EUROPEAN CENTRAL BANK

EUROSYSTEM

Working Paper Series

Fédéric Holm-Hadulla, Sebastiaan Pool

Interest rate control and the transmission of monetary policy

No 3048

Abstract

We study how short-term interest rate volatility affects the transmission of monetary policy. To identify exogenous changes in volatility, we exploit the pronounced heteroskedasticity visible in the time-series of euro area short-term rates over the past two and a half decades. Interacting the exogenous variation in volatility with high-frequency-identified monetary policy shocks, we find that increases in volatility dampen the effects of monetary policy on output and prices. This dampening effect is visible already at the earlier stages of transmission, including in the pricing and volume of bank lending.

JEL: E44; E52; E58

Keywords: Interest Rate Volatility, Monetary Policy Transmission, Monetary Policy Implementation

Non-technical summary

The day-to-day discourse on monetary policy typically centers on the policy stance decisions central banks take at regular and fairly frequent intervals. By contrast, the modalities by which central banks implement monetary policy tend to attract much less attention and mostly operate behind the scenes. But recent years have seen an exception to this rule. In particular, major shifts in financial structures and policy conduct have prompted many central banks to reopen some of the subtle but often consequential questions around how to best design their 'operational framework' for implementing monetary policy.

In the current paper, we link the two domains – stance and implementation – and ask how they interact in shaping the transmission of monetary policy to the economy. A core function of the implementation framework is to steer short-term money market rates in line with the desired monetary policy stance – a function usually referred to as interest rate control. But, since these short-term rates are determined in a market setting, they may fluctuate also absent changes in the stance, for instance owing to temporary shifts in commercial banks' demand for liquidity. Central banks, in turn, can influence the scope for such non-policy induced volatility, for example, in setting the differential between the rate at which they lend reserves to commercial banks and the rate at which they accept commercial bank deposits.

Our paper highlights an important reason for keeping the scope for short-rate volatility within reasonable bounds. In particular, we present empirical evidence showing that higher short-rate volatility significantly dampens the transmission of monetary policy to key macroeconomic variables, including economic activity and prices. Moreover, the dampening effect of short-rate volatility shows up already at earlier stages of monetary policy transmission in that it also renders bank lending rates and volumes markedly less responsive to monetary policy.

These findings connect to a long-standing literature on how volatility affects the response of economic agents to shocks. In particular, the dampening effect of short-rate volatility on monetary policy transmission echoes the classic Dixit (1991) result that uncertainty around relevant state variables creates an 'option value of the status quo'. As a result, it may be optimal for economic agents not to adjust in response to aggregate shocks, so long as these shocks remain within certain bounds. A core condition for this type of inertia to apply is that economic agents incur costs in adjusting the variables over which they optimize. Consistent with our empirical finding that the dampening effect of volatility arises already at the bank-lending stage of transmission, the respective literature indeed emphasizes adjustment costs as a key feature of lending behavior.

1 Introduction

The day-to-day discourse on monetary policy typically centers on the policy stance decisions central banks take at regular and fairly frequent intervals. By contrast, the modalities by which central banks implement monetary policy tend to attract much less attention and mostly operate behind the scenes. But recent years have seen an exception to this rule. In particular, major shifts in financial structures and policy conduct have prompted many central banks to reopen some of the subtle but often consequential questions around how to best design their 'operational framework' for implementing monetary policy.

In the current paper, we link the two domains – stance and implementation – and ask how they interact in shaping the transmission of monetary policy to the economy. A core function of the implementation framework is to steer short-term money market rates in line with the desired monetary policy stance – a function usually referred to as interest rate control (Bindseil, 2016). But, since these short-term rates are determined in a market setting, they may fluctuate also absent changes in the stance, for instance owing to temporary shifts in commercial banks' demand for liquidity. Central banks, in turn, can influence the scope for such non-policy induced volatility via the parameters of the operational framework, for instance in setting the differential between the rate at which they lend reserves to commercial banks and the rate at which they accept commercial bank deposits. Narrowing this differential, *ceteris paribus*, reduces the leeway for money market rates to fluctuate – albeit at the cost of crowding out money market activity in the very short-term segment.

Our paper highlights an important reason for keeping the scope for short-rate volatility within reasonable bounds. In particular, we present empirical evidence showing that higher short-rate volatility significantly dampens the transmission of monetary policy to key macroeconomic variables, including economic activity and prices.

To identify exogenous variation in the degree of interest rate control exerted by the ECB, we exploit a set of regime shifts in the implementation of monetary policy that have taken place since the introduction of the euro. These regime shifts partly resulted from technical adjustments to make monetary policy implementation more efficient and partly came as a byproduct of the ECB expanding its monetary policy toolkit. What they have in common is that they generated stark differences in how short-rate volatility has fluctuated over time. This, in turn, is a key enabling factor for the heteroskedaticity-based identification strategy proposed by Rigobon (2003), which we hence apply in the current paper. In a second step, we then incorporate the instrumented variation in short-rate volatility into a standard local projections models (Jordà, 2005) and interact it with high-frequency-identified monetary policy shocks, following Jarociński and Karadi (2020) and building on the Euro Area Monetary Policy Event-Study Database of Altavilla et al. (2019a).

The estimates show that bank credit, real activity, and inflation become substantially less responsive to monetary policy as short-rate volatility rises. At the trough of the GDP response, the impact of a given monetary policy shock declines by around half (in absolute terms) for each standard-deviation increase in short-rate volatility. The response of the GDP deflator to a monetary policy shock is cut by a similar degree if volatility rises by one standard deviation. Moreover, the dampening effect of short-rate volatility shows up already at earlier stages of monetary policy transmission in that it also renders bank lending rates and volumes markedly less responsive to monetary policy. We find these results to be robust to extensions of our baseline model to account for potential confounders, such as financial stress and the size of the ECB's balance sheet, and to variations in the design of the heteroskedasticity-based identification strategy.

The weakening of transmission implies that monetary policy has to respond more forcefully to a given shock driving inflation away from target and may thus lead policy rates to hit their lower bounds more often. Moreover, we observe that volatility itself is heteroskedastic: in periods with high *average* volatility, it also tends to vary a lot from month to month. This further complicates monetary policy: central banks cannot simply adjust to a regime of high volatility and weak transmission. Instead, at each point in time, it is uncertain which volatility/transmission regime prevails. This, in turn, raises the risk of over- or underreactions to shocks.

Related literature. Our results connect to a long-standing literature on how volatility affects the response of economic agents to shocks. In particular, the dampening effect of short-rate volatility on monetary policy transmission echoes the classic Dixit (1991) result that uncertainty around relevant state variables creates an 'option value of the status quo'. As a result, it may be optimal for economic agents not to adjust in response to aggregate shocks, so long as these shocks remain within certain bounds.[1] A core condition for this type of inertia to apply is that economic agents incur costs in adjusting the variables over which they optimize. Consistent with our empirical finding that the dampening effect of volatility arises already at the bank-lending stage of transmission, the respective literature indeed emphasizes adjustment costs as a key feature of lending behavior (see Gerali et al. (2010) and the references cited therein).[2]

Intuitively, banks for example would reflect a surprise change in short-term money market rates – over and above what was priced in prior to the respective policy announcement – into their lending decisions. But in the presence of high money market volatility induced by factors other than monetary policy, they partly discount any observed change in the relevant money market rates for lack of certainty on whether it constitutes a genuine shift or just transitory fluctuations for

[1]As discussed in Hansen (1999) and Vavra (2014), higher uncertainty may not only widen the optimal 'range of inertia' but also increase the amplitude of swings in the relevant state variable. As a consequence, it is a priori ambiguous whether higher uncertainty raises or lowers the frequency with which realizations in the state variable break out of the range of inertia and therefore whether it raises or lowers the optimal frequency of changes in the respective choice variable. At the same time, our analysis focuses on the propensity *of a given shock size* to trigger adjustments in other variables, which should have an unambiguous, inverse, relation to the prevailing degree of volatility in the Dixit (1991) logic.

[2]From a broader perspective, our findings also have a relevant parallel to those in Ehrmann and Fratzscher (2007), who show empirically that more predictable monetary policy communication triggers a stronger market response to changes in such communication. In particular, higher volatility in our setting means that the stance signal resulting from communication events becomes less of a predictor for actual short-term money market dynamics. Woodford (2005) provides a theoretical rationale for why higher predictability may support central banks in meeting their macroeconomic stabilization objectives.

which it would not be worth incurring the fixed cost of changing the loan conditions of their (prospective) customers. The resultant dampening then also feeds through to the response of output and prices.

Second, our paper contributes to the literature on how the choice of instruments to implement monetary policy shapes its broader economic effects (see Poole (1970) for an early contribution and Bindseil (2014) for a critical review). This issue has acquired renewed prominence in recent years, as central banks have started to reverse part of the massive balance sheet expansions that took place over the previous low-inflation period. This process has given rise to debate on both: the likely impacts of balance sheet normalization (Acharya et al., 2023; Diamond et al., 2024; Altavilla et al., 2025); and the balance sheet size that central banks should aim for in a new steady state (Arce et al., 2020; Lane, 2023; Schnabel, 2023). Balance sheet size, in turn, is a major determinant of short-term interest rate volatility (Afonso et al., 2022; Lopez-Salido and Vissing-Jorgensen, 2023), implying that our findings may also contribute to assessing the trade-offs involved in these choices.

