Evidence about the transmission of monetary policy

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Olli-Matti Laine

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Foreword

I conducted this research while working at the Bank of Finland. I would like to thank the former Head of the Monetary Policy and International Economy division, Hanna Freystätter, PhD, for the opportunity to dedicate time to this project. She initially proposed the topic of the second study of this dissertation. I am also grateful to the Head of Research, Docent Esa Jokivuolle, PhD, and the Head of Monetary Policy and Research Department, Docent Juha Kilponen, PhD, for their valuable guidance and comments during this research project.

I express my sincere gratitude to my supervisor at the Tampere University, Hannu Laurila, PhD. His inspiring teaching in my undergraduate studies was one of the reasons for choosing economics as my major instead of accounting and finance.

I would also like to thank the preliminary examiners of this dissertation, Professor of Economics at the University of Jyväskylä, Juha Junttila, PhD, and Docent Juuso Vataja, PhD.

I have received comments and suggestions from many others as well. The idea for the first study was born in the Bank of Finland's dealing room. The head of monetary policy implementation, Niko Herrala, MSc, noted that it might be worthwhile to decompose stock market movements into movements of risk premia, risk-free rates and dividend expectations. In taking up the challenge, I received many good comments from my colleagues Mikael Ahovuo, MSc, Jukka Lähdemäki, MSc, Rasmus Rannikko, MSc, and Kristian Tötterman, MSc. I am also grateful to Gene Ambrocio, PhD, for comments at the early stage of the research project. As the work progressed Markus Sihvonen, PhD, offered valuable insights on the existing literature, the dividend discount formula and the pricing of dividend future contracts. I am also grateful to the Professor of Money, Banking, and Finance at the University of Illinois, George Pennacchi, PhD, for his valuable comments on the pricing of dividend futures. Regarding the applied econometrics, I received good comments from Jaakko Nelimarkka, PhD. I thank Professor of Finance at the University of Turku, Mika Vaihekoski, PhD, for his comments at the GSF Winter Workshop in Finance. In addition, I want to thank Juho Alasalmi, MSc, and Professor of Economics at the BI Norwegian Business School, Fabio Canova, PhD, and Samu Kärkkäinen, PhD, and Jukka Ilomäki, PhD, and Lauri Vilmi, PhD, for commenting the first article.

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I wish to thank Gregory Moore, JD, for revising the language of the manuscript.

I want to thank my family as well. I am grateful to my parents for all their support. I believe that they have had a major impact on this piece of work accidentally; my initial interest in monetary policy and financial markets began as a teenager probably after some discussion with my mother Hannele Laine, MSc; academic endeavours have always felt natural thanks to my father, Docent Heikki-Jussi Laine, M.D. Finally, I want to thank my dear wife Essi Laine, M.A. Having a linguist at home has been an asset while writing the research articles. She has read many parts of my manuscripts and improved my English. Most importantly, she has looked after my mental health and occasionally asked me to step out of my study.

Helsinki, February 2022

Olli-Matti Laine

List of original publications

I. Laine, O.-M. (forthcoming). Monetary Policy and Stock Market Valuation. *International Journal of Central Banking*. The article is printed here with the permission of the International Journal of Central Banking, where the article will be published in 2023.

II. Laine, O.-M. (2021). The Effect of Targeted Monetary Policy on Bank Lending. *Journal of Banking and Financial Economics*, 1 (15), 25–43. The article is reprinted here under the Creative Commons Attribution 4.0 International License.

III. Laine, O.-M. (2020). The effect of the ECB's conventional monetary policy on the real economy: FAVAR-approach. *Empirical Economics*, 59(6), 2899-2924. The article is reprinted here under the Creative Commons Attribution 4.0 International License.

ABSTRACT

This doctoral dissertation analyses the transmission of monetary policy. It applies a variety of empirical methods to study how conventional and unconventional monetary policy measures transmit to different macroeconomic and financial variables.

The first article analyses the effect of monetary policy on the term structure of stock market risk premia. The implied term structure is solved in a novel way utilizing equity analysts' dividend forecasts and dividend future prices. The results show that monetary policy affects risk premia differently at different discounting horizons. Monetary policy easing lowers the short-horizon premia and raises the long-horizon premia. The effect on the average risk premium is positive.

The second article studies the effect of targeted longer-term refinancing operations on bank lending. The results suggest that these targeted operations stimulated bank lending to firms. However, no evidence about a positive effect on lending to households is found.

