisions. Starting in January 2015, monetary policy meetings shifted to a six-week cycle. I use OIS rates and STOXX indices around monetary policy meetings from the Euro Area Monetary Policy Database (see Altavilla et al. 2019 for details).

Monetary policy shocks are identified using a high-frequency strategy. The main assumption is that a "monetary policy shock" leads to a negative co-movement of interest rates and stock prices, while an "information shock" results in a positive co-movement (Jarociński and Karadi 2020; Nakamura and Steinsson 2018; Bauer and Swanson 2023).[4] A contractionary pure policy shock raises interest rates and lowers expected dividends, resulting in lower stock prices. If stock prices increase despite monetary tightening, it suggests a more optimistic central bank assessment through the policy announcement.

Monetary policy surprises, in the baseline, are gauged by the 2-year OIS rate change during the window spanning the press release and press conference. Two reasons underlie the choice of 2-year OIS rates: firstly, they align with the paper's focus on the 2-year horizon; and secondly, this time horizon encompasses both conventional and unconventional monetary policy shocks. Panel (a) of Figure A1 displays time series of these shocks. They are roughly evenly split between expansionary and contractionary impulses.

[4] The additional identification assumptions include the absence of other relevant news within the specified time window. The utilization of tick data and a narrow window around the release of the monetary policy decision, statement reading, and subsequent Q&A in the press conference helps mitigate this potential issue.

2.2 Daily data

Inflation rates y-o-y flash surprises. Since 2005, the euro area has provided monthly flash estimates of the Harmonized Index of Consumer Prices (HICP). I gathered information on HICP inflation flash release dates from Refinitiv. The primary sample period covers all regularly scheduled HICP inflation flash releases from November 30, 2004, to September 29, 2023. In total, the sample includes 225 HICP inflation flash release dates, with only two, August 31, 2006, and April 30, 2020, coinciding with a monetary policy announcement by the Governing Council. Occasionally, additional macroeconomic indicators like GDP growth and the unemployment rate were released alongside the HICP inflation flash news. Within the sample, there are 99 HICP inflation flash release dates that also include the euro area unemployment rate. The HICP inflation flash surprise is the difference between the actual value of the HICP inflation flash data and the median of analysts' forecasts obtained from the Refinitiv poll conducted a few days before the data release (refer to panel (b) of Figure A1).

Unemployment rate surprises. I collect released and forecasted euro-area unemployment rates from Refinitiv Economic Indicators each month. The sample period is from April 1, 2004, to October 2, 2023. The unemployment rate surprise is the disparity between the actual rate and the median of analysts' forecasts from the Refinitiv poll conducted a few days prior to data release (refer to panel (c) of Figure A1).

GDP growth rate y-o-y flash surprises. The preliminary flash estimate dates, along with actual and forecasted annual GDP growth, are sourced from Refinitiv Economic Indicators. The sample period extends from April 29, 2016, to July 31, 2023. The GDP growth surprise is defined as the difference between the actual value of the GDP year-on-year growth flash data and the median of analysts' forecasts obtained from the Refinitiv poll conducted a few days before the data release (refer to panel (d) of Figure A1).

Inflation-linked swap rates. As a measure of inflation expectations I use forward ILS rates, in line with existing literature for the euro area and other major economies (e.g. Gürkaynak et al. 2010). Usually, the swap contract is linked to a non-seasonally adjusted consumer price index (CPI). In the euro area, this index is the HICP excluding tobacco (HICPxT). The euro-area ILS market is widely considered the most liquid globally, conferring significant advantages over alternative indicators of inflation expectations. Notably, ILS are traded at a considerably higher frequency than survey data, rendering them more suitable for analyzing reactions to macroeconomic news. I obtained euro-area ILS rates (EUSWI) from Bloomberg. The sample period covers June 23, 2004, to October 4, 2023.

Risk-free interest rates. The risk-free rate curve in the euro area is proxied by the Overnight Index Swaps (OIS) term structure. In the euro market, OISs predominantly use the euro overnight index average (EONIA) rate, which is a weighted average of interest rates on unsecured overnight loans in the euro area interbank market. Since October 2, 2019, the EONIA has been replaced by the euro short-term rate (€STR), reflecting the wholesale euro unsecured overnight borrowing costs of euro-area banks. I obtained daily OIS rates (OIEUR) from Refinitiv. Rates for OIS with maturities of 1 year and 2 years have been available since October 7, 1999, and January 16, 2001, respectively. Rates exceeding 3 years have been accessible since August 15, 2005.

2.3 Quarterly data

Expected GDP growth rates. GDP growth expectations are drawn from Consensus Economics surveys, which are released in January, April, July, and October of each year.[5] Professionals participating in the survey are tasked with providing GDP growth forecasts for the current year as well as for each of the subsequent five years. Additionally, they are asked

[5]From April 2005 to April 2014, Consensus Economics offered semiannual data on long-term GDP growth forecasts, released in April and October annually. I transformed these series into quarterly frequency using interpolation.

to predict the average GDP growth rate for the period spanning six to ten years ahead.

Quarterly surveys yield fixed-event forecasts, meaning each expectation pertains to a specific calendar year, regardless of the survey's collection time (e.g., either April or October). For instance, expectations for a given year, when formulated in April, necessitate a forecast of GDP growth rate for at least the subsequent 9 months; when formulated in October, they require a forecast for only 3 months.

Following a standard approach in the literature (e.g. Dovern et al. 2012), I approximate the fixed-horizon forecast for the 1-year, 1-year ahead GDP growth rate ($g_{1y1y|t}$) as an average of the fixed-event forecast for the one-year ahead GDP growth rate ($g_{1y|t}$) and the two-year ahead GDP growth rate ($g_{2y|t}$), weighted by their share in the forecasting horizon. The 1-year, 1-year ahead forecast for quarter q and year y is given by:

$$g_{1y1y|t} = \frac{4-q+1}{4} g_{1y|t} + \frac{q-1}{4} g_{2y|t}.$$

Similarly, the 5-year, 5-year ahead forecast for quarter q and year y is given by:

$$g_{5y5y|t} = \frac{4-q+1}{4} g_{5y|t} + \frac{q-1}{4} g_{6+y|t},$$

where $g_{5y|t}$ is the 5-year ahead GDP growth rate forecast and $g_{6+y|t}$ denotes average GDP growth rate between six and ten years ahead.

