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Synchronous but Dissimilar: Heterogeneity in the Bank
Lending Channel Across the Euro Area

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This study investigates how monetary policy is transmitted to bank lending in euro area countries and whether there are differences across countries. The findings suggest that the peak effects consistently occur within 12 to 18 months after monetary policy actions. However, the size of the effects varies considerably across countries, and the effects are asymmetric as expansionary monetary policy impacts are larger. We find that these results have been stable over the past several years.

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Monetary policy effects have a time lag

How fast and how strongly monetary policy affects important variables in the economy is subject to considerable uncertainty.



Transmission varies across countries

The transmission of monetary policy to bank lending is synchronous across euro area countries with effects peaking around 12 to 18 months, but impacts differ markedly in terms of magnitude. Moreover, expansionary monetary policy affects bank lending more strongly than contractionary monetary policy.



Why it matters

Policy makers need timely information on the dynamic effect of monetary policy on macroeconomic variables. Our findings suggest that transmission of monetary policy actions to bank lending is country-specific, asymmetric and stable over the past several years.

Opinions expressed by the authors of studies do not necessarily reflect the official viewpoint of the Oesterreichische Nationalbank or the Eurosystem.

Challenges for Monetary Policy Transmission in a Changing World Network (ChaMP)

This paper contains research conducted within the network “Challenges for Monetary Policy Transmission in a Changing World Network” (ChaMP). It consists of economists from the European Central Bank (ECB) and the national central banks (NCBs) of the European System of Central Banks (ESCB).

ChaMP is coordinated by a team chaired by Philipp Hartmann (ECB), and consisting of Diana Bonfim (Banco de Portugal), Margherita Bottero (Banca d’Italia), Emmanuel Dhyne (Nationale Bank van België/Banque Nationale de Belgique) and Maria T. Valderrama (Oesterreichische Nationalbank), who are supported by Melina Papoutsi and Gonzalo Paz-Pardo (both ECB), 7 central bank advisers and 8 academic consultants.

ChaMP seeks to revisit our knowledge of monetary transmission channels in the euro area in the context of unprecedented shocks, multiple ongoing structural changes and the extension of the monetary policy toolkit over the last decade and a half as well as the recent steep inflation wave and its reversal. More information is provided on its [website](#).

Synchronous but Dissimilar: Heterogeneity in the Bank Lending Channel Across the Euro Area^{*}

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Abstract

The following study analyzes the transmission of monetary policy to bank lending in the euro area. It focuses particularly on the heterogeneity in the speed and strength of propagation of monetary policy shocks across euro area countries. We employ instrumental variables, based on common euro area wide high-frequency identified monetary policy surprises in combination with local projections, to infer the dynamic effect of monetary policy on bank lending. Local projections are estimated for loan growth to non-financial corporations, allowing for a flexible specification that accounts for potential asymmetries in responses to easing versus tightening monetary policy shocks. The robustness of inferred impulse response functions is examined during the pandemic and across different phases of the monetary policy normalization that began in 2022, including both hiking and easing cycles and extending to the end of 2024. We find that the response of bank lending to monetary policy shocks is generally synchronous across euro area countries in that impulse responses peak after 12 to 18 months. However, the magnitude of peak effects varies considerably pointing to pronounced country heterogeneity in the bank lending channel. Moreover, we find significant asymmetries in country-specific bank lending responses wherein monetary policy tightening shocks are generally less strongly propagated suggesting downward rigidity in the bank lending channel. Taken together our results help policymakers quantify the heterogeneity of common euro-area monetary policy transmission to bank lending across member countries.

Keywords: Monetary policy transmission, Bank lending channel, Country heterogeneity
JEL: E52, F45, G21

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Disclaimer: The views expressed in this paper do not necessarily reflect those of the Oesterreichische Nationalbank or the Eurosystem.

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Non-technical summary

How fast and how strongly monetary policy affects important variables in the economy such as economic activity and inflation is subject to considerable uncertainty. However, policymakers need timely knowledge on how monetary policy actions affect the economy. This study contributes to a growing body of empirical literature that analyzes how strongly and how quickly monetary policy impacts the economy in the euro area.

Recent empirical studies have found that there are differences in the effect of monetary policy on economic activity and prices across euro area countries. However, the timing it takes for these effects to occur seems broadly synchronous in euro area countries. Typically, the peak effects on measures of economic activity materialize within one to two years after initial monetary policy actions. In contrast, the transmission of monetary policy might not be as strongly synchronous across countries when it comes to financial variables like bank credit.

This study provides estimates for the effects of monetary policy and analyzes how strongly and how quickly monetary policy affects bank lending in euro area countries. Moreover, it considers the possibility that the effects are asymmetric wherein expansionary and contractionary monetary policy may lead to different impacts on bank lending. Lastly, it also analyzes whether the effects have been stable over the past several years from the onset of the COVID-19 pandemic on.

The findings suggest that there are marked differences in monetary policy transmission to bank credit across euro area countries. The effects are up to three times as large in the most affected country compared to the least affected one. Hence, monetary policy does not affect bank lending in the same way in different countries. However, the timing of the response in bank lending appears very much synchronous with the peak effects occurring after 12 to 18 months. Importantly, the results suggest that effects on bank lending are highly asymmetric. Expansionary monetary policy affects bank lending more strongly than contractionary monetary policy. Lastly, the impact of monetary policy appears largely stable over the past several years. This suggests that the transmission of monetary policy to bank lending has not materially changed in the last decade.

1. Introduction

The speed and strength of monetary policy transmission to macroeconomic variables such as economic activity and inflation is subject to considerable uncertainty. Consequently, the European Central Bank's (ECB) 2025 monetary policy strategy statement reconfirms and emphasizes a medium-term orientation of policy actions in view of time lags in monetary policy transmission.¹ Nevertheless, policy makers require timely estimates of the dynamic responses of the economy to monetary policy shocks. This study contributes to the growing empirical literature on the speed and strength of monetary policy transmission in the euro area. In particular, we focus on the bank lending channel, examining the presence of cross-country heterogeneity, asymmetric responses and potential structural breaks.

Heterogeneity in the transmission of monetary policy across euro area countries has been documented extensively in the empirical literature (see e.g. Ciccarelli et al., 2013; Hafemann and Tillmann, 2020; Corsetti et al., 2022; Mandler et al., 2022; Mandler and Scharnagl, 2023). A common finding of these studies is that the transmission of euro area monetary policy to prices and output differs across countries in terms of the magnitude (strength) of dynamic effects. However, the timing (speed) of monetary policy transmission to output and prices appears broadly synchronous across euro area countries, with peak effects typically materializing between one year and two years after the initial shock.

In contrast, this result of strong synchronization across countries observed for macroeconomic variables may not hold when it comes to the transmission of monetary policy to financial variables such as credit and bank lending. Indeed, evidence presented by Hafemann and Tillmann (2020) for aggregate bank lending on the level of euro area countries suggests marked cross-country differences in the transmission of monetary policy to bank lending. While monetary policy is transmitted forcefully to credit growth in some countries, it appears to have only limited or insignificant effects in others. Hence, we investigate whether the transmission of euro area monetary policy to bank credit operates homogeneously across euro area countries, both in terms of its speed and strength, and whether this transmission mechanism has changed in recent years in light of potential structural breaks arising from the COVID-19 pandemic and the European Central Bank's recent monetary policy hiking cycle amid abundant central bank liquidity.

Relatedly, a comprehensive meta-analysis by Bajzik et al. (2023) suggests that estimated effects of monetary policy transmission on bank credit strongly depend on the empirical identification strategy employed. Therefore, following the current best practice outlined by Wolf (2020), we use external instruments for identification. Specifically, we employ high-frequency identified monetary policy surprises as proxy and instrumental variables to assess the robustness of our results across alternative specifications. Impulse response functions are estimated using local projections of bank lending on exogenous

¹See https://www.ecb.europa.eu/mopo/strategy/strategy-review/ecb.strategyreview202506_strategy_overview.en.html or Lane, P. R. (2024), "Monetary policy under uncertainty", speech at the Bank of England Watchers Conference 2024, Kings College London, November 25th 2024.

monetary policy shocks over a 24-month horizon, considering a medium-term time frame of monetary policy transmission. In addition, we provide empirical evidence on cross-country variation in instrument relevance used in LP-IV exercises. Finally, we consider asymmetric transmission of monetary policy tightening and easing shocks as an additional dimension of heterogeneity.

Our results suggest pronounced heterogeneity in the strength of euro area monetary policy transmission to bank credit for non-financial corporations. The peak magnitude of the impulse responses to easing and tightening shocks is more than three times as large in the most affected country compared to the least affected one. However, we find that, regarding the speed of transmission, impulse response functions are broadly synchronous across countries with dynamic impacts peaking one to two years after the realization of an initial shock. Our results are also indicative of asymmetric transmission of monetary policy shocks, particularly in smaller member countries. Lastly we do not find significant differences in impulse responses when considering different cut-off dates for the data. This suggests that the transmission mechanism of euro area monetary policy to bank lending for non-financial corporations has remained broadly stable over the past decade.

The rest of the paper is structured as follows: Section 2 reviews the related literature; Section 3 presents the empirical strategy and discusses the proposed partial-identification approach; Section 4 describes the data, Section 5 discusses the results, and Section 6 concludes. A detailed appendix provides additional information on the data, external instruments and various robustness checks.

2. Literature review

Heterogeneity in the transmission of monetary policy in the euro area has been studied on several occasions in the literature, with earlier research on the subject in Ciccarelli et al. (2013) and more recently published papers in Hafemann and Tillmann (2020), Corsetti et al. (2022), Mandler et al. (2022), Mandler and Scharnagl (2023) and the latest working papers in Barigozzi et al. (2024) and Zlobins (2025).

Ciccarelli et al. (2013) analyze quarterly data from 2002 until 2011 and use a vector autoregressive (VAR) model for a panel of twelve countries. They base their identification strategy on using the overnight rate as monetary policy instrument and the Cholesky decomposition with a specific ordering of variables. The authors demonstrate that monetary policy transmission has changed in response to the financial crisis of 2008-2009. Subsequent studies typically rely on so-called external instruments to identify monetary policy shocks.

High-frequency monetary policy surprises have a long history and have been studied for the US in Gürkaynak et al. (2005). Here, the authors use high-frequency event studies to investigate the effects of monetary policy on asset prices. Gertler and Karadi (2015) build on this idea and identify monetary policy surprises applying the methodology developed in Stock and Watson (2012) and Mertens and Ravn (2013) in a VAR framework containing economic and financial variables. For the euro area, Altavilla et al. (2019)

present a decomposition into the target factor for the press release window and the timing, forward guidance, and quantitative easing factors in the press conference window. The authors construct a database (EA-MPD, Euro Area Monetary Policy Event-Study Database), which is also employed in Jarociński and Karadi (2020). In that study, the authors distinguish between monetary policy shocks (characterized by negative comovement with stock prices) and central bank information shocks (associated to positive comovement with stock prices), in a structural VAR.

Hafemann and Tillmann (2020) adopt an approach similar to Gertler and Karadi (2015) by estimating a structural VAR model in which they identify a common monetary policy shock for the euro area based on an external instrument, i.e., changes in the 10-year German government bond yield on the ECB’s Governing Council meeting days. Using local projections to analyze country-specific responses, they identify broadly synchronous effects across countries for output (industrial production) and prices. However, they point to differences for variables related to financial transmission channels, such as stock prices and bank lending. Their sample includes monthly data from January 2002 to October 2016 for ten euro area member countries, ending around the time the ECB introduced negative interest rate policy (NIRP).

While Faia and Karau (2021) focus on the risk-taking channel of monetary policy, their work is worth mentioning here because they also employ a two-step procedure (as in Gertler and Karadi, 2015; Hafemann and Tillmann, 2020). In the first step, they identify a structural monetary policy shock using a high-frequency instrument in a proxy SVAR, and in the second step, they apply a local projection framework to study the dynamic effects and include the interactions with macroprudential policy.

Corsetti et al. (2022) also follow the identification strategy using a high-frequency instrument, combining it with a dynamic factor model applied to quarterly euro area data from 1999 to 2016. The motivation for using the factor model is to incorporate a broader set of real and nominal macroeconomic indicators as well as selected financial market variables. Although their impulse response functions are not explicitly used to quantify the effects on credit growth — i.e., the bank lending channel — they illustrate how the observed heterogeneity in the impulse responses of macroeconomic and financial variables can be attributed to country-specific institutional factors, such as the share of adjustable-rate mortgages, home-ownership rates, and labor market rigidities.

Mandler et al. (2022) examine cross-country differences in the transmission of conventional monetary policy shocks in France, Germany, Italy, and Spain using a Bayesian vector autoregressive (BVAR) model. The estimation of their model covers quarterly data between 1999Q1 and 2014Q3. The authors choose a sample period that contains only small effects of unconventional monetary policy measures. The identification approach relies on sign restrictions. They find a stronger output response in Germany, but a weaker effect on the price level. In Spain, the price level reacts more strongly than in other countries, but the output effect is smaller than in Germany. Mandler et al. (2022) also perform an analysis of the determinants of the effects on output and the price level, with a particular focus on Germany. They find that the stronger effect on output is driven by a higher share

of manufacturing and more competition in the banking system, also the strong focus on exports strengthens the transmission. Moreover, a more flexible labor market in Germany makes quantities rather than prices more responsive to monetary policy. The authors conclude, in line with the existing literature, that although there is heterogeneity across countries, the effects of monetary policy shocks on output and prices tend to be small.

Barigozzi et al. (2024) recently released the EA-MD-QD dataset for the euro area and ten main member countries. In their working paper, the authors estimate a common component vector autoregression (CC-VAR) model. Their identification strategy is also based on Gertler and Karadi (2015). They decided to use the two-year euro area interest rate as policy indicator and high frequency changes in the one-year swap rate as instrument based on the F -statistic. A key finding from the working paper is that real variables appear to be more driven by a common factor, while inflation rates also substantially react to country-specific factors.

Zlobins (2025) is a most recent working paper on the lags and strength of monetary policy transmission in the euro area. The author estimates a Bayesian vector autoregressive model and applies various identification strategies, including common high-frequency instruments and also narrative sign restrictions. The sample covers monthly data from January 2002 to October 2023 a period that encompasses the recent hiking cycle but not the easing cycle which began in June 2024. Furthermore, he estimates a version with time-varying parameters and finds that the impact of euro area monetary policy on output has been broadly stable over time, but the reaction of inflation has been stronger and more persistent in recent tightening cycles.

There is also a strand of literature that analyzes the transmission of monetary policy with micro-level credit register data from AnaCredit. Altavilla et al. (2024b) analyze the external finance premium based on individual loan contracts to euro area firms. Their findings indicate that loan pricing reflects cross-country differences and is influenced by firms balance sheet characteristics. Moreover, the authors document asymmetries in the transmission mechanism between policy rate hikes and quantitative easing measures. A similar setup involving AnaCredit data is also employed in Altavilla et al. (2024a). In this study, the authors identify a climate risk-taking channel by taking into account whether firms have higher or lower carbon emissions. Both of the before mentioned papers employ high-frequency monetary policy surprises from the EA-MPD (see Altavilla et al., 2019). A recent working paper by Vilerts et al. (2025) finds that loan contracts in AnaCredit linked to shorter-maturity risk-free rates exhibit a stronger monetary policy pass-through. The previously mentioned studies offer valuable insights at a highly disaggregated level. However, these works primarily focus on price-related effects and largely overlook loan growth dynamics.

Local projections with instrumental variables have become a preferred tool for structural identification in modern macroeconomics; see Jordà et al. (2015) and Jordà and Taylor (2025) for a recent review. They facilitate the inclusion of asymmetries that may exist in the bank lending channel for different countries. They are also easier to extend to a panel setting than a VAR model. Another clear advantage of local projections is that

they are more robust to misspecification.

Beyond structural modeling, the recent empirical literature offers various identification strategies for estimating parameters and recovering structural shocks in vector autoregressive (VAR) models. Common approaches impose short-run zero or sign restrictions on the contemporaneous impact matrix, or long-run zero restrictions derived from theoretical considerations. While these methods provide identification grounded in macroeconomic theory, the restrictions they impose on the system of variables may be excessively stringent (see, e.g., Nakamura et al., 2024). Furthermore, relying solely on sign restrictions typically yields limited identifying information and the application of a valid external instrument provides the ideal case of identification (Wolf, 2020).