Finally, the paper adds a new angle to the broader literature on how the transmission of monetary policy is influenced by the cyclical or structural features of an economy, including for instance: the state of its business, credit and interest rate cycles (Peersman and Smets, 2005; Tenreyro and Thwaites, 2016; Jordà et al., 2020; Alpanda et al., 2021); the level and composition of private sector debt (Becker and Ivashina, 2014; Alpanda and Zubairy, 2019; Holm-Hadulla and Thürwächter, 2021; Corsetti et al., 2022); or the degree of household inequality (see McKay and Wolf (2023) for an overview).

The remainder of this paper is organized as follows. Section 2 describes the empirical model, data, and identification strategy. Section 3 presents the results, Section 4 examines the robustness of our main findings, and Section 5 concludes.

2 Empirical methodology

2.1 Model

The empirical estimates rely on a standard local projections model using high-frequency financial market data to identify exogenous variation in monetary policy interest rates.[3] We derive this exogenous variation from changes in the 1-month OIS rate around ECB policy communication events, using the Euro Area Monetary Policy Event-Study Database, EA-MPD (Altavilla et al., 2019a). Following Jarociński and Karadi (2020), we designate as genuine monetary policy shocks only those events at which the 1-month OIS rate moved in the opposite direction than stock prices (measured by the Euro Stoxx 50 index).[4]

We then interact the monetary policy shocks with the volatility in overnight unsecured money market rates measured over the respective ECB reserve maintenance period preceding the shock (and instrumented as described in Section 2.3). Maintenance periods essentially mark the interval between two ECB Governing Council meetings; but they start a few days after a meeting, when a ECB policy rate decision is implemented, and last until the next maintenance period begins. Governing Council meetings took place (with a few exceptions) at a monthly frequency through 2014 and have moved to a six-weekly frequency since then. By considering only fluctuations within a given maintenance period, the analysis narrows in on the concept of rate controllability, defined as overnight rates remaining broadly stable in the ab-

[3]The identification strategy is based on the assumption that observed changes in market rates over narrow time windows around central bank communication events are unlikely to reflect news on the economy and instead can be plausibly characterized as an exogenous innovation in monetary policy. See, for instance, Barakchian and Crowe (2013), Cochrane and Piazzesi (2002), Gertler and Karadi (2015), Gürkaynak et al. (2005), Kuttner (2001), Nakamura and Steinsson (2018) and Piazzesi and Swanson (2008) for the U.S. and Altavilla et al. (2019a), Andrade and Ferroni (2021), Auer et al. (2021) and Jarociński and Karadi (2020) for the euro area.

[4]This identification based on cross-asset correlations accounts for the possibility that central bank communication provides signals not only on monetary policy but also on the state and prospects of the economy. The latter type of signal may blur the genuine impact of monetary policy on key outcome variables.

sence of changes in the policy stance. Vice versa, it allows us to avoid confounding effects related to the volatility arising from changes in the policy stance itself, as the latter take place only between maintenance periods. While policy-induced volatility may also affect monetary policy transmission (Tillmann, 2020), our interest is in the 'noise' arising in the implementation of a given policy stance, not in the uncertainty surrounding the stance signal itself.[5]

Formally, the empirical specification consists of the following local projections model (Jordà, 2005):

$$Y_{t+h} = \beta_{0,h} + \left(\beta_{1,h} + \beta_{2,h}\hat{\sigma}_{t-1}\right) S_t + \beta_{3,h}\hat{\sigma}_{t-1} + \gamma_h \sum_{p=1}^{2} X_{t-p} + \varepsilon_{t+h} \tag{1}$$

where t, and h denote the month and impulse response function (IRF) horizon, respectively, and p denotes the number of lags included in the set of control variables X_{t-p}. The vector of dependent variables Y_{t+h} comprises the 3-month OIS rate, bank lending rates to households and firms along with the respective volumes, as well as real GDP and the GDP deflator as the key macro outcome variables. The explanatory variables of main interest are the monetary policy shock S_t and its interaction with the instrumented standard deviation of the unsecured overnight rate, $\hat{\sigma}_{t-1}$.

The vector of controls X_{t-p} includes the first two lags of each dependent variable, as well as the monetary policy shock, the instrumented standard deviation of the unsecured overnight rate and their interaction. These lags help purge the high-frequency surprises from potential serial correlation and thereby strengthen

[5]In addition, by calculating volatility over the preceding maintenance period, the analysis conditions transmission on the environment prevailing in the run-up to the respective monetary policy shock. This step further helps in ensuring that $\hat{\sigma}$ primarily captures pre-existing volatility conditions, rather than some endogenous feedback from the monetary policy shock itself. Since March 2004, each maintenance period has started on the Wednesday in the week after the monetary policy press conference, held on Thursdays. Hence, short-rate volatility measured over the previous maintenance period also includes four business days after the monetary policy press conference. However, the results presented in Section 3 are robust to interacting short-rate volatility in the penultimate maintenance period leading up to the press conference ($\hat{\sigma}_{t-2}$) with the monetary policy shock S_t. The results are available upon request.

identification (Ramey, 2016). Further, X_{t-p} includes the the EUR-USD exchange rate and global commodity prices to capture changes in the external environment.

As the final element of X_{t-p}, we control for two sets of factors that are likely to co-move with both, short-rate volatility and macroeconomic conditions and are therefore key to sharpening the identification strategy. The first is financial market stress, which we measure by the 'composite indicator of systemic stress' (CISS, developed by Hollo et al. (2012)). The second is the presence of non-standard monetary policy, which we proxy with shocks to 5-year and and 10-year OIS rates, also sourced from the EA-MPD (as part of the robustness checks, we consider the excess liquidity generated by the ECB's credit- and quantitative-easing programs as an alternative proxy for non-standard monetary policy). Importantly, we do not only include these variables as separate controls, but also their interactions with the monetary shock S_t. This helps us avoid that the estimated conditioning impact of short-rate volatility on transmission is confounded by systematic differences in how monetary policy transmits in crisis times or when standard policy-rate shocks coexist with non-standard monetary policy.

2.2 Data

The model is estimated on monthly data spanning from January 2002 to February 2020. Like Altavilla et al. (2019a), we exclude the first two years after the introduction of the euro because the financial market data used to construct the monetary policy shock is very noisy over this period. Additionally, we omit the period following the outbreak of COVID-19, given its unusual impact on the economy. Monetary policy shocks are constructed by converting high-frequency changes in market interest rates around ECB press conferences into a monthly frequency. In months without a monetary policy event, we set the shock to zero; in months with more than one event we sum up the observations. In total, the sample contains 218 press conferences and the corresponding reserve maintenance periods.

Short-rate volatility is measured by the standard deviation of unsecured overnight money market rates, measured at daily frequency. We base these calculations on the Euro Overnight Index Average (EONIA) rate until October 2019 and the euro short-term rate (€STR) thereafter, in line with the transition from the former to the latter on regulatory grounds.

GDP and the deflator are linearly interpolated from quarterly to monthly frequency so as to match the frequency of the monetary policy press conferences. Interest rates, the short-rate standard deviation, and the monetary policy shock are expressed in percentages, while all other variables enter in log-levels. For ease of interpretation, all variables interacted with the monetary policy shock are expressed as z-scores, with zero mean and unit variance. Table 4 in Annex 6 list the descriptive statistics and sources of the variables used in the analysis.

2.3 Identification of exogenous changes in rate volatility

2.3.1 Strategy

A challenge in estimating equation 1 is that the observed variation in short-rate volatility may not be exogenous to the other variables in the system. For instance, short-rate volatility may not just influence how monetary policy transmits to macroeconomic variables, such as output and prices, but macroeconomic conditions may also feed back into the degree of short-rate volatility observed over a certain period. Moreover, notwithstanding the rich set of controls in equation 1, we cannot rule out the existence of unobserved confounders that correlate with short-rate volatility and the main outcome variables.

To address these potential sources of endogeneity, we exploit the heteroskedasticity-based identification strategy proposed by Rigobon (2003). The basic intuition is that, in a system of simultaneous equations, distinct shifts in the volatility of one type of shock can be used to recover the structural parameters of other equations. In the canonical example of a supply and demand system, a shift from low to high volatility in the supply shocks, for instance, would imply that the observed outcomes become more clustered around the demand curve. This shifts in observed outcomes can be exploited to identify the location and shape of the demand curve.