3 Baseline policy rule specification

I assume that market participants believe that the central bank follows a simple policy rule:

$$i_t = \rho_t i_{t-1} + \alpha_t i_t^* + \phi_t(\pi_t - \pi_t^*) + \tau_t(g_t - g_t^*) + \mu_t, \tag{1}$$

where i_t represents the policy rate at time t, $\pi_t - \pi_t^*$ denotes the deviation of HICP inflation from its equilibrium target at time t, i_t^* is the equilibrium nominal short-term interest rate, and $g_t - g_t^*$ is the real output growth gap. All coefficients are time-varying. The parameters ϕ_t, τ_t, and α_t represent the policy rates' responses to the inflation gap, output growth gap, and equilibrium nominal short-term interest rate, respectively. Finally, the variable μ_t represents a monetary policy shock that is exogenous to the policy rule. Introducing $\rho_t \neq 0$ incorporates Taylor's specification of a policy rule with partial adjustment, reflecting the potential for interest rate smoothing. This particular policy rule aligns with the specifications commonly employed in empirical macroeconomics literature (e.g. Taylor 1993, Orphanides 2003, and Taylor and Williams 2010).

In the baseline, I consider a fixed horizon: the 1-year forward rate for the year ahead, for several reasons. Fixed horizon forecasts are more easily comparable over time than those where the forecast horizon varies throughout the year and may become shorter than the policy horizon. For instance, inflation forecasts from professional forecasters usually pertain to different horizons for the inflation rate in a given calendar year, constituting fixed event forecasts.

Additionally, fixed horizon forecasts are more practical for policy purposes. Central banks, for example, are typically most interested in inflation expectations one to two years ahead, as this is the horizon where monetary policy will exert its greatest impact. They may have less use for shorter forecasts (e.g. Svensson 1997).

The perceived monetary policy at a 1-year, 1-year-ahead horizon at time t can be expressed using the following equation:

$$i_{1y1y|t} = \hat{\rho}_t i_{1y|t} + \hat{\alpha}_t i_{5y5y|t} + \hat{\phi}_t(\pi_{1y1y|t} - \pi_{5y5y|t}) + \hat{\tau}_t(g_{1y1y|t} - g_{5y5y|t}) + e_{1y1y|t}, \tag{2}$$

where $i_{1y1y|t}$, $i_{5y5y|t}$, and $i_{1y|t}$ represent OIS rates at the 1-year, 1-year-ahead, at the 5-year, 5-year-ahead, and at the 1-year horizon as of time t. $\pi_{1y1y|t}$ and $\pi_{5y5y|t}$ are the ILS rates at the 1-year, 1-year-ahead and at the 5-year, 5-year-ahead horizon as of time t. The expected output gap ($g_{1y1y|t} - g_{5y5y|t}$) is given by the difference between the forecasters' average GDP growth in the euro area expected 1-year ahead ($g_{1y1y|t}$) and 5-years, 5-year ahead ($g_{5y5y|t}$) respectively. The error term $e_{1y1y|t}$ contains the expectation of the future monetary policy shock at the 1-year, 1-year ahead, denoted as $E_t\mu_{1y1y}$, as well as any measurement and specification errors affecting the OIS rates at the 1-year, 1-year-ahead.[6]

Long-run variables in equation (1) operate at a 5-year, 5-year horizon, with long horizons for OIS and ILS rates being more susceptible to changes in term premia than their 1-year, 1-year ahead counterparts. Empirical research indicates that a relaxation of monetary policy, even through conventional tools in standard conditions (namely when interest rates are *not* at or near the effective lower bound), generally leads to a decline in both the term premiums of long-term Treasury bonds and the credit spreads on corporate bonds (Hanson and Stein 2015, Gertler and Karadi 2015, Gilchrist et al. 2015). This suggests that monetary policy influences capital market risk premiums, possibly through mechanisms involving risk-taking or a reach-for-yield phenomenon.

Before recovering the perceived monetary policy rule from regression (2), in the following sections, I first address the problem of omitted variable or simultaneous causality bias that can affect estimates. Subsequently, I demonstrate that most of the variability in OIS rates can be explained by regression (2).

3.1 Endogeneity bias

The first concern is that regression (2) may be affected by endogeneity bias, as the perceived effects of monetary policy shocks on inflation could influence estimates. In this section, I demonstrate that endogeneity bias may be less severe when focusing on dates of HICP inflation flash releases, as shocks to fundamentals are more likely to drive most of the het-

[6]The derivation of (2) assumes financial markets perceive the monetary policy rule parameters as highly persistent, or formally, as time-varying martingale parameters orthogonal to other shocks, e.g., $E_t\phi_{t+h} = \hat{\phi}_t$ and $E_t\phi_{t+h}v_{t+h} = \hat{\phi}_t E_t v_{t+h}$ for any macro variable v_t. This assumption is common in the literature on time-varying parameter models (e.g., Bauer et al. 2022 and Primiceri 2005).

eroskedasticity on these dates.

Let $|\widetilde{\pi}_t|$ be the absolute value of a 3-day change in ILS rates around the HICP inflation flash release date t, including the ILS rates on the trading days before and after t. I estimate the following equation:

$$|\widetilde{\pi}_t| = a + k_{-4}R_{t-4} + k_{-2}R_{t-2} + k_0 R_t + k_{+2}R_{t+2} + k_{+4}R_{t+4} + \sum_{x=1}^{7} b_x |\widetilde{\pi}_{t-x}| + d_t + w_t + m_t + e_t, \quad (3)$$

where R_{t+x} is a dummy variable indicating HICP flash release x days from t.[7] The analysis focuses on a $\{-4, +4\}$ day window around HICP flash releases. Therefore, k_0 represents the differential intercept term on the HICP flash release date t and indicates how much the change in ILS rates on HICP flash releases differs from the mean change (namely the Constant a in eq. 3). For $x \neq 0$, the k_x-coefficients measure the differential ILS changes from the mean x days from t. d_t, w_t, and m_t denote day-of-week, week-of-year, and month dummies.