The third article examines the effects of conventional monetary policy during the 2008 financial crisis and the era of ultra-low interest rates. Several earlier studies conclude that the effects of conventional monetary policy shocks stayed almost the same during and after the financial crisis. Revisiting this research question, the findings suggest that the impulse response functions of industrial production and unemployment changed drastically after the financial crisis.

Keywords: monetary policy, stock market, bank lending, time-varying effects

JEL: E44, E51, E52, E58, E6, G12, G21

TIIVISTELMÄ

Tässä väitöskirjassa analysoidaan rahapolitiikan välittymistä. Useita erilaisia empiirisiä menetelmiä hyödyntäen tutkitaan, kuinka tavanomainen ja epätavanomainen rahapolitiikka välittyvät eri makrotaloudellisiin ja rahoituksellisiin muuttujiin.

Ensimmäisessä osatyössä analysoidaan rahapolitiikan vaikutusta osakemarkkinoiden riskipreemioiden aikarakenteeseen. Piilevä aikarakenne ratkaistaan uudenlaisella tavalla hyödyntämällä osakeanalyytikoiden osinkoennusteita sekä osinkofutuureja. Tulokset osoittavat, että rahapolitiikka vaikuttaa eri diskonttaushorisonttien riskipreemioihin eri tavoin. Rahapolitiikan keventäminen alentaa lyhyen horisontin riskipreemiota ja nostaa kaukaisten horisonttien riskipreemioita. Vaikutus keskimääräiseen riskipreemioon on positiivinen.

Toisessa osatyössä tarkastellaan kohdennettujen pitempiaikaisten jälleenrahoitusoperaatioiden vaikutusta pankkien luotonantoon. Tulokset viittaavat siihen, että kohdennetut operaatiot kasvattavat pankkien luotonantoa yrityksille. Positiivisesta vaikutuksesta kotitalouslainanantoon ei kuitenkaan löydy näyttöä.

Kolmannessa osatyössä tarkastellaan tavanomaisen rahapolitiikan vaikutuksia vuoden 2008 finanssikriisin ja äärimmäisten alhaisten korkojen aikana. Useassa aiemmassa tutkimuksessa tullaan tulokseen, että tavanomaisten rahapolitiikkasokkien vaikutukset pysyivät lähes ennallaan koko finanssikriisin ajan ja sen jälkeen. Kolmannessa osatyössä tarkastellaan tutkimuskysymystä uudelleen. Tulokset viittaavat siihen, että teollisuustuotannon ja työttömyyden impulssivastefunktiot muuttuivat voimakkaasti finanssikriisin jälkeen.

Avainsanat: rahapolitiikka, osakemarkkinat, pankkien lainananto, ajassa muuttuvat vaikutukset

JEL: E44, E51, E52, E58, E6, G12, G21

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1 Introduction

High inflation, unemployment and economic fluctuations adversely affect people's well-being (e.g. Di Tella, MacCulloch and Oswald, 2001, 2003). With strong empirical evidence to these effects, modern central banks typically enjoy mandates to pursue price stability and high employment. The actions central banks take in pursuit of such targets is called monetary policy. Monetary policy has conventionally meant setting the level of short-term interest rates, and more recently, for example quantitative easing, forward guidance and targeted lending operations – tools that were previously considered unconventional. The conduct of monetary policy requires an understanding of how these policy measures transmit to the economy, so it is hardly surprising that monetary policy transmission mechanism is the subject of extensive study.[1]

Since the 2008 global financial crisis drove monetary policymakers to the effective lower bound of interest rates, central banks have tried to achieve their policy goals by influencing long-term interest rates and various risk premia. To accomplish this, they have, for example, bought long-term bonds, communicated about the future path of short-term interest rates and targeted their refinancing operations.[2]

Unusual circumstances and unconventional monetary policy measures have created high demand for empirical research on the transmission of these policy measures. The last time interest rates were as low as during the past decade was in the 1930s during the Great Depression and run-up to WWII. Many of today's novel monetary policy tools have been implemented without prior knowledge of their effectiveness. At the same time equity prices have surged on many stock exchanges, and some commentators have expressed worries that central banks are fuelling stock market bubbles or overheating markets. However, the previous literature offers limited evidence as to how monetary policy affects the required rates of return used to discount future expected dividends.