In applying this approach to equation 1, we benefit from the fact that the ECB's modalities for implementing monetary policy have undergone several regime changes since the start of the euro; and these regime changes, mostly as a side effect, have generated stark differences in how short-rate volatility has fluctuated from period to period – a key enabling factor for the heteroskedaticity-based identification strategy. In particular, we define five unique regimes, which we summarize here and describe in more detail in Section 2.3.2 (see Table 1 for an overview and Figure 1 for the history of short-rate volatility across regimes; readers familiar with the history of ECB monetary policy implementation may choose to go straight to Section 2.3.3):

- the first change pertained to a technical adjustment that removed the scope for speculative bidding in ECB operations; it led to a strong decline not only in average short-rate volatility within each maintenance period (from 12 bps to 7 bps), but also in its variance across maintenance periods (from 0.62 bps to 0.22 bps);

- the second regime change took place at the height of the Global Financial Crisis and was followed by a renewed rise in volatility, to 13 bps on average, driven by the prevailing financial turmoil and notwithstanding the ECB's move to a more elastic liquidity provision in its refinancing operations; along with the average, the month-on-month variance of short-rate volatility spiked up again, to 0.59 bps;

- the third regime change brought about a steep drop in volatility to an average of 3 bps, driven by the introduction of very long-term refinancing operations which generated ample liquidity conditions and thereby insulated unsecured overnight rates from short-term shifts in the supply and demand for reserves; this also led the variance of short-rate volatility to collapse to 0.10 bps;

- the fourth regime change, coinciding with the ECB's introduction of quantitative easing, further suppressed short-rate volatility, to an average of 1 bps, owing to the abundant supply of central bank reserves; with short-rate volatility essentially flat around zero, also its variance across maintenance periods edged down further to 0.01 bps.

Taken together, these regime shifts generate pronounced heteroskedasticity in short-rate volatility over the sample period, in that the degree to which σ fluctuates across maintenance periods varies markedly from regime to regime. Moreover, this variation is substantially more pronounced than that recorded for the volatility of the key macro variables in the model (see Table 5 in the Annex). Accordingly, exploiting heteroskedasticity in short-rate volatility emerges as a plausible strategy to identify the system in equation 1.

Table 1: Regimes definition and description

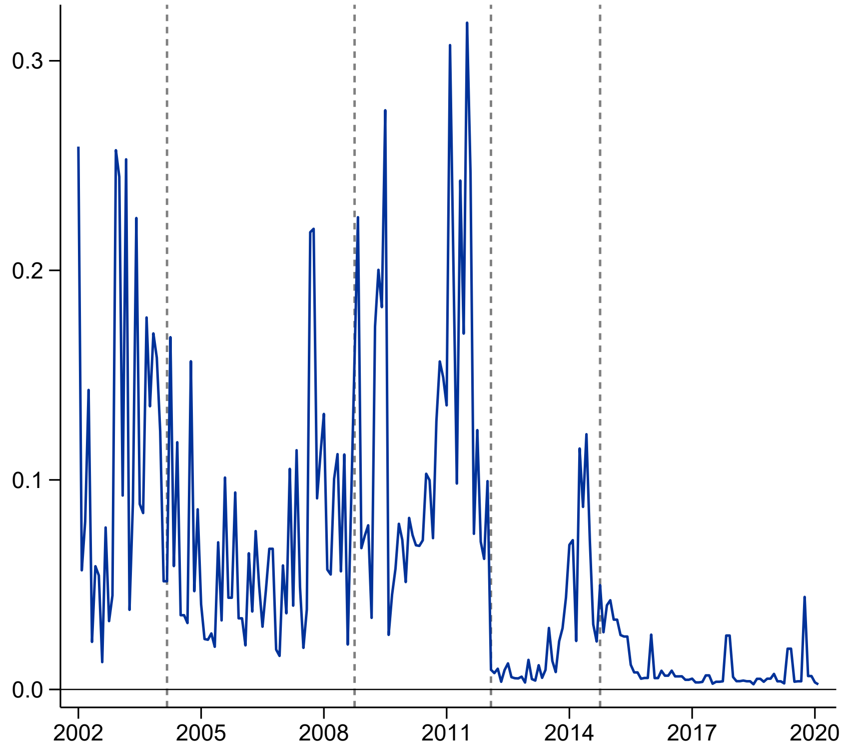
Definition	Description	Start	End	Variance of σ_t
(1) Pre-tech. change	High volatility regime before technical change in monetary policy implementation framework	Jan-99	Feb-04	0.62
(2) Pre-FRFA	Moderate volatility regime before outbreak of global financial crisis	Mar-04	Sep-08	0.22
(3) Post-FRFA	High volatility regime, partially dampened by FRFA in tender procedures	Oct-08	Jan-12	0.59
(4) Ample reserves	Mild volatility regime in an abundant reserves system	Feb-12	Feb-15	0.10
(5) Abundant reserves	Low volatility regime following the start of the outright asset purchase programme	Mar-15	Feb-20	0.01

Note: variance of σ_t expressed in basis points (bps).

2.3.2 Regimes

The first regime change was announced in January 2003 and implemented in March 2004. At the time, the ECB decided on two adjustments to the Eurosystem's framework for monetary policy implementation. First, the timing of the reserve maintenance periods was adapted to consistently start on the settlement day of the main refinancing operation (MRO) following the Governing Council meeting at which the monetary policy stance was decided. Second, the maturity of the MROs was shortened from two weeks to one week. Together, these two changes aimed to stabilize conditions in euro area money markets by removing the scope for speculative bidding

Figure 1: Short-rate volatility across the volatility regimes (y-axis: percentage points; x-axis: years)



Notes: the standard deviation of the short-rate (€STR or EONIA) is computed over the maintenance period. The dotted lines mark the 4 regime changes on January 2002, October 2008, February 2012 and March 2015.

throughout the maintenance period when policy rate changes were expected.[6]

The second regime change occurred in October 2008, at the peak of the Global Financial Crisis, when the ECB switched from a variable-rate tender procedure with a predetermined allotment amount for its refinancing operations to a fixed-rate tender procedure in which banks' bids were fully allotted. This adjustment in the ECB's tender procedure meant that the availability of liquidity for banks moved from scarce and largely at the ECB's discretion to potentially unlimited and more directly determined by the banking system's demand for reserves. The more elastic availability of liquidity reduced the probability that banks would end up short of liquidity at the end of a business day and be compelled to resort to the euro area money markets, thereby preventing liquidity conditions from deteriorating and triggering a potential credit crunch. At the same time, this shift in the modalities for providing central bank reserves was dominated by the heightened banking sector stress after the collapse of Lehman Brothers. The latter, in turn, led liquidity conditions in euro area money markets to deteriorate rapidly and hence put upward pressure on short-rate volatility.[7]

The third regime change was initiated in December 2011 when the ECB announced the first series of long-term refinancing operations (LTROs), spanning a period of up to three years. Prior to this, the ECB offered only 1-week and 3-month lending operations besides some ad-hoc 6- and 12-month operations. However, the intensification of stress in sovereign debt markets in the second half of 2011, coupled with high levels of uncertainty, increasingly hampered euro area banks' access to market-based funding. The introduction of operations with a term of up to three years, at the same pricing as the standard shorter-term operations, led to a surge in banks' reserve holdings with the ECB, adding hundreds of billions in liquidity.[8]

---

[6]See the ECB press release entitled "Changes to the Eurosystem's operational framework for monetary policy".

[7]The stress in the financial system could be an important confounding factor, which motivates some of the robustness checks in Section 4.

[8]See Box 3 of the ECB's Economic and Monetary Developments Report titled Impact of the two three-year longer-term refinancing operations and Darracq-Paries and Santis (2015) for more information.

The first operation was settled at the end of December 2011 and the second operation at the end of February 2012. In February, the ECB also reduced the minimum reserve requirement ratio from 2% to 1%, reducing the reserve requirement by approximately €104 billion. Together, these measures substantially improved liquidity conditions in euro area money markets, which was reflected in a steep decline in short-rate volatility.[9]

The final regime change occurred in March 2015, when the ECB started its first large-scale asset purchase program to further ease the monetary policy stance in the proximity of the effective lower bound on its key interest rates. While motivated by stance considerations, the adoption of large-scale purchases marked a *de facto* qualitative shift in policy conduct, from a demand-driven system (in which banks determine the amount of liquidity they borrow from the ECB), to a supply-driven system (in which the central bank directly injects liquidity by buying financial assets with reserves). The large volume of central bank reserves that banks received in exchange for (intermediating) asset sales to the ECB in turn created abundant liquidity conditions. This insulated overnight rates even further from shorter-term supply or demand shocks in money markets and consequently bringing short-rate volatility near zero on a sustained basis in this last regime.

2.3.3 Implementation and validity of the instrument

As regards implementation of the heteroskedasticity-based identification strategy proposed by Rigobon (2003), we follow the steps proposed by Lewbel (2012). We start by estimating a separate first-stage regression for each of the endogenous variables:

$$\sigma_{t-1} = \alpha_0^1 + \alpha_1^1 \sum_{p=1}^{2} X_{t-p} + \varepsilon_t^1, \tag{2}$$

$$\sigma_{t-1} S_t = \alpha_0^2 + \alpha_1^2 \sum_{p=1}^{2} X_{t-p} + \varepsilon_t^2. \tag{3}$$

[9] As visible from Figure 1, short-rate volatility collapsed to levels slightly above zero at the beginning of this regime. Over time, however, banks gradually repaid their LTRO borrowings, causing short-rate volatility to increase again, while remaining low by historical standards.

Next, we compute the residual of both regressions $\hat{\varepsilon}_t^1 = \sigma_{t-1} - \hat{\alpha}_0^1 - \hat{\alpha}_1^1 \sum_{p=1}^{2} X_{t-p}$ and $\hat{\varepsilon}_t^2 = \sigma_{t-1} S_t - \hat{\alpha}_0^2 - \hat{\alpha}_1^2 \sum_{p=1}^{2} X_{t-p}$ and define Z_t as a vector containing four dummy variables each identifying one of the regime changes. We then compute: $(Z_t - \bar{Z})\hat{\varepsilon}_t^1$ and $(Z_t - \bar{Z})\hat{\varepsilon}_t^2$, where $\bar{Z}$ denotes the sample mean of Z_t, and use those terms, together with X_t, as instruments for σ_{t-1} and $\sigma_{t-1} S_t$.

