

Holm-Hadulla, Frédéric; Thürwächter, Claire

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Heterogeneity in corporate debt structures and the transmission of monetary policy

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Fédéric Holm-Hadulla, Claire Thürwächter

Heterogeneity in corporate debt structures and the transmission of monetary policy

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Abstract

We study how differences in the aggregate structure of corporate debt financing affect the transmission of monetary policy. Using high-frequency financial market data to identify monetary policy shocks in a panel of euro area countries, we find that: bond finance dampens the overall response of firm credit to monetary policy shocks in economies with a high initial share of bond- relative to bank-based finance; this effect weakens, and may even reverse, in economies with a low share of bond financing; and the dampening effect of a larger bond financing share also attenuates the ultimate impact of monetary policy on economic activity. These findings point to corporate bond markets acting as a “spare tire” in situations when bank lending contracts.

JEL Classification: E44, E52, G21, G23.

Keywords: Firm Financing Structure, Bank Lending, Corporate Bonds, High-Frequency Identification, Local Projections.

Non-technical summary

We study how the relative role of corporate bonds versus bank loans in the debt structure of firms affects the transmission of monetary policy shocks to the economy. To this end, we extend a standard empirical macro model with variables measuring the volume and cost of credit to firms. Estimating this model in a panel of euro area countries, and using high-frequency methods to identify monetary policy shocks, we find that the corporate debt financing structure is highly consequential for monetary policy transmission. In particular, a greater role of bond financing goes along with a weaker response of firm credit and economic activity to monetary policy shocks. Accordingly, bond finance acts as a substitute source of credit when bank lending contracts and, vice versa, it is crowded out when bank lending expands. At the same time, this substitution function requires the initial bond financing ratio in the economy to be relatively high: in economies with a low bond financing ratio, the adjustment patterns weaken and may even reverse so that bond finance instead reinforces the overall response of firm credit to monetary policy shocks.

These patterns point to two countervailing mechanisms emphasised in the related theoretical literature. The substitution pattern at high bond finance ratios is consistent with the bank lending view of transmission, according to which monetary policy triggers shifts in loan supply that go beyond its impact on broader financing conditions in the economy. The amplification pattern at low bond finance ratios is consistent with state-contingent firm preferences for different financing instruments, as the flexibility afforded by bank lending (e.g. in renegotiating credit contracts in response to weakening business prospects) becomes more valuable in tight and less valuable in loose monetary conditions. The reversal in the impact of monetary policy on the bond finance ratio depending on the initial financing structure suggests that a sufficiently large initial reliance on bond finance is necessary for the substitution effect to dominate.

Our findings imply that differences in firm financing structures constitute an important source of geographical heterogeneity in the transmission of monetary policy across the euro area. Such heterogeneity may complicate efforts to ensure a broadly uniform monetary policy stance throughout the economy. And it highlights the need for convergence not only in the economic but also the financial structures across different parts of the currency union, including via a broadening of firm access to direct market finance as envisaged in the context of the EU Capital Markets Union.

1 Introduction

The theory and practice of corporate finance draw a sharp distinction between bank loans and corporate bonds as sources of firm credit. Among the distinctive features, bank loans are typically easier to renegotiate or restructure in case a borrower falls on hard times, whereas corporate bonds tend to be less exposed to macroeconomic disturbances that induce shifts in loan supply. These differences in the underlying economics of different debt-financing instruments in turn may impact their behaviour in response to monetary policy shocks. But the direction of the impact is ambiguous and recent empirical evidence on the interaction between the structure of corporate debt and the transmission of monetary policy is scarce. At the same time, the issue has acquired renewed relevance as many advanced economies have experienced a marked shift in the financing mix of firms, leading to an increasingly relevant role of corporate bonds relative to bank loans over the past decade (see, e.g., [Adrian et al. \(2013\)](#) and [De Fiore and Uhlig \(2015\)](#)).

Against this background, the current paper studies whether the financing structure of firms matters for the transmission of monetary policy to the economy. To this end, we subdivide the question into three distinct but integral aspects and set up an empirical model that allows us to analyse them simultaneously. The first aspect is whether the two debt financing instruments differ in their response to monetary policy shocks. The second aspect is whether any such differences become more or less accentuated depending on the financing structure prevailing prior to the shock. And the third aspect is whether these differences in the responsiveness and their dependence on the initial financing structure alter the ultimate impact of monetary policy on key macroeconomic aggregates. Our empirical approach consists of a local projections setup for a panel of euro area countries, using high-frequency surprises to identify monetary policy shocks and allowing the impact of these shocks on firm credit and economic activity to differ with the bond-to-loan finance ratio prevailing prior to the shock.

Besides their obvious policy relevance, our research question plays into a rich, but highly contested, literature on the interaction between corporate finance and monetary policy transmission. As to the first aspect, a prominent strand of this literature has argued that bank balance sheet frictions may render loan supply more responsive to monetary policy shocks than other sources of debt finance ([Kashyap et al. \(1993\)](#)). As a consequence, bond issuance may mitigate or even counteract the impact of monetary-policy induced shifts in loan supply on the overall amount of credit available to firms. However, the subsequent literature has also highlighted alternative mechanisms pointing in the opposite direction. In particular, the inherent

advantages of bank lending – in terms of the screening and monitoring of borrowers (Diamond (1984)), as well as the flexibility to renegotiate the terms of a credit contract in case of need (Berlin and Mester (1992)) – may become more valuable when monetary conditions tighten and, consequently, collateral values decline and business prospects deteriorate (Bolton and Freixas (2006)).¹ Hence, monetary policy tightening shocks may shift the preferred debt financing structure of firms towards bank loans and, via this channel, trigger a stronger contraction in corporate bonds than in loan volumes.²

The resultant theoretical ambiguity also extends to the second aspect of whether the relative responsiveness of different financing instruments depends on the existing financing structure of the respective economy. For instance, changes in firm financing structures entail fixed costs – *e.g.* for an entirely bank-dependent firm, the process of building a presence in global capital markets is resource- and time-consuming. This mechanism, in turn, would weaken the capacity of bond markets to act as a substitute financing option when loans contract and it would become more relevant the higher is the reliance on bank relative to bond finance in an economy.³ However, also here, countervailing effects are conceivable: for instance, investors in international bond markets may strive for some geographic diversification, which may provide the opportunity for firms in countries with small bond markets to issue at relatively favourable rates due to the limited supply from other, competing issuers (Coerdacier and Martin (2009)).

Taken together, the theoretical predictions on whether loans or bonds are more responsive to monetary policy shocks and how this relative responsiveness would change with the relative prevalence of different financing instruments are therefore not clear-cut. And, since these two aspects constitute integral elements of the transmission process, this theoretical ambiguity carries over to the third aspect, consisting in the overall implications of the corporate financing structure for the effects of monetary policy on the economy.

In view of these ambiguous theoretical predictions, the current paper provides an empirical analysis of

¹For further, closely related, analysis on the specific role of banks in screening and monitoring borrowers, see also Holmstrom and Tirole (1997); Diamond (1991); Rajan (1992); and on the differences in bank- versus direct market-based finance, see Dewatripont and Maskin (1995); Berlin and Mester (1992); Boot et al. (1993). While useful to highlight key conceptual differences, the distinction between different types of credit is approximate: in practice, bank lending relationships often involve several contracting parties, *e.g.* via syndicated loans; loans may also be sold, either directly or in securitized form; bond markets also perform monitoring functions *etc.*

²This mechanism may be reinforced by composition effects: to “qualify” for bond market access, firms have to meet certain credit quality standards and monetary policy may shift the dividing line between qualifying and non-qualifying firms in a procyclical manner (for a related argument, see De Fiore and Uhlig (2011, 2015) and Crouzet (2017)).

³There are further mechanisms that may inhibit changes in firm financing structures. For instance, existing banking relationships may lead to some lock-in as borrowers and lenders try to capitalize on the informational monopoly established over time (Sharpe (1990); Rajan (1992); Kashyap and Stein (1994); Berger and Udell (1995, 2002)).

how the structure of corporate debt financing affects the transmission of monetary policy. As the specific context of the analysis, we consider a panel of euro area countries which, for two reasons, provides a particularly suitable setting to study this question. First, the shift in the relative importance of bond- vs. bank-based finance has been particularly pronounced ([ECB \(2016\)](#)). In particular, the ratio of corporate bonds to bank loans in the overall stock of euro area firm debt has almost doubled since its trough in 2007. Second, the euro area exhibits a high degree of heterogeneity in the relative role of these different financing instruments across countries ([Rodríguez-Palenzuela et al. \(2013\)](#)). For instance, since the start of the euro, the average ratio of bond- over bank-based finance has ranged from less than 0.1 in Greece and Spain to 0.5 in France. This variation in both, the time-series and cross-sectional dimension, offers ample scope for empirical analysis. Moreover, given these countries belong to one currency union, it allows us to clearly identify a common monetary policy stance, without requiring assumptions about the comparability of measures taken by different central banks.

Our econometric model starts from the standard set of variables typically considered in the monetary economics literature, including indicators for real activity, inflation and policy-controlled short-term interest rates. It then adds a set of variables measuring corporate financing volumes, broken down by bank loans and debt securities, as well as a set of controls that may influence the relative role of these financing instruments across countries and over time. In terms of estimation, we follow [Jordà \(2005\)](#)'s local projections method, which offers a flexible approach to model how the interaction of monetary policy with the financing structure of firms alters its effects on key macroeconomic aggregates. For the identification of monetary policy shocks, we resort to high-frequency surprises in key financial market prices around ECB policy events based on a comprehensive dataset, recently published by [Altavilla, Brugnolini, Gürkaynak, Motto and Ragusa \(2019\)](#).

Our findings show that the corporate debt financing structure is highly consequential for monetary policy transmission. In particular, we find that a greater role of bond financing goes along with a weaker overall response of firm credit to monetary policy shocks and this dampening effect also attenuates the impact of these shocks on economic activity. The underlying mechanism is that bond finance acts as a substitute source of credit when bank lending contracts (and, vice versa, it is crowded out when bank lending expands). But this substitution function requires the initial bond financing ratio in the economy to be relatively high. In fact, in economies with a low bond financing ratio, the adjustment patterns reverse so that bond finance amplifies the contraction (expansion) in bank lending in response to a policy tightening (loosening). As a

consequence, larger bond markets act as a buffer that dampens the response of debt financing to firms in response to policy shocks and this buffering function is also reflected in a weaker contractionary effect of the monetary tightening on output. Overall, the findings thus are consistent with the interpretation of corporate bond markets as acting as a “spare tire” in situations when banks cut back on lending.

Related literature. These findings add to a rapidly growing empirical literature that has sprung from the shift in firm financing structures observed since the crisis and touches upon some of the aspects we address in the current paper. Regarding the first aspect, how the response to monetary policy differs across bank versus bond-finance, [Becker and Ivashina \(2014\)](#) also find evidence for the substitution function of bond finance in firm level data, thus confirming earlier findings by [Kashyap et al. \(1993\)](#) and [Kashyap et al. \(1996\)](#), as well as the pattern we observe for economies with a high bond financing ratio. Further, several recent studies, *e.g.* including [Lhuissier and Szczerbowicz \(2018\)](#), [Arce et al. \(2017\)](#) and [Grosse-Rueschkamp et al. \(2019\)](#), distinguish between conventional and unconventional monetary policy (or zoom in on specific unconventional measures such as the ECB’s corporate sector purchase program) and find differences in the response patterns across financing instruments. Finally, [Crouzet \(2019\)](#) presents a structural model that incorporates the countervailing mechanisms by which monetary policy shocks may affect firm financing structures and tests these mechanisms in a panel of US firms.