We follow the recent literature by employing high-frequency identified monetary policy surprises as external instruments, combined with local projections that offer assumption-light identification. The use of external instruments allows for partial identification and provides a framework that is less assumption-intensive than fully identified approaches, which impose restrictions on the contemporaneous impact matrix of a VAR model. We favor local projections because they are less biased under model misspecification compared to SVAR-IV methodologies (Olea et al., 2024). This exercise reveals heterogeneous effects of a common monetary policy stance that are often not specifically considered or masked when analyzing the credit impulse with responses to monetary policy shocks at the aggregate euro area level.

3. Methodology

3.1. Structural representation

We examine how monetary policy shocks, affecting the euro area as a whole, impact bank lending growth to non-financial corporations at the level of individual euro area countries. Thus, our aim is to infer the dynamic response of bank lending to monetary policy shocks in the medium-term transmission horizon up to two years after an initial shock occurred.

We motivate our research question within the framework of the New Keynesian model, which formalizes the structural relationships among key macroeconomic variables. These variables are represented through a system of structural equations. Following Gertler and Karadi (2015), we consider a medium-scale VAR that features an interest rate with 12-month maturity as the monetary policy indicator, thereby capturing the stance of monetary policy during periods of forward guidance. Furthermore, the country VARs include variables for industrial production, the price level as measured by the headline inflation index according to the Harmonized Index of Consumer Prices (HICP) and an external finance premium that reflects variation in private-sector borrowing cost in bond markets. As in Mandler and Scharnagl (2020), we include euro area aggregate time series of industrial production and HICP as endogenous variables to account for the possibility that the monetary policy reaction function is incorrectly estimated if national variables for economic activity and inflation are not representative of the euro area as a whole.

Moreover, the inclusion of euro area aggregate variables provides a parsimonious way to account for the possibility of spillovers from aggregate euro area dynamics.²

For each country c and month t , the relationship between macro-economic variables of economic activity, inflation, private borrowing cost and monetary policy is characterized by a system of structural equations:³

$$A_c Y_{c,t} = \alpha_c + B_{c,1} Y_{c,t-1} + \dots + B_{c,p} Y_{c,t-p} + \lambda'_c W_{c,t} + u_{c,t}, \quad (1)$$

where $Y_{c,t}$ is a k -dimensional vector of endogenous variables, $W_{c,t}$ denotes a set of contemporaneous control variables featuring the national unemployment rate, the EUR/USD spot exchange rate, a commodity price index and a dummy variable that indicates the onset of the COVID-19 pandemic. The vector $u_{c,t} \sim N(0, \sigma_c^{2'} I^k)$ denotes idiosyncratic mutually uncorrelated error terms (structural innovations to the system), $\sigma_c^2 \in \mathbb{R}^{k \times 1}$ is a vector of variances of structural innovations, $I^k \in \mathbb{R}^{k \times k}$ denotes the identity matrix and $\alpha_c \in \mathbb{R}^{k \times 1}$ is a vector of intercepts.⁴ The system permits a structural interpretation of relationships among the k macro-economic variables, with innovations in each equation interpreted as structural shocks, provided that

$$u_{c,t}^i \perp\!\!\!\perp u_{c,t}^j \text{ for } i \neq j \wedge (i, j) \in [1, 2, \dots, k] \Rightarrow u_{c,t}^{mp} \perp\!\!\!\perp u_{c,t}^{\neq mp}. \quad (2)$$

Hence, the structural innovation in the monetary policy equation is contemporaneously exogenous with respect to all other structural innovations in the system. Any violation of contemporaneous exogeneity would, by definition, preclude a structural interpretation.⁵

²See Sections 4 and Appendix A for details on the macro-economic variables and transformations applied. We consider alternative specifications for the lag length, for instance optimizing by the Akaike Information Criterion (AIC), but ultimately adopt 12 monthly lags, in line with common practice in the literature. Employing other selection criteria, such as the Bayesian Information Criterion (BIC), which typically suggests a shorter lag length, yields very similar results. A comparable finding for the euro area aggregate is reported in Zlobins (2025).

³See, for instance, Mandler and Scharnagl (2020) for an analysis of bank lending shocks using country-specific BVARs for France, Germany, Italy and Spain or Bijsterbosch and Falagiarda (2014) for an investigation of credit supply and economic activity in country-specific time-varying parameter VARs. Guarda and Jeanfils (2012) likewise investigate the contributions of financial shocks to the evolution of real activity on the level of euro area members using country-specific VARs and Georgiadis (2014) studies monetary policy transmission to output and prices in (among others) individual euro area countries employing identical country-specific VARs just as motivated in Equation (1) above.

⁴To account for the impact of the COVID-19 pandemic, the matrix $W_{c,t}$ includes a dummy variable for the period from February 2020 to May 2020. Several recent empirical studies provide vindications for accounting for this period with dummy variables. Braun et al. (2024) for instance employ a dummy variable covering a much longer time frame in estimating a VAR to infer the effect of monetary policy shocks in the UK and include a pandemic dummy for the period from March 2020 to August 2022. Similarly, Ng (2021) shows that impulse responses comparable in shape and magnitude to those obtained before 2020 can be recovered by treating the COVID-19 period as an exogenous level shift in variables. Our results are robust to excluding observations from the pandemic period, as suggested by Baumeister and Hamilton (2023).

⁵We want to stress the conceptual distinction between the statistical concept of structural innovations in the monetary policy equation associated with a given VAR and the time series object of common currency monetary policy surprises which, by definition, do not vary across geographical units. Throughout this study, we are concerned with the former concept of structural innovations to the monetary policy equation implied by a given country-specific VAR structure. For this purpose, we employ nominal measures of monetary policy surprises that have been extensively used as external instruments in monetary policy VAR studies. A formal definition is provided by Equations (1),(2) and (4).

Hence, from the latent structural form in Equation (1), we obtain the implied observable reduced-form system:

$$Y_{c,t} = A_c^{-1}\alpha_c + A_c^{-1}B_{c,1}Y_{c,t-1} + \dots + A_c^{-1}B_{c,p}Y_{c,t-p} + \lambda'_c A_c^{-1}W_{c,t} + A_c^{-1}u_{c,t}, \quad (3)$$

which can be written more compactly as

$$Y_{c,t} = \mu_c + S_{c,1}Y_{c,t-1} + \dots + S_{c,p}Y_{c,t-p} + \theta_c W_{c,t} + \varepsilon_{c,t}, \quad (4)$$

with reduced-form residuals being normally distributed $\varepsilon_{c,t} \sim N(0, \Sigma_c)$ and variance-covariance matrix Σ_c as implied by the dense matrix of contemporaneous coefficients A_c in Equation (1) above.

We employ high-frequency identified monetary policy surprises as instrumental variables for $(u_{c,t}^{mp})$ to infer the impact of monetary policy shocks on bank lending to non-financial corporations in each country considered using local projections to infer impulse response functions.⁶ The following section discusses the identification approach in more detail.

3.2. Identification approach

When inferring the monetary policy shock in Equation (1), we encounter the well-known identification problem: the matrix of contemporaneous coefficients A_c cannot be uniquely identified from observational data without imposing additional structure on the system. Since our aim of identifying monetary policy shocks only requires partial identification, we refrain from full identification of the system in Equation (1) by imposing restrictions on A_c^{-1} . Instead, we rely on external instrumental variables for identification.

While there is no consensus on identifying restrictions, identification uncertainty in monetary policy VARs is even more pronounced at the level of individual euro area countries, which we are concerned with. More precisely, we do not wish to a priori rule out that a monetary policy shock contemporaneously affects all variables in the system in Equation (1), even when using data on a monthly frequency by imposing a Cholesky identification scheme or sign/zero restrictions. This reasoning is consistent with, for example, Nakamura et al. (2024), who discuss similar concerns for VARs involving potentially fast-moving financial variables. Thus, we employ high-frequency identified monetary policy surprises, inferred from price changes of financial assets within short time windows around monetary policy announcements, that satisfy the exogeneity condition with respect to the information set observable to economic agents. Hence, monetary policy surprises capture unanticipated variation in monetary policy and provide a proxy variable for monetary policy shocks. Moreover, high-frequency identified monetary policy surprises can be

⁶Plagborg-Møller and Wolf (2021) show that VAR impulse responses and dynamic responses estimated via local projections are closely related, in particular, VAR impulse responses can be estimated using local projections and vice versa. While this equivalence generally holds for unrestricted lag structures, we consider 12 lags of monthly data in the estimation of impulse response functions and expect effects from higher-order lags to be negligible in the proposed specifications.

employed as instrumental variables for structural monetary policy innovations in the set-up proposed in Equation(1), as they reflect common currency area monetary policy (see Montiel Olea et al., 2021, for the methodology on inference with an external instrument). Moreover, given they serve as valid instrumental variables, they permit to retrieve explicit time series estimates of $\kappa_c u_{c,t}^{mp}$ for $\kappa_c \in \mathbb{R}$. Indeed, the application of a valid external instrument in conjunction with LP-IV or SVAR-IV estimation provides the ideal case of identification (Wolf, 2020). Therefore, we compare local projection results in specifications wherein common euro area monetary policy surprises are directly employed to LP-IV exercises. Moreover, retrieve explicit estimates of $\kappa_c u_{c,t}^{mp}$ to provide empirical guidance on instrument relevance. More detailed explanations are provided in Section 3.4.

3.3. Selection of the instrumental variable

We require an instrumental variable, denoted z_t , in order to implement the methodological steps above. This requires z_t to adhere to the relevance and exogeneity conditions outlined below:

$$E[u_{c,t}^{mp}|z_t] \neq 0 \quad (\text{relevance}) \quad (5)$$

$$E[u_{c,t}^{\neq mp}|z_t] = 0 \quad (\text{exogeneity}) \quad (6)$$

Given an external instrument that satisfies both conditions, we can employ it as proxy variable for monetary policy shocks and use it to infer estimates of structural monetary policy innovations (see e.g., Montiel Olea et al. (2021), Ramey (2016) or for an empirical related application Albrizio et al. (2020)). We use high-frequency identified monetary policy surprises from the Euro Area Monetary Policy Event-Study Database (EA-MPD) (see Altavilla et al., 2019; Jarociński and Karadi, 2020) as instrumental variables. These surprises are constructed from high-frequency movements in yields, stock prices and interest rates in a narrow event window surrounding ECB monetary policy announcements, and are used to identify unanticipated changes in monetary policy.

To this end, we consider euro area monetary policy surprises as particularly useful instrumental variables for the application at hand for two reasons. First, regarding Equation (5), one might argue that the high-frequency movements in financial variables only constitute a proxy variable for the structural innovation in the monetary policy equations in country-specific VARs. However, we deem them to be valid instruments for $u_{c,t}^{mp}$ because they captures surprises in the common currency area monetary policy. Additionally, the application of a weak instrument in external identification approaches has been widely discussed, and potential weakness can be addressed by adjusting confidence intervals when inferring impulse responses accordingly, such that this concern should not, in principle, jeopardize such analyses. Second, we are naturally concerned with the validity of Equation (6). However, euro area monetary policy does not systematically respond to the country-specific macroeconomic conditions (see, e.g., Jiménez et al., 2014, 2022),

which leads us to expect no systematic correlation between the country-specific structural non-monetary policy innovations and the proposed instrumental variables.

For the empirical results presented in the main text, we employ the change in the 12-month overnight interest swap (OIS) rate around monetary policy announcements retrieved from the Euro Area Monetary Policy Event-Study Database (EA-MPD), which matches the maturity of the monetary policy indicator employed in the VARs alongside the *monetary policy factor* by Jarociński and Karadi (2020) that purges high-frequency movement in interest rate from possible monetary policy information effects.⁷

3.4. LP-IV and recovering structural innovations

We present results comparing the estimates of impulse response functions using monetary policy surprises as proxy variables and two-stage least squares local projection (LP-IV) as in Jordà et al. (2015). The high-frequency identified monetary policy surprises aim to measure the unanticipated component of monetary policy (i.e., they measure the structural innovation to the monetary policy equation), in that, when using them as external instruments, we can recover structural variation of monetary policy by regressing the observed VAR reduced-form residuals on the external instrument (see Montiel Olea et al. (2021), Miranda-Agrippino and Ricco (2021)) and retrieve time series of structural innovations up to proportionality constant. This approach permits to explicitly identify an estimate of time series of structural monetary policy time series and allows to compare the relevance of monetary policy instruments in first-stage regressions across countries. We deem this exercise as informative for two reasons. Firstly, as outlined above, it permits to identify estimates of variation in structural innovations to monetary policy in the VARs of Equation (1) which can be compared across countries. Secondly, it provides us with a sanity check with regards to the first-stage regressions employed in LP-IV methodology that likewise employs the set of endogenous variables and the instrument in regressions on the variable to which the shock of interest pertains. Appendix Section Appendix E presents pooled regression results alongside F-statistics for first-stage regression of reduced-form residuals on the two time series of high-frequency surprises employed. Through implementing the above steps we find evidence for stable instrument relevance across countries that is suggestive of homogeneous transmission of monetary policy to country-specific SVARs. The resulting time series of structural innovations (up to scale) are presented in Figures F.11 and F.12 in the appendix alongside the *monetary policy factor* by Jarociński and Karadi (2020). A Block-Bootstrap analysis using 5.000 iterations with minimum block length of 12 months of the first-stage regressions of reduced-form monetary policy residuals in Equation(4) on the instrumental variables has been performed. Mean F-statistics lie between 1.6 and 3.4 for the *monetary policy factor* and between 7.5 and 10.8 for the surprise in 12-month OIS rates. This suggests that the high-frequency change in the 12-month OIS

⁷Refer to Appendix D for a discussion of the monetary policy indicator and Appendix E for a discussion of the choice of the external instrument. Additionally, appendix Figures F.9 and F.10 present results in Section 5 employing the high-frequency surprises in 3-month OIS rates retrieved from EA-MPD for robustness.

rate is a better suited instrument in subsequent LP-IV estimation for the application at hand.

3.5. Local projections allowing for (a)symmetric transmission

The rationale for employing local projections follows from the perspective of minimizing erroneous inference due to identification and estimation uncertainty. Local projection analysis has been shown to yield more reliable inference on impulse responses with regards to unbiasedness of the (local) parameter estimates (implying a trade-off with parameter estimate variance) and is more robust to possible misspecification of the underlying VAR system (see, e.g., Jordà and Taylor, 2025). As argued above, we want to take identification uncertainty seriously, and therefore prioritize unbiasedness of our parameter estimates over narrow inferred confidence intervals. Recent contributions such as Olea et al. (2024) demonstrate that local projections can correctly cover confidence intervals even when the underlying model is strongly misspecified due to the double robust property of the estimator. In addition, any other estimation approach that is to produce even lower parameter estimation variance than the local projection estimate in large samples, must impose additional restrictions on the transmission mechanism (that is, the impulse response function) or the identification of the shocks and on the contemporaneous covariance matrix Σ_c (Olea et al., 2025). Finally, local projections have been shown to yield superior estimation uncertainty with regard to bias in longer-term horizons stretching up to several months as well and the entailed confidence intervals are robust as compared to those implied by VAR identification schemes even under mild model misspecification Olea et al. (2025).

In a first step, for results shown in Section 5.1 we employ the above-mentioned monetary policy surprises retrieved from EA-MPD (z_t) as proxy variables for structural innovations to monetary policy ($u_{c,t}$).

The estimation equation for local projections using the high-frequency identified monetary policy surprise and first-stage fitted values in LP-IV specifications respectively (both denoted as $\hat{u}_{c,t}^{mp}$) is given by:

$$y_{c,t+h} - y_{c,t} = \alpha_{c,h} + \beta_{c,h}^* \hat{u}_{c,t}^{mp} + \psi_{c,h} y_{c,t-1} + \sum_{j=1}^{12} \gamma_{c,h} W_{c,t-j} + \rho_{c,h} X_{c,t} + \epsilon_{c,t+h}, \quad (7)$$

with $\beta_{c,h}^*$ the dynamic country-specific effect of monetary policy shocks on bank lending as the coefficient of interest.