In line with the methodology outlined by Lewbel (2012), we include the regime variables Z_t in our set of controls X_t. Likewise, also any of the variables from equation (1) could be incorporated into the vector Z_t. However, given potential endogeneity between the variables within the control set X_t and the set of macroeconomic outcome variables Y_{t+h}, we take a more prudent approach and, at first, exclude all elements of X_t from Z_t. In our robustness tests, we then relax this assumption and demonstrate that our results remain robust when the full set of control variables is included in Z_t. We estimate our instrumental variable regression using a generalized method of moments (GMM) estimator and Newey-West standard errors.[10]

Table 2: Results for Pagan and Hall (1983) general test for heteroskedasticity

PH-value GDP	p-value GDP	PH-value Deflator	p-value Deflator
1.51	0.82	3.92	0.42

Note: the test statistics are derived from the general heteroskedasticity test applied following an instrumental variable regression as proposed by Pagan and Hall (1983), based on equation (1) for $h = 0$ with regimes z_t serving as instruments. Short-rate volatility and its interaction with monetary policy shocks are instrumented using probabilistic instruments identified through heteroskedasticity, following the methodology outlined by Rigobon (2003).

The key assumptions for the estimator to be valid are $Cov(Z_t, \varepsilon_{t+h} {\varepsilon_t^{1/2}}) = 0$ and $Cov(Z_t, (\varepsilon_t^{1/2})^2) \neq 0$. The first assumption states that the common components shared by the errors of equations (1), (2) and (3) are homoskedastic, *i.e.*: their variance is uncorrelated with the regime variables (Z_t). Based on the Pagan and Hall (1983) test, we do not reject the null hypothesis of homoskedasticity in the error terms ε_t^1 and ε_t^2, neither for GDP, nor the deflator (Table 2). The second

[10] We implement the heteroskedasticity-based estimator for linear regression models containing an endogenous regressor using the Stata package developed by Baum and Lewbel (2019).

assumption states that the squared errors terms $(\varepsilon_t^1)^2$ and $(\varepsilon_t^2)^2$ are correlated with the regime variables Z_t. This is necessary to avoid weak instrument problems, which would result in imprecise estimates with large standard errors. The corresponding test essentially reduces to a conventional heteroskedasticity test on the two first-stage regressions, equations (2) and (3). As also visible in Table 3, we reject the null hypothesis of homoskedasticity at a 1% significance level for both regressions, speaking against weak-instrument problems.

Table 3: Results for Breusch–Pagan/Cook–Weisberg test for heteroskedasticity

Equation (2)		Equation (3)	
χ^2-statistic	p-value	χ^2-statistic	p-value
67.10	0.00	71.32	0.00

Note: the test statistics are based on the Breusch and Pagan (1979) and Cook and Weisberg (1983) test for heteroskedasticity, based on the two first-stage regressions, Equation (2) and Equation (3), with the regimes z_t as exogenous instruments.

3 Results

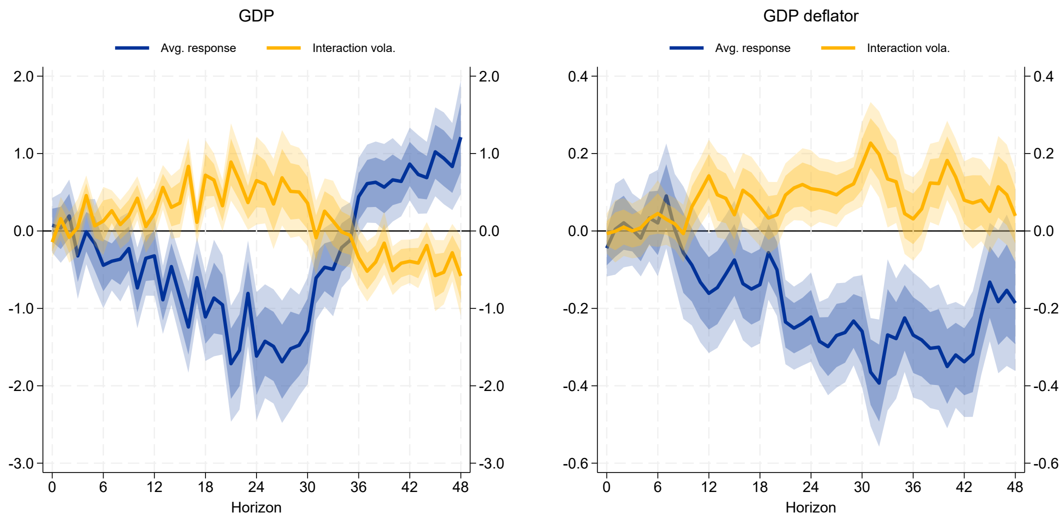
Figure 2 presents the impact of an exogenous policy rate hike on GDP and the deflator at average levels of short-rate volatility (in blue) and how this impact changes with each one-standard-deviation exogenous increase in short-rate volatility as estimated by the interaction-term coefficient (in yellow).[11]

The impulse responses for activity and prices at average volatility levels follow the typical U-shaped pattern found in the related literature. GDP contracts with a lag and reaches a trough after roughly two years. Thereafter, GDP gradually converges back to its initial level. Also in line with typical patterns, the GDP deflator follows suit with some additional delay. The point estimate becomes consistently significant after one-and-a-half years and reaches a trough a year later before slowly returning towards its initial level. The scale of the deflator response is similar to that found in

[11] The IRFs are scaled to a 25bps peak response in the 3-month OIS rate and we allow for a delay of up to 12 months in the 3-month OIS rate's peak response to the monetary policy shock.

recent studies using a similar estimation strategy (see, e.g., Jarociński and Karadi (2020) and Holm-Hadulla and Thürwächter (2021)), whereas the GDP response is somewhat more pronounced.

Figure 2: IRFs following a monetary policy shock conditional on short-rate volatility to GDP and the GDP deflator (y-axis: percentages; x-axis: months)

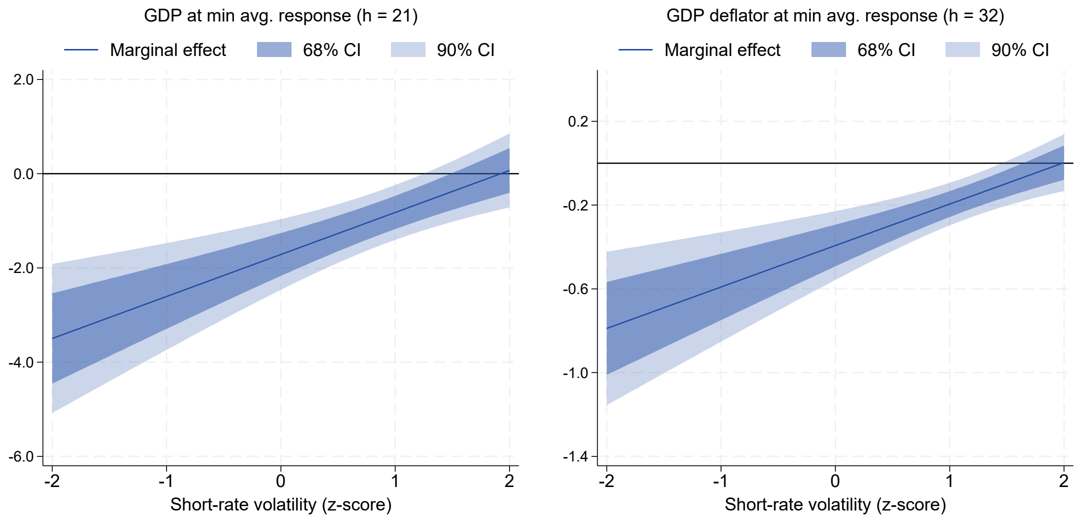


Note: the IRFs are based on equation (1) and normalized to a 25bps peak response in the 3-month OIS rate and a one standard deviation increase in €STR volatility. The dark- (light-)shaded area shows the 90% (68%) confidence interval.

The significant positive coefficients on the interaction term indicate that rate volatility has a dampening effect on the real and nominal effects of the monetary policy shock.[12] This dampening effect is economically relevant: for instance, at the trough of the GDP response, the impact of a given monetary policy shock declines by almost half (in absolute terms) for each one-standard-deviation increase in short-rate volatility (Figure 3). For the deflator, the dampening effect is comparable in that the decline is also cut by almost 50% if volatility rises by one standard deviation. For both variables, a two-standard-deviation increase in short-rate volatility is sufficient to render the response insignificant.

[12]We also estimated our model replacing the GDP deflator by seasonally adjusted headline or core inflation using the harmonised index for consumer prices (HICP) and the corresponding index excluding energy and food prices. Also for these variables, our key finding of volatility dampening the transmission of monetary policy to inflation remains intact.