Regarding the second aspect, whether differences of monetary policy transmission depend on the financing structure prevailing prior to the shock, the literature is much scarcer. In a methodologically similar setup to ours, [Grjebine et al. \(2018\)](#) study how corporate debt structures affect macroeconomic outcomes; but their focus is on the pace of recovery after recessions, rather than the transmission of monetary policy. A potentially interesting, related strand of work is the paper by [Darmouni et al. \(2019\)](#), which shows that, in the euro area, the responsiveness of firm share prices to monetary policy shocks intensifies the higher is their reliance on bond finance, whereas this relationship is absent in the US. The empirical designs of the two papers, focused on stock price reactions of individual firms in [Darmouni et al. \(2019\)](#) as opposed to the overall macro transmission across different economies in ours, is too far apart to try and square these different findings. However, efforts to bridge the more micro corporate-finance based perspectives with the analysis of macro transmission appear as a very interesting avenue for further research.

To the best of our knowledge, our paper is the first to simultaneously explore how the response to monetary policy differs across corporate debt instruments, how these responses are shaped by the debt financing

structure prevailing prior to the shock, and how these differences ultimately affect the macroeconomic implications of a given monetary policy shock.

The remainder of the paper proceeds as follows. Section 2 introduces the model and identification strategy. Section 3 describes the data and key stylized facts. Section 4 presents our main results followed by a battery of robustness checks in Section 5. Section 6 concludes.

2 Econometric methodology

2.1 Model

We estimate the dynamic response of key macroeconomic aggregates and firm financing variables to a monetary policy shock via local projections (Jordà (2005)). These consist in a set of separate panel regressions for each of the dependent variables and each horizon. The baseline model to estimate the impulse response functions (IRFs) is:

$$Y_{i,t+h} = \alpha_{i,h} + (\beta_{0,h} + \beta_h(B/L)_{i,t-1}) \text{shock}_t^{IR} + \gamma_h \sum_{p=1}^2 X_{i,t-p} + \theta_h \sum_{p=1}^2 \bar{X}_{t-p} + \varepsilon_{i,t+h} \quad (1)$$

where subscripts i , t , and h denote the country, month, and IRF horizon, respectively, and subscript p denotes the number of lags included in the sets of control variables $X_{i,t-p}$ and \bar{X}_{t-p} .

The dependent variables $Y_{i,t+h}$ comprise real GDP and the GDP deflator, as well as bank loans to non-financial corporations and the ratio of corporate bonds over bank loans (henceforth referred to as the bond finance ratio). The former two variables serve as measures of economic activity and the general price level and the latter two as measures of the volume and structure of corporate debt financing in each country and month. The dependent variables also include a policy-controlled short-term interest rate, which is the same across countries and which helps assess the plausibility of our proxy for monetary policy shocks.⁴ All variables enter in log-levels, except the bond finance ratio, which is expressed as 100 times the notional stocks of bonds over loans, as well as the shock, which is expressed in percentage points, and the policy rate,

⁴The underlying logic is that changes in policy-controlled short-term interest rates constitute the first step in the transmission of standard monetary policy shocks. So, a significant and relevant response in these variables to the shock-variable are a minimum criterion for the latter to be a plausible proxy for exogenous changes in the monetary policy stance; see, *e.g.*, Gertler and Karadi (2015).

which is expressed in percent.⁵

The key explanatory variables of interest are the monetary policy shock ($shock_t^{IR}$) and its interaction with the bond finance ratio $(B/L)_{i,t}$. As Section 2.2 explains in detail, the shock series is extracted from high-frequency changes in key money market interest rates around ECB monetary policy events. The coefficient $\beta_{0,h}$ captures the response in each of the dependent variables in period $t+h$ to an exogenous monetary policy tightening in period t . The coefficient β_h shows whether and how this impact changes depending on the financing structure prevailing in the period prior to the shock. Taken together these coefficients summarize the combined impact of the monetary policy shock at horizon h conditional on the prevailing bond finance ratio as:

$$\frac{\partial Y_{i,t+h}}{\partial shock_t^{IR}} = \beta_{0,h} + \beta_h(B/L)_{i,t-1}. \quad (2)$$

Further, the model includes a set of control variables to capture common and country-specific influences on the economy over time, denoted by \bar{X}_{t-p} and $X_{i,t-p}$, respectively. The common influences consist of euro area real GDP and the GDP deflator, as well as the EUR-USD exchange rate and global commodity prices to capture the external environment. Moreover, \bar{X}_{t-p} includes the “loan supply indicator” (LSI) constructed by Altavilla, Darracq Pariès and Nicoletti (2019) and the “composite indicator of systemic stress” (CISS) by Kremer et al. (2012). Both variables are motivated by the specific timing of the observed shifts in firm financing structures in the euro area: the trend towards a greater relative role of bond-based finance started from the dislocations in the banking system and the protracted period of financial stress that followed the Global Financial Crisis (Adrian et al. (2013); De Fiore and Uhlig (2015); Crouzet (2017)). Over the same period, also the monetary policy stance underwent marked changes, as major central banks sought to contain the fallout from the crisis. Accordingly, the specific conditions prevailing over the post-crisis period may be correlated with both, monetary policy and the relevant outcome variables in $Y_{i,t+h}$. By including the CISS and LSI in the specifications, with the former capturing financial market stress and the latter aggregate loan supply conditions in the euro area, we control for this possibility and thereby further sharpen the identification. We also include lags of the euro area wide policy interest rate and the shock in \bar{X}_{t-p} . In particular, the inclusion of the lagged shock is important to purge the high-frequency surprises from serial correlation, which further underpins the identifying assumption that the shock represents unanticipated changes to

⁵By including both, bank loans $L_{i,t}$ in logs and the bond finance ratio $(B/L)_{i,t}$ we ensure that the overall debt financing volume ($V_{i,t} = L_{i,t} + B_{i,t}$) from these two sources is reflected in the regressions, as approximated by $\log V_{i,t} \approx \log L_{i,t} + (B/L)_{i,t}$.

monetary policy (Ramey (2016)).

Besides lags of the variables included in $Y_{i,t+h}$, the country-specific controls $X_{i,t-p}$ also comprise two composite financing cost measures, one on bank borrowing and one on bond financing of firms, to capture the relative attractiveness of these different instruments; and they include the lagged interactions of the shock with the bond finance ratio. Last, we saturate the model with a set of country-fixed effects $\alpha_{i,h}$.⁶

Our estimations use Driscoll and Kraay (1998) standard errors that allow for both serial correlation, which is an inherent to local projection estimates, and spatial dependence across countries. Our baseline regressions include $p = 2$ lags of the variables comprised in $X_{i,t-p}$ and \bar{X}_{t-p} .⁷

Taken together, this baseline specification closely corresponds to the standard model set-up used in much of the related literature to study the impact of monetary policy and augments it with additional variables to capture corporate financing volumes. Besides including the key variables of interest from a macroeconomic perspective, this close correspondence has the additional benefit of allowing us to benchmark our findings against other studies. Due to its flexible functional form, the local projections method is as a suitable modelling framework to assess how the corporate debt structure affects the transmission of monetary policy.⁸

2.2 Identification strategy and estimation

High-frequency identification. Our identification strategy follows an increasingly active strand of the monetary economics literature using high-frequency changes in financial market variables around events of central bank communication as a measure of exogenous changes in the monetary policy stance (see Ramey (2016) for a review of this literature). The basic rationale of this high-frequency identification strategy (HFI) is that, provided the time window around these events is sufficiently narrow, it is plausible to attribute the observed changes in interest rates to monetary policy – and to exclude that they reflect any other confounding factors that may be simultaneously correlated with the policy stance and the macroeconomic or financial outcome variables.

Following this logic, which essentially treats the policy events as a controlled experiment, HFI has been applied in a host of papers studying the impact of monetary policy on the US economy and financial mar-

⁶We also tested specifications including a deterministic time trend, which left our main results unaffected however.

⁷We experimented with different lag numbers, but also found this to be inconsequential for our estimates.

⁸For a similar modelling strategy, see Auer et al. (2019), who study the link between corporate leverage and monetary policy transmission.

kets (Kuttner (2001); Cochrane and Piazzesi (2002); Gürkaynak et al. (2005); Piazzesi and Swanson (2008); Barakchian and Crowe (2013); Gertler and Karadi (2015); Nakamura and Steinsson (2018)). More recently, a nascent literature has adopted this approach to similar applications in the euro area context (Lhuissier and Szczerbowicz (2018); Andrade and Ferroni (2018); Jarocinski and Karadi (forthcoming); Altavilla, Brugnolini, Gürkaynak, Motto and Ragusa (2019); Auer et al. (2019)).

The data underlying our HFI approach comes from the Euro Area Monetary Policy Event-Study Database (EA-MPD), constructed by Altavilla, Brugnolini, Gürkaynak, Motto and Ragusa (2019) and published on the ECB website.⁹ The EA-MPD computes intra-day changes in a broad set of financial market variables around the time the ECB's Governing Council communicates and explains its monetary policy decisions. These communication events follow a preset schedule and format in that: the Governing Council meets in regular intervals and the meeting dates are publicly known well in advance;¹⁰ at 1.45 pm on the day of the meeting, a press release is published on the ECB website announcing the Governing Council's monetary policy decision (which may also consist in an announcement that the policy configuration remains unchanged); and at 2.30 pm a press conference takes place, at which the President first reads out a written statement explaining the rationale of the decision, followed by a question- and answer-session with journalists. The EA-MPD calculates three sets of surprises, one for the publication of the press release, one over the period of the press conference, and one for both events together. Taken together, the EA-MPD constitutes a very comprehensive, detailed, and publicly available database available for high-frequency identification of monetary policy shocks in the euro area.¹¹

Definition of monetary policy shocks. In our baseline estimations, we use the change in a short-term interest rate over the entire event window, *i.e.* from before the press release to after the press conference, as the basis for calculating the monetary policy shocks. In doing so, we take into account an important

⁹https://www.ecb.europa.eu/pub/pdf/annex/Dataset_EA-MPD.xlsx

¹⁰According to the regular schedule, the Governing Council has a monetary policy meeting every six weeks. This regular meeting schedule has changed twice since the introduction of the euro in 1999. Initially, the meetings took place twice per month, whereas after November 2001 and until December 2014 the regular monetary policy meetings followed a four-week rhythm. Further, on a few occasions the Governing Council deviated from the regular meeting schedule to respond to acute crisis events. A list of all the events included in the EA-MPD and a discussion of irregularities in the scheduling and practice of the Governing Council meetings is laid out in the Appendix to Altavilla, Brugnolini, Gürkaynak, Motto and Ragusa (2019).

¹¹For further detail, see Altavilla, Brugnolini, Gürkaynak, Motto and Ragusa (2019). A similarly rich dataset with high-frequency surprises around monetary policy events for the euro area has been provided by Jarocinski and Karadi (forthcoming). The main reason for us to rely on the EA-MPD is that the latter is regularly updated, provides a more detailed breakdown of the policy surprises into those deriving from the press release and those deriving from the press conference, and includes a broader set of financial market variables.

insight into the anatomy of monetary policy surprises deriving from the recent literature: when central banks talk about policy, market participants do not only receive a signal on whether and how the central bank is going to adjust its policy instruments, but also on how the central bank assesses economic prospects. If market participants in turn assume that the central bank commands over superior information on the state and prospects of the economy, this may lead them to revise their own economic assessment. Thus, depending on which type of signal dominates, investors may draw very different inferences on the current and future stance of monetary policy resulting in different constellations of financial market adjustments around the event.

For instance, an unexpected interest rate increase may be accompanied by a decline in stock prices if market participants perceive the central bank decision as a true monetary policy tightening, which engenders an expected contraction of economic activity; in this case, the negative cross-asset correlation between interest rates and stock prices would qualify the high-frequency surprise as a genuine monetary policy shock. Alternatively, a positive interest rate surprise may also be accompanied by rising stock prices if the rate increase is interpreted as a sign of the central bank's superior information, suggesting that economic prospects are more buoyant than previously thought; in this case, the surprise would constitute what the literature has come to refer to as an information shock.