To assess potential asymmetries between restrictive tightening and expansionary easing monetary policy shocks, we estimate a modified specification of the local projection specification that distinguishes between positive (tightening) and negative (easing) monetary policy shocks, such that we have:⁸

⁸See for example, Kho (2025) for an empirical investigation of asymmetric transmission of monetary policy to private bank deposit interest rates in the ECB's monetary policy hiking cycle as of 2022 using local projections.

$$\hat{u}_{c,t}^{mp(+)} = \begin{cases} \hat{u}_{c,t}^{mp} & \text{if } \hat{u}_{c,t}^{mp} \geq 0, \\ 0 & \text{otherwise.} \end{cases} \quad (8)$$

An analogous definition applies to $\hat{u}_{c,t}^{mp(-)}$. We estimate the equation for local projections in long-differences, which is given by

$$y_{c,t+h} - y_{c,t} = \alpha_{c,h} + \beta_{c,h}^{P*} \hat{u}_{c,t}^{mp(+)} + \beta_{c,h}^{N*} \hat{u}_{c,t}^{mp(-)} + \psi_{c,h} y_{c,t-1} + \sum_{j=1}^{12} \gamma_{c,h}' W_{c,t-j} + \rho_{c,h}' X_{c,t} + \epsilon_{c,t+h} \quad (9)$$

for the purpose of inferring the dynamic response of bank lending to positive and negative monetary policy shocks ($\beta_{c,h}^{P*}$ and $\beta_{c,h}^{N*}$) considering a medium-term transmission horizon of 24 months. We include a vector of contemporaneous control variables X_t comprising a dummy variable indicating the COVID-19 pandemic period from February 2020 to May 2020, the unemployment rate, the Euro to US-Dollar exchange rate and a commodity price index as well as all endogenous variables in Equations (1) and (4). The matrix $W_{c,t-j}$ features the other endogenous variables in Equations (1) and (4), namely the 12-month OIS rate as monetary policy indicator, industrial production, the HICP and the external finance premium each included with 12 lags. For a more detailed description of data sources and transformations applied, please refer to Appendix A. As shown above, we augment the estimation equations of local projections by including one lag of the dependent variable $y_{c,t-1}$. According to Olea et al. (2025), lag-augmented confidence intervals are more robust than conventional VAR confidence intervals, and given the strong persistence observed in the dependent variable, this extension of the specification is well justified.⁹

⁹While the baseline lag-augmented local projections do not employ Newey-West standard errors, the results shown in Figures F.5, F.7, F.6 and F.8 are obtained from a specification that includes lag augmentation and incorporates Newey-West standard errors.

4. Data

The dataset covers nine euro area countries, as well as the euro area aggregate, spanning the period from January 2005 to August 2024. The reduced-form VAR models include industrial production and the HICP (both in logarithms) for each country and the euro area aggregate, a corporate bond risk premium measured by the option-adjusted spread (OAS) of the ICE BofA Euro High Yield Index and the 12-month overnight index swap (OIS) rate as the monetary policy indicator. In addition, the VAR models feature a set of exogenous control variables: the national unemployment rate as defined by Eurostat, the CRB Commodity Index (comprising 19 categories of raw materials), the EUR/USD exchange rate, and a dummy variable that captures the onset of the COVID-19 pandemic.¹⁰ The dependent variable is the year-on-year growth rate of total bank lending to non-financial corporations by domestic monetary financial institutions in the euro area. The data are neither seasonally nor calendar adjusted. Figure 1 below highlights the heterogeneity in bank lending to non-financial corporations across euro area countries.¹¹

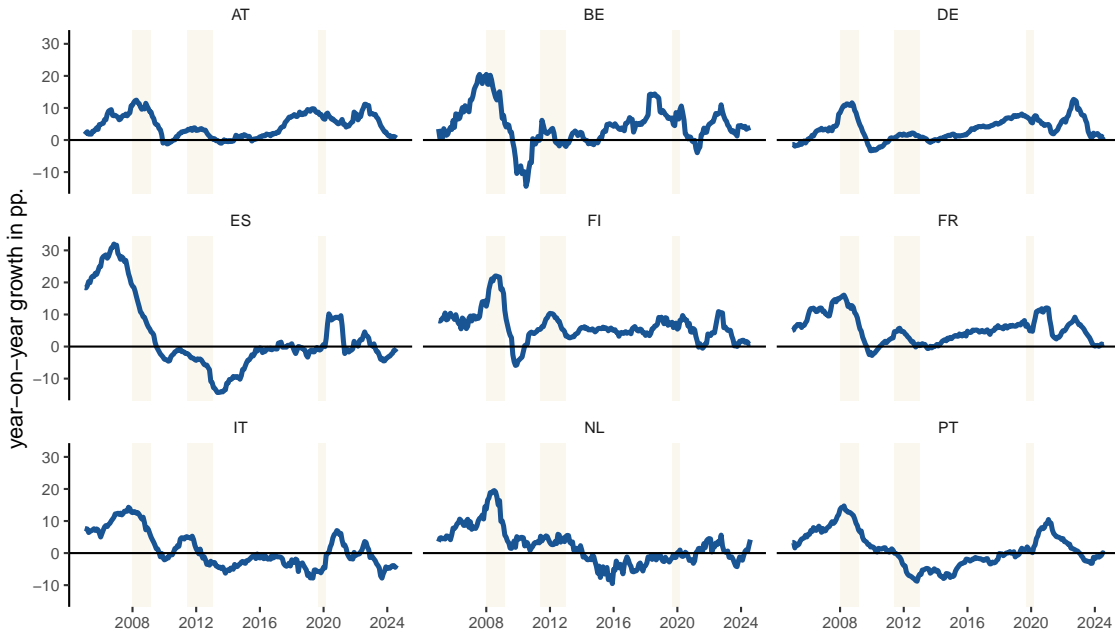


Figure 1: Heterogeneity in bank lending to non-financial corporations. Note: Fig. 1 reports the year-on-year growth rates of total bank lending to non-financial corporations. The shaded regions indicate recessions as defined by the Euro Area Business Cycle Network. Recessions are identified on a quarterly basis, and the first month of each quarter is used as the reference date for the concerned quarter.

¹⁰See Appendix A for detailed information on the variables employed, and Appendix C for a comprehensive discussion of the choice of the nominal interest rate as the monetary policy indicator.

¹¹The presence of international lending relationships could necessitate to account for foreign business cycles in addition to domestic and aggregate euro area dynamics. However, the results remain largely unchanged when restricting the sample to domestic bank lending to non-financial corporations, as is shown in Appendix Figures F.1, F.4, F.2 and F.4. Using AnaCredit data, Altavilla et al. (2024b) document that more than 96% of bank-firm lending relationships in the euro area are domestic. This share remains elevated at 84% when applying volume-weighting of bank-firm relationships. We regard this as the main reason for the similarity of results presented in Appendix Figures F.3 and F.4.

5. Results

The following section presents the main empirical results of the local projections outlined above. First, we discuss the baseline specification for the impulse responses of bank lending to non-financial corporations in Section 5.1 and summarize the country-specific results regarding speed and strength of monetary policy transmission. In the following we initially present results for symmetric specifications as in Equation (7) followed by results that consider differential effects of monetary policy easing and tightening as shown in Equation (9). Next, Section 5.2 compares results to LP-IV estimates, while Section 5.3 explores robustness to alternative sample periods.

5.1. Speed and strength of transmission

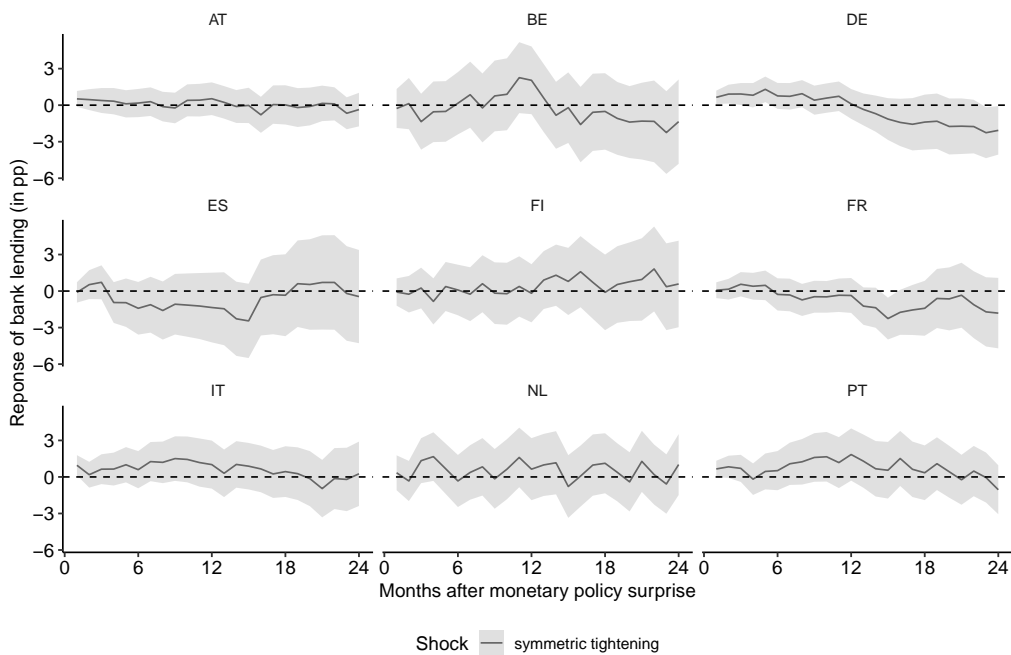


Figure 2: Symmetric impulse response functions of bank lending growth to non-financial corporations to a 25 bps monetary policy surprise. Note: Fig 2 reports IRFs with 90% confidence intervals for symmetry in monetary policy surprises. The results are obtained by using the *monetary policy factor* by Jarociński and Karadi (2020) in local projections with 12 lags of endogenous variables and contemporaneous controls, consistent with the VAR specification described above.

Figure 2 presents impulse response estimates to 25 basis point tightening surprises in the *monetary policy factor* by Jarociński and Karadi (2020) alongside 90% confidence intervals. Reported results are estimated by the linear, symmetric specification as given in Equation (7). Hence, we report IRFs for monetary policy tightening solely due to symmetry. Point estimates generally show small effects of monetary policy tightening surprises and are insignificant for all countries. Figure 3 below presents similar insignificant results for local projections employing 12-month OIS surprises. Just as in Hafemann and Tillmann (2020) who focus on aggregate bank loans, we find Null results of symmetric specifications of monetary policy shocks in Italy, the Netherlands, Portugal and France as well as Finland for most of the transmission horizon considered.

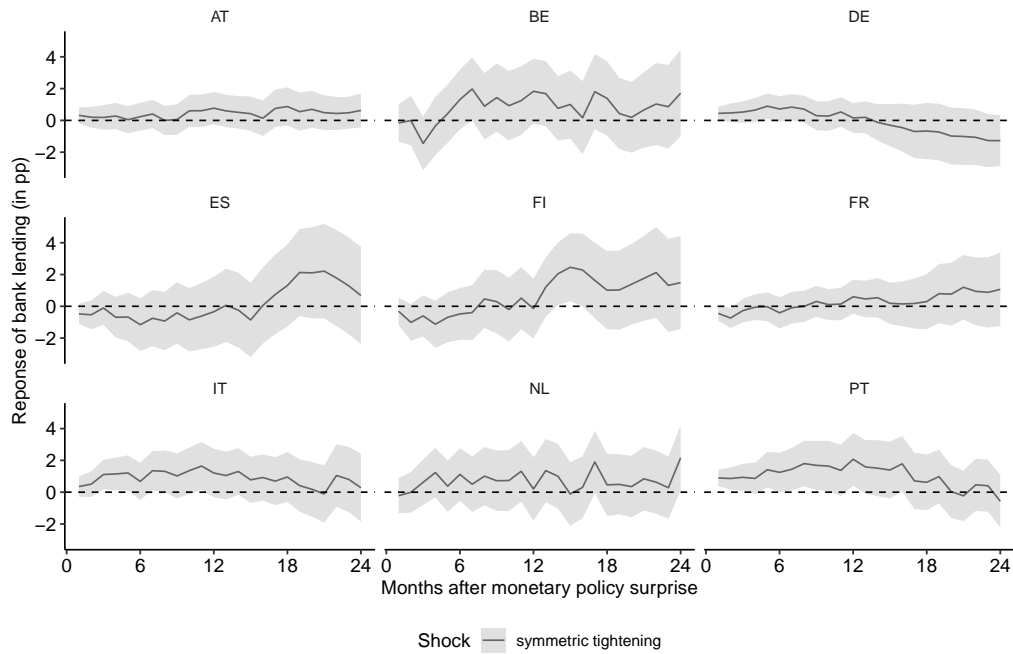


Figure 3: Symmetric impulse response functions of bank lending growth to non-financial corporations to a 25 bps monetary policy surprise. Note: Fig 3 reports IRFs separately with 90% confidence intervals for symmetry in monetary policy surprises. The results are obtained by using the surprise in the 12-month OIS rate in local projections with 12 lags of endogenous variables and contemporaneous controls, consistent with the VAR specification described above.

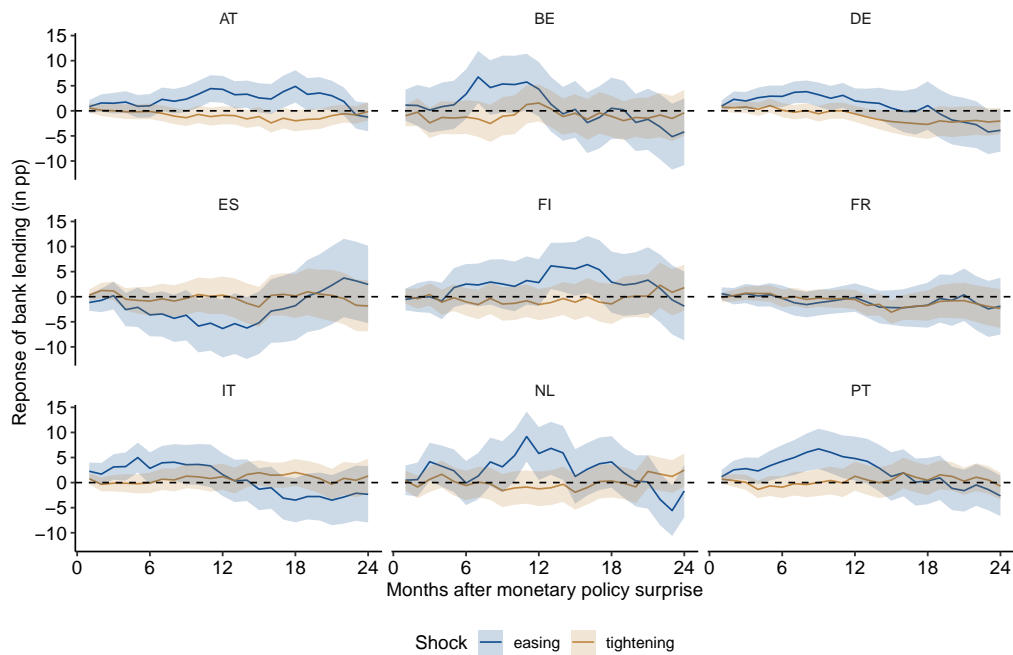


Figure 4: Asymmetric impulse response functions of bank lending growth to non-financial corporations to a 25 bps monetary policy surprise. Note: Fig 4 reports IRFs separately with 90% confidence intervals for easing and tightening surprises. The results are obtained by using the *monetary policy factor* by Jarociński and Karadi (2020) in local projections with 12 lags of endogenous variables and contemporaneous controls, consistent with the VAR specification described above.

Figures 4 and 5 present asymmetric impulse responses of bank lending growth to non-financial corporations, following monetary policy easing and tightening surprises. First, symmetric specifications yield null results when evaluated with 90% confidence intervals while asymmetric specifications show expected signs especially for easing. Second, we find substantial heterogeneity in the responses of bank lending to monetary policy, both in terms of magnitude and peak effects. Third, although the magnitudes of impulse responses vary considerably across countries, the peak response typically occurs around eight to twelve months after the monetary policy surprises. Fourth, the degree of asymmetry varies across countries. For instance, estimates for Germany suggest theory-consistent and more pronounced effects of monetary policy easing. In contrast, for Spain, we cannot reject the null hypothesis of zero effects for tightening and find significant and unexpected results of monetary easing at a horizon of around 12 months. This is again consistent with a similar prior finding in the analysis of monetary policy transmission to aggregate bank lending in Hafemann and Tillmann (2020). Likewise, we find small and insignificant effects in symmetric (and asymmetric) results for France. While peak responses vary considerable, they generally occur approximately 12 to 18 months after initial monetary policy surprises.