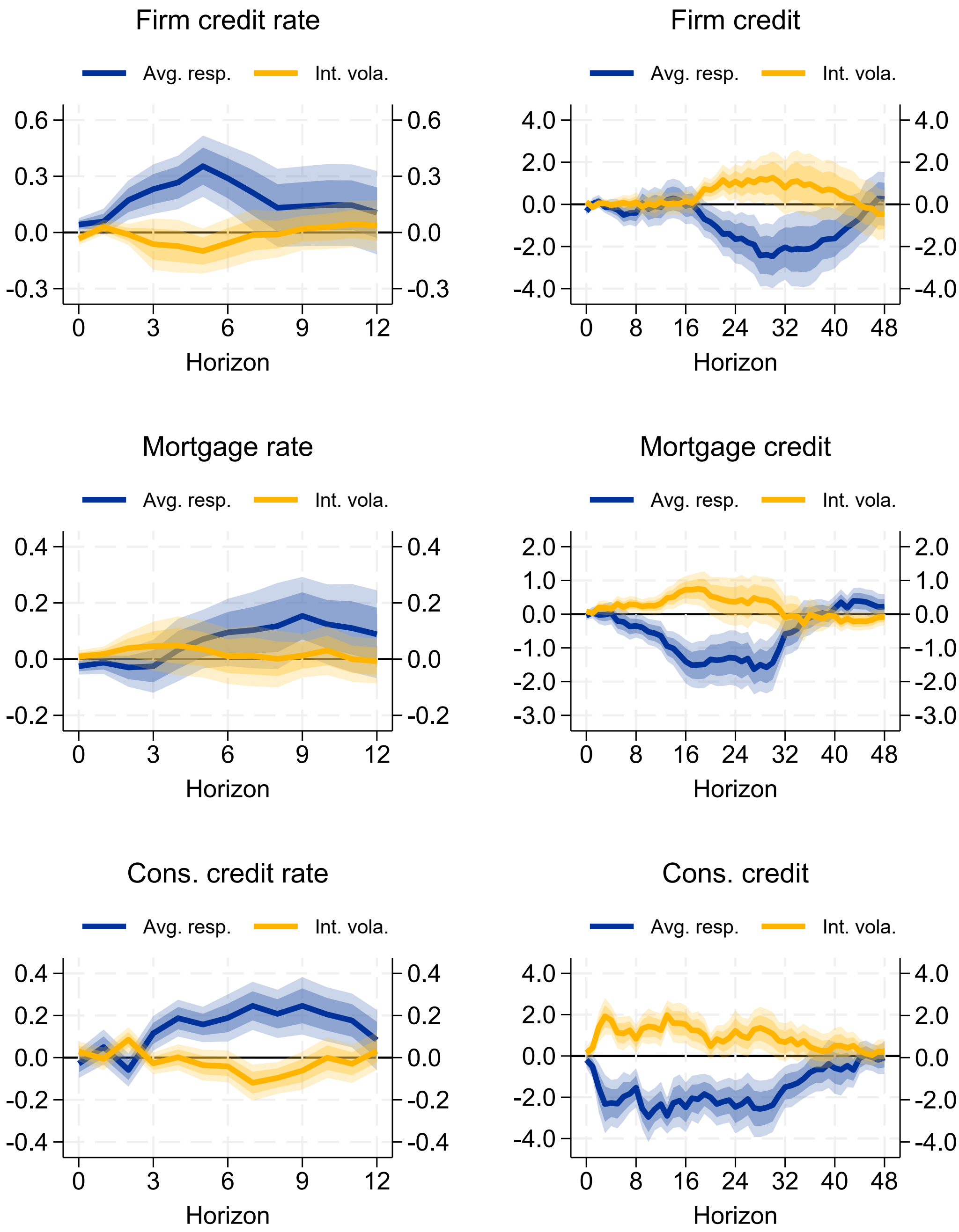
Figure 3: Marginal effect of short-rate volatility on the impact of monetary policy (y-axis: percentages)



Note: the marginal effects are computed at the trough of the average response shown in Figure 2. The dark- (light-)shaded area shows the 90% (68%) confidence interval.

The dampening effect of short-rate volatility shows up already at earlier stages of transmission. In particular, we detect this pattern also in the response of bank intermediation which, in the euro area, is key to linking financial markets to the real economy. The impact of monetary policy on lending rates to firms takes a few months to build up and peaks after around half a year (Figure 4, top left). Again, this impact becomes weaker, the higher is the volatility in overnight rates (although we reject the null of no dampening effect only at a 68% confidence level). For credit to households, the picture is more diverse: the rates on consumer credit display the same qualitative pattern as those on loans to firms, whereas mortgage rates show no clear significant response to the policy shock (neither on average, nor conditional on changes in short-rate volatility). A plausible explanation for the latter result is that mortgage rates in large parts of the euro area are linked to market reference rates of significantly longer maturities, rather than the short tenor underlying the shocks used in the current analysis.

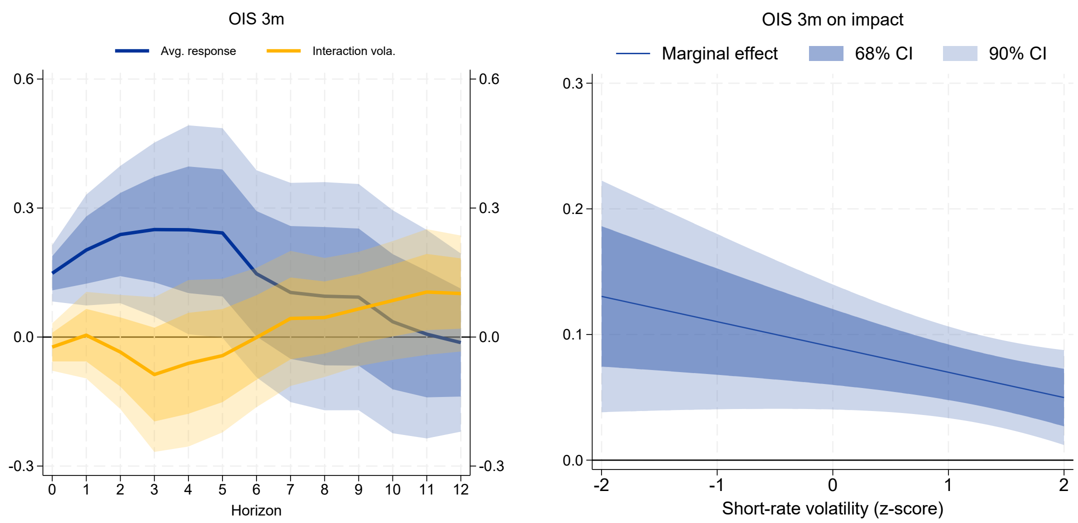
Figure 4: IRFs following a monetary policy shock conditional on bank lending rates and volumes (y-axis: LHS percentage points; RHS percentages; x-axis: months)



Note: the IRFs are based on equation (1) and normalized to a 25bps peak response in the 3-month OIS rate and a one standard deviation increase in €STR volatility. The dark- (light-)shaded area shows the 90% (68%) confidence interval.

Loan volumes, on average, contract in response to the monetary policy tightening (in the case of firms, after a brief expansion in the first few months).[13] Loans reach a trough after around two and a half years, for all categories (including household mortgages and consumer credit as well as loans to firms), and then return to their pre-shock levels. Like for the other variables in the system, the impact is attenuated by higher short-rate volatility, which reduces, for instance, the absolute impact on household mortgage loans by around one-third after 18 months.[14]

Figure 5: Impact of monetary policy shocks and short-rate volatility on the 3-month OIS rate (y-axis: percentage points; x-axis: LHS months)



Note: the IRFs in the left panel are based on equation (1) and normalized to a 25bps peak response in the 3-month OIS rate and a one standard deviation increase in €STR volatility. The marginal effects in the right panel are estimated on impact ($h = 0$). The dark- (light-)shaded area shows the 90% (68%) confidence interval.

At the initial stage of transmission, instead, the patterns are less clear-cut, indi-

[13]A possible interpretation for the brief expansion of corporate loan volumes is that intermediated credit (such as bank loans) offers greater flexibility (for instance to renegotiate existing contracts) than market-based finance and that these flexibility benefits become more valuable as monetary conditions tighten and hence contribute to a less favorable cyclical outlook; see e.g. Holm-Hadulla and Thürwächter (2021).

[14]It is is noteworthy that, for mortgages, a significant volume response coincides with an insignificant rate response to the monetary policy tightening. A potential channel is that banks do not (only) rely on rates but also on the credit standards they apply to their (prospective) customers when adjusting to macroeconomic shocks (Altavilla et al., 2019b).

cating that the dampening effect of volatility on transmission largely originates in the banking system. The monetary policy shock, at average volatility levels, pushes up short-term money market rates on impact and the response builds up further over a six-month horizon, before reverting to zero and temporarily undershooting at longer horizons (Figure 5, blue IRF in left panel).[15] The estimates provide some indications that, as volatility rises, the response is again attenuated. However, the interaction coefficient is insignificant (yellow IRF), and the gradient of the monetary policy impact on the 3-month OIS rate conditional on volatility is much flatter than for instance for the macro variables (Figure 5, right panel). Hence, even absent discernible changes in the money-market response to monetary policy shocks, banks seem to become less inclined to adjust their lending behavior if these shocks are surrounded by elevated levels of non-policy-related noise.