As visible from Figure 1, this distinction is of major practical relevance in the euro area context (and, as shown by [Jarocinski and Karadi \(forthcoming\)](#), also in the US context). For instance, over the sample considered in our empirical analysis, almost 40% of high-frequency surprises fall under the information shock category (see upper right and lower left quadrant). Since our aim is to estimate the effects of genuine monetary policy shocks, our baseline analysis instead only considers the shocks in the other two quadrants, so as to sharpen the identification approach.

Choice of short-term rate. The specific variable used in our baseline to extract the monetary policy shocks is the 1-month OIS rate. As the EA-MPD contains surprises for a broad range of risk-free interest rates rates, including a range of maturities from 1-week to 30-years for the OIS and in the absence of strong conceptual reasons to favour one over the other point on this spectrum, we are left with several degrees of freedom in choosing between them.

We thus discipline our choice by imposing two criteria. The first is to ensure a relatively close mapping from the shock variable to the policy variable we included in our dynamic model. Since, in line with much

of the related literature, we choose the policy variable to be the 3-month OIS rate, a natural choice would be to also use a shock located on the same or a similar point on the term structure. The second criterion is that, within a reasonably narrow segment around the 3-month maturity, we choose a shock that has the highest statistical fit for the 3-month OIS rate after controlling for all other variables in equation (1). This is important to ensure that, also from a statistical perspective, there is a close mapping from the shock variable to the policy variable.

Figure 1: Stock price and policy rate surprises

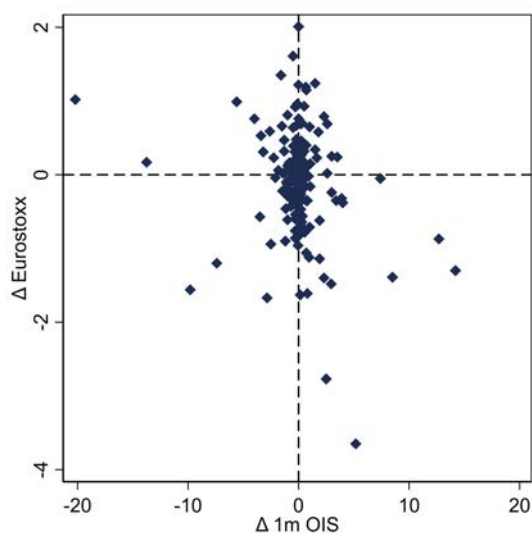
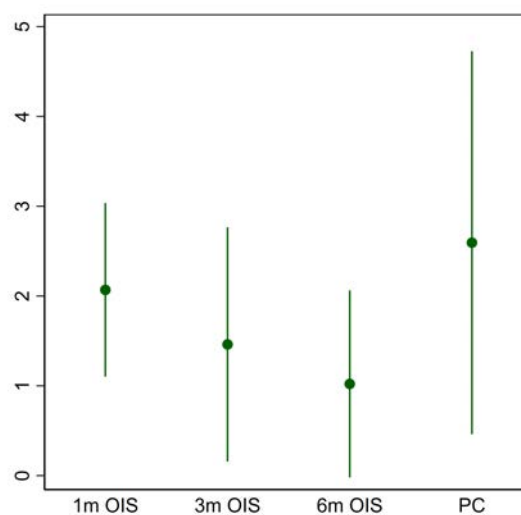


Figure 2: Response of policy rate on impact



Note: Surprises in the Eurostoxx 50 and the 1-month OIS rate in Figure 1 are in percentage points and basis points respectively. The response in Figure 2 is scaled to a 100 basis point tightening shock in the respective high-frequency surprise at the average bond finance ratio and is shown in percentage points. "PC" refers to the first principal component of the 1-week to 2-year OIS rates. The range shows the 95% confidence interval.

Regressing the residuals of the 3-month OIS rate on different shocks from the EA-MPD, the 1-month OIS surprise emerges as the most suitable among a range of candidates (see Figure 2). Its coefficient is estimated with a markedly higher precision than that of the 3-month OIS rate shock and the principal component (PC) of different rates; and for the 6-month OIS rate shock, the estimated 95%-confidence interval includes zero. In summary, our baseline definition of $shock_t^{IR}$ thus consists of the monetary policy shock component of the surprises in the 1-month OIS rate. In Section 5.2, we test for the robustness of our key findings to this choice.¹²

¹²A time-series plot of the high-frequency surprises that we consider in our baseline model and the robustness section can be

3 Data and stylized facts

We estimate the model on monthly data over the sample period from January 2003 to April 2019.¹³ At the cross-sectional level, we include a panel of 10 euro area countries, which together account for 96% of euro area GDP.¹⁴ As the monetary policy variable, we use the monthly averages of daily observations of the 3-month OIS rate. GDP and the deflator are interpolated from quarterly to monthly frequency, so as to match the frequency of the firm financing variables.¹⁵ Loans and bond finance to non-financial corporations are the notional stocks, which correct for changes in the financing volumes that arise from valuation changes so that we capture only changes in the stock stemming from transactions. Moreover, the loan data is adjusted for securitisation, cash pooling, and sales and we restrict the counterparty sector to residential non-financial corporations.

For the control variables, the GDP and deflator series at the euro area level are interpolated from quarterly to monthly frequency in the same procedure as the corresponding country level data. The cost of loan finance is included as a spread of the bank lending rate for loans to corporations on new business vis-à-vis a risk-free rate. As risk-free rate we choose the 1-year OIS rate so as to broadly match the maturity and average rate fixation profile of the bulk of loans in our sample.¹⁶ The monthly series for the costs of bond finance in each country are based on [De Santis \(2018\)](#) and are constructed by aggregating up the spreads on individual bonds purged of bond- and issuer-specific characteristics.¹⁷ The variables capturing the external environment are the EUR-USD exchange rate, averaging daily observations over each month, and an encompassing commodity price index, which is provided by the IMF. For the CISS we also average the

found in [Figure 10](#) in [Appendix A](#).

¹³We follow the recommendation of [Altavilla, Brugnolini, Gürkaynak, Motto and Ragusa \(2019\)](#) and exclude the years from 1999 to the end of 2001 from the sample due to noise and sparse quotes in the intraday OIS data during the first years of the euro. A further sample restriction is imposed by the LSI which is only available from 2003 onward. Estimating the model on monthly data allows us to exploit the availability of loan and bond volumes at that frequency. The series are retrieved from the Statistical Data Warehouse (SDW) of the ECB, unless noted otherwise.

¹⁴The countries are Austria, Belgium, Germany, Spain, Finland, France, Ireland, Italy, Netherlands, and Portugal.

¹⁵For interpolation we employ the [Chow and Lin \(1971\)](#) method. The monthly variables along which we interpolate are industrial production (excluding construction) and HICP for real GDP and the GDP deflator respectively. Industrial production data for Ireland are from Eurostat.

¹⁶Over our sample, around 60% of euro area loans to non-financial corporations have a maturity of up to one year or an interest rate reset within one year if the initial maturity is above one year. Our results are unaffected when instead computing the spread against the 2-year OIS rate.

¹⁷Among these characteristics, [De Santis \(2018\)](#) also includes the expected default frequency (EDF) of the issuer, which allows for the construction of an excess bond premium. By contrast, the cost of bond finance measure used in our analysis does not include EDF in the list of controls since differences in credit risk exposures may be one of the features that distinguish bond- from bank-based finance.

daily values per month and for the LSI we linearly interpolate the quarterly measure to monthly frequency. The macroeconomic variables, financing volumes, and the commodity index are seasonally adjusted.

In constructing the monetary policy shock, we directly convert the daily high-frequency surprises to monthly frequency.¹⁸ As Ramey (2016) points out, any time-averaging of the high-frequency surprises introduces serial correlation, which is undesirable given the shock, by definition, should be unpredictable.¹⁹ In Section 5.2, we however test for the robustness of our findings under an alternative frequency conversion that weights the shocks according to the dating of the monetary policy decision within the month.

Table 1: Summary statistics

	Mean	Std.dev.	Min	Max
Monetary policy shock				
1m OIS	0.06	2.58	-20.20	14.20
Dependent variables				
GDP	81,509	75,244	13,342	271,457
Deflator	94.37	6.49	79.11	106.59
Loans	382,405	329,030	27,695	1,079,412
Debt securities	91,603	122,737	1,706	643,114
Bond finance ratio	26.65	15.86	1.28	62.06
3m OIS	1.12	1.47	-0.37	4.33
Control variables				
GDP (EA)	854,472	41,737	773,412	943,743
GDP deflator (EA)	94.62	5.91	82.83	104.64
Loan spread	1.94	0.97	0.19	6.35
Cost of bond finance	-0.10	1.68	-9.18	18.10
Loan supply indicator	-0.03	1.43	-4.02	4.08
CISS	0.19	0.17	0.03	0.77
EUR-USD	1.27	0.12	1.05	1.58
Commodity price index	127.97	34.92	61.44	198.08
Observations	1,877			

Note: The shock is in bps; GDP, loans, and debt securities are in million EUR; the 3-month OIS rate, loan spread, and cost of bond finance are in percent. GDP is expressed in real terms with base year 2015; the deflator and commodity prices are indexed to 2015.

¹⁸If there are two events in one month we sum up the observations. If there is no event our shock measure carries a zero-value.

¹⁹As discussed in Gertler and Karadi (2015), a rationale for using weighted monthly averages of surprises around policy events in constructing the shock is that this accounts for the fact that an event taking place later in a month will have less scope to transmit to the monthly variables than one taking place early in the month.

Table 1 presents summary statistics of the variables that enter our model. The data exhibit rich variation, also reflecting the pronounced size differences across euro area countries, visible for instance in GDP and loan volumes. Important for our analysis is that the relative size of bank- to bond-based finance also covers a wide range. In particular, the minimum bond finance ratio of only 1.28 indicates that for some countries loans represent the dominant source of debt to firms while at the other end of the distribution, a value of 62.06 implies that for some countries more than a third of firms' debt finance is directly sourced from credit markets.

Figure 3: Cross-country distribution of bond finance ratio

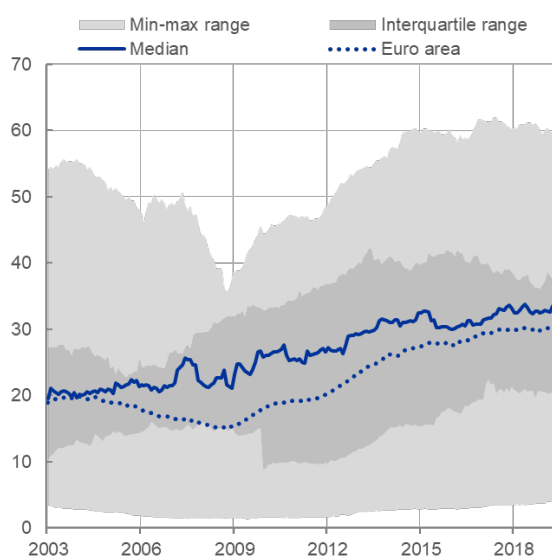
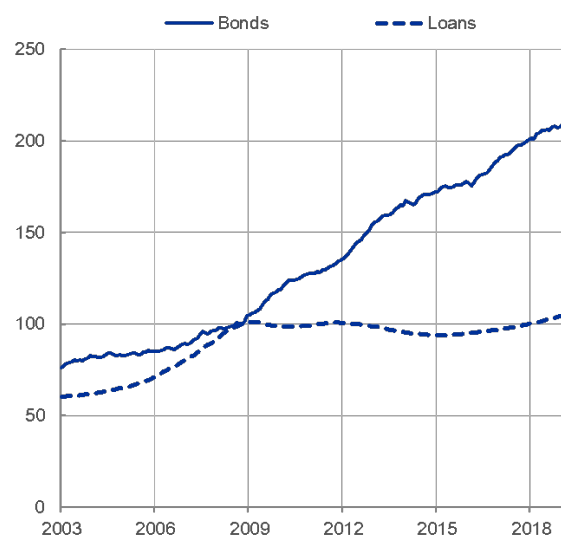


Figure 4: Euro area corporate bond and loan volumes



Note: The drop in the interquartile range at the end of 2009 in Figure 3 is due to a shift in the distribution when data for Irish bond finance becomes available. The series in Figure 4 are indexed to October 2008, which corresponds to the trough in the euro area bond finance ratio depicted in Figure 3.