Table 1 presents estimates of equation (3) spanning from August 31, 2005, to September 29, 2023. Short-term inflation expectations exhibit systematically higher volatility surrounding HICP flash releases compared to the days immediately before and after. Coefficients k_0 for maturities up to 3 years suggest that absolute changes in ILS rates on flash release dates are, on average, 1-7 basis points larger than the average absolute change ("Constant"). For instance, the volatility on flash releases for 1-year and 2-year ILS is respectively three and two times larger than the average. In the days immediately following the HICP flash release, volatility tends to be lower than the average, while in the day immediately preceding, it is not statistically different from the average.[8]

This result suggests that around data release dates traders learn the most about where inflation is heading. This is in line with Bahaj et al. (2023) who find that the volatility of

[7] I focus on 3-day changes in ILS rates to tackle challenges related to illiquidity, infrequent trading, and the over-the-counter (OTC) nature of these instruments. Notably, Bahaj et al. (2023) demonstrate that UK ILS prices incorporate new information within two to three days.

[8] In Appendix A.2, I demonstrate that the results remain robust even after accounting for other concurrent macro information releases that have the potential to influence ILS rate volatility around HICP flash releases.

Table 1: ILS around HICP flash releases

VARIABLES	(1) $\widetilde{\pi}^{1y}$	(2) $\widetilde{\pi}^{2y}$	(3) $\widetilde{\pi}^{3y}$	(4) $\widetilde{\pi}^{4y}$	(5) $\widetilde{\pi}^{5y}$	(6) $\widetilde{\pi}^{10y}$
k_{-4}	-0.004	-0.005	-0.000	-0.002	-0.002	0.003
	(0.007)	(0.004)	(0.003)	(0.003)	(0.003)	(0.002)
k_{-2}	0.014	0.009	0.004	0.003	0.001	0.001
	(0.009)	(0.005)	(0.004)	(0.004)	(0.003)	(0.002)
k_0	0.065***	0.023***	0.011**	0.005	0.006	0.002
	(0.015)	(0.008)	(0.006)	(0.004)	(0.004)	(0.003)
k_{+2}	-0.039***	-0.018***	-0.015***	-0.007**	-0.006*	0.002
	(0.008)	(0.005)	(0.004)	(0.003)	(0.003)	(0.003)
k_{+4}	-0.013**	-0.009**	-0.006*	-0.003	-0.003	0.000
	(0.006)	(0.004)	(0.003)	(0.002)	(0.002)	(0.002)
Constant	0.017***	0.012***	0.009***	0.009***	0.008***	0.007***
	(0.003)	(0.002)	(0.001)	(0.001)	(0.001)	(0.001)
Observations	4,040	4,135	4,227	4,220	4,135	3,788
R^2	0.420	0.384	0.393	0.372	0.373	0.333
day-of-week FE	Y	Y	Y	Y	Y	Y
week-of-year FE	Y	Y	Y	Y	Y	Y
month FE	Y	Y	Y	Y	Y	Y

Notes: Huber-White robust standard errors in parentheses *** p<0.01, ** p<0.05, * p<0.1. The sample period spans from August 31, 2005, to September 29, 2023.

ILS rates in UK is noticeably higher around data release dates. Although liquidity shocks may also be higher during those dates, they argue that the major difference should be that actual fundamentals are revealed on these dates.

3.2 Is regression (2) consistent with the data?

Another concern regarding the recovery of perceived monetary policy from regression (2) is its poor fit with the data, suggesting that risk-free interest rates may depend on factors beyond inflation and the output gap.

Table 2 presents estimated regression coefficients for (2) across the entire sample period, controlling for macroeconomic news releases coinciding with the HICP inflation flash release and EURO STOXX 50 volatility. Columns 1 and 2 of Table 2 report the OLS estimation of regression (2) using Bartlett/Newey-West standard errors with five lags. All coefficients have the expected sign and are statistically different from zero. The average R^2 is 0.98, indicating a very good fit with the data.

As a robustness exercise, columns 3 and 4 report the 2SLS estimates, where the inflation gap is instrumented using forecaster surprises on the HICP inflation flash release, defined as the difference between the flash estimate of the euro-area headline inflation and the median forecasts in the Refinitiv survey. Forecasts are provided a few days before the estimation of the HICP inflation. The Kleibergen-Paap F-statistic of excluded instruments comfortably exceeds critical values according to Stock and Yogo (2005) for all considered horizons. All coefficients estimated with OLS or 2SLS are statically not different from each other.[9]

All in all, the results so far suggest that focusing solely on the HICP inflation flash dates offers a credible identification strategy for the direct effects of inflation surprises on short-term inflation expectations. The short time span between the inflation shock and the revision of expected inflation ensures that the revisions are not a response to other transmission channels of the inflation surprise.

[9] These results maintain robustness even under the assumption that all long-run variables are derived from forecaster surveys or considering a 1-year, 9-year ahead horizon for financial variables (see Table A2 in the Appendix).

Table 2: 1-year forward Taylor rules at the 1-year ahead horizon with macro news

VARIABLES	(1) 1Y1Y OLS	(2) 1Y1Y OLS	(3) 1Y1Y 2SLS	(4) 1Y1Y 2SLS
$\hat{\alpha}$	0.364***	0.344***	0.378***	0.351***
	(0.041)	(0.038)	(0.037)	(0.036)
$\hat{\rho}$	0.682***	0.693***	0.650***	0.673***
	(0.039)	(0.037)	(0.038)	(0.035)
$\hat{\tau}$	0.076***	0.060***	0.051	0.043
	(0.024)	(0.023)	(0.041)	(0.039)
$\hat{\phi}$	0.331***	0.362***	0.505***	0.484***
	(0.105)	(0.101)	(0.131)	(0.122)
GDP growth surprise		-0.014	0.011	-0.016
		(0.022)	(0.075)	(0.074)
Unemployment surprise		-0.239	-0.227	-0.303
		(0.190)	(0.214)	(0.209)
VSTOXX		0.006*		0.007**
		(0.003)		(0.003)
Observations	217	217	217	217
R^2	0.983	0.983		
month FE	Y	Y	Y	Y
KP F-test			35.05	40.22

Notes: Bartlett/Newey-West standard errors with five lags are displayed in parentheses. *** p<0.01, ** p<0.05, and * p<0.1. The sample period spans from August 31, 2005, to September 29, 2023. The inflation gap is instrumented using forecaster surprises on the HICP inflation flash release, defined as the difference between the flash estimate of the euro-area headline inflation and the median forecasts in the Refinitiv survey.