Taken together, our estimates point to a substantial dampening effect of short-rate volatility on the output and price response to a monetary policy shock; and this dampening effect shows up already at the bank-based stage of monetary policy transmission. This pattern, in turn, is consistent with the classic Dixit (1991) result that uncertainty widens the optimal range of inertia whenever adjustments to past decisions are costly.

4 Robustness

To test the robustness of our findings, we make four adjustments to our baseline model and identification approach.

First, we include as additional controls the aggregate excess liquidity holdings of euro area banks (scaled by total bank assets) and their interaction with the monetary policy shock (see Annex, Figure 7, first row). In the baseline specification, we account for potential collinearity between short-rate volatility and the ECB's use of unconventional monetary policy measures by including shocks to the 5- and

[15]Since financial variables tend to respond to monetary policy more swiftly than activity and prices, we zoom in on a 12-month horizon here. The full IRFs are in Figure 6 available in the Annex.

10-year OIS rates as additional covariates (along with their interactions with the standard policy rate shock). However, the level of excess liquidity may, in its own right, affect financial intermediation and thereby feed through to macroeconomic outcomes (Acharya et al., 2023; Diamond et al., 2024; Altavilla et al., 2025); and significant changes in banks' excess liquidity holdings coincided with the regime shifts in monetary policy implementation used for identification in the above analysis and the concomitant changes in short-rate volatility. To guard against omitted variable bias, we hence directly control for excess liquidity and its interaction with the monetary shock.

Second, we exclude short-rate volatility from the list of endogenous regressors, while retaining the interaction between short-rate volatility and the monetary policy shock as an endogenous variable (Figure 7, second row). Since it appears implausible that the constitutive terms of the interaction are exogenous while the interaction itself is not, we instrument both in our baseline estimations. However, it is prudent practice to restrict the number of endogenous variables (Angrist and Pischke, 2009), which is why we consider it as a robustness check here.

Third, we re-estimate the model by incorporating all elements of X_t into Z_t, with the exception of the constant term (Figure 7, third row). This approach follows the default assumption in the Lewbel (2012) estimator and amounts to exploiting potential heteroskedasticity in all control variables for identification. Our baseline exploits heteroskedasticity driven by distinct institutional changes in the history of euro area monetary policy implementation and thus has parallels to identification strategies based on quasi-experiments. By adding the Lewbel default-specification, we extend this method with a more agnostic 'internal-instruments-based' approach.

Finally, we sequentially merge each of the four regime changes in Z_t with the preceding one to assess their relative importance in identifying the impact of short-rate volatility on macroeconomic outcome variables (Figure 8).[16] While omitting a regime may reduce the strength of the instruments, it ensures that the results are not driven by any particular event.

[16] In other words, the first of these robustness checks merges baseline regimes 1 and 2, the second merges baseline regimes 2 and 3, *et cetera*.

Throughout these robustness checks, our main findings remain intact. Even after including excess liquidity and its interaction with the policy shock, we continue to observe a dampening effect of short-rate volatility of similar magnitude as in the baseline, which is striking given the high correlation between excess liquidity and short-rate volatility (Figure 7, first row). Treating the level of short-rate volatility as exogenous is inconsequential for the estimates (second row), and adding all covariates to Z_t likewise leaves the point estimates largely unchanged, while strongly improving their precision (third row). Finally, the specifications merging regimes show that none of the individual regime changes chosen for identification materially drive our results (Figure 8).

5 Conclusion

This study provides evidence on how short-term interest rate volatility affects the transmission of monetary policy. By leveraging the heteroskedasticity in the time series of euro area money market rates, we isolate exogenous variation in volatility and interact it with high-frequency identified monetary policy shocks. Our findings reveal that increased short-rate volatility dampens the effect of monetary policy on activity and prices; and this dampening effect is evident across the entire transmission chain – from key money market rates to bank lending rates and volumes. Our results are robust across various model specifications.

For the calibration of monetary policy implementation frameworks, our findings highlight the effective transmission of monetary policy as an important reason for keeping the scope for short-rate volatility within reasonable bounds. At the same time, the benefit of greater rate controllability needs to be traded-off against other desirable features, such as preserving some room for money market activity, which speaks against fully muting any volatility in key short-term market rates.

References

Acharya, V. V., Chauhan, R. S., Rajan, R., and Steffen, S. (2023). Liquidity dependence and the waxing and waning of central bank balance sheets. *National Bureau of Economic Research*, 31050.

Afonso, G., Giannone, D., La Spada, G., and Williams, J. C. (2022). Scarce, abundant, or ample? A time-varying model of the reserve demand curve. *Federal Reserve Bank of New York*, Staff report 1019.

Alpanda, S., Granziera, E., and Zubairy, S. (2021). State dependence of monetary policy across business, credit and interest rate cycles. *European Economic Review*, 140 Article 103936.

Alpanda, S. and Zubairy, S. (2019). Household debt overhang and transmission of monetary policy. *Journal of Money, Credit and Banking*, 51(5):1265–1307.

Altavilla, C., Brugnolini, L., Gürkaynak, R. S., Motto, R., and Ragusa, G. (2019a). Measuring euro area monetary policy. *Journal of Monetary Economics*, 108:162–179.

Altavilla, C., Paries, M. D., and Nicoletti, G. (2019b). Loan supply, credit markets and the euro area financial crisis. *Journal of Banking & Finance*, 109 Article 105658.

Altavilla, C., Rostagno, M., and Schumacher, J. (2025). When banks hold back: credit and liquidity provision. *ECB Working Paper*, No. 2025/3009.

Andrade, P. and Ferroni, F. (2021). Delphic and Odyssean monetary policy shocks: Evidence from the euro area. *Journal of Monetary Economics*, 117:816–832.

Angrist, J. D. and Pischke, J.-S. (2009). *Mostly harmless econometrics: An empiricist's companion*. Princeton University Press.

Arce, O., Nuno, G., Thaler, D., and Thomas, C. (2020). A large central bank balance sheet? Floor vs corridor systems in a new Keynesian environment. *Journal of Monetary Economics*, 114:350–367.

Auer, S., Bernardini, M., and Cecioni, M. (2021). Corporate leverage and monetary policy effectiveness in the euro area. *European Economic Review*, 140 Article 103943.

Barakchian, S. M. and Crowe, C. (2013). Monetary policy matters: Evidence from new shocks data. *Journal of Monetary Economics*, 60(8):950–966.

Baum, C. F. and Lewbel, A. (2019). Advice on using heteroskedasticity-based identification. *The Stata Journal*, 19(4):757–767.

Becker, B. and Ivashina, V. (2014). Cyclicality of credit supply: Firm level evidence. *Journal of Monetary Economics*, 62:76–93.

Bindseil, U. (2014). *Monetary policy operations and the financial system.* Oxford University Press.

Bindseil, U. (2016). Evaluating monetary policy operational frameworks. In *Speech at the Jackson Hole Conference*, volume 31.

Breusch, T. S. and Pagan, A. R. (1979). A simple test for heteroscedasticity and random coefficient variation. *Econometrica*, 47(5):1287–1294.

Cochrane, J. H. and Piazzesi, M. (2002). The Fed and interest rates—a high-frequency identification. *American Economic Review*, 92(2):90–95.

Cook, R. D. and Weisberg, S. (1983). Diagnostics for heteroscedasticity in regression. *Biometrika*, 70(1):1–10.

Corsetti, G., Duarte, J. B., and Mann, S. (2022). One money, many markets. *Journal of the European Economic Association*, 20(1):513–548.

Darracq-Paries, M. and Santis, R. A. D. (2015). A non-standard monetary policy shock: The ECB's 3-year LTROs and the shift in credit supply. *Journal of International Money and Finance*, 54:1–34.

Diamond, W., Jiang, Z., and Ma, Y. (2024). The reserve supply channel of unconventional monetary policy. *Journal of Financial Economics*, 159 Article 103887.

Dixit, A. (1991). Analytical approximations in models of hysteresis. *The Review of Economic Studies*, 58(1):141–151.

Ehrmann, M. and Fratzscher, M. (2007). Communication by central bank committee members: Different strategies, same effectiveness? *Journal of Money, Credit and Banking*, 39(2-3):509–541.

Gerali, A., Neri, S., Sessa, L., and Signoretti, F. M. (2010). Credit and banking in a DSGE model of the euro area. *Journal of Money, Credit and Banking*, 42:107–141.

Gertler, M. and Karadi, P. (2015). Monetary policy surprises, credit costs, and economic activity. *American Economic Journal: Macroeconomics*, 7(1):44–76.

Gürkaynak, R. S., Sack, B., and Swanson, E. (2005). The sensitivity of long-term interest rates to economic news: Evidence and implications for macroeconomic models. *American Economic Review*, 95(1):425–436.

Hansen, P. S. (1999). Frequent price changes under menu costs. *Journal of Economic Dynamics and Control*, 23(7):1065–1076.

Hollo, D., Kremer, M., and Lo Duca, M. (2012). CISS - a composite indicator of systemic stress in the financial system. *ECB Working Paper Series No. 1426.*

Holm-Hadulla, F. and Thürwächter, C. (2021). Heterogeneity in corporate debt structures and the transmission of monetary policy. *European Economic Review*, 136 Article 103743.

Jarociński, M. and Karadi, P. (2020). Deconstructing monetary policy surprises—the role of information shocks. *American Economic Journal: Macroeconomics*, 12(2):1–43.