Meanwhile, the relevance of bond finance has clearly increased over time, both in individual countries and in the euro area as a whole. In the cross-country distribution presented in Figure 3, several observations stand out. First, the wide range of the bond finance ratio across countries appears to be a persistent characteristic of the euro area; and not only the min-max range covers a broad spectrum, but also the interquartile range points to pronounced and persistent cross-country differences. Second, the prevalence of bond finance has become increasingly important, especially since the global financial crisis after which the euro area average, as well as the median and the interquartile and min-max ranges of the distribution have risen markedly.

As visible from Figure 4, the time-series patterns of the bond finance ratio are primarily driven by the steady increase in outstanding bond volumes over the sample period. In particular, since the onset of the financial crisis bond finance has shown a marked ascent and has more than doubled, whereas the outstanding volume of loans plateaued.

4 Baseline results

Our estimates show that the two corporate debt financing instruments differ in their response to monetary policy shocks and that these responses are shaped by the financing structure prevailing prior to the shock. In economies with high bond finance ratios, the contraction (expansion) in bank lending after a monetary policy tightening (loosening) is buffered by an increase in bond relative to loan volumes. This buffering function weakens, and may even reverse, for lower initial bond finance ratios. The resultant differences in the responses of overall firm credit also translate into the macro transmission of the shock in that the impact of monetary policy on activity is weaker in economies with higher bond finance ratios.

In the following, we describe these findings in greater detail, starting with the impulse response functions (IRF) to the monetary policy shock at the average bond finance ratio in the sample; then moving to the interaction terms that capture how the IRFs change when the bond finance ratio rises by one percentage point; and, finally, presenting the combined slope and interaction effect as a function of the bond finance ratio (as per equation (2)).

Estimates at the average bond finance ratio. As visible from the first column of Figure 5, the IRFs at the average bond finance ratio display the typical transmission patterns of monetary policy shocks. The short-term policy rate responds contemporaneously to the tightening shock and builds up gradually reaching a peak of 1.7 percentage points after four months; it then reverts and shows some temporary undershooting, which is also a common feature in many macro models. Loans contract and the response becomes significant around two years after the tightening shock, while the response path of the bond finance ratio remains insignificant throughout the horizons, implying that bonds fall in roughly equal proportions to loans. The decline in credit is also accompanied by falls in real GDP and prices. GDP declines gradually to a trough of -6.4% 30 months after the shock; and prices, measured by the GDP deflator, respond with a lag relative to activity, which also in line with typical transmission regularities. The trough response of the deflator is -1.4% and is attained

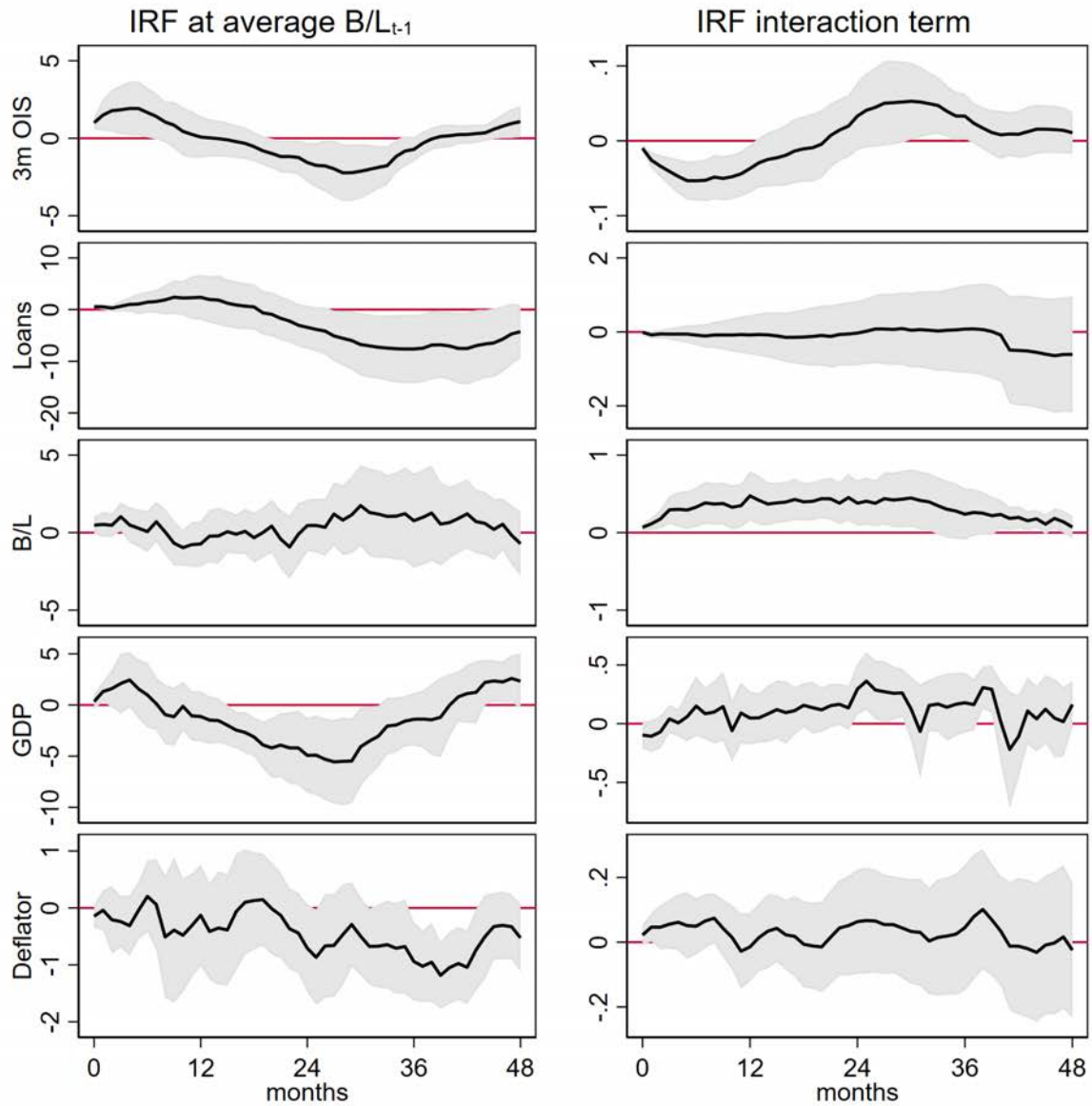
after 39 months.

These estimates are broadly in line with the related literature, also from a quantitative perspective. For instance, scaling the GDP response to a shock leading to a 100 basis point (bps) peak increase in the OIS rate, we obtain a trough effect of -3.8%, which falls within the upper part of the range of estimates for real activity in the US context, as reviewed by [Ramey \(2016\)](#). The corresponding trough in the deflator, of -0.8%, points to a fairly muted transmission to prices, compared to the transmission to real activity. Besides replicating familiar patterns from the related literature, it is also worth noting that the IRFs from our model display a very smooth pattern, considering that local projections usually produce rather jagged responses due to the direct estimation of coefficients at each horizon.

Interactions with bond finance ratio. As visible from the second column of [Figure 5](#), which plots coefficients $\hat{\beta}_h$ on the interaction term in equation (1), the response path of these variables materially changes depending on the prevailing bond finance ratio. As the initial level of the bond finance ratio increases, the build up of the policy rate is slightly muted and the OIS peak response declines by 0.04 percentage points. For changes in loan volumes as well as the GDP deflator the interaction term is not statistically significant. By contrast, the response of the debt finance ratio to the shock significantly varies with the bond finance ratio prevailing prior to the shock. In particular, higher initial bond finance ratios are associated with a significant increase in that ratio in response to the shock. Similarly, the interaction term for the GDP response reveals a relative increase in the IRF at higher initial levels of the bond finance ratio.

These findings suggest that the debt financing structure alters the transmission of monetary policy to firm credit and the macroeconomy. As the prevalence of market-based finance increases, corporate bonds offer an alternative source of credit that acts as a substitute for bank lending when monetary policy tightens; and, vice versa, it is crowded out when bank lending expands in response to a monetary policy easing. The availability of an alternative source of firm credit is also reflected in real activity as a shift towards greater bond finance goes along with a weaker transmission of monetary policy shocks to GDP.

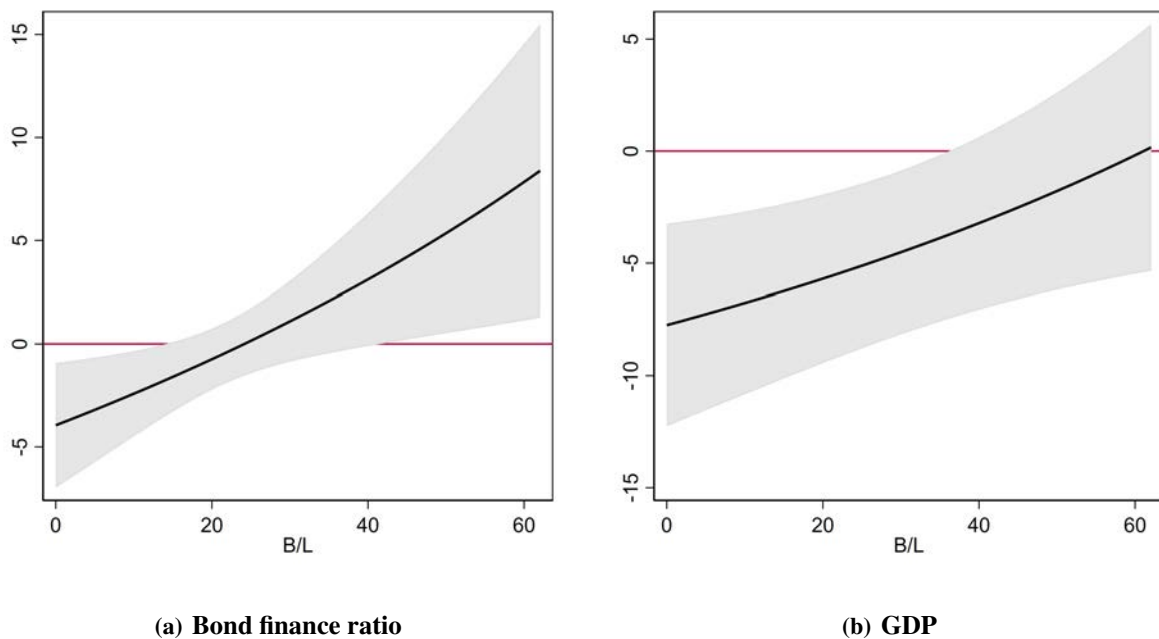
Figure 5: Impulse responses (baseline)



Note: The IRFs in the left column are normalised to a 100 bps impact response in the 3-month OIS rate. The responses are measured in percent except for the 3-month OIS rate and the bond finance ratio where it is in percentage points. The grey area is the 90% confidence interval.

Combined effects. Combining the coefficients for the slope and interaction term, we trace out the response to the shock over the full spectrum of bond finance ratios for a fixed projection horizon in Figure 6.²⁰ As dependent variables, we focus on the bond finance ratio and real GDP since they exhibit relevant and significant differences in the IRFs of the interaction term. As the projection horizon, we choose $h = 24$, as most of the transmission has materialised after two years. All values are normalised to a 100 bps impact response of the policy rate.²¹

Figure 6: Impulse responses across bond finance ratios (baseline)



Note: The IRFs are normalised to a 100 bps impact response in the 3-month OIS rate at each point of the spectrum of bond finance ratios. The response horizon is $h = 24$. For the bond finance ratio the response is in percentage points and for GDP in percent. The grey area is the 90% confidence interval.