Figure 5: Impulse response functions of bank lending growth to non-financial corporations to a 25 bps monetary policy surprise. Note: Fig 5 reports IRFs separately with 90% confidence intervals for easing and tightening surprises. The results are obtained by using high-frequency changes in the 12-month OIS rate in the monetary policy event window retrieved from the EA-MPD in local projections with 12 lags of endogenous variables and contemporaneous controls, consistent with the VAR specification described above.

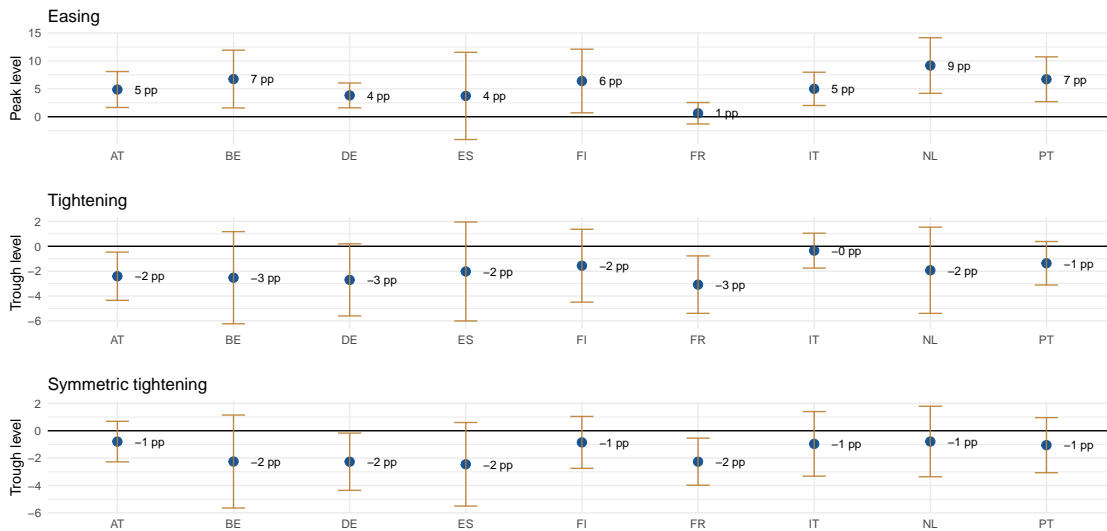


Figure 6: Strength of transmission from IRFs of bank lending growth to non-financial corporations. Note: Fig. 6 reports the peak response to a 25 basis point easing surprise and the trough response to a 25 basis point tightening surprise. The bars denote 90% confidence intervals at the horizon where the peak (or trough) effect occurs. The results are obtained by using the *monetary policy factor* by Jarociński and Karadi (2020) as an external shock.

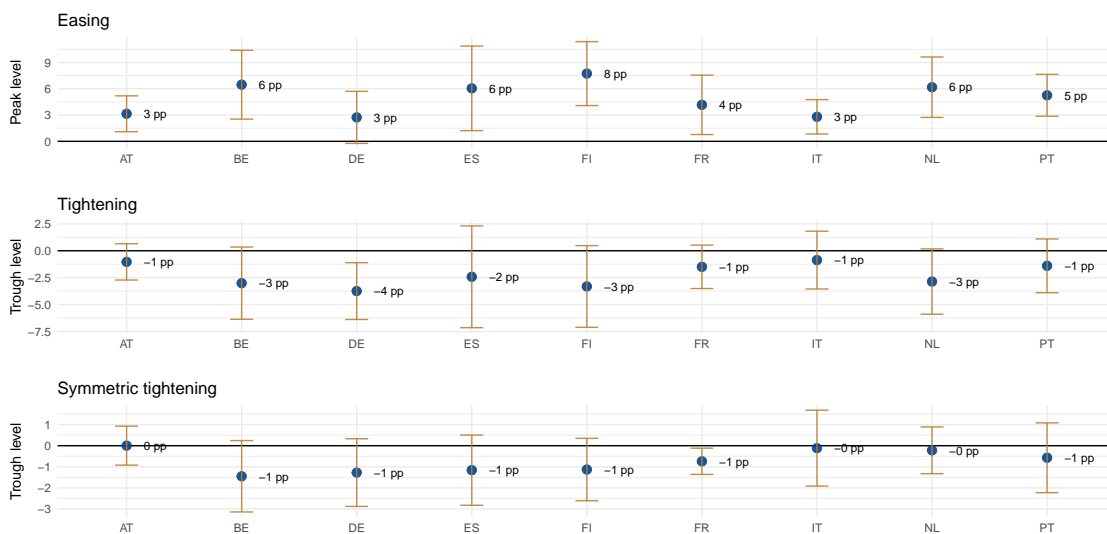


Figure 7: Strength of transmission from IRFs of bank lending growth to non-financial corporations. Note: Fig. 7 reports the peak response to a 25 basis point easing surprise and the trough response to a 25 basis point tightening surprise. The bars denote 90% confidence intervals at the horizon where the peak (or trough) effect occurs. The results are obtained by using high-frequency changes in the 12-month OIS rate in the monetary policy event window retrieved from the EA-MPD.

Figures 6 and 7 display the point estimates of peak cumulative impacts, summarizing the results shown in Figures 2, 4, 3 and 5 above in a more compact form. We find peak impacts of bank lending growth to non-financial corporations in response to 25 basis points tightening shocks of around three percentage points. The peak impacts of equivalent easing shocks fall within the range between one and nine percentage points, suggesting differences between the effects of easing and tightening shocks for most countries. Austria and Finland

constitute notable exceptions, as monetary policy easings yield markedly larger effects on bank lending to non-financial corporations in terms of peak effects. Relative to the recent empirical literature (Hafemann and Tillmann, 2020; Bijsterbosch and Falagiarda, 2014), the peak impact of easing shocks in the Netherlands appear overstated, even when accounting for the relatively wide confidence intervals.

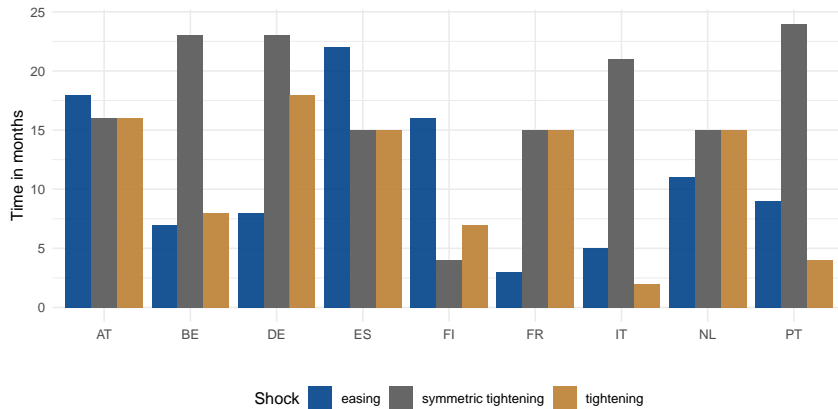


Figure 8: Speed of transmission from IRFs of bank lending growth to non-financial corporations. Note: Fig. 8 reports the time (in months) to the peak response to a 25 basis point easing surprise and the time to the trough response to a 25 basis point tightening surprise. The results are obtained by using the *monetary policy factor* by Jarociński and Karadi (2020) as an external shock.

In a similar vein, Figures 8 and 9 summarize the speed of transmission in terms of months to the peak impact, rather than the magnitude of peak impact. Monetary policy shocks typically reach their peak effect of bank lending 12 to 18 months after the initial occurrence of a shock in most countries. Estimates of time to peak vary considerably across symmetric and asymmetric specifications as prior results highlight. The difference between time to peak for for easing and tightening surprises is especially pronounced in France though qualitatively different across specification employing the *monetary policy factor* and the 12-month OIS rate. Results for Italy also differ substantially in this regard.

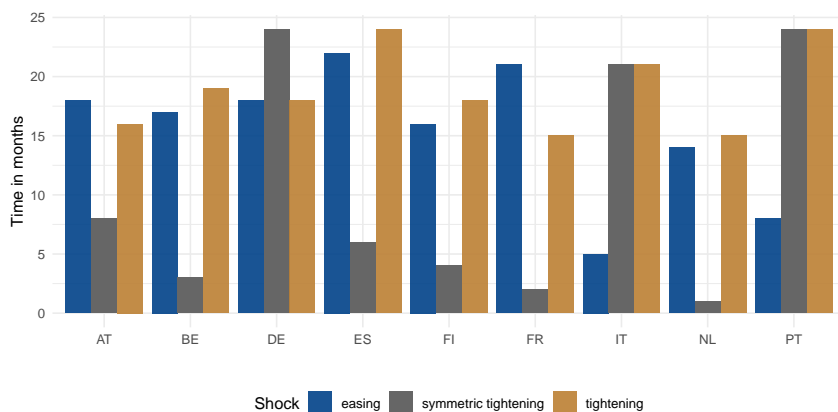


Figure 9: Speed of transmission from IRFs of bank lending growth to non-financial corporations. Note: Fig. 9 reports the time (in months) to the peak response to a 25 basis point easing surprise and the time to the trough response to a 25 basis point tightening surprise. The results are obtained by using high-frequency changes in the 12-month OIS rate in the monetary policy event window retrieved from the EA-MPD.

For Austria, monetary policy easing yields larger positive lending responses while contraction only shows mild negative effects. Results for both asymmetric specifications are similar in Belgium, showing hump-shaped IRFs for monetary policy easing while the impacts are markedly less persistent in the specification employing the *monetary policy factor*. For Germany, results across both specifications are more similar, with roughly symmetric shapes of impulse responses. However, the effects of monetary policy tightening seem to only peak after around 18 months. The specification employing the 12-month OIS surprises yields more symmetric results. Spain constitutes a prominent outlier across all specifications presented in our analysis. While contractionary monetary policy does not produce significant effects, expansionary monetary shows up in counterintuitive negative effects on bank lending up to an estimation horizon of one and a half years. As in most other countries, expansionary monetary policy generates stronger than contractionary measures in Finland and France as well. While the impact of expansionary measures is estimated to be insignificant in France for the specification using the *monetary policy factor*. Impulse responses for monetary policy tightening are essentially flat in Italy as well. In case of the *monetary policy factor* specification, expansionary measures show larger effects that abate after 12 months. Lastly, monetary policy easing indicates a more pronounced bank lending channel in the Netherlands and Italy. Contractionary measures show more limited effects.

Taken together, we find evidence of notable heterogeneity in bank lending in response to monetary policy across countries. Moreover, impulse responses suggest non-linear reactions wherein monetary policy contractions and expansion are transmitted differentially to bank lending. Generally, asymmetry in the above results mainly shows up in larger responses to expansionary versus contractionary monetary policy surprises. This results suggest downward rigidity in bank lending in response to common euro area monetary policy. Although these results suggest marked heterogeneity and non-linearity across euro area countries, the impulse response functions point to synchronization in the bank lending channel. While peak responses to easing and tightening are different across countries, impulse responses build up within several months after initial monetary policy actions and abate within approximately one and a half years time. This results also underpins the longer-term effects of monetary policy, showing that its effects on bank lending materialize with a considerable lag.

The following Section 5.2 presents the above results in comparison to LP-IV methodology.

5.2. Robustness to LP-IV

As outlined in Section 3.3, monetary policy surprises based on EA-MPD data provide a particularly useful set of instrumental variables for our application at hand. Since monetary policy is common across the euro area the proposed instruments are relevant by construction as the measure unanticipated changes in financial variables around in a narrow window around monetary policy announcements. Moreover, common euro area monetary policy does not systematically react to country-specific macroeconomic conditions (Jiménez et al., 2022) which provides a further argument supporting the exogeneity condition. Figures 10 and 11 compare symmetric and asymmetric specifications from the preceding chapter with local projection estimation Equation (9), using external instruments by Jarociński and Karadi (2020) and 12-month OIS surprises in LP-IV estimation (Jordà et al., 2015). Results of high-frequency surprises are similar across the two external instruments for symmetric specifications. Larger differences arise in the asymmetric LP-IV estimates. Again, symmetric specifications yield null results under 90% confidence intervals and asymmetric results are more pronounced and yield significant estimates in most countries. Generally, impulse response function estimates using the *monetary policy factor* as external instrument produce noticeably larger confidence intervals while results employing the 12-month OIS rate are aligned with results prior sections. Point estimates employing the 12-month OIS rate are very much similar to those obtained from using the *monetary policy factor* which is not contaminated by central bank communication effects. Asymmetric results resemble results for bank lending aggregates reported by Hafemann and Tillmann (2020) for monetary policy easing shocks in Spain, Finland and Austria, although we employ different external instruments and a sample covering the steepest monetary policy hiking cycle conducted by the ECB since 2022. Spain presents also represents an outlier in LP-IV estimates with Null effects of contractionary monetary policy and counterintuitive impacts of expansionary monetary policy (again, consider the same striking result reported by Hafemann and Tillmann (2020) for aggregate bank loans). Moreover, our results suggest stronger effects of ECB policy on bank lending to non-financial corporations when accounting for the possibility of asymmetry in the bank lending channel than suggested for aggregate bank lending to the non-financial sector reported in the aforementioned empirical contribution. Overall, the comparison reinforces our earlier results, showing that the estimated impulse responses generally show the expected, theory-consistent signs, while magnitudes differ across countries. Peak impacts typically materialize one to two years after a monetary policy shock. Although the magnitude of the peak dynamic effects varies across external instruments employed by several percentage points, the timing of the peak effect is very similar across specifications.

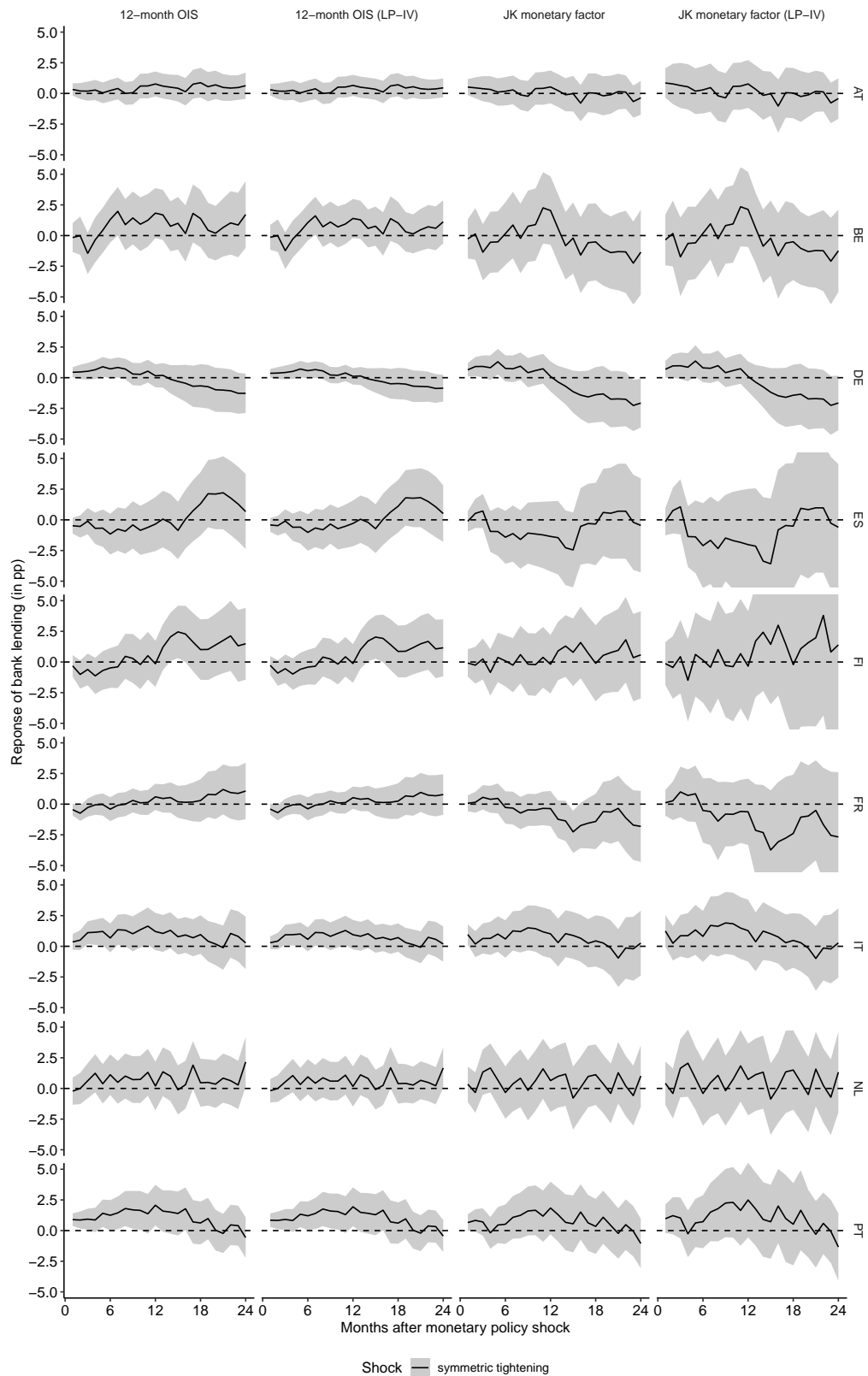


Figure 10: Symmetric impulse responses of bank lending to a 25 basis point monetary policy shock. Note: Fig. 10 presents a comparison to LP-IV estimates and shows 90% confidence intervals. The results are obtained by local projection exercises based on the methodology as described above. Alternative estimates are derived by directly incorporating the external instrument in local projections (LP-IV).