Jordà, Ò. (2005). Estimation and inference of impulse responses by local projections. *American Economic Review*, 95(1):161–182.

Jordà, Ò., Schularick, M., and Taylor, A. M. (2020). The effects of quasi-random monetary experiments. *Journal of Monetary Economics*, 112:22–40.

Kuttner, K. N. (2001). Monetary policy surprises and interest rates: Evidence from the fed funds futures market. *Journal of Monetary Economics*, 47(3):523–544.

Lane, P. R. (2023). Central bank liquidity: a macroeconomic perspective. *Welcome address at the ECB Conference on Money Markets*.

Lewbel, A. (2012). Using heteroscedasticity to identify and estimate mismeasured and endogenous regressor models. *Journal of Business & Economic Statistics*, 30(1):67–80.

Lopez-Salido, D. and Vissing-Jorgensen, A. (2023). Reserve demand and balance sheet run-off. *Federal Reserve Board Working Paper*.

McKay, A. and Wolf, C. K. (2023). Monetary policy and inequality. *Journal of Economic Perspectives*, 37(1):121–144.

Nakamura, E. and Steinsson, J. (2018). High-frequency identification of monetary non-neutrality: the information effect. *The Quarterly Journal of Economics*, 133(3):1283–1330.

Pagan, A. and Hall, D. (1983). Diagnostic tests as residual analysis. *Econometric Reviews*, 2(2):159–218.

Peersman, G. and Smets, F. (2005). The industry effects of monetary policy in the euro area. *The Economic Journal*, 115(503):319–342.

Piazzesi, M. and Swanson, E. T. (2008). Futures prices as risk-adjusted forecasts of monetary policy. *Journal of Monetary Economics*, 55(4):677–691.

Poole, W. (1970). Optimal choice of monetary policy instruments in a simple stochastic macro model. *The Quarterly Journal of Economics*, 84(2):197–216.

Ramey, V. A. (2016). Macroeconomic shocks and their propagation. *Handbook of macroeconomics*, 2:71–162.

Rigobon, R. (2003). Identification through heteroskedasticity. *Review of Economics and Statistics*, 85(4):777–792.

Schnabel, I. (2023). Back to normal? Balance sheet size and interest rate control. *Speech given at an event organised by Columbia University and SGH Macro Advisors, New York*.

Tenreyro, S. and Thwaites, G. (2016). Pushing on a string: US monetary policy is less powerful in recessions. *American Economic Journal: Macroeconomics*, 8(4):43–74.

Tillmann, P. (2020). Monetary policy uncertainty and the response of the yield curve to policy shocks. *Journal of Money, Credit and Banking*, 52(4):803–833.

Vavra, J. (2014). Inflation dynamics and time-varying volatility: New evidence and an Ss interpretation. *The Quarterly Journal of Economics*, 129(1):215–258.

Woodford, M. (2005). Central bank communication and policy effectiveness. *National Bureau of Economic Research*, 31050.

6 Annex

Table 4: Descriptive statistics

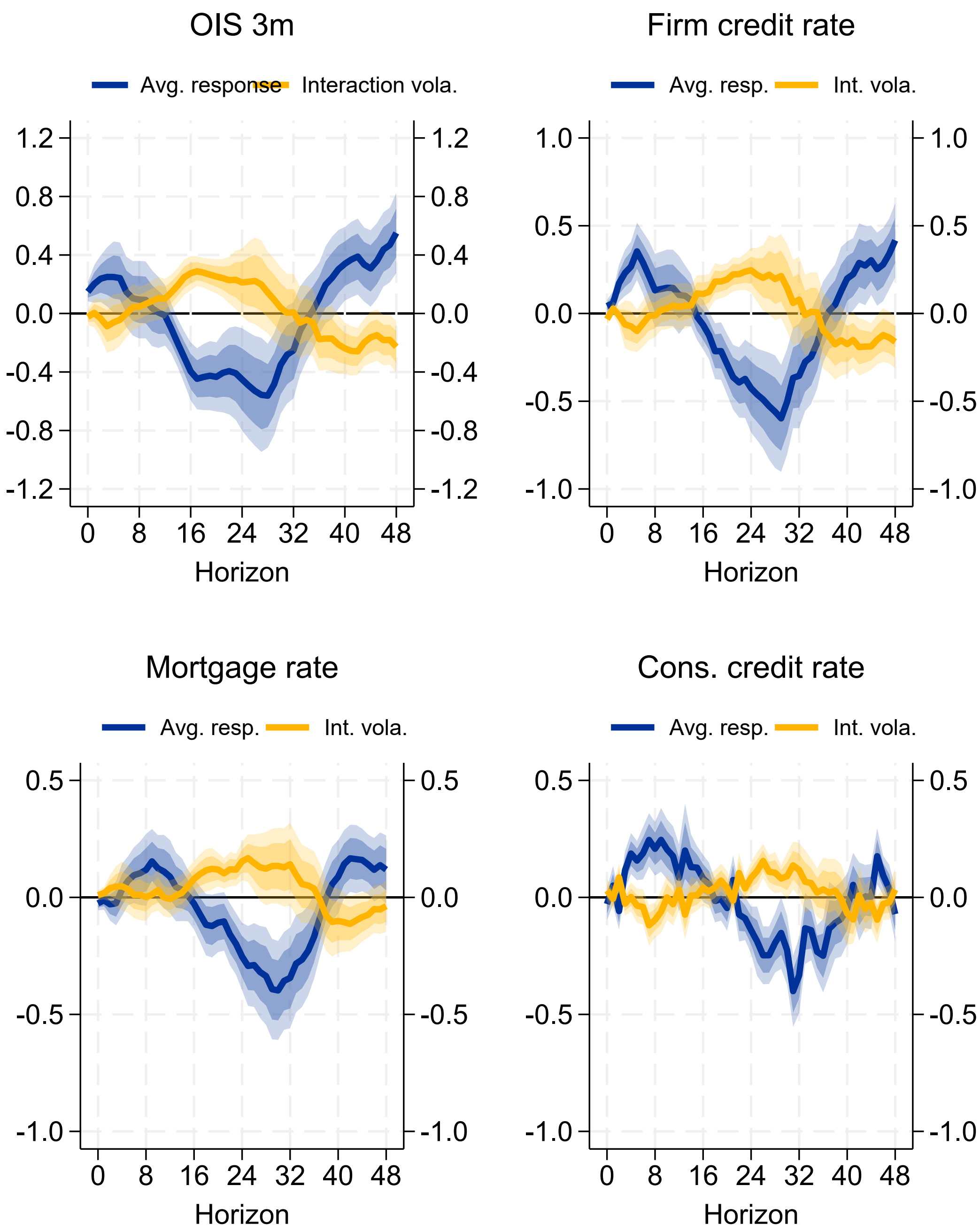
	Mean	Median	Std.Dev.	Min	Max	Source
Dependent variables						
Euro area real GDP (EUR bn)	855	854	48	768	959	Eurostat National Accounts
Euro area GDP Deflator (index)	31	32	2	27	36	Eurostat National Accounts
Household loans for house purchase (EUR bn)	3,493.23	3,726.10	674.77	2,037.65	4,559.30	ECB Balance Sheet Items
Household loans for consumption (EUR)	595.71	604.01	58.88	477.88	719.73	ECB Balance Sheet Items
NFC credit (EUR bn)	4,123.03	4,329.30	618.25	2,907.14	4,877.12	ECB Balance Sheet Items
3-month OIS (percent)	1.17	0.45	1.52	-0.47	4.33	Reuters
Household rate for house purchase (percent)	3.43	3.59	1.20	1.39	5.91	ECB MFI Interest Rate Statistics
Household rate for consumption (percent)	6.75	6.89	0.74	5.26	8.18	ECB MFI Interest Rate Statistics
NFC rate (percent)	3.01	2.84	1.21	1.34	5.77	ECB MFI Interest Rate Statistics
Control variables						
Commodity price index	126.77	119.49	34.07	61.89	200.81	IMF
EUR-USD rate	1.24	1.24	0.14	0.86	1.58	Statistical data warehouse ECB
CISS composite index	0.19	0.11	0.17	0.02	0.80	Statistical data warehouse ECB
Excess liquidity over total bank assets (percent)	1.45	0.41	2.06	- 0.00	6.12	ECB and Balance Sheet Items
Shock variables						
1-month OIS (percent)	0.00	0.00	0.03	-0.20	0.14	Altavilla et al. (2019a)
5-year OIS (percent)	0.00	0.00	0.02	-0.12	0.13	Altavilla et al. (2019a)
10-year OIS (percent)	0.00	0.00	0.02	-0.09	0.12	Altavilla et al. (2019a)
Std. Dev. €STR (percent)	0.06	0.04	0.07	0.00	0.32	ECB

Table 5: Variances across regimes

Regime	σ	GDP	Deflator
Pre-tech. change	0.621	0.007	0.018
Pre-FRFA	0.222	0.117	0.082
Post-FRFA	0.590	0.027	0.007
Ample reserves	0.095	0.008	0.010
Abundant reserves	0.008	0.083	0.039