While the combined impact, at the average bond finance ratio of around 27, is close to zero and statistically insignificant, a nuanced response pattern emerges when moving away from the average. In fact, at very low bond finance ratios, bonds contract relative to bank loans, whereas the opposite response emerges

²⁰The range of the bond finance ratios over which we evaluate the responses goes from 0 to 62.1, which is the maximum value in our sample. The minimum ratio in the sample is 1.3, but for presentational reasons we extend the estimated range to a bond finance ratio of zero.

²¹The impact response in the policy rate may vary at the respective bond finance ratio which leads to the slight non-linear shape in the graphs.

for high bond finance ratios. Likewise, the GDP contraction gets smaller as the prevalence of bond finance increases and at high levels of the bond finance ratio, the impact of the shock even becomes insignificant.

The combined impact estimates at the higher end of the bond finance spectrum thus favour the bank lending view of monetary policy transmission, by which direct market finance mitigates the credit contraction caused by the bank lending response to monetary policy shocks; and this buffering function of market finance is also reflected in a weaker responsiveness of activity in economies with a high prevalence of bond finance. At the same time, the impact estimates at the lower end of the bond finance spectrum suggest that this buffering function requires bond finance to exceed a certain threshold. By contrast, below that threshold the response pattern reverses. This in turn may reflect that, in economies dominated by bank-dependent firms, the fixed cost of building a presence in bond markets may act as an obstacle to a greater shift to market-based finance, whereas bilateral lending relationships with banks become more valuable as monetary conditions tighten. Overall, this pattern implies that the contraction in credit in response to monetary policy tightening shocks is amplified in economies with low prevalence of bond finance (and vice versa for easing shocks).²²

5 Robustness

To test the robustness of our estimates, we modify the baseline specification in two directions. First, we control for additional types of cross-country heterogeneity that may affect monetary policy transmission and potentially be correlated with the financing structure. Second, we vary the way in which monetary policy is measured and the monetary policy shocks are constructed. As we describe in greater detail below, these modifications leave our main findings intact.

5.1 Further sources of cross-country heterogeneity

Our results may be prone to omitted variable bias if countries with high or low bond finance ratios also differ in other aspects that are relevant to monetary policy transmission. In particular, two aspects deserve attention

²²The finding that monetary policy effects on output vanish at high bond finance ratios is striking, but it warrants two qualifications. First, it is based on coefficients that are uniform across observations, whereas heterogeneity in the estimated impact solely results from applying these uniform coefficients to different values of the bond finance ratio as the conditioning variable. Thus, it is not possible to infer from Figure 6 that monetary policy shocks do not affect output in specific economies at the upper end of the spectrum of bond finance ratios as the uniform coefficients may not capture all aspects of heterogeneity across economies. Second, while the upward sloping pattern in Figure 6 is a robust finding, some of the modifications to our baseline specification presented in Section 5 yield substantial negative combined effects also at high bond finance ratios.

in this regard.

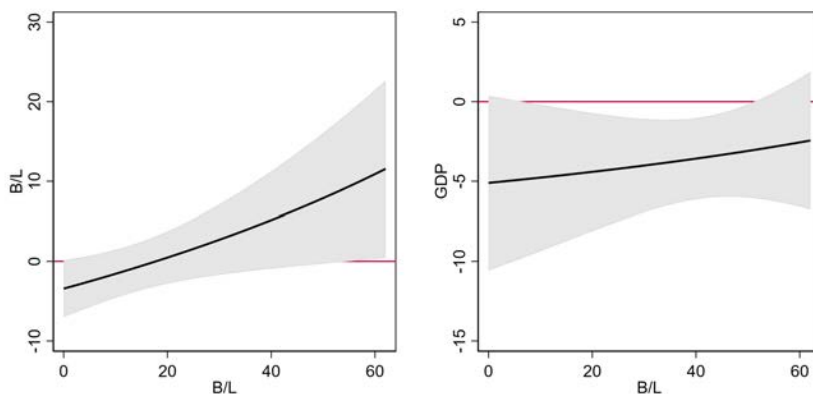
The first relates to differences in firm-size structures. In some euro area countries, such as Germany and France, the universe of firms ranges from small and medium-sized enterprises (SMEs) to very large corporates. In others, especially in Southern Europe, the firm-size structure is instead dominated by SMEs. As it is typically only large corporates that have access to bond markets, these differences are also mirrored in the aggregate bond finance ratio of each country. At the same time, large and small firms also tend to differ in their response to monetary policy shocks, in that the latter tend to be more strongly affected (Oliner and Rudebusch (1996), Gertler and Gilchrist (1994) and Crouzet and Mehrotra (2018)). So the differential transmission of monetary policy shown in Figure 6(a) and 6(b) may not just reflect differences in the financing structures, but also in firm-size structures.

The second aspect pertains to the maturity structure of bank lending, which may also affect the transmission of monetary policy (Gertler and Gilchrist (1994) and Christiano et al. (1996)). In some countries, again including Germany and France, a fairly large share of bank loans to firms have a medium- to longer-term maturity, whereas in others, such as Italy or Portugal, short-term loans make up a greater share of bank lending to firms.²³ Again, these structural cross-country differences may impinge on monetary policy transmission. For instance, a higher share of short-term loans may allow firms to more flexibly adjust their external borrowing and hence accelerate the response to a shock compared to a situation in which a large part of the outstanding stock of loans, at each point in time, is tied up in long-term contracts.²⁴

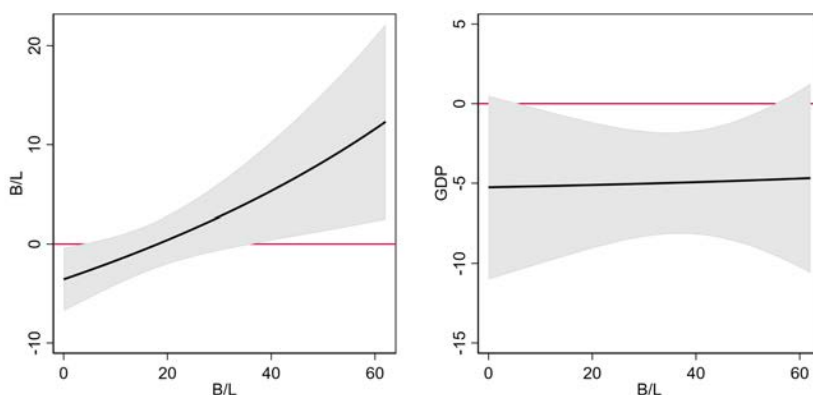
²³In our sample, the average share of loans to non-financial corporations with an initial maturity of less than one year is around 20% for Germany and France and 39% and 32% for Italy and Portugal, respectively.

²⁴Indeed, when subdividing the loan-variable in equation (1) into short- and long-term loans, we are able to confirm this presumption (see Figure 11 in Annex B, where short-term loans are defined as loans with an initial maturity of up to one year): the response of short-term loans is stronger than that of long-term loans and the interaction term shows that the responsiveness of short-term loans increases further as the bond finance ratio rises, again confirming the substitution function of bond markets. These findings suggest that the maturity composition of bank loans may be relevant.

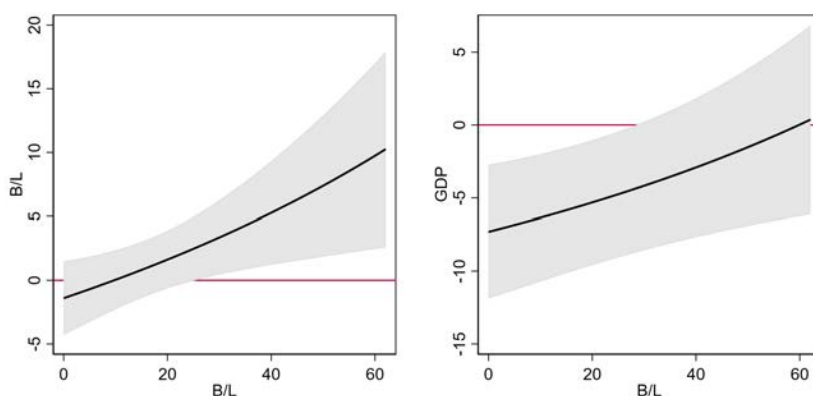
Figure 7: Impulse responses across bond finance ratios (robustness: cross-country heterogeneity)



(a) Share of large firms



(b) Share of large firms in value added



(c) Maturity of bank lending

Note: The IRFs are normalised to a 100 bps impact response in the 3-month OIS rate at each point of the bond finance spectrum. The response horizon is $h = 30$ for panels (a) and (b) and $h = 24$ for panel (c). For the bond finance ratio the response is in percentage points and for GDP in percent. The grey area is the 90% confidence interval.

To account for these points, we thus allow the transmission of shocks to differ not only with the bond finance ratio but also with the firm-size structure and the maturity structure of bank lending. To this end, we select a set of firm-size and maturity proxies and, for each of them, create a dummy variable that equals one if a country's median exhibits a value that is higher than or equal to the sample median and zero otherwise. We then interact the respective dummy with the monetary policy shock and add this interaction along with the dummy variable to the regression equation (1). As firm-size proxies, we use the share of non-financial firms with more than 250 employees in the universe of non-financial firms and, as an alternative, the share of value added of these firms in total value added in the business economy of the respective country.²⁵ As maturity proxy we use the ratio of short-term over long-term loans, with the former consisting of all loans with an initial maturity of up to one year and the latter of all other loans.²⁶

Our main findings are robust to these model extensions (Figures 7(a) to 7(c)). In particular, in all three specifications the response of the bond finance ratio to a monetary policy shock increases and the contractionary effect of the shock on GDP is attenuated, the higher is the bond finance ratio prevailing prior to the shock.²⁷ The estimates are somewhat less precise than in the baseline as, for instance, the slight contraction in the bond finance ratio at the lower end of the spectrum loses significance in some of the alternative specifications. But the baseline patterns overall remain intact, thus suggesting that differences in the industry and loan maturity structure across countries do not drive our results.

5.2 Alternative monetary policy indicators and shocks

As is standard in the related literature, we use a short-term interest rate as our indicator of the monetary policy stance. While appropriate for the initial years of our sample, this indicator might be an overly narrow representation of the stance over the latter part of the sample, when short-term interest rates approached their lower bound and the ECB resorted to unconventional monetary policy measures, such as quantitative

²⁵The data for the number of large firms is obtained from Eurostat and the share of value added of large firms is retrieved from the OECD data base. The data is available annually from 2005 except for Finland where the first observation on the share of value added is 2006. The latest available data for the share of value added is 2016 except for Spain where it is 2017 and for the number of large firms the latest available data is 2017 except for Belgium, Finland, Ireland, and Italy where it is 2016.

²⁶We obtain data for loans to the domestic sector by maturity from an internal ECB data base as these series are not publicly available. In contrast to total loans, these series are not adjusted for cash pooling, sales, and securitisation. However, when estimating our baseline model using aggregate loans data that is not adjusted for these items our results are unchanged.

²⁷For the firm size estimates the horizon at which the impulse response is evaluated is set to $h = 30$ as the through response in GDP at the average bond finance ratio is realized slightly later than in the baseline model (see Figure 12 and 13 in Appendix B for the full set of IRFs).

easing, to inject additional accommodation. We therefore test the model in alternative specifications that accommodate this change in policy conduct – first, by including a broader measure of the monetary policy stance in the regressions and, second, by allowing the coefficient on the policy shocks to differ across sample sub-periods.