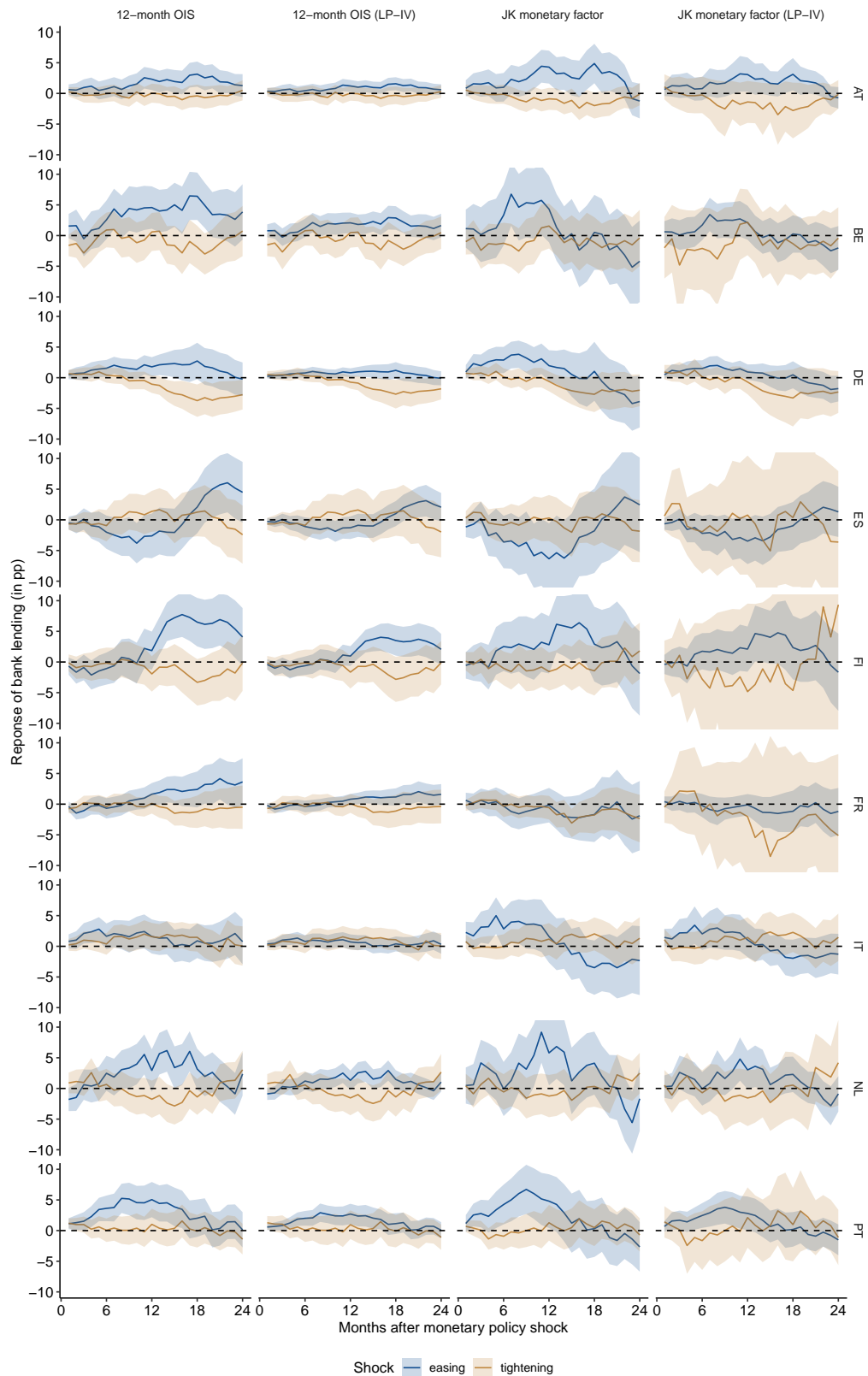


Figure 11: Asymmetric impulse responses of bank lending to a 25 basis point monetary policy shock. Note: Fig. 11 presents a comparison to LP-IV estimates and shows 90% confidence intervals. The results are obtained by local projection exercises based on the methodology as described above. Alternative estimates are derived by directly incorporating the external instrument in local projections (LP-IV).

5.3. Robustness to alternative sample periods

The effects of monetary policy might not only vary across regions but also across time. Zlobins (2025) for instance, explicitly consider time-varying effects of monetary policy in a recent empirical study for the euro area. In addition to the above discussion of different instruments and methodology, the following section presents results for different end points of our sample.

Figures 12, 13, 14 and 15 present the baseline estimates from Section 5.1 for various subsamples. To assess robustness, we replicate the exercise in Section 5 by restricting the data to observations up to January 2018, prior to the onset of the COVID-19 pandemic and the associated supply chain disruptions and inflationary pressures. In addition, we examine subsamples ending in January 2020, immediately before the recent monetary policy hiking cycle, and in October 2023, corresponding to its peak. This exercise provides insights into the stability of monetary policy passthrough to bank lending in recent years, potentially susceptible to multiple structural breaks in macroeconomic dynamics. Recent findings by Zlobins (2025) indicate that monetary policy transmission to key macroeconomic variables in the euro area has remained broadly stable, based on an extensive Bayesian VAR analysis with time-varying parameters. Impulse response function estimates do not produce significant results in symmetric specifications (with small negative effect of contractionary monetary policy in France after approximately 12 to 18 months as the only exception in Figure 12). In Figure 13 monetary policy easing exhibits small but positive effects that quickly subside when restricting the endpoint of the sample to 2018 Q1 while effects become markedly more pronounced when including data up until 2024 Q3 in the case of Finland. In asymmetric specifications employing the 12-month OIS rate shown in Figure 15, estimates of the impact of expansionary monetary policy on bank lending to non-financial corporations become noticeably more pronounced when extending the estimation sample. Broadly speaking, the shape of impulse response functions does not materially change across estimation periods particularly for asymmetric specifications with some exceptions. While our methodological framework does not allow for the explicit incorporation of time-varying effects, we deem the broad alignment of impulse response stability with those of Zlobins (2025) as supportive of our underlying modeling assumptions and specifications. Similarly, our estimates suggest that the passthrough of monetary policy shocks to bank lending has remained largely stable during the recent monetary hiking cycle. This holds despite the unprecedented volume of excess central bank reserves in the banking system.

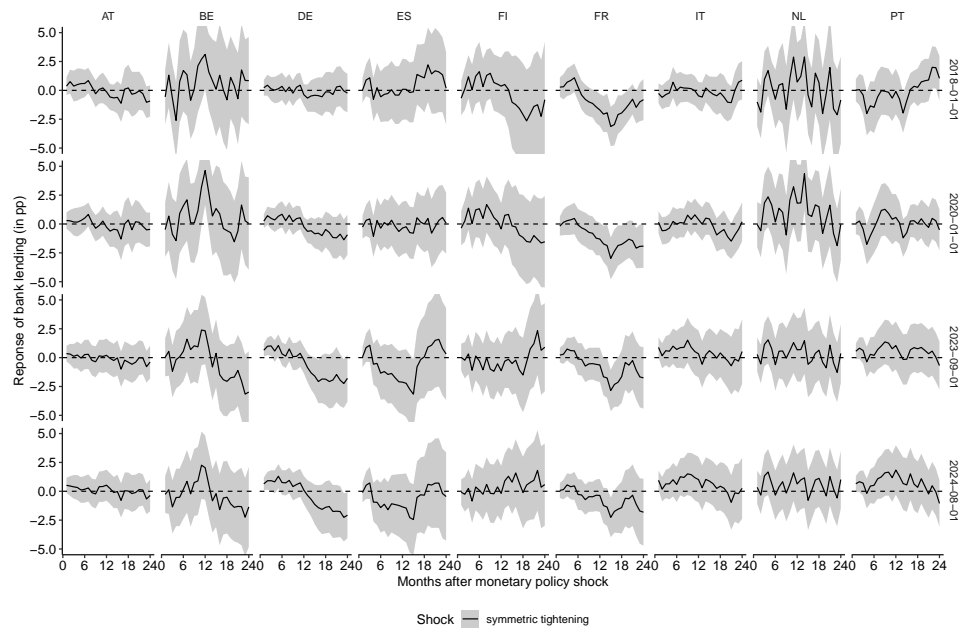


Figure 12: Symmetric impulse responses with extending sample over time with 90% confidence intervals. The results are obtained by using the *monetary policy factor* by Jarociński and Karadi (2020) as an external instrument.

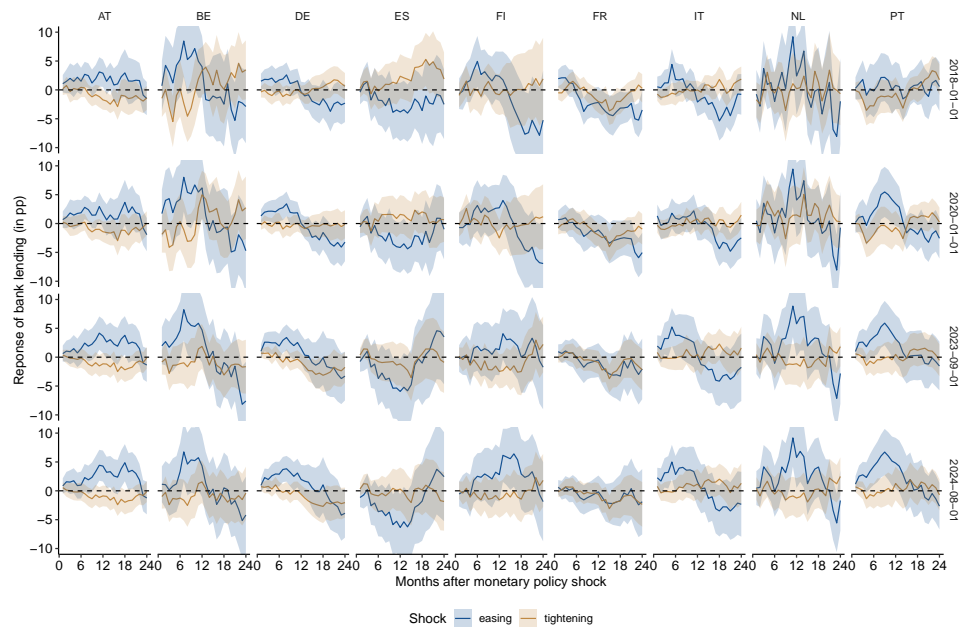


Figure 13: Asymmetric impulse responses with extending sample over time with 90% confidence intervals. The results are obtained by using the *monetary policy factor* by Jarociński and Karadi (2020) as an external instrument.

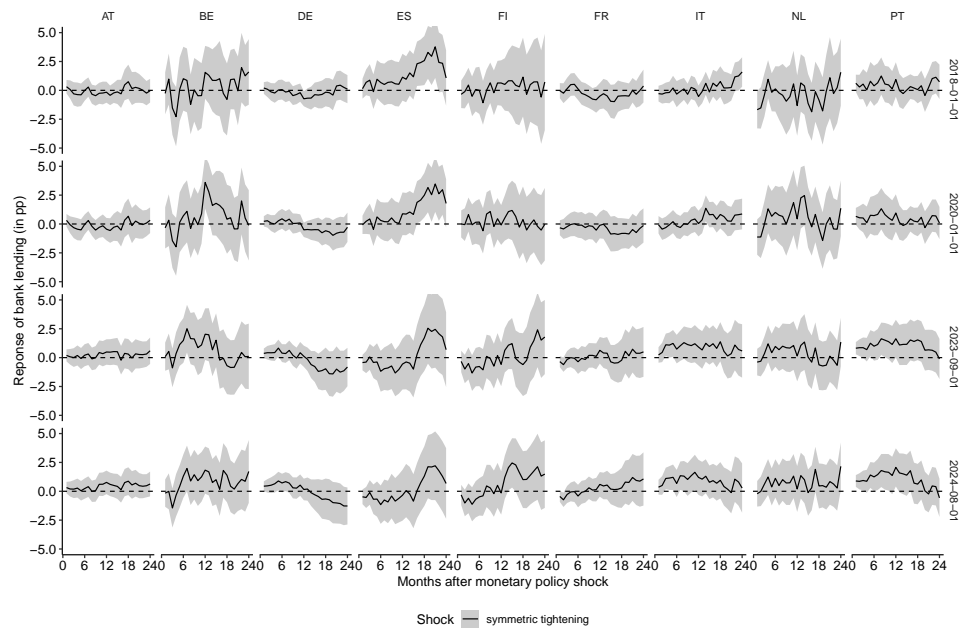


Figure 14: Symmetric impulse responses with extending sample over time with 90% confidence intervals. The results are obtained by using high-frequency changes in the 12-month OIS rate in the monetary policy event window retrieved from the EA-MPD.

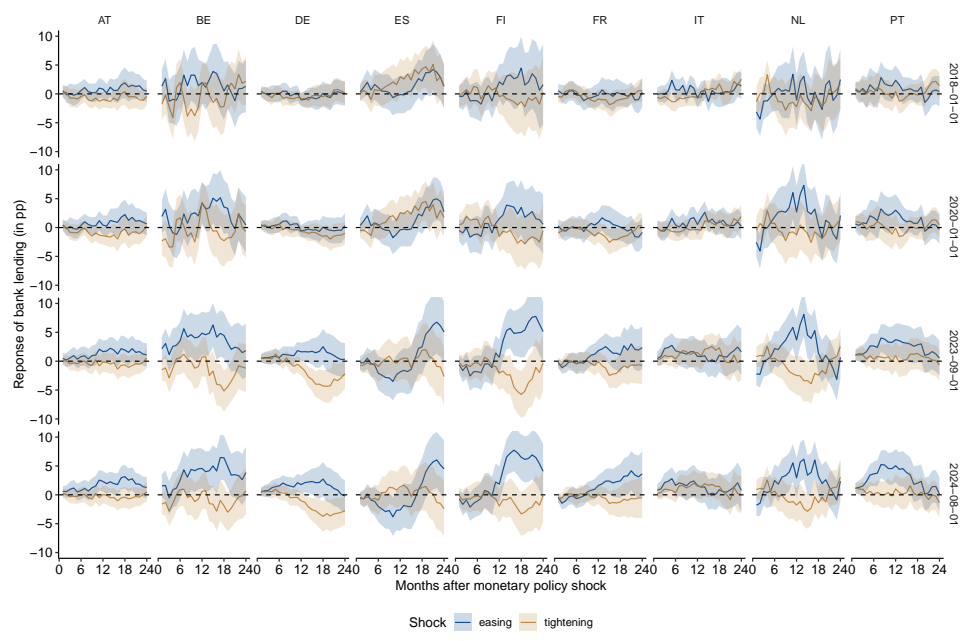


Figure 15: Asymmetric impulse responses with extending sample over time with 90% confidence intervals. The results are obtained by using high-frequency changes in the 12-month OIS rate in the monetary policy event window retrieved from the EA-MPD.