4 Perceived monetary policy reaction function

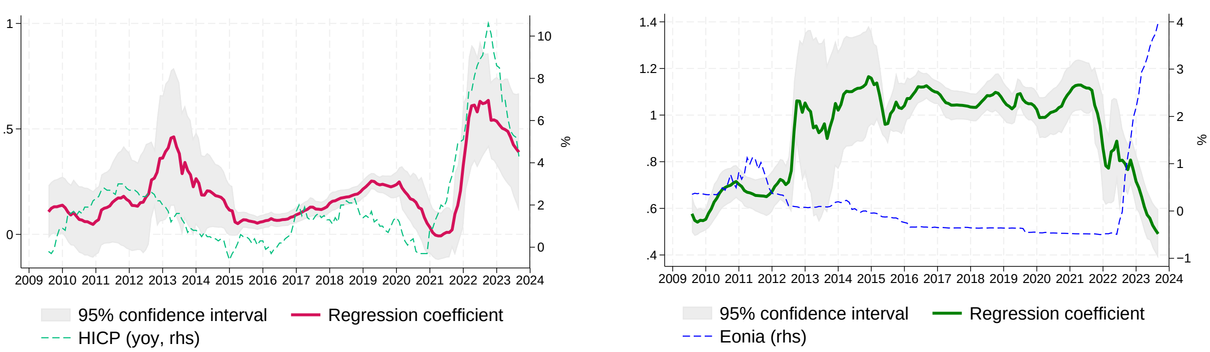
This section centers on estimating the time-varying coefficients of regression (2) exclusively on the HICP inflation flash realease dates.[10] The identifying assumption is that spikes in variance around these dates result solely from the emergence of fundamental news about inflation.

[10]Results remain robust when considering a 3 or 5-day window around HICP inflation flash releases.

Figure 1 illustrates the pattern of the estimated coefficients $\hat{\phi}_t$ and $\hat{\rho}_t$ in regression (2) at the 1-year, 1-year ahead horizon using rolling OLS regression with Huber-White robust standard errors. Specifically, I use a 4-year rolling window to get time-varying coefficient values. The sample period extends from August 31, 2005, to September 29, 2023. Carvalho et al. (2021), for instance, advocate for OLS estimation of monetary policy rules, asserting that the impulse response functions produced by the model under the policy rule estimated by single-equation OLS are close to the true impulse response functions.

Figure 1: 1Y1Y Taylor rule estimated coefficients

(a) $\hat{\phi}_t$ (b) $\hat{\rho}_t$



Notes: Coefficients in (2) are estimated through 4-year rolling windows OLS on HICP inflation flash release dates, employing Huber-White robust standard errors. The EONIA reported in the figure has adopted a revised methodology since October 2, 2019, incorporating €STR plus a fixed spread of 8.5 basis points.

Panel (a) depicts a significant and sustained upward shift in the ECB's perceived responsiveness to inflation since early 2022. The substantial and prolonged surge in inflation prompted a vigorous and remarkably swift monetary policy response by the ECB. Notably, annual inflation increased by 4 percentage points, rising from 4.9 percent to 8.9 percent, between December 2021 and July 2022 (dashed lines in panel (a) of Figure 1). The 1-year ahead ILS rates also surged by 5 percentage points, increasing from 3 percent to 8 percent. A second surge in $\hat{\phi}_t$ occurred between 2011 and 2012 when inflation remained above 2%. In April 2011, euro area HICP inflation reached 2.8%, and the 1-year ahead ILS rates were 2.5%, prompting the ECB to implement its first interest rate rise since 2008. A further increase in key interest rates took place in July 2011, signaling "upside risks to price stability."[11] The monetary policy response came in the summer of 2012 with the announcement of the Outright Monetary Transactions (OMT) to address the sovereign debt crisis, which had shifted from a tail risk for inflation to a substantial threat to price stability.[12]

As highlighted by Ferroni et al. (2023), what is essential for stabilizing inflation is not only the immediate response of the interest rate to current inflation at the beginning of a tightening cycle, but also the expectation that the cumulative increase in the interest rate over the entire cycle is sufficiently robust. Panel (b) in Figure 1 illustrates that inertial coefficient $\hat{\rho}_t$ moved from 0.7 to 1 in mid-2012. Only with the onset of the rise in policy interest rates in the summer of 2022 did the inertia coefficient begin to decline, aligning with the ECB Governing Council's data-dependent and meeting-by-meeting approach since the tightening of monetary policy.[13]

Overall, the patterns of coefficients in Figure 1 suggest that the perceived responsiveness of monetary policy to inflation varies significantly over time, reflecting the ECB's changing concerns regarding price stability risks versus financial and other risks. It tends to be high during monetary tightening cycles when ECB policy is perceived to be data-dependent, and low during easing cycles and times of heightened economic and financial uncertainty.

Bauer et al. (2022) find that the Fed's inflation gap coefficient is less meaningful than the output gap coefficient, given the stability of inflation near the two percent target and limited supply shocks. However, Appendix A.4 discloses that, excluding some episodes, estimated output gap coefficients $\hat{\tau}_t$ predominantly remained close to zero in the euro area. It is noteworthy that the ECB's single mandate, outlined in Article 105(1) of the Treaty establishing the ECB, is different from the Fed's dual mandate. The ECB's "primary objective... shall be

[11] See the Introduction statement to the press conference in July 2011 (`https://www.ecb.europa.eu/press/pressconf/2011/html/is110707.en.html`).

[12] See e.g. the speech by Mario Draghi at the ECB Forum on Central Banking in June 2018 (`https://www.ecb.europa.eu/press/key/date/2018/html/ecb.sp180619.en.html`)

[13] The reduction in the inertial component in the Taylor rules is also observed in the ECB's survey of monetary analysts since July 2022 (Bernardini and Lin 2023).

to maintain price stability."[14] Furthermore, as instances arise where $\hat{\rho}_t$ exceeds one, leading to undefined long-run perceived responses ($\hat{\phi}_t/(1\text{-}\hat{\rho}_t)$), the focus of my analysis will be on the short-run perceived responses, $\hat{\phi}_t$.

5 $\hat{\phi}_t$ and monetary policy credibility

Having analyzed the changes in $\hat{\phi}_t$ on HICP inflation flash release days and across the business cycle, the subsequent phase involves examining the response of long-term inflation expectations to monetary policy, contingent on the perceived monetary policy responsiveness to inflation.