To implement the former approach, we combine the 3-month OIS interest rate with a “shadow rate” which approximates the short-term interest rate that, based on broader yield curve constellations, would be expected to prevail in the absence of a lower bound. Following [Hartmann and Smets \(2018\)](#), we deploy a summary statistic of five shadow rate estimates, which is less sensitive to the specific model from which each individual shadow rate is derived. Specifically, we extract the first principle component of the euro area shadow rates by [Kortela \(2016\)](#), [Krippner \(2015\)](#), [Lemke and Vladu \(2017\)](#) (using an adaptive and a monotonic lower bound specification), and [Wu and Xia \(2017\)](#).²⁸ Up until December 2011, which we define as the starting point of the period in which the ECB engaged in unconventional monetary policy, we use the 3-month OIS rate and afterwards we add the cumulative change from the principal component to it.²⁹ This combined measure replaces the policy rate in our baseline model. Our second approach is slightly more agnostic in that it accounts for the presence of unconventional monetary policy via a dummy that is zero through November 2011 and one afterwards. This dummy is included in the regressions along with its interaction with the monetary policy shock, thus allowing for differences in shock transmission during the two sample periods.

Finally, we modify the construction of monetary policy shocks based on high-frequency surprises, which is essential to our identification approach. As laid out in Section 2.2, we carefully discipline our choice of the 1-month OIS rate surprises that we use as a measure of monetary policy shocks in our baseline estimation. However, to assure that this choice does not impact our findings we also test alternative measures. First, we vary the maturity of the shock, by instead: (i) using surprises in the 3-month OIS, which directly corresponds to the policy indicator in the model; and (ii) in line with [Nakamura and Steinsson \(2018\)](#), constructing a composite measure based on the first principle component of the surprises in the 1-week, 1-month, 3-month, 6-month, 1-year, and 2-year OIS rates. Second, we change the way we aggregate daily surprises to monthly frequency: in the baseline, the shock in each month consists of the unweighted sum of surprises taking

²⁸Since the measure by [Wu and Xia \(2017\)](#) starts only in September 2004 we extend it with the EONIA rate prior to that.

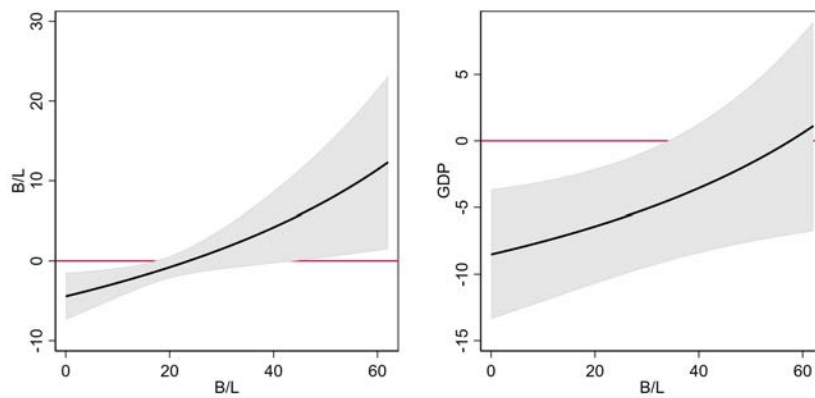
²⁹In December 2011, the ECB’s Governing Council decided on the first round of longer-term refinancing operations to banks which, given the very long (three-year) maturity, constituted a break with prior practice ([ECB \(2011\)](#)).

place over this period; but a monetary policy event that takes place at the beginning of the month has more time to propagate to other variables than a decision that is dated at the end of a month; to ensure that our baseline aggregation approach does not disregard relevant information, we thus construct weighted shocks, as proposed by [Gertler and Karadi \(2015\)](#), that reflect the timing of the monetary policy event. In particular, for each month, we construct the weighted average of the surprise in the 1-month OIS rate from the current and the previous month with the weights reflecting the number of days that have elapsed within each month that have elapsed up until the policy event relative to the total number of days in the month.³⁰

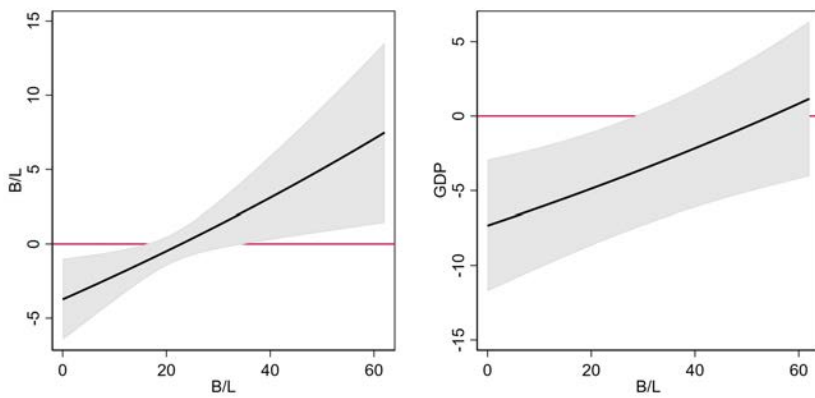
Again, the robustness checks confirm our main findings. In the specifications with alternative policy indicators, the results point to a significant reduction of bonds relative to loans in response to the shock when bond finance is low and an expansion when the prevalence of bond finance rises. The contraction in GDP again becomes smaller for higher levels of the initial bond finance ratio and loses significance at the upper end of the spectrum (see [Figures 8\(a\) and 8\(b\)](#)); the IRFs for the entire model are available in [Appendix B](#). Likewise, the alternative approaches to construct monetary policy shocks yield very similar point estimates, while in some cases exhibiting slightly lower precision (see [Figure 9](#)). Taken together, we can therefore also rule out that our particular measure of the monetary policy stance or the approach to construct policy shocks determines our findings.

³⁰As a concrete example: policy events took place on 12 September 2019 and on 24 October 2019. So, the value for the weighted policy shock for October 2019 is computed as 24/31 times the plain shock recorded in September plus 7/31 times the plain shock recorded in October. Further, by including lagged values of the weighted shock in the estimations, we purge them from the serial auto-correlation that this weighting scheme introduces; see also [Ramey \(2016\)](#) for a discussion of this issue.

Figure 8: Impulse responses across bond finance ratios (robustness: alternative monetary policy indicators)



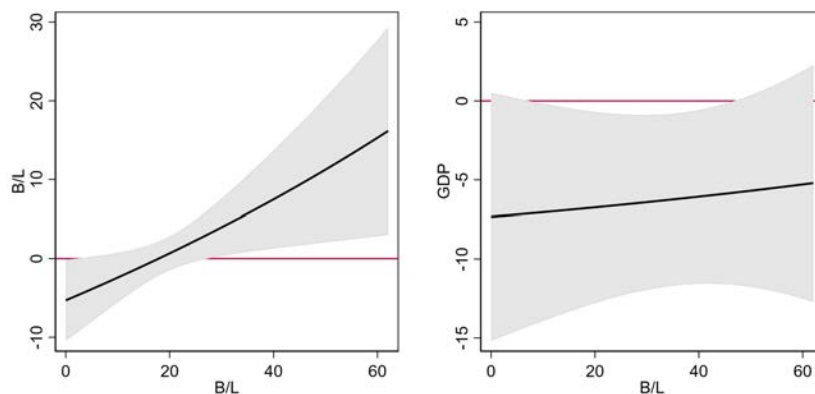
(a) Shadow rate



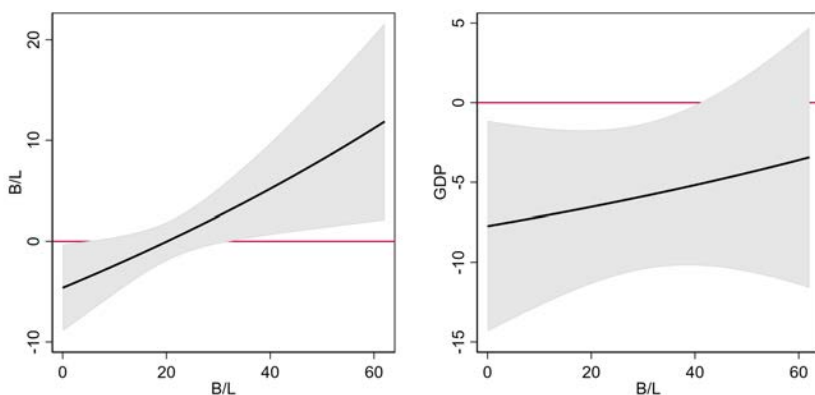
(b) Unconventional policy period

Note: The IRFs are normalised to a 100 bps impact response in the 3-month OIS rate at each point of the bond finance spectrum. The response horizon is $h = 24$. For the bond finance ratio the response is in percentage points and for GDP in percent. The grey area is the 90% confidence interval.

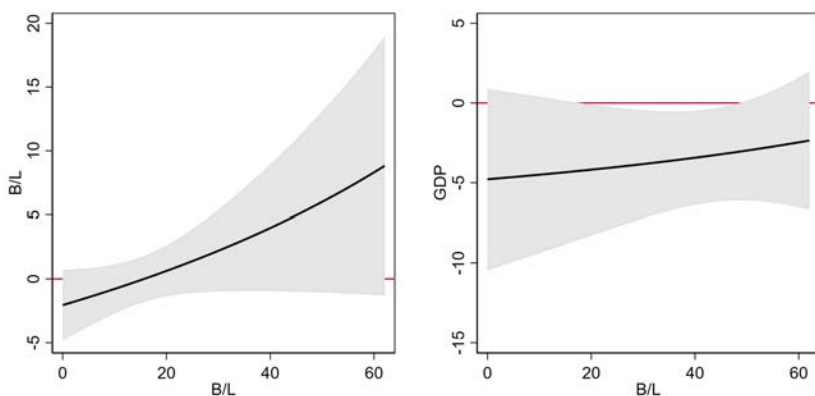
Figure 9: Impulse responses across bond finance ratios (robustness: alternative shocks)



(a) 3-month OIS



(b) Principal component



(c) Weighted surprises

Note: The IRFs are normalised to a 100 bps impact response in the 3-month OIS rate at each point of the bond finance spectrum. For the panels (a) and (b) the response horizon is $h = 24$ and $h = 30$ for panel (c). For the bond finance ratio the response is in percentage points and for GDP in percent. The grey area is the 90% confidence interval.

6 Conclusion

In this paper, we have studied how the relative role of corporate bonds versus bank loans in the debt structure of firms affects the transmission of monetary policy shocks to the economy. To this end, we have extended a standard empirical macro model with variables measuring the volume and cost of credit to firms. Estimating this model in a panel of euro area countries, and using high-frequency methods to identify monetary policy shocks, we find that the corporate debt financing structure is highly consequential for monetary policy transmission. In particular, a greater role of bond financing goes along with a weaker response of firm credit and economic activity to monetary policy shocks. Accordingly, bond finance acts as a substitute source of credit when bank lending contracts and, vice versa, it is crowded out when bank lending expands. At the same time, this substitution function requires the initial bond financing ratio in the economy to be relatively high. By contrast, in economies with a low bond financing ratio, the adjustment patterns weaken and may even reverse so that bond finance reinforces the overall response of firm credit to monetary policy shocks.

These patterns point to two countervailing mechanisms emphasised in the related theoretical literature. The substitution pattern at high bond finance ratios is consistent with the bank lending view of transmission, according to which monetary policy triggers shifts in loan supply that go beyond its impact on broader financing conditions in the economy. The amplification pattern at low bond finance ratios is consistent with state-contingent firm preferences for different financing instruments, as the flexibility afforded by bank lending (e.g. in renegotiating credit contracts in response to weakening business prospects) becomes more valuable in tight and less valuable in loose monetary conditions. The reversal in the impact of monetary policy on the bond finance ratio depending on the initial financing structure suggests that a sufficiently large initial reliance on bond finance is necessary for the substitution effect to dominate.