In general, the findings reported in Section 5 also holds for the subsamples considered. Symmetric specifications yield small and insignificant effects while local projections that permit asymmetric transmission yield larger and more often times significant and expected dynamic effect estimates.

6. Conclusion

We study the transmission of common euro area monetary policy to bank lending to non-financial corporations on the level of individual euro area countries. We propose country-specific specifications in the spirit of Gertler and Karadi (2015) and employ high-frequency identified measures of monetary policy surprises as external instruments in local projections. We link the reduced-form representation of country-specific macroeconomic structures to common euro area monetary policy surprises to evaluate instrument relevance across countries in LP-IV estimation. Subsequently, we compare results of direct applications of high-frequency identified monetary policy surprises to LP-IV estimates. Moreover, we assess the robustness of inferred impulse response functions across time by evaluating their stability in sample including and excluding the COVID-19 period and the recent monetary policy hiking cycle. We cannot reject the null hypothesis of a stable passthrough of monetary policy shocks to bank lending since the onset of the COVID-19 pandemic and the ensuing inflationary shock and monetary policy hiking cycle in line with recent empirical findings. Moreover, we explicitly consider the possibility of asymmetries in the bank lending channel on the level of individual euro area countries.

Our main findings are threefold: First, while peak impacts of monetary policy measures show considerable variation across countries, transmission is broadly synchronous across the common currency area in that the impact on bank lending peaks between one and two years after the initial monetary policy shocks, consistent with the medium-term orientation of monetary policy. Second, despite similar timing, magnitudes differ markedly across countries; the largest peak effect is up to five times as large in the most affected country compared to the least affected one, implying substantial cross-country heterogeneity in the sensitivity of bank lending dynamics to monetary policy. Third, our results generally indicate pronounced asymmetries in the transmission of monetary policy shocks. While the euro area aggregate suggests roughly symmetric transmission of monetary policy shocks, at the country level, expansionary monetary policy shocks propagate faster and more forcefully than contractionary shocks. Taken together, these results indicate that ECB monetary policy affects bank lending in a strong but possibly asymmetric manner.

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Appendix A. Variable definitions and sources

Variable	Unit	Transformation	Source	Availability
<i>Model variables</i>				
Bank-lending to non-financial corporations	index of notional stock, total (year-on-year growth rate)	neither seasonally nor working-day adjusted	ECB SDW	2005M01-2024M08
Industrial production	index of total	$\log(\cdot)*100$	Eurostat	2005M01-2024M08
Harmonized Index of Consumer Prices	index of total	$\log(\cdot)*100$	Eurostat	2005M01-2024M08
12-month Overnight Interest Rate Swap (OIS 12-month)	annualized percentage interest rate	$\text{id}(\cdot)$	ECB SDW	2005M01-2024M08
Option-Adjusted Spread (OAS) of the ICE BofA Euro High Yield Index	index	arithmetic average across daily observations	FRED St. Louis	2005M01-2024M08
COVID-19 Dummy (2020M02-2020M05)	dummy variable	$\text{id}(\cdot)$	NBER definition	2005M01-2024M08
Commodity CRB Index	index	$\text{id}(\cdot)$	Bloomberg	2005M01-2024M08
USD-EUR exchange rate	exchange rate (spot rate)	$\text{id}(\cdot)$	FRED St. Louis	2005M01-2024M08
Unemployment rate	percentage	$\text{id}(\cdot)$	Eurostat	2005M01-2024M08
<i>External instruments</i>				
Jarociński and Karadi (2020) <i>monetary policy factor (MP median)</i>	basis points	$\text{sum}(\cdot)$ by month or set to zero if missing	retrieved from personal website	2005M01-2024M08
Δ OIS 12-month (<i>Monetary Event Window</i>)	basis points	$\text{sum}(\cdot)$ by month or set to zero if missing	Euro Area Monetary Policy Event-Study Database (EA-MPD)	2005M01-2024M08

Table A.1: Variable definitions and sources

Appendix B. Transformations of external instruments

The external instrumental variables are aggregated to monthly frequency because the original data are available at daily frequency, while the dynamic relationships in the VAR are modeled at a monthly frequency. The dataset provided by the Euro Area Monetary Policy Event-Study Database (EA-MPD) contains 33 instances of multiple monetary policy shocks across 273 monthly observations.¹² We aggregate monetary policy shocks to a monthly frequency discarding any within-month dynamics by simply summing shock realizations within a given month. For months without any identified monetary policy shocks, the value of the external instrument is set to zero.

¹²As retrieved on May 30, 2025.

Appendix C. Monetary Policy transmission and the New Keynesian model

In the theoretical foundation of the New Keynesian model with nominal rigidities, the monetary authority can influence longer-term real interest rates by committing to a path for the short-term nominal policy rate, due to the presence of nominal frictions, such as price stickiness. Together with the expectations hypothesis of the term structure, the approximate return on an m -period nominal bond is given by Gertler and Karadi (2015) as

$$i_t^m = E_t \frac{1}{m} \left\{ \sum_{j=0}^{m-1} i_{t+j} \right\} + \xi_t^m. \quad (\text{C.1})$$

Accounting for inflation yields the real return on the riskless bond:

$$i_t^m - E_t \pi_t^m = E_t \frac{1}{m} \left\{ \sum_{j=0}^{m-1} (i_{t+j} - \pi_{t+j}) \right\} + \xi_t^m, \quad (\text{C.2})$$

where i_t denotes the nominal bond yield, π_t the period-to-period inflation, ξ_t^m the term premium and E_t the expectations operator at time t .

Using Equations (C.1) and (C.2), we find that the monetary authority can influence longer-term real interest rates and thereby affect economic activity and inflation by steering the short-term nominal interest rate (subject to the degree of nominal rigidities), assuming ξ_t^m constant in a local region around the steady state. When financial markets are not frictionless, private borrowing rates will exceed the riskless government bond real return in Equation (C.2) by an *external finance premium* (τ_t^m) (see Gertler and Karadi, 2015) such that

$$i_t^{mp} = i_t^m + \tau_t^m. \quad (\text{C.3})$$

The *excess return on a private bond* over government bond holdings is given by

$$\chi_t^m = \xi_t^m + \tau_t^m. \quad (\text{C.4})$$

While a perfectly competitive and frictionless baseline of the macro-economy implies no significant effects of changes in nominal short-term monetary policy interest rates, Gertler and Karadi (2015) find strong evidence that monetary policy influences the *external finance premium* in Equation (C.2). Monetary policy can, by influencing short-term nominal interest rates, introduce sizable movements in long-term refinancing costs due to the reaction of term premia and credit spreads (Gertler and Karadi, 2015). In our estimation equations aggregate growth in bank lending is a function of long-term real interest rates, and depends on monetary policy interest rates, an external finance premium and inflation.

One may argue that the effectiveness of the monetary authority in influencing longer-term real interest rates—and hence output and inflation—depends on the conduct of monetary policy and the set of available instruments. To the extent that conventional

monetary policy implemented through adjustments of nominal short-term interest rates, is complemented by unconventional monetary policy tools (e.g., explicitly targeting longer-term maturities on the yield curve by providing information about future conduct), one could expect more pronounced effects of monetary policy. Since we are concerned with heterogeneous monetary policy transmission in the euro area following the Great Financial Crisis, we consider the impact of unconventional monetary policy tools on bank lending to non-financial corporations, and their varying effectiveness, to be central for assessing the total effect of monetary policy. We abstain from proposing a modeling approach that incorporates the explicit effects of unconventional monetary policy such as Boeckx et al. (2019) and employ high-frequency identified monetary policy surprises by Jarociński and Karadi (2020) that explicitly identify monetary policy surprises accounting for central bank information effects. In this regard, the high-frequency change in 12-month OIS rates could be contaminated by central bank information effects and the application of instruments by Jarociński and Karadi (2020) serves as a robustness check to employing 12-month OIS rate surprises.

Appendix D. Choice of the monetary policy indicator

To obtain guidance on the monetary policy indicator, we examine high-frequency changes in the Overnight Index Swap (OIS) curve for the euro area money market around monetary policy announcements (sourced from the Euro Area Monetary Policy Event-Study Database), as reported in the tables below. This provides a market-based and timely indicator for the effect of monetary policy announcements in the euro area (see Altavilla et al., 2019).

When regressing changes in longer-term nominal interest rates around monetary policy announcements on surprises in shorter-term maturities, we find statistically significant and strong positive correlations across all ECB announcement windows available in the EA-MPD. The comovement of surprises in very short-term overnight index swap rates to a medium-term maturity (5-year OIS rate) is strongest during the press conference window following the announcements of monetary policy interest rates. This finding supports Equation (C.1) by indicating that the monetary authority is able to steer longer-term nominal interest rates through its announcements of short-term nominal policy interest rates (consistent with a largely parallel shift in nominal interest rates across OIS maturities).

Variation in the 1-year OIS rate is strongly correlated with movements in the longer-term 5-year OIS rate. Contemporary comovement in interest rates during monetary policy announcements is notably smaller for the 5-year and the 1-week OIS rates. Taken together, these observations motivate our choice of the 1-year OIS rate as the monetary policy indicator variable, thereby also following Gertler and Karadi (2015) and striking a balance by employing an interest rate of medium-term maturity.¹³

¹³While a comparison of first-stage F-statistics across OIS maturities, following the approach of Gertler and Karadi (2015), would suggest using a very short-term indicator—typically between one and three months—we employ an interest rate that comoves more closely with longer-term nominal yields during

	<i>Dependent variable: Change in OIS 5Y</i>		
	Monetary event	Press release	Press conference
Change in OIS 1W	0.297* (0.158)	0.208** (0.103)	0.660* (0.393)
Observations	117	117	117
R ²	0.029	0.033	0.024
Adjusted R ²	0.021	0.026	-0.015
F-Statistic (df = 1; 117)	3.523	4.099	2.83

Source: EA-MPD. * $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

Table D.2: Regression results for $\Delta OIS_5Y = \beta \Delta OIS_1W + \epsilon$ around monetary policy announcements using data from 2011-08-04 to 2024-07-18.

	<i>Dependent variable: Change in OIS 5Y</i>		
	Monetary event	Press release	Press conference
Change in OIS 1Y	0.857*** (0.059)	0.763*** (0.058)	0.987*** (0.063)
Observations	117	117	117
R ²	0.645	0.597	0.676
Adjusted R ²	0.642	0.593	0.673
F-Statistic (df = 1; 117)	212.6	173.0	244.2

Source: EA-MPD. * $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

Table D.3: Regression results for $\Delta OIS_5Y = \beta \Delta OIS_1Y + \epsilon$ around monetary policy announcements using data from 2011-08-04 to 2024-07-18.

We employ the 12-month OIS rate as monetary policy indicator, as its movements around monetary policy announcements are most closely aligned with shifts in longer-term interest rates that steer investment decisions. However, longer-term interest rates may be contaminated by the term premium ξ_t^m (Corsetti et al., 2022). Hence, we strike a balance by choosing the 1-year overnight index swap (OIS) rate as monetary policy indicator and compare results with the *monetary policy factor* by Jarociński and Karadi (2020).

monetary policy event window, as argued above. Figures F.9 and F.10 present employing surprises in 3-month OIS rates for robustness.

Appendix E. First-stage regressions

Recall the definition of observed reduced-form residuals in Equation (4), expressed as a linear combination of latent structural shocks. Without loss of generality, let the monetary policy indicator variable be ordered first (i.e., $u_t^1 = u_t^{mp}$). Then the corresponding reduced-form residual can be written as

$$\varepsilon_{c,t}^1 = \varepsilon_{c,t}^{mp} = A_c^{-1} u_t = \sum_{j=1}^k A_{c,1,j}^{-1} u_{c,t,j} = A_{c,1,1}^{-1} u_{c,t}^{mp} + \sum_{j=2}^k A_{c,1,j}^{-1} u_{c,t,j}, \quad (\text{E.1})$$

and let

$$E[u_{c,t}^{mp} z_t] = \phi_c \text{ with } \phi_c \in \mathbb{R}^+. \quad (\text{E.2})$$

Consequently, let $A_{c,1,1}^{-1}$ denote the contemporaneous impact of monetary policy shocks in the monetary policy equation. By the exogeneity condition, we have

$$z_t \perp\!\!\!\perp u_t^{\neq mp} \Rightarrow E[u_t^{\neq mp} z_t] = 0 \text{ and } E[\varepsilon_{c,t}^{mp} z_t] = A_{c,1,1}^{-1} \phi_c = \kappa_c.$$

Since the structural shocks $u_{c,t}$ are not observed, the previous expressions cannot be used directly. However, the reduced-form residuals $\varepsilon_{c,t}$ are observable, and the external instrument z_t can be used to recover variation in the reduced-form innovations $\varepsilon_{c,t}^{mp}$ that is solely driven by the structural shock $u_{c,t}^{mp}$ up to scale (Gertler and Karadi, 2015; Montiel Olea et al., 2021). This identification strategy is valid under the representations of Equations (4) and (1) and the assumptions in Equations (5) and (6).

Thus, the projection of observable $\varepsilon_{c,t}^{mp}$ onto the external instrument z_t identifies the structural monetary policy shock $u_{c,t}^{mp}$ up to a scale factor κ_c .

Given a valid instrument, we can recover an estimate of the structural monetary policy shock up to a scale factor in Equation (1) from observational data in Equation (4) (see, e.g., Miranda-Agrippino and Ricco, 2021). In a first step, we obtain reduced-form residuals from the monetary policy equation in each VAR by estimating Equation (4) by ordinary least squares (i.e., by partialling out lagged endogenous variables from the monetary policy indicator).

$$\varepsilon_{c,t} = Y_{c,t} - \Xi_c X_{c,t} \quad (\text{E.3})$$

$$\Xi_c = (X_t^T X_t)^{-1} X_t^T Y_{c,t} \quad (\text{E.4})$$

$$X_t = \{\mathbf{1}, Y_{c,t-1}, \dots, Y_{c,t-12}, W_{c,t}\} \quad (\text{E.5})$$

We have $\varepsilon_{c,t} \sim N(0, \Sigma_c)$ where Σ_c is full rank with non-zero off-diagonal elements (i.e., implying contemporaneous covariance of innovations $\varepsilon_{c,t}$), such that we do not observe the structural shocks of interest but only linear combinations. As discussed above, the reduced-form residual related to monetary policy is a linear combination of all structural innovations of the system of macroeconomic variables.

$$\varepsilon_{c,t}^{mp} = A_{c,1,1}^{-1}u_{c,t}^{mp} + \sum_{j=2}^k A_{c,1,j}^{-1}u_{c,t,j} \quad (\text{E.6})$$

Only ε_t is observed, while the remaining terms in the linear combination represent latent variables. In order to estimate unobserved $u_{c,t}^{mp}$ we employ an instrumental variable that is relevant to the structural monetary policy shocks and exogenous to all other structural innovations $u_{c,t}^{\neq mp}$. Table E.4 below displays the coefficients of first-stage regressions and F-statistics using external instruments and the reduced-form residuals in Equation (1).

	<i>Dependent variable: Reduced-form residual ε_t^{mp}</i>	
	Jarocinski and Karadi (2020)	ΔOIS12M
z_t	0.363*** (0.063)	0.715*** (0.046)
Degrees of freedom	2,123	2,123
R ²	0.016	0.101
Adj. R ²	0.015	0.101
F Stat.	33.8	240
Min. F-Stat. (country)	2.3 (AT)	21.6 (FI)

* $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

Table E.4: Pooled instrumental variable regression results of $\varepsilon_{c,t}^{mp} = \beta_c z_t + \epsilon_{c,t}$. Note: Variables employed as instrumental variables in the first-stage regression pooled across countries include Jarociński and Karadi (2020) *monetary policy factor* estimate and the change of OIS 12-month interest swap rate in the monetary policy event window as retrieved from the EA-MPD. The sample covers 2005-01-01 to 2024-08-01.