Persistent and significant inflation surprises challenge monetary policy. If not addressed with appropriate actions, they can lead to long-term inflation expectations detaching from the central bank's target for investors, businesses, and households. It is crucial to note that as inflation deviations from the target increase, the link between monetary policy credibility and perceived aggression in addressing inflation becomes even more vital. In cases of significant inflation surprises, market participants closely observe the central bank's actions and responsiveness. If the central bank is seen as insufficiently aggressive in its measures, it can erode credibility and exacerbate monetary policy challenges.

Carvalho et al. (2023) find that the degree of expectations anchoring depends on the endogenous link between long-term expectations and short-term forecast errors. Stronger monetary policy responses reduce self-referentiality–beliefs determining inflation, which, in turn, determines beliefs (in the language of Marcet and Sargent 1989). This leads firms to choose a forecasting model with lower sensitivity to new information (lower gain) given other firms' behavior.[15]

[14] The ECB's definition of price stability evolved from annual increases below 2% to "inflation rates below, but close to 2% over the medium term," and in July 2021, it revised the objective to "2% over the medium term," considering deviations from this target as equally undesirable.

[15] In the euro area data, Miccoli and Neri (2019) investigate the relationship between inflation shocks and expectations measured by inflation-linked swap contracts. They find that negative inflation surprises led to a significant drop in expectations, but this effect diminished after the introduction of the ECB's Asset Purchase Programme. Additionally, García and Werner (2021) observe that the euro area experienced a

To assess monetary policy credibility, I analyze the responsiveness of long-term inflation expectations to short-term expectations, contingent on the market's perceived reaction to the inflation gap. Specifically, I estimate the following equation:

$$\pi_{5y5y|t} = \xi + \beta\pi_{jy|t} + \chi[\hat{\phi}_{t|t-1} \times \pi_{jy|t}] + \theta\hat{\phi}_{t|t-1} + u_t \qquad j \in \{1, 2\}, \tag{4}$$

where $\pi_{5y5y|t}$ is 5-year, 5-year ahead ILS rates and $\hat{\phi}_{t|t-1}$ represents the most recent market-perceived 1-year, 1-year ahead interest rate response to the inflation gap on the HICP inflation flash release before time t. For short-term inflation expectations, I consider either $\pi_{1y|t}$ or $\pi_{2y|t}$, representing 1-year and 2-year ILS rates, respectively. The interaction term, $\chi[\hat{\phi}_{t|t-1} \times \pi_{jy|t}]$, accounts for potential non-linearities in the pass-through from short- to long-term inflation expectations, where β, χ, ξ, and θ are constant coefficients. The total pass-through on the 5-year, 5-year ahead inflation expectation is determined by $\partial\pi_{5y5y|t}/\partial\pi_{jy|t} = \beta + \chi\hat{\phi}_{t|t-1}$.[16]

weaker anchoring of expectations during the post-2013 period of low inflation, as also indicated by the greater impact of inflation news on market-based measures of inflation compensation during this period.

[16]To address the generated regressor nature of $\hat{\phi}_t$, I calculate the Huber-White standard errors using 5,000 bootstrap replications. This approach allows for robust estimation and inference, considering the potential uncertainty associated with the generated regressor.

Table 3: Monetary policy credibility

VARIABLES	(1) 1Y	(2) 1Y	(3) 1Y	(4) 2Y	(5) 2Y	(6) 2Y
β	0.34***	0.35***	0.16***	0.59***	0.59***	0.30***
	(0.05)	(0.01)	(0.01)	(0.07)	(0.02)	(0.02)
χ	-0.60***	-0.60***	-0.24***	-0.96***	-0.96***	-0.40***
	(0.09)	(0.02)	(0.02)	(0.13)	(0.03)	(0.03)
Observations	170	3,687	3,687	170	3,687	3,687
R^2	0.329	0.328	0.991	0.453	0.454	0.991
$\beta + \chi \mathbb{E}\hat{\phi}_{t\|t-1}$	0.04	0.04***	0.04***	0.10**	0.11***	0.09***
s.e.	0.03	0.01	0.01	0.05	0.01	0.01
Year-Month FE	N	N	Y	N	N	Y

Notes: Huber-White standard errors in parentheses *** p<0.01, ** p<0.05, * p<0.1. Standard errors are computed using 5,000 bootstrap replications. The sample period in columns 1 and 4 includes only the dates of HICP inflation flash releases. The sample period in the other columns spans from September 1, 2009, to September 26, 2023.

The OLS β coefficients in Columns 1-3 of Table 3 indicate positive associations between 1-year and long-term inflation expectations, with values of 0.3 and 0.2 without and with year-month fixed effect, respectively. However, the mitigating effect of monetary policy responsiveness is evident through the negative interaction coefficient χ in eq. (4), suggesting that a more aggressive policy stance perceived by financial markets dampens the pass-through from short to long-term inflation expectations.

To assess overall pass-through from short- to long-term inflation expectations, I then calculate the coefficient $\beta + \chi \mathbb{E}\hat{\phi}_{t|t-1}$, where $\mathbb{E}\hat{\phi}_{t|t-1}$ is the average value of $\hat{\phi}_{t|t-1}$ in the period 2022-2023. Notably, a one percentage point increase in 1-year ILS rates corresponds to 4 basis point increases in 5-year, 5-year ILS rates. This implies that the average level of $\hat{\phi}_t$ during the recent monetary policy tightening results in an overall pass-through very close to zero, highlighting an increased credibility of the central bank to anchor long-term expectations. Columns 4-6 of Table 3 show that the findings for the 2-year ILS rates align with those for the 1-year ILS rates.

6 Can monetary policy affect $\hat{\phi}_t$?

In this section, I show that the central bank's perceived responsiveness to inflation is influenced by actual ECB decisions when financial markets anticipate high inflation.

Figure 1 suggests that changes in monetary policy responsiveness to inflation are more likely in the state of the world with significant changes in inflation expectations. To assess whether $\hat{\phi}_t$ actually responds to monetary policy surprises in a state-contingent manner without specifying an underlying multivariate dynamic system, I follow Jordà (2005) and Stock and Watson (2018) using the local projections method.

$$\hat{\phi}_{t+h} = \delta^h + \gamma_1^h \mu_t \times (1 - high_t) + \gamma_2^h \mu_t \times high_t + \varrho^h high_t + \varsigma^h \hat{\phi}_{t-1} + u_{t+h} \tag{5}$$

where μ_t signifies monthly monetary surprises, normalized so that a positive unit value corresponds to a 25 basis points contractionary shock (see Section 2), and $high_t$ is a dummy variable equalling one when inflation expectations are high, namely when the 1-year, 1-year ahead ILS rates are above the 75th percentile, and zero otherwise.[17] The regressions control for lagged $\hat{\phi}_t$ to address serial correlation in the perceived policy rule coefficient.