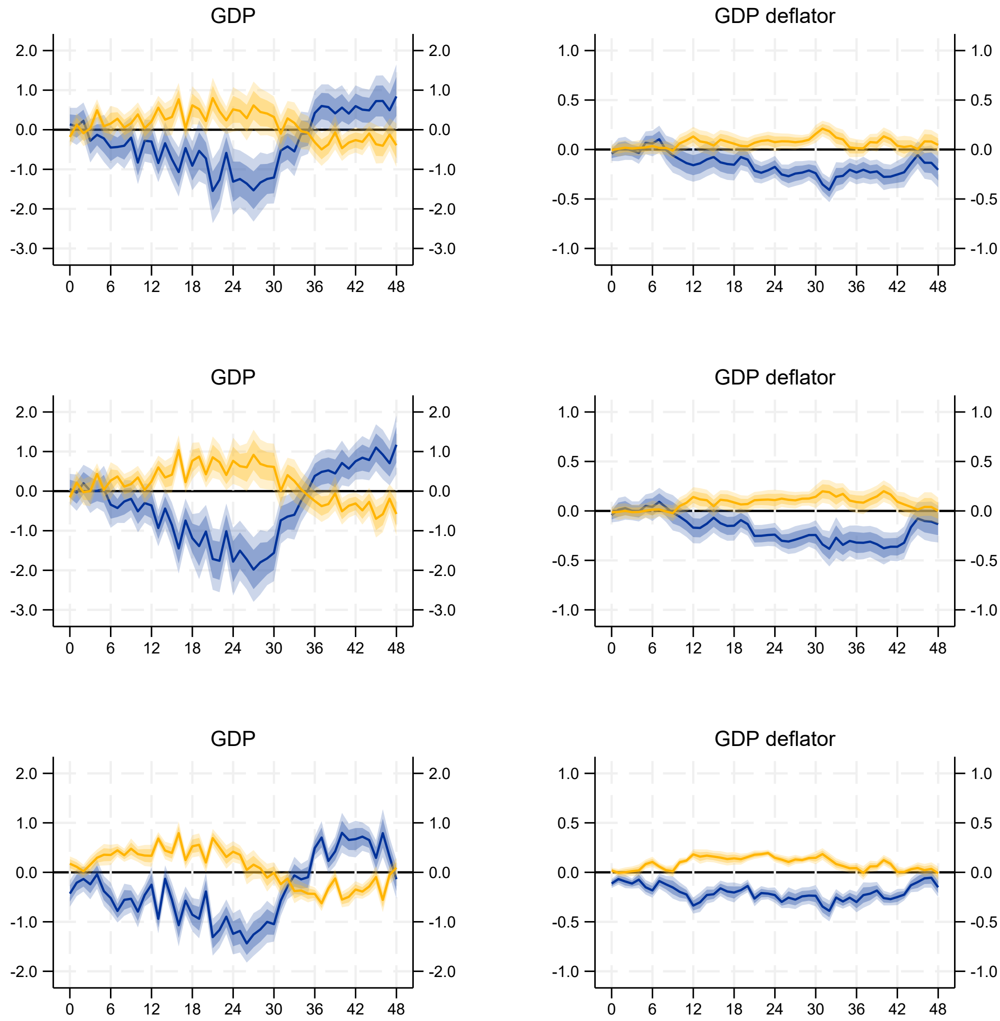
Note: variances expressed in basis points (bps).

Figure 6: IRFs following a monetary policy shock conditional on short-rate volatility (y-axis: percentage points; x-axis: months)



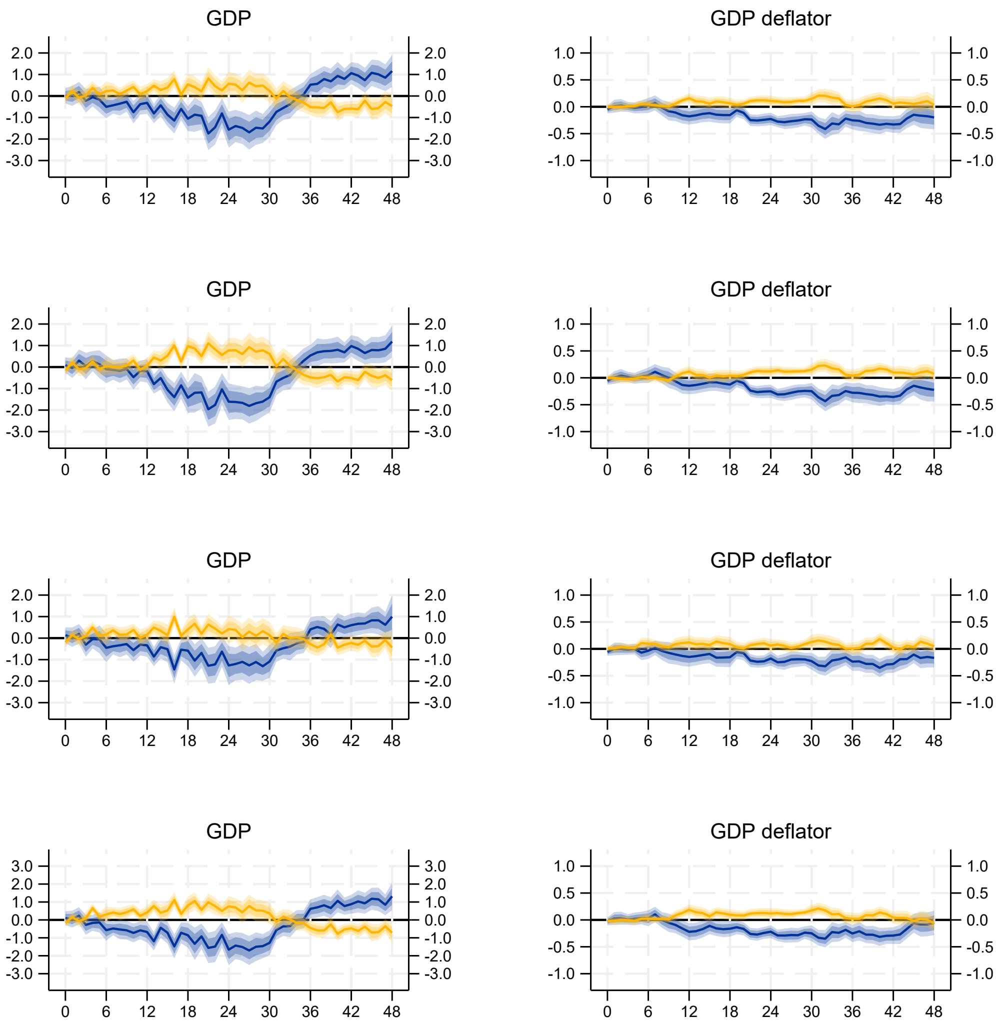
Note: the IRFs are based on equation (1) and normalized to a 25bps peak response in the 3-month OIS rate and a one standard deviation increase in €STR volatility. The dark- (light-)shaded area shows the 90% (68%) confidence interval.

Figure 7: IRFs following a monetary policy shock conditional on short-rate volatility (y-axis: percentages; x-axis: months)



Note: the IRFs are based on equation (1), modified as described in Section 4, and normalized to a 25bps peak response in the 3-month OIS rate and a one standard deviation increase in €STR volatility. Short-rate volatility and the interaction between short-rate volatility and the monetary policy shocks are instrumented by probabilistic instruments identified through heteroskedasticity following Rigobon (2003). The IRFs also control for banks' excess liquidity holdings as a share of total assets (row 1), include short-rate volatility as an exogenous variable (row 2), and apply the the default Lewbel (2012) estimator by including all elements of X_t in Z_t (row 3). The dark-(light-)shaded area is the 90% (68%) confidence interval. €STR volatility is expressed as a z-score.

Figure 8: IRFs following a monetary policy shock conditional on short-rate volatility (y-axis: percentages; x-axis: months)



Note: the IRFs are based on equation (1), modified as described in Section 4, and normalized to a 25bps peak response in the 3-month OIS rate and a one standard deviation increase in €STR volatility. Short-rate volatility and the interaction between short-rate volatility and the monetary policy shocks are instrumented by probabilistic instruments identified through heteroskedasticity following Rigobon (2003). The IRFs successively merge adjacent regimes (with row 1 showing the specification merging regime 1 and 2 etc.). The dark- (light-)shaded area is the 90% (68%) confidence interval. €STR volatility is expressed as a z-score.

Acknowledgements

We thank Bianca Brunori for excellent research assistance. Earlier versions of this work have benefited from comments and suggestions by Carlo Altavilla, Christophe Kamps, Peter Karadi, Wolfgang Lemke, Massimo Rostagno, Julian Schumacher, Johannes Tischer, Oreste Tristani, Thomas Vlassopoulos, and seminar participants at the ECB.
The views in this paper are those of the authors and do not necessarily reflect the views of the European Central Bank (ECB) or the Eurosystem.

Fédéric Holm-Hadulla
European Central Bank, Frankfurt am Main, Germany; email: federic.holm-hadulla@ecb.europa.eu

Sebastiaan Pool
European Central Bank, Frankfurt am Main, Germany; email: sebastiaan.pool@ecb.europa.eu

Postal address 60640 Frankfurt am Main, Germany
Telephone +49 69 1344 0
Website www.ecb.europa.eu

PDF ISBN 978-92-899-7245-1 ISSN 1725-2806 doi:10.2866/5877858 QB-01-25-110-EN-N