Table E.4 above presents the results when partialling out lagged endogenous variables from the VAR in Equation (4) from the monetary policy indicator and regressing them on the two high-frequency monetary policy surprises. The high-frequency monetary policy surprises by Jarociński and Karadi (2020) yield a low F-statistic in the case of Austria while the minimum F-statistic for 12-month OIS surprises exceeds twenty in the case of Finland. This result provides an additional motivation for employing 12-month OIS rates alongside the *monetary policy factor* by Jarociński and Karadi (2020).

Appendix F. Further results

Appendix F.1. Restriction to domestic lending (symmetric)

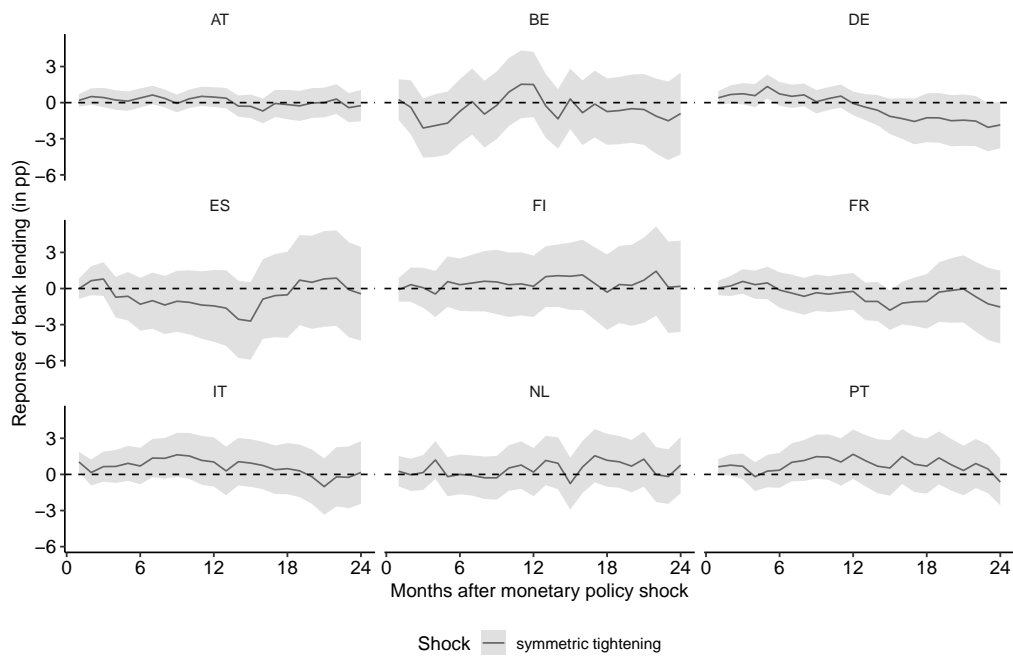


Figure F.1: Symmetric impulse response functions of bank lending growth to non-financial corporations to a 25 bps monetary policy surprise in the *monetary policy factor*. Note: Fig. F.1 presents the result for bank lending to domestic non-financial corporations only.

Appendix F.2. Restriction to domestic lending (asymmetric)

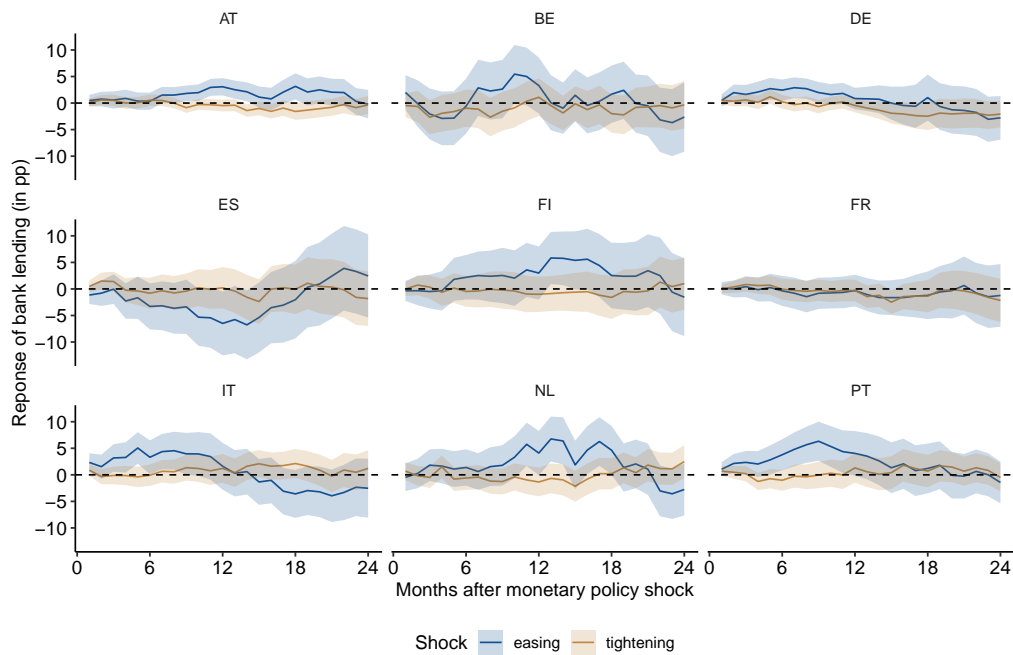


Figure F.2: Asymmetric impulse response functions of bank lending growth to non-financial corporations to a 25 bps monetary policy surprise in the *monetary policy factor*. Note: Fig. F.2 presents the result for bank lending to domestic non-financial corporations only.

Appendix F.3. Restriction to domestic lending (symmetric)

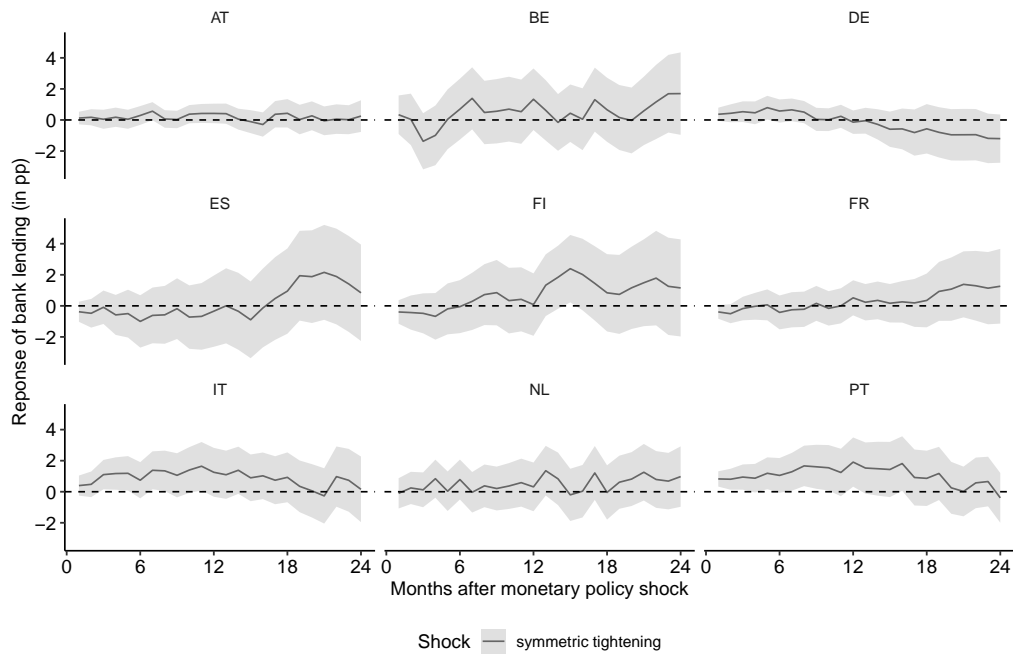


Figure F.3: Symmetric impulse response functions of bank lending growth to non-financial corporations to a 25 bps monetary policy surprise in the 12-month OIS rate. Note: Fig. F.3 presents the result for bank lending to domestic non-financial corporations only.

Appendix F.4. Restriction to domestic lending (asymmetric)

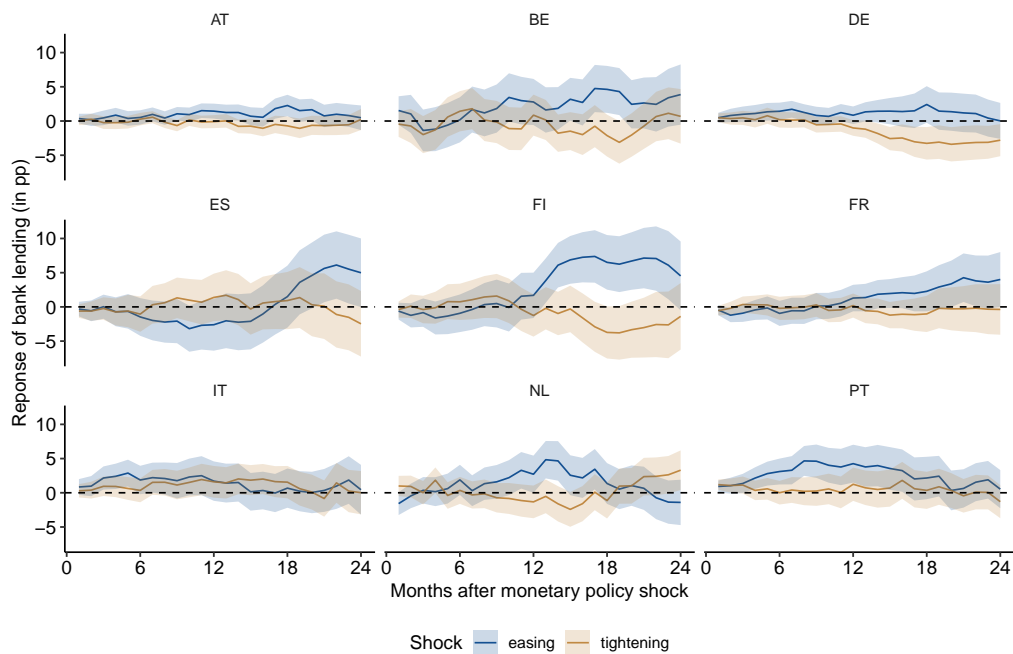


Figure F.4: Asymmetric impulse response functions of bank lending growth to non-financial corporations to a 25 bps monetary policy surprise in the 12-month OIS rate. Note: Fig. F.4 presents the result for bank lending to domestic non-financial corporations only.

Appendix F.5. Newey-West standard errors

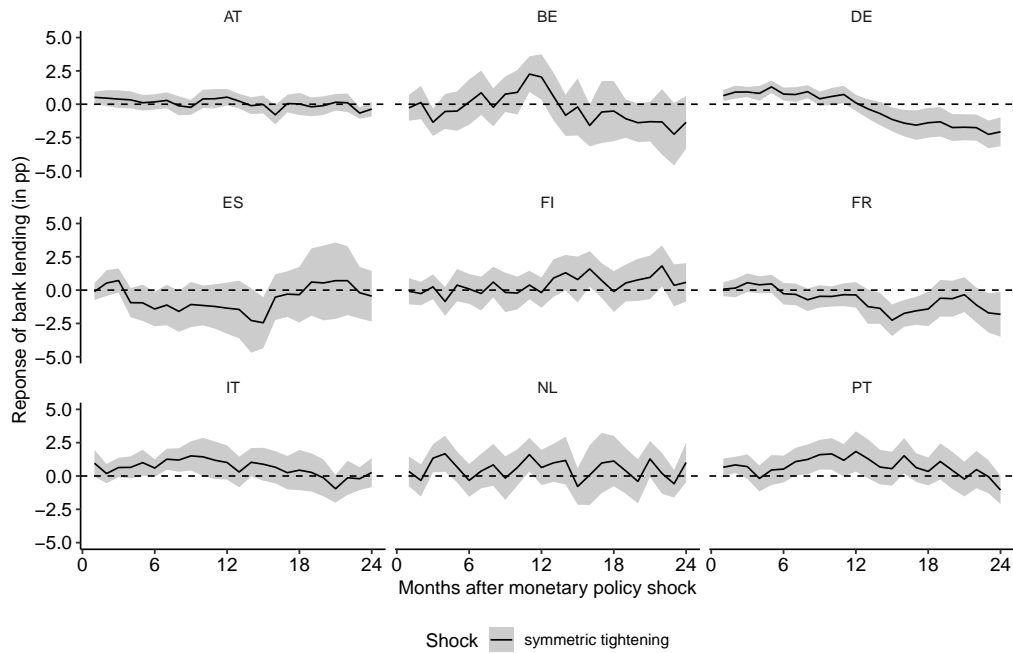


Figure F.5: Symmetric impulse response functions of bank lending growth to non-financial corporations to a 25 bps monetary policy surprise employing the *monetary policy factor*. Note: Fig. F.5 presents the IRFs as shown in Section 5.1 with Newey-West corrected standard errors using 13 lags and one lag of the dependent variable for lag augmentation.

Appendix F.6. Newey-West standard errors (asymmetric)

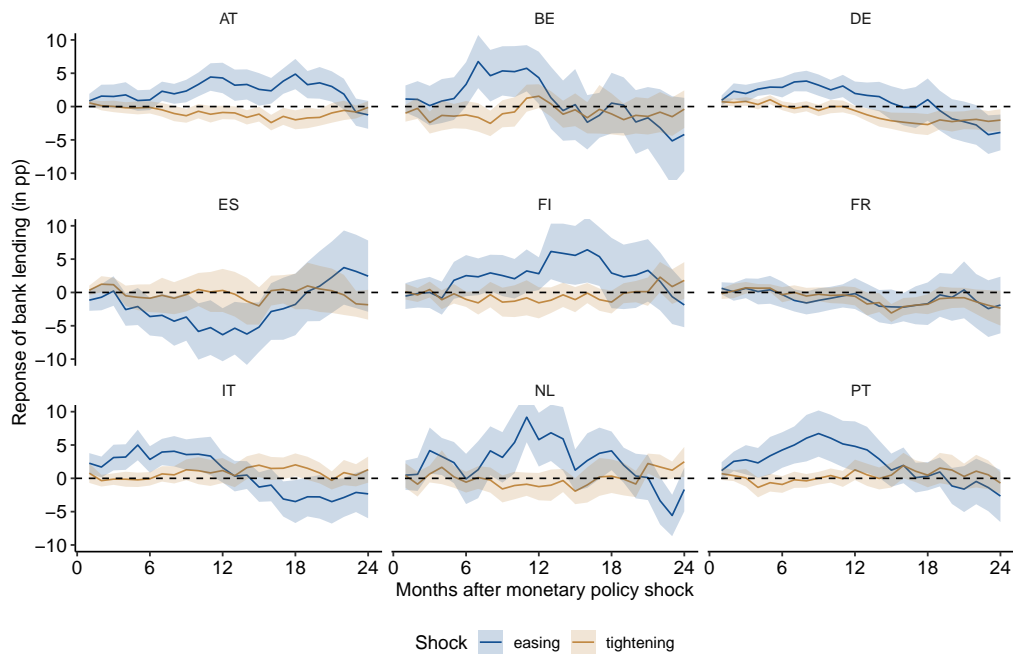


Figure F.6: Asymmetric impulse response functions of bank lending growth to non-financial corporations to a 25 bps monetary policy surprise employing the *monetary policy factor*. Note: Fig. F.6 presents the IRFs as shown in Section 5.1 with Newey-West corrected standard errors using 13 lags and one lag of the dependent variable for lag augmentation.