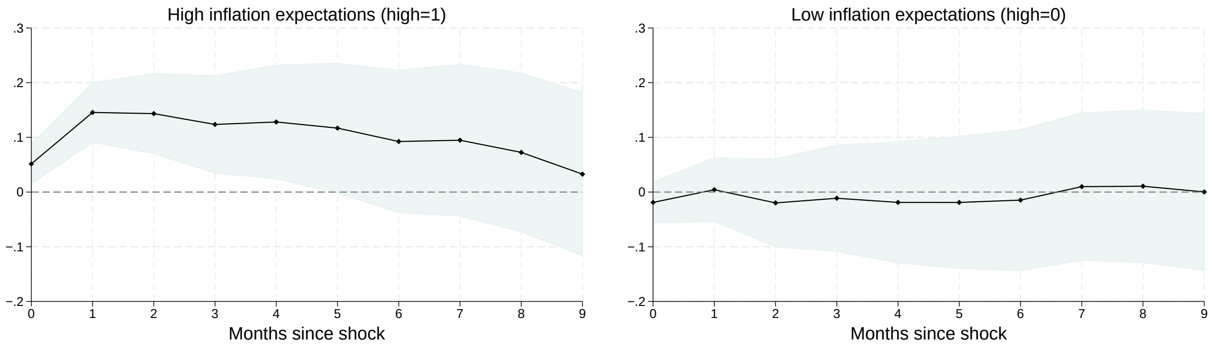
The local projections model is estimated based on the OLS over the entire sample for $\hat{\phi}_t$, spanning from August 2009 to September 2023. I account for serial correlation in the error terms by using the Newey-West standard errors with $1.5 \times h$ lags. I set μ_t to zero when no policy surprises are available.

The left panel of Figure 2 illustrates impulse responses of $\hat{\phi}_t$ following a contractionary monetary shock when high anticipated inflation is present. A 25 basis points monetary policy tightening has a positive and two-quarter persistent effect on $\hat{\phi}_t$, reaching a peak of 0.15 in the first month after the shock. This change accounts for an economically significant amount of the perceived monetary policy responsiveness to inflation and is about one standard deviation

[17]The dummy $high_t$ equals 1 for 3 months in 2009, 4 months in 2011, 3 months in 2012, and for 11 months in 2022, and 9 months in 2023.

of $\hat{\phi}_t$. Conversely, the right panel of Figure 2 shows that a monetary policy shock does not appear to affect $\hat{\phi}_t$ when inflation expectations are relatively low.

Figure 2: $\hat{\phi}_t$ responses to a 25bps Monetary policy shock

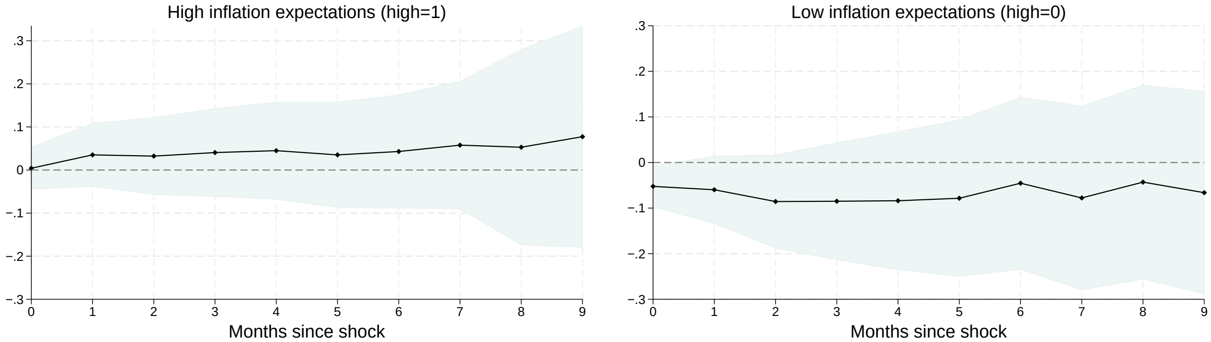


Notes: This figure shows the state-dependent impulse responses of $\hat{\phi}_t$ to a contractionary monetary policy shock of 25 basis points. The coefficients γ_1^h and γ_2^h (black lines) are estimated using equation (5). The light gray shaded areas represent 90% confidence bands, respectively, constructed using Newey-West estimators with $1.5 \times h$ lags.

The above findings suggest that the sensitivity of $\hat{\phi}_t$ to monetary shocks is state-dependent, contingent upon financial market participants' inflation expectations. In an anticipated high inflation environment, a monetary tightening influences how financial markets perceive the trajectory of future monetary policy relative to inflation, potentially impacting the reaction function. Conversely, in an expected low inflation environment, monetary surprises seem less effective in altering the monetary policy reaction function. This could be attributed to the prolonged period of the zero bound on short-term nominal interest rates during the sample period, as well as the ECB's forward guidance on interest rates since 2013 (see panel (b) of Figure 1).

To assess whether information shocks can also affect $\hat{\phi}_t$ in a state-contingent manner, Figure 3 presents estimated values of γ_1^h and γ_2^h when considering an information shock instead of a monetary shock. These coefficients are statistically not different from zero, indicating that the values of $\hat{\phi}_t$ effectively capture the financial markets' expected monetary policy reaction function.

Figure 3: $\hat{\phi}_t$ responses to a 25bps Information shock



Notes: This figure shows the state-dependent impulse responses of $\hat{\phi}_t$ to a contractionary information shock of 25 basis points. I utilize a one-month lead of $\hat{\phi}_t$ to accommodate for at least one month between the monetary policy shock and the HICP inflation flash release dates. The coefficients γ_1^h and γ_2^h (black lines) are estimated using equation (5). The light gray shaded areas represent 90% confidence bands, respectively, constructed using Newey-West estimators with $1.5 \times h$ lags.

7 Conclusions

This paper introduces new time-varying estimates of the monetary policy rule perceived by financial market participants on preliminary inflation release dates, namely when inflation-linked swap rates' volatility is relatively higher. The identifying assumption is that news about inflation is disseminated in a lumpy manner, with specific days determined by the HICP inflation release dates.