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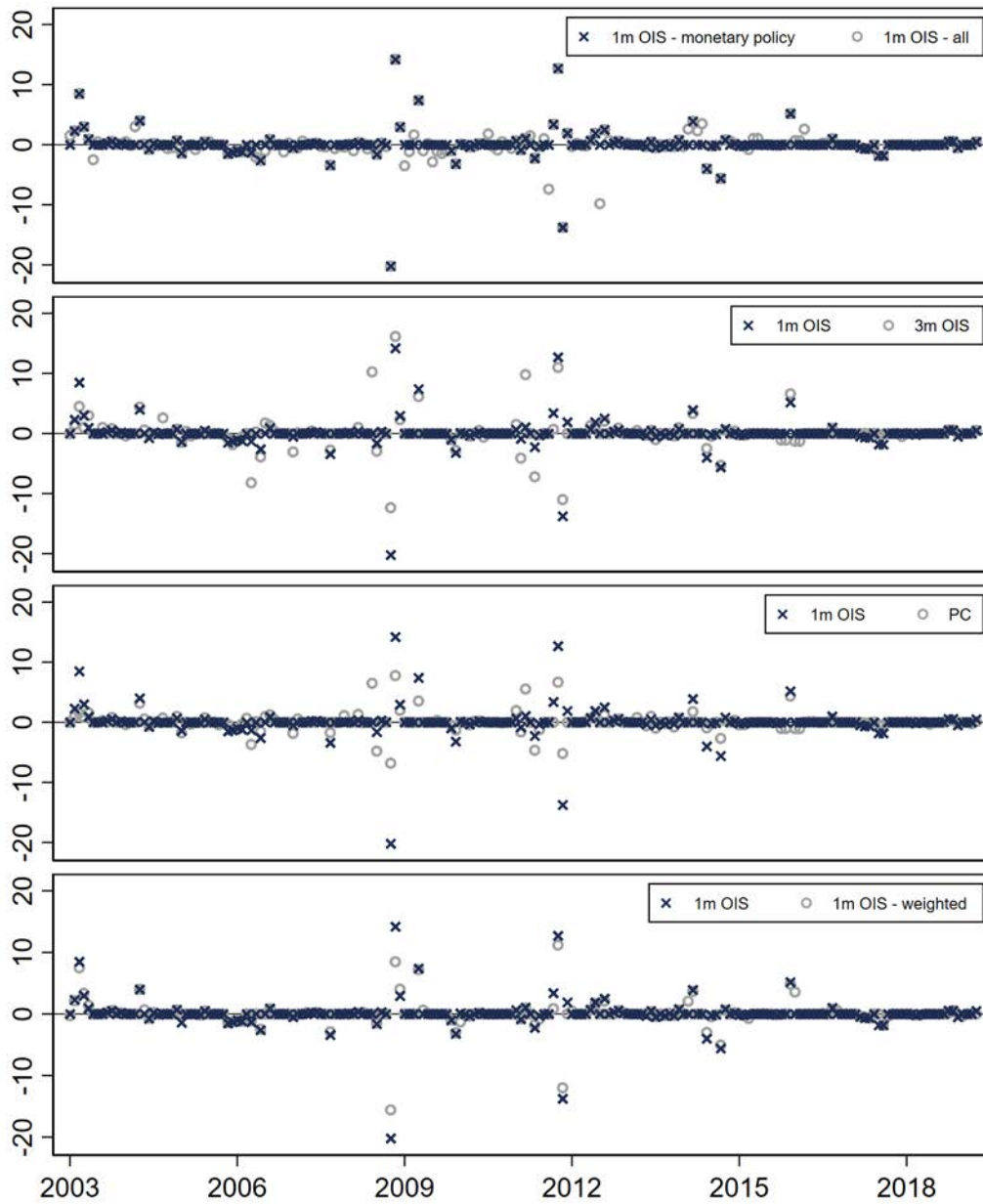
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Appendices

A High-frequency surprises

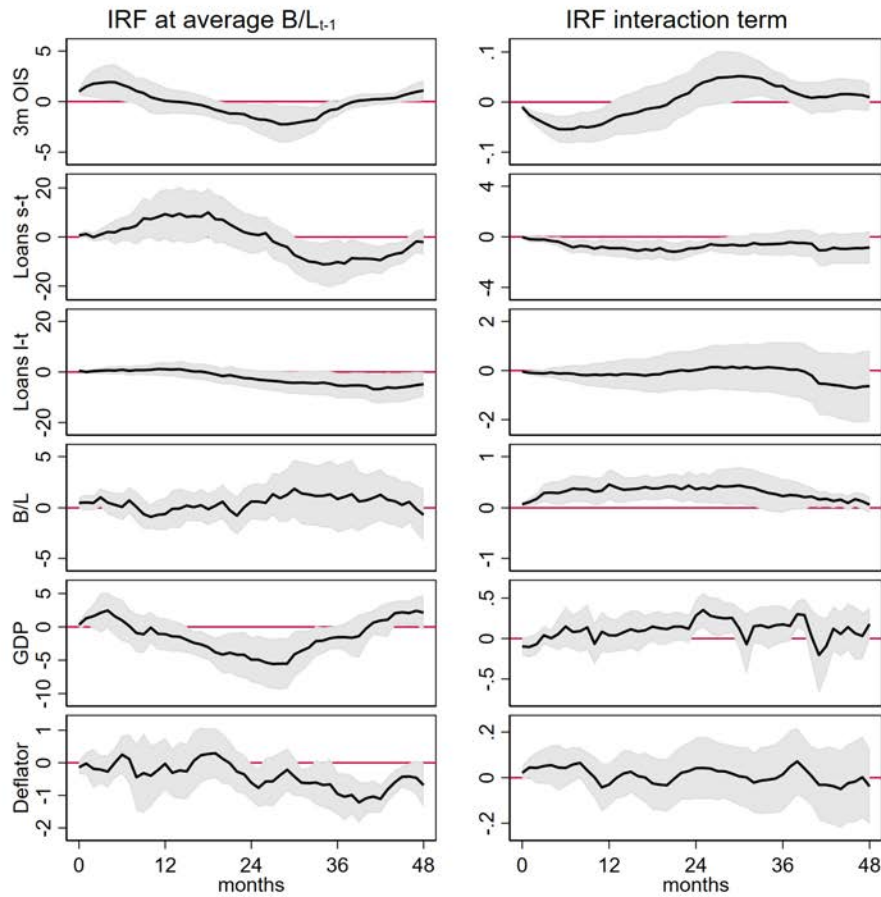
Figure 10: Time series of shocks



Note: All shocks except for "1m OIS - all" have been identified through the cross-asset correlation of the surprise between the respective interest rate and the stock price. The series with the x-marker represents the shock from our baseline model. The shocks are in bps.

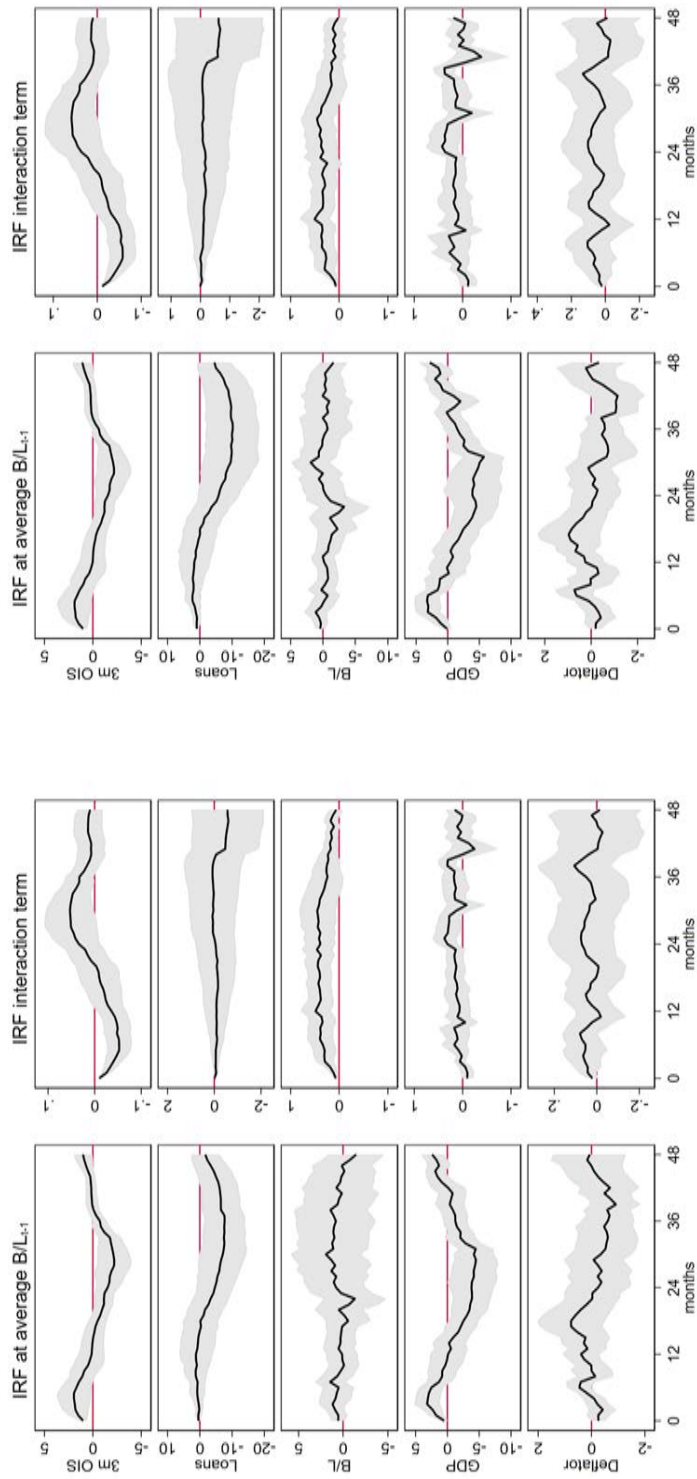
B Robustness

Figure 11: Impulse responses (maturity breakdown of loans)



Note: The IRFs in the left column are normalised to a 100 bps impact response in the 3-month OIS rate. The responses are measured in percent except for the 3-month OIS rate and the bond finance ratio where it is in percentage points. The grey area is the 90% confidence interval.

Figure 12: Impulse responses (robustness: cross-country heterogeneity – 1)

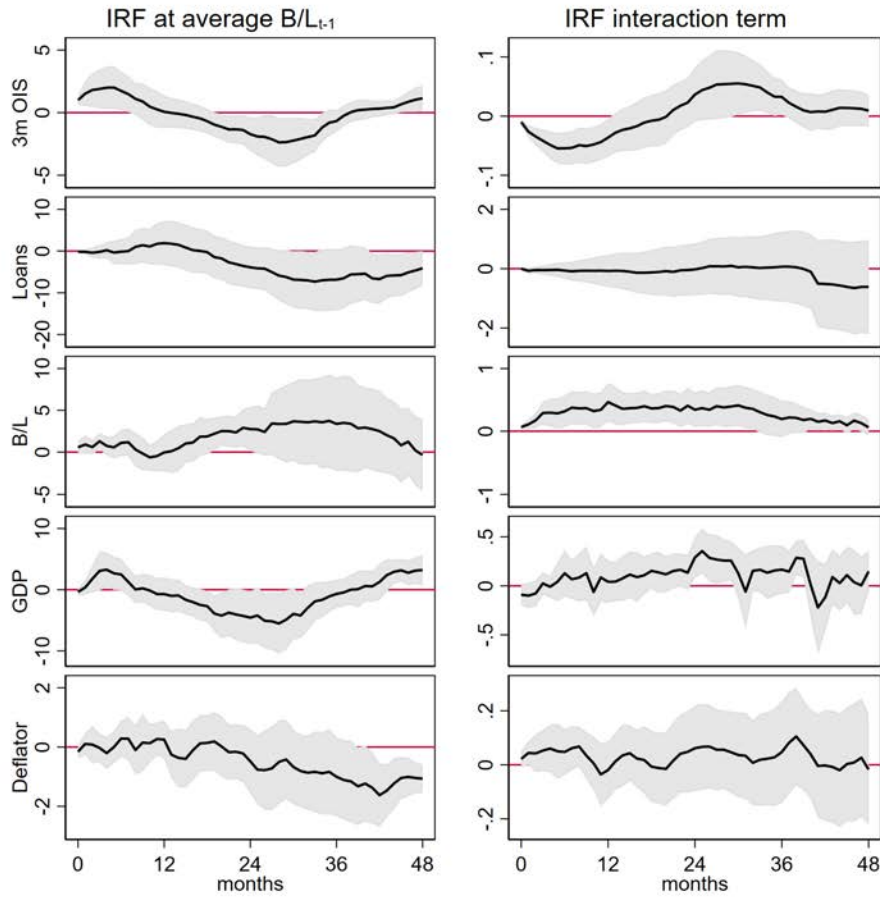


(a) Share of large firms

(b) Share of large firms in value added

Note: The IRFs in the left column are normalised to a 100 bps impact response in the 3-month OIS rate. The responses are measured in percent except for the 3-month OIS rate and the bond finance ratio where it is in percentage points. The grey area is the 90% confidence interval.