Appendix F.7. Newey-West standard errors

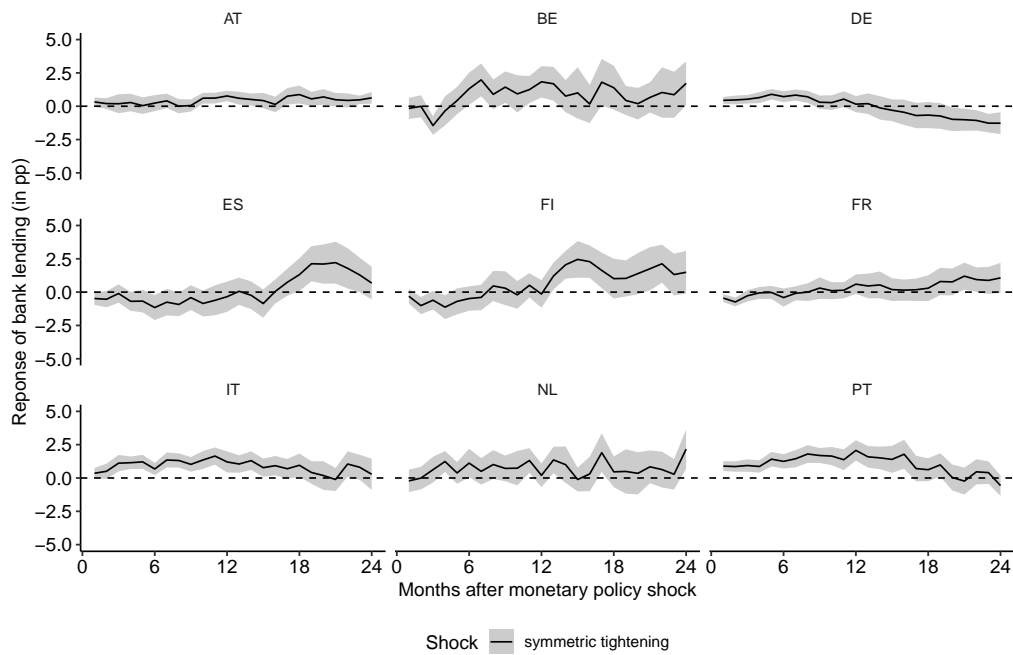


Figure F.7: Symmetric impulse response functions of bank lending growth to non-financial corporations to a 25 bps monetary policy surprise employing the 12-month OIS rate. Note: Fig. F.7 presents the IRFs as shown in Section 5.1 with Newey-West corrected standard errors using 13 lags and one lag of the dependent variable for lag augmentation.

Appendix F.8. Newey-West standard errors (asymmetric)

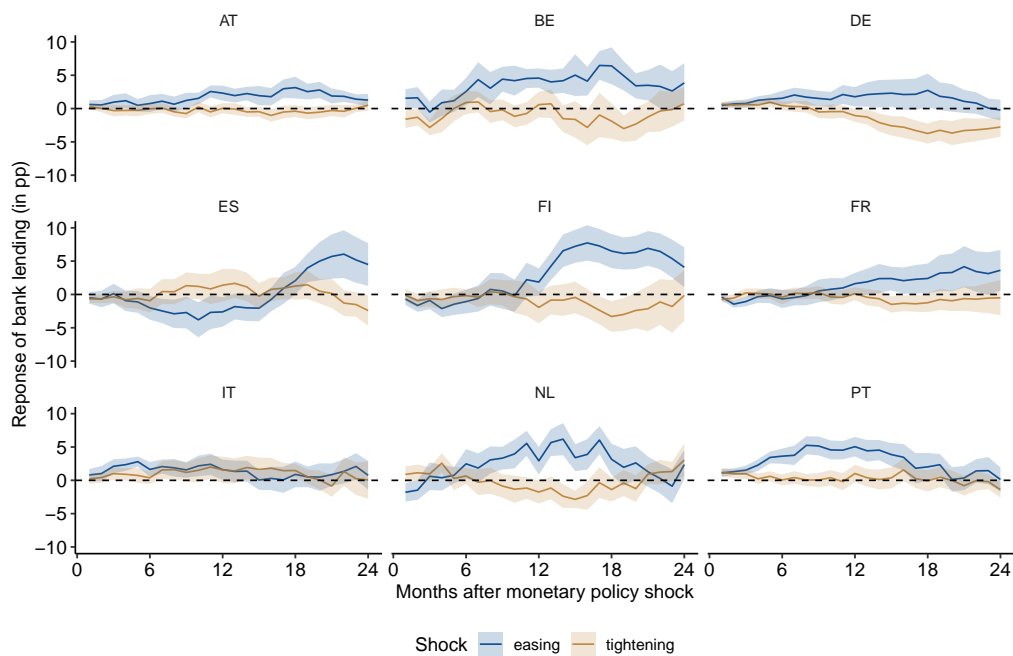


Figure F.8: Asymmetric impulse response functions of bank lending growth to non-financial corporations to a 25 bps monetary policy surprise employing the 12-month OIS rate. Note: Fig. F.8 presents the IRFs as shown in Section 5.1 with Newey-West corrected standard errors using 13 lags and one lag of the dependent variable for lag augmentation.

Appendix F.9. Symmetric result employing 3-month OIS rate

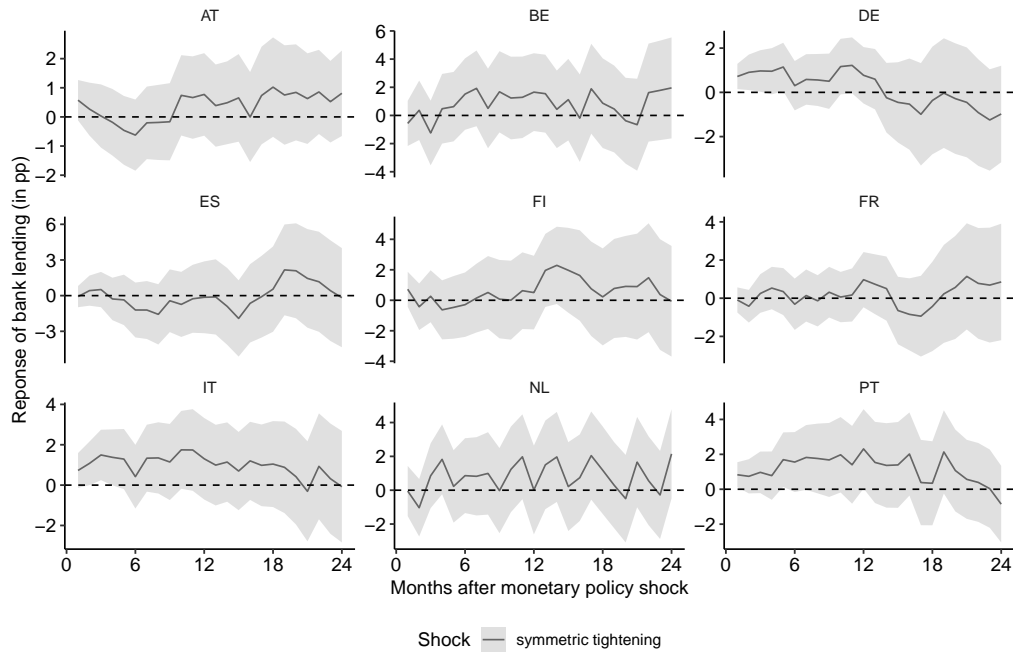


Figure F.9: Symmetric impulse response functions of bank lending growth to non-financial corporations to a 25 bps monetary policy tightening surprise. Note: Fig. F.9 presents the IRFs as shown in Section 5 employing the high-frequency surprise in the 3-month OIS rate as instrumental variable.

Appendix F.10. Asymmetric result employing 3-month OIS rate

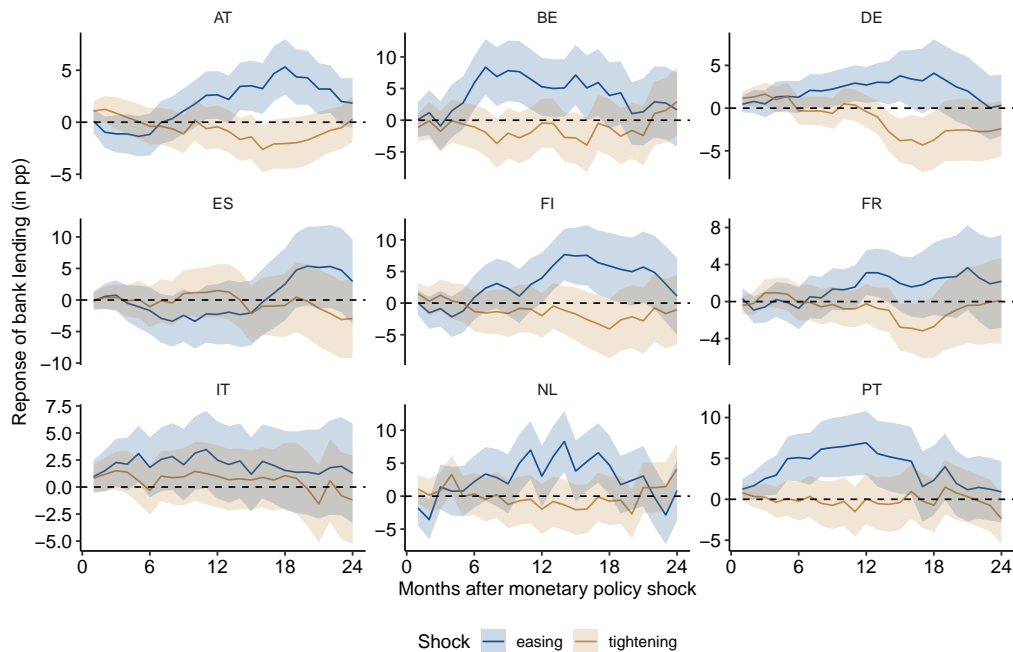


Figure F.10: Asymmetric impulse response functions of bank lending growth to non-financial corporations to a 25 bps monetary policy surprise. Note: Fig. F.10 presents the IRFs as shown in Section 5 employing the high-frequency surprise in the 3-month OIS rate as instrumental variable.

Appendix F.11. Time series of monetary policy shocks (Jarociński and Karadi, 2020)

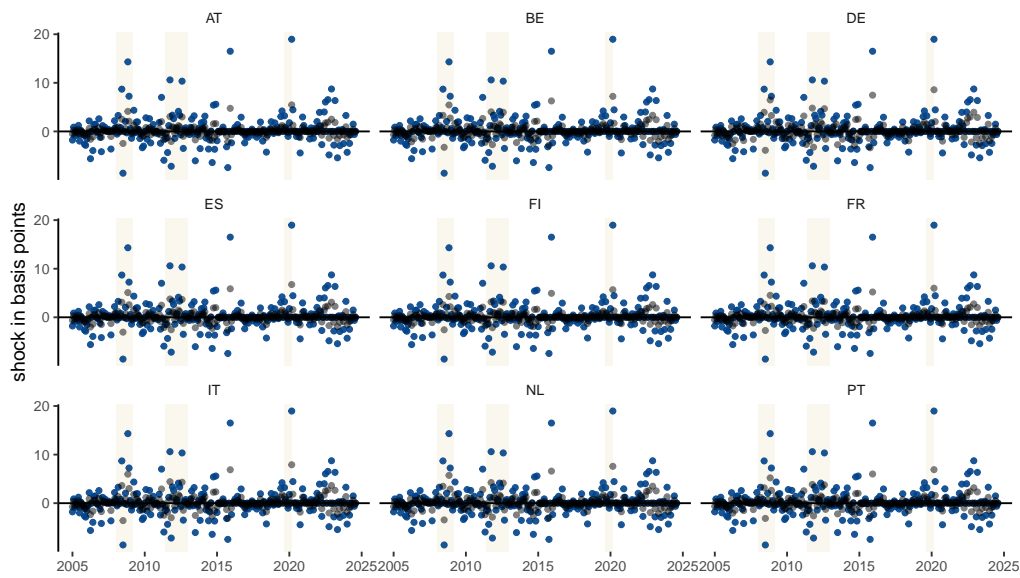


Figure F.11: The Figure compares the time series of monetary policy surprises by Jarociński and Karadi (2020) *monetary policy factor* (shown in blue) to the time series of structural innovations to monetary policy as described in Section 3.4 using the *monetary policy factor* as instrumental variable.

Appendix F.12. Time series of monetary policy shocks (12-month OIS rate)

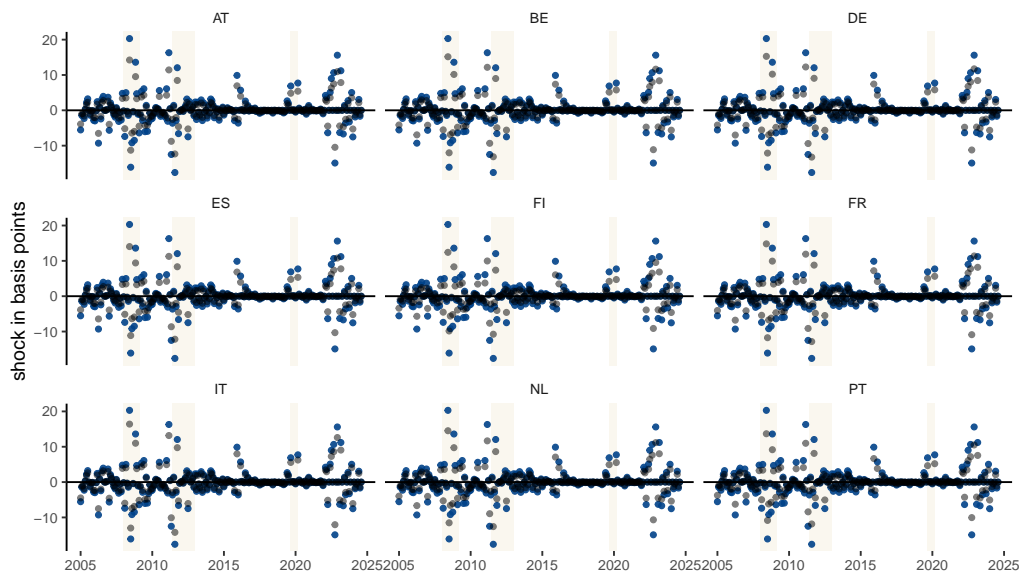


Figure F.12: The Figure compares the time series of changes in the 12-month OIS rate around monetary policy announcements retrieved from EA-MPD (shown in blue) to the time series of structural innovations to monetary policy as described in Section 3.4 using the changes in the 12-month OIS rate as instrumental variable.

Appendix F.13. Time series of monetary policy shocks (Jarociński and Karadi, 2020)

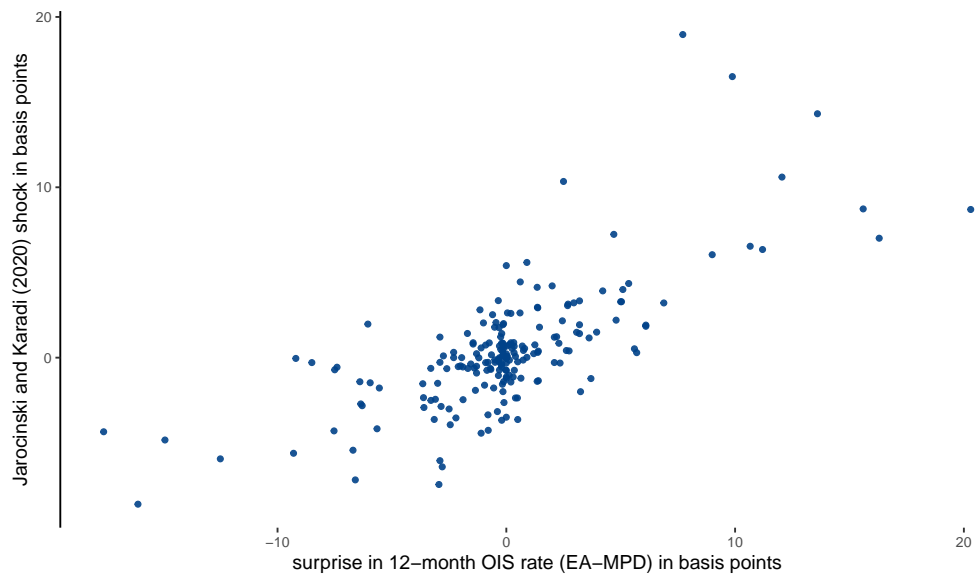


Figure F.13: The Figure compares the monetary policy surprises by Jarociński and Karadi (2020) *monetary policy factor* to the surprise in the 12-month OIS rate retrieved from EA-MPD as a scatter plot.

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