With the estimates of the perceived monetary policy rule, I document three relevant facts for monetary policy and asset pricing. First, the perceived responsiveness of monetary policy to inflation varies significantly over time, reflecting the ECB's changing concerns regarding price stability risks. It tends to be high during monetary tightening cycles when ECB policy is perceived to be data-dependent, and low during easing cycles and times of heightened economic and financial uncertainty. Second, the sensitivity of the perceived responsiveness of monetary policy to monetary shocks is state-dependent, meaning it depends on financial market participants' inflation expectations. In an environment with expected high inflation, a monetary tightening increases the sensitivity of the perceived path of future monetary policy per unit of inflation. However, in an environment with expected low inflation, monetary surprises do not appear to be as effective in altering the perceived monetary policy reaction function. Third, a stronger perceived monetary policy response to inflation indicates strong monetary policy credibility as it reduces the sensitivity of long-term inflation expectations to short-term expectations.

These findings suggest that well-anchored inflation expectations are more likely in a high inflation environment with a perceived strong monetary policy response to inflation deviations from the target. However, heightened sensitivity of interest rates to inflation could amplify volatility in financial markets. Thus, policymakers should carefully adjust monetary policy to anchor long-term inflation expectations while minimizing excessive fluctuations in financial markets and economic activity. Further research on how monetary policy should balance these aspects is necessary.

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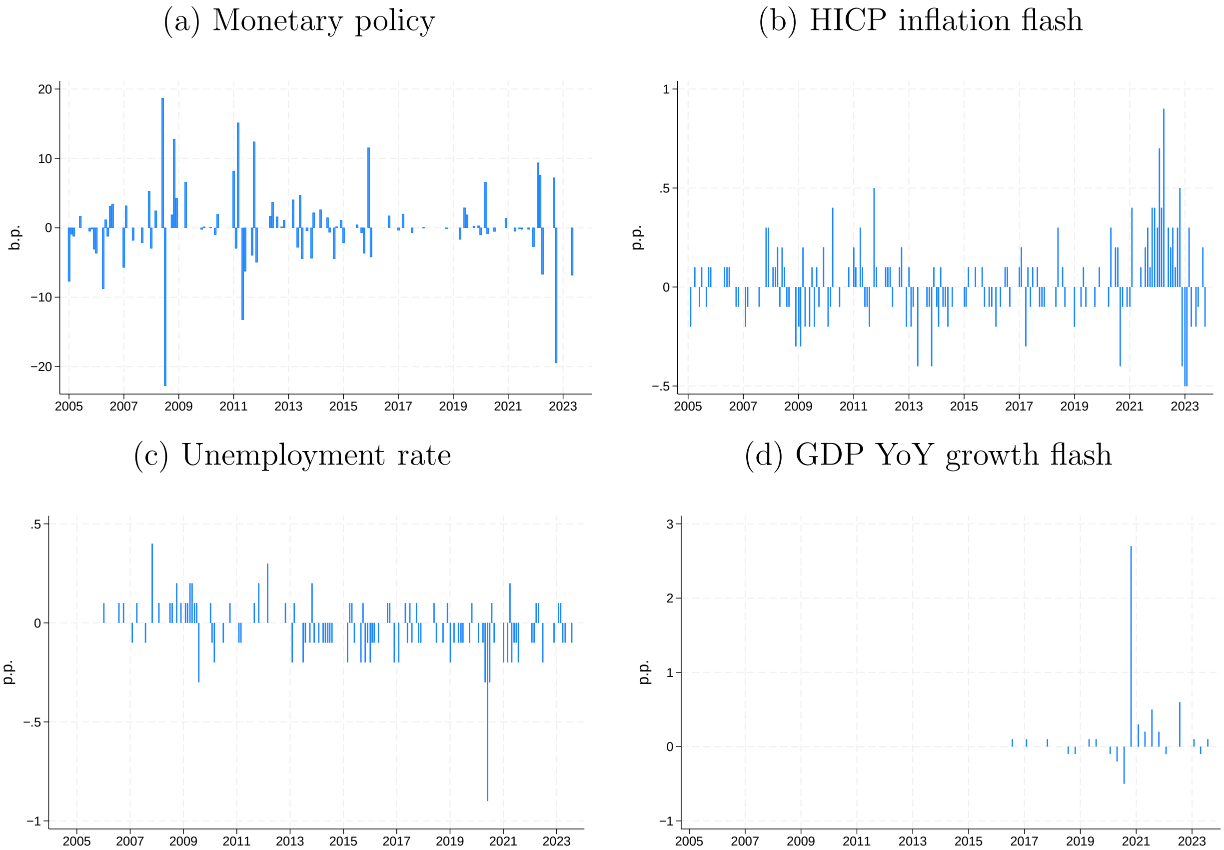
Appendix

A.1 Macroeconomic news release surprises

Panel (a) of Figure A1 depicts monetary policy surprises (1-year OIS rate changes) from the Euro Area Monetary Policy Event-Study Database (EA-MPD) Altavilla et al. (2019). The "Monetary-Event Window" captures the change in the median quote from 13:25-13:35 before the press release to 15:40-15:50 after it. These interest-rate surprises may not solely be monetary policy shocks, as central bank announcements can also convey information about the economy to market participants. In Figure A1, I employ a decomposition of monetary policy surprises by Jarociński and Karadi (2020).

Panels (b)-(d) of Figure A1 respectively present HICP inflation flash, Unemployment rate, and GDP yoy growth surprises. These are computed as the deviation between actual values and the median of analysts' forecasts obtained from the Refinitiv poll conducted a few days before the data release.

Figure A1: Euro-area macro surprises



Sources: Refinitiv and Euro Area Monetary Policy Database.
Notes: The time series of high-frequency ECB monetary policy shocks are displayed over the actual estimation period and are standardized such that a positive unit value represents a one-standard deviation contractionary shock.

A.2 Robustness of ILS rate volatility estimations

The short-term ILS rate volatility around HICP inflation flash releases may be influenced by concurrent information releases within the 3-day window or base effects. To address this, I estimate (3), excluding the two concurrent dates of ECB Governing Council policy decisions (August 31, 2006, and April 30, 2020), and adding a dummy equal to one on the first working day of each month. This dummy controls for the base effect in ILS contracts, as these contracts link with inflation three months before their start. Additionally, I incorporate surprise announcements in unemployment rates and annual preliminary GDP growth rates on the HICP inflation flash release day among eq. (3) controls. If there is no surprise on a given day, the variable is set to zero.