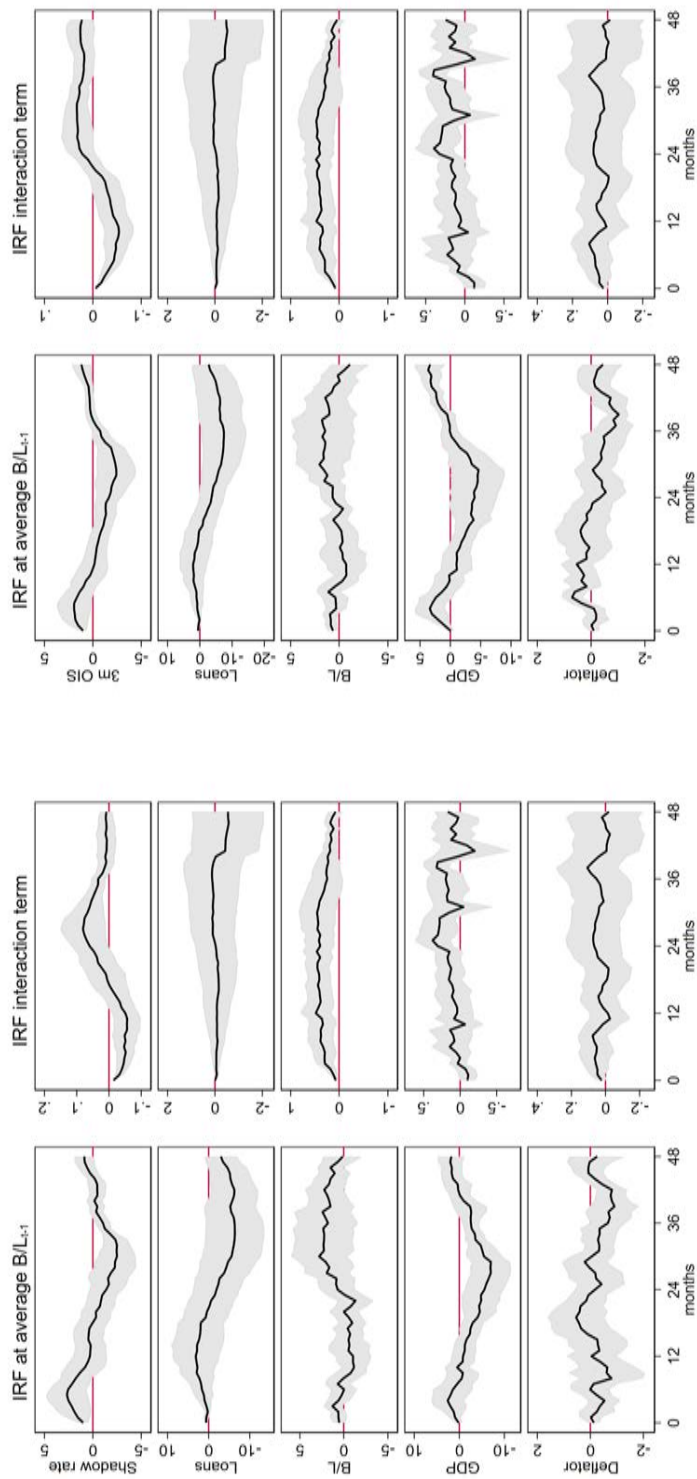
Figure 13: Impulse responses (robustness: cross-country heterogeneity – 2)



(c) Maturity of bank lending

Note: The IRFs in the left column are normalised to a 100 bps impact response in the 3-month OIS rate. The responses are measured in percent except for the 3-month OIS rate and the bond finance ratio where it is in percentage points. The grey area is the 90% confidence interval.

Figure 14: Impulse responses (robustness: alternative monetary policy indicators)

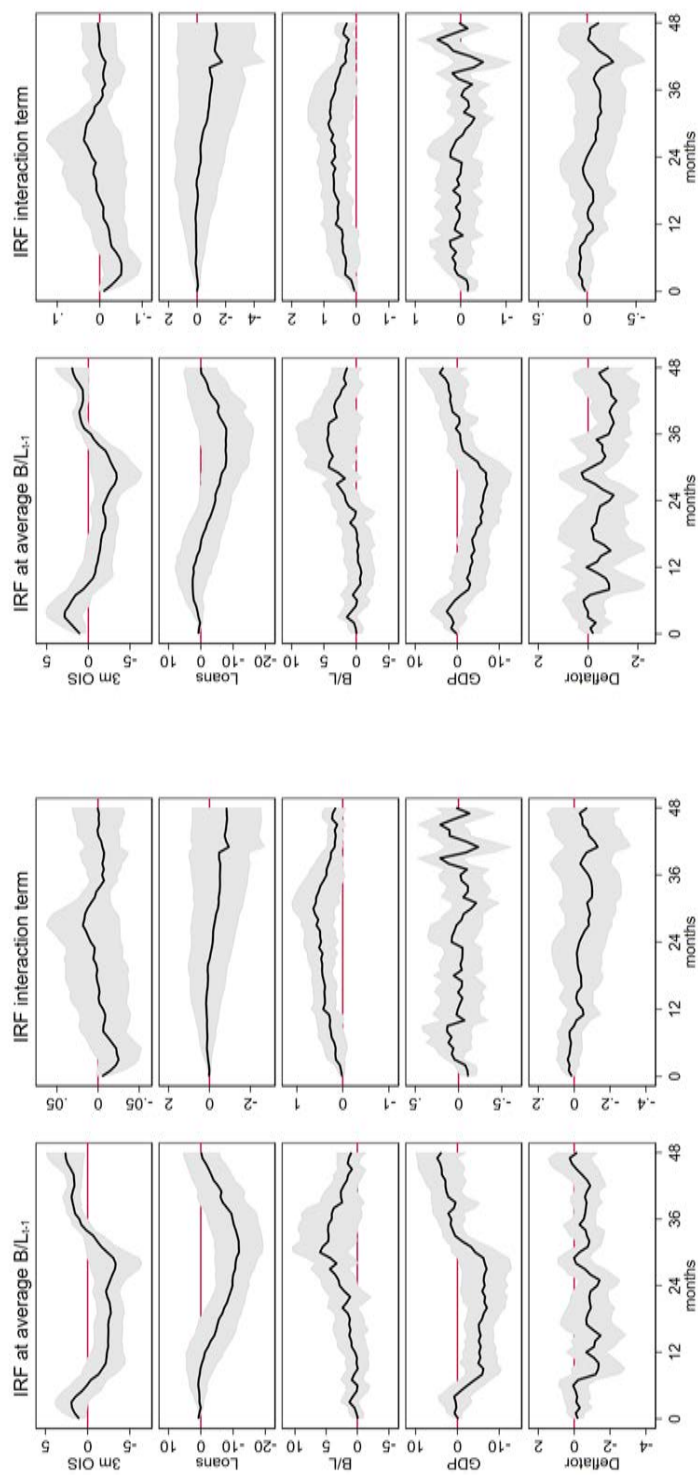


(a) Shadow rate

(b) Unconventional policy period

Note: The IRFs in the left column are normalised to a 100 bps impact response in the 3-month OIS rate. The responses are measured in percent except for the 3-month OIS rate and the bond finance ratio where it is in percentage points. The grey area is the 90% confidence interval.

Figure 15: Impulse responses (robustness: alternative shocks – 1)

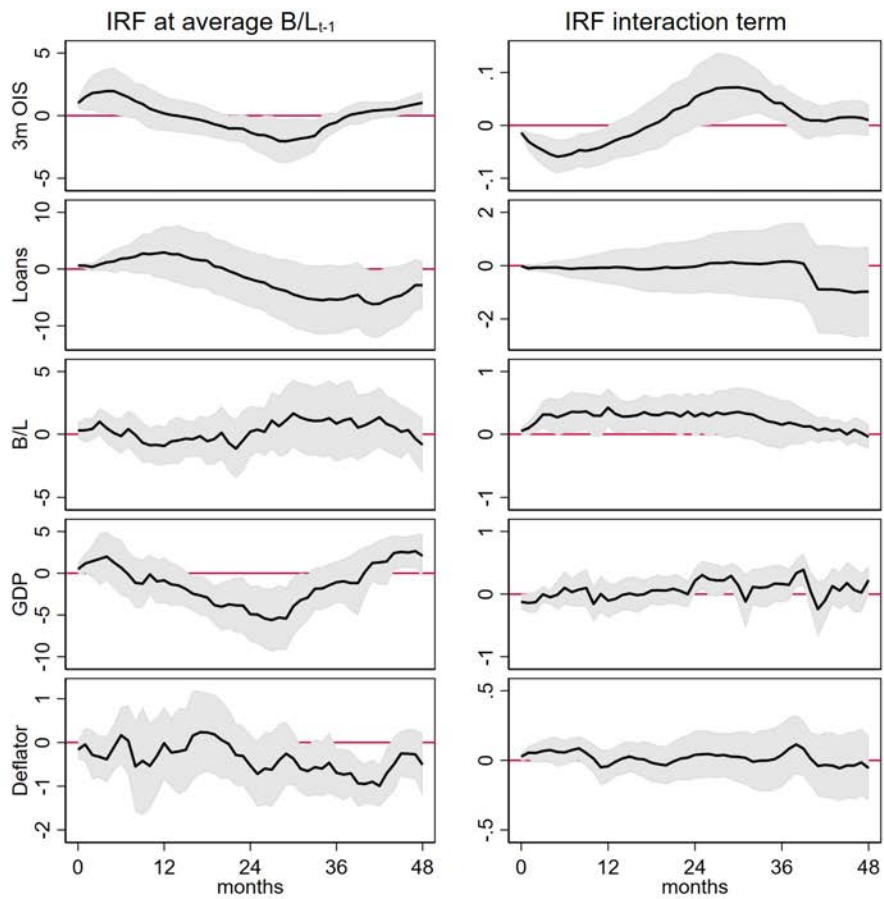


(a) 3-month OIS

(b) Principal component

Note: The IRFs in the left column are normalised to a 100 bps impact response in the 3-month OIS rate. The responses are measured in percent except for the 3-month OIS rate and the bond finance ratio where it is in percentage points. The grey area is the 90% confidence interval.

Figure 16: Impulse responses (robustness: alternative shocks – 2)



(c) Weighted surprises

Note: The IRFs in the left column are normalised to a 100 bps impact response in the 3-month OIS rate. The responses are measured in percent except for the 3-month OIS rate and the bond finance ratio where it is in percentage points. The grey area is the 90% confidence interval.

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Fédéric Holm-Hadulla

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

Claire Thürwächter

Institute for International Economic Studies, Stockholm University, Stockholm, Sweden; email: claire.thurwachter@iies.su.se

© European Central Bank, 2020

Postal address 60640 Frankfurt am Main, Germany

Telephone +49 69 1344 0

Website www.ecb.europa.eu

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