Table A1 reports the estimated k-coefficients in eq. (3). The change in ILS rates on HICP inflation flash release days (k_0) for short-term maturities exceeds the average values (Constant). Moreover, the table confirms that inflation expectation volatility tends to be lower in the subsequent days. Overall, controlling for the coincidence of other information during the HICP inflation flash release dates does not affect the main result in Table 1.

Table A1: ILS rates around HICP inflation flash controlling for other macro news releases

VARIABLES	(1) $\widetilde{\pi}^{1y}$	(2) $\widetilde{\pi}^{2y}$	(3) $\widetilde{\pi}^{3y}$	(4) $\widetilde{\pi}^{4y}$	(5) $\widetilde{\pi}^{5y}$	(6) $\widetilde{\pi}^{10y}$
k_{-4}	-0.003	-0.004	-0.000	-0.001	-0.002	0.003
	(0.007)	(0.004)	(0.003)	(0.003)	(0.003)	(0.002)
k_{-2}	0.008	0.007	0.004	0.004	0.002	0.001
	(0.009)	(0.005)	(0.004)	(0.004)	(0.003)	(0.002)
k_0	0.073***	0.029***	0.015**	0.007	0.008**	0.004
	(0.016)	(0.008)	(0.006)	(0.005)	(0.004)	(0.003)
k_{+2}	-0.004	0.001	-0.003	-0.000	-0.001	0.005*
	(0.007)	(0.005)	(0.004)	(0.004)	(0.003)	(0.003)
k_{+4}	-0.007	-0.011***	-0.008***	-0.005*	-0.004*	-0.001
	(0.007)	(0.004)	(0.003)	(0.003)	(0.003)	(0.002)
Constant	0.015***	0.011***	0.009***	0.008***	0.007***	0.005***
	(0.003)	(0.002)	(0.001)	(0.001)	(0.001)	(0.001)
Observations	3,754	3,833	3,927	3,913	3,834	3,508
R^2	0.360	0.344	0.355	0.355	0.363	0.339
day-of-week FE	Y	Y	Y	Y	Y	Y
week-of-year FE	Y	Y	Y	Y	Y	Y
month FE	Y	Y	Y	Y	Y	Y
macro-news control	Y	Y	Y	Y	Y	Y
base FE	Y	Y	Y	Y	Y	Y

Notes: Huber-White robust standard errors in parentheses *** p<0.01, ** p<0.05, * p<0.1. This table excludes variations that occur during a Governing Council monetary policy decision meeting. Additionally, it adds unemployment rate and annual GDP growth rate surprises among controls.

A.3 Robustness of regression (2)

In this analysis, the inflation gap is defined as the difference between the 1-year, 1-year ahead ILS rates and Consensus forecasters' expectations on the 5-year, 5-year ahead inflation rate. Additionally, the long-term interest rate is considered at the 1-year, 9-year ahead horizon. Results persist across these alternative specifications.

Table A2: 1-year forward Taylor rules at the 1-year ahead horizon with macro news and other long-run variables

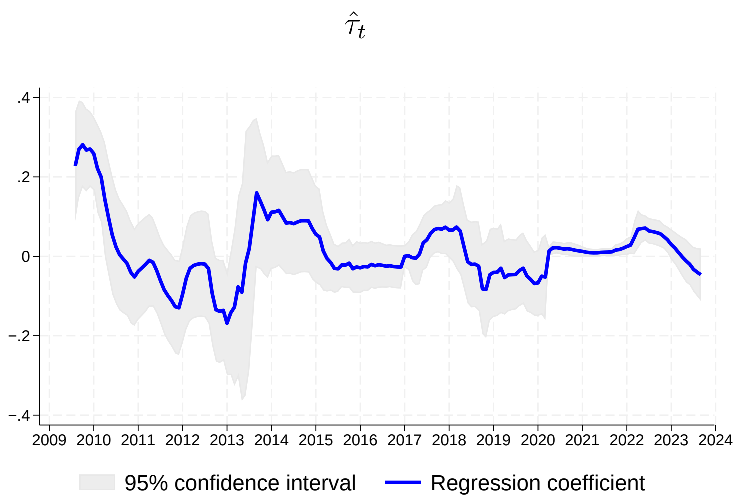
VARIABLES	(1) 1Y1Y OLS	(2) 1Y1Y 2SLS	(3) 1Y1Y 2SLS	(4) 1Y1Y 2SLS
$\pi_{1y1y} - \pi_{5y5y}$	0.304***	0.461***	0.471***	0.443***
	(0.100)	(0.117)	(0.118)	(0.109)
i_{1y9y}	0.274***	0.249***	0.246***	0.222***
	(0.044)	(0.043)	(0.043)	(0.043)
i_{1y}	0.716***	0.688***	0.688***	0.713***
	(0.040)	(0.035)	(0.035)	(0.033)
$g_{1y1y} - g_{5y5y}$	0.086**	0.072*	0.067	0.058
	(0.033)	(0.041)	(0.042)	(0.040)
GDP growth surprise			0.005	-0.029
			(0.077)	(0.075)
Unemployment surprise			-0.189	-0.291
			(0.215)	(0.206)
VSTOXX				0.008**
				(0.003)
Observations	217	217	217	217
month FE	N	N	N	N
KP F-test		37.72	37.03	41.67

Notes: Bartlett/Newey-West standard errors with five lags are displayed in parentheses. *** p<0.01, ** p<0.05, and * p<0.1. The sample period spans from August 31, 2005, to September 29, 2023. The inflation gap is instrumented using forecaster surprises on the HICP inflation flash release, defined as the difference between the flash estimate of the euro-area headline inflation and the median forecasts in the Refinitiv survey.

A.4 The output-gap coefficient

Figure A2 displays the estimated output gap coefficient ($\hat{\tau}_t$) for 1-year, 1-year ahead OIS rates in equation (2). Confidence intervals are estimated using Huber-White robust standard errors. Excluding certain episodes, estimated output gap coefficients ($\hat{\tau}_t$) mainly hovered near zero in the euro area.

Figure A2: 1Y1Y output gap coefficients



Notes: $\hat{\tau}_t$ in (2) is estimated through a 4-year rolling OLS on HICP inflation flash release dates, employing Huber-White robust standard errors.

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