This dissertation contributes to the literature by providing evidence about the effects of monetary policy. The first article "Monetary Policy and Stock Market Valuation" (Study I) studies the effect of monetary policy on the equity risk premium. The second article "The Effect of Targeted Monetary Policy on Bank Lending" (Study II) analyses the effectiveness of the longer-term targeted refinancing operations on bank lending. The final article "The effect of the ECB's conventional monetary policy on the real economy: FAVAR-approach" (Study III) studies the transmission of conventional monetary policy shocks during the 2008 global financial crisis.

[1] See e.g. the well-known reviews of Christiano, Eichenbaum and Evans (1999) and Boivin, Kiley and Mishkin (2010).

[2] Rostagno, Altavilla, Carboni, Lemke, Motto, Guilhem and Yiangou (2021) describe the monetary policy conducted by the European Central Bank (ECB). Bernanke (2020) discusses these novel monetary policy tools from the perspective of the Federal Reserve (Fed).

The dissertation has the following structure. Section 2 reviews the literature. Section 3 explains some methodological choices and provides an overview of the data. Section 4 summarises the three original articles included in the dissertation. Section 5 concludes.

2 Literature review

2.1 Monetary policy and equities

As noted by Sellin (2001), the discussion of monetary policy impacts on equities straddles monetary economics and finance. Different fields pose slightly different questions, creating a somewhat scattered literature. In general, asset pricing or finance studies focus on explaining and predicting stock returns – the monetary policy discussion is largely incidental. In contrast, monetary economists and macroeconomists are interested in the transmission of monetary policy per se.

It is widely acknowledged that expected returns vary over time (e.g. Fama and French, 1989; Cochrane, 2008).[3] Variables such as dividend yield or term spread predict stock returns (for rational or irrational reasons). A well-known rational explanation for the predictability of stock returns is provided by Campbell and Cochrane (1999).[4]

Patelis (1997) analyses the role of monetary policy in predicting stock returns. He starts his analysis by including monetary policy indicators such as the federal funds rate in an otherwise standard long-horizon predictive regression that follows Fama and French (1989). His results show that tighter monetary policy predicts lower expected returns. The regression coefficients, however, decline in absolute value, indicating higher expected short-horizon returns in the future. He then incorporates the first difference of federal funds rate into a VAR model similar to that of Campbell and Ammer (1993) and arrives at similar conclusions. It should be noted that the "long horizon" considered by Patelis is just two years, i.e. one may only draw conclusions regarding equity risk premia at relatively short horizons.[5]

Bernanke and Kuttner (2005) also apply the VAR framework of Campbell and Ammer (1993). They include a monetary policy surprise as an exogenous variable in their VAR model to study the dynamic response of excess equity return to an unexpected tightening of monetary policy. The

[3] The cross-section of stock returns forms a separate strand in the finance literature. The model by Fama and French (2015), which serves as a benchmark empirical model for explaining the cross-section of stock returns, includes a market factor, as well as size, book-to-market, profitability and investment factors.

[4] Campbell and Cochrane (1999) assume that investors have time-varying subsistence level, added to the standard power utility function. When consumption declines in a recession, the curvature of the utility function rises. Thus, prices of risky assets fall and expected returns rise.

[5] The value of a stock is the present value of all its expected future dividends. As the expected excess returns of stock vary over time, it is reasonable to think that equity risk premia used to discount the future expected dividends vary over discounting horizons. The discount rates for the first two years may behave differently than discount rates elsewhere in the term structure of discount rates.

results show that contractionary monetary policy generates an immediate decline in equity prices followed by a period of higher-than-normal excess returns. Bernanke and Kuttner report the responses for the first two years after the monetary policy shock meaning that, like in Patelis (1997), the interpretation regarding equity risk premia for long horizons is not possible.

While Bernanke and Kuttner (2005) provide *indirect* evidence that monetary policy tightening raises the short-horizon equity risk premia by looking at realized excess returns, others attempt *direct* assessment of the equity premium. For example, Claus and Thomas (2001) solve the (average) equity premium implied by analysts' cash flow expectations and study its variation over time. The advantage of implied risk premia over regression-based expected excess returns is that the researcher does not need to make assumptions about the econometric model. A potential disadvantage is that it may be difficult to find data about dividend forecasts that represent the true expectations of market participants.

Implied risk premia and regression-based premia may also have different interpretations. The dividend forecasts of analyst or market participants may be biased. Therefore, implied premia may differ systematically from the expected returns given by a "true model" that generates the excess returns. This is not a disadvantage if the dividend forecasts used represent true dividend expectations and the purpose is analysing risk premia rather than forecasting excess returns. Investors may have biased expectations, so implied premia may represent the investors' required risk premia better than the predictive regressions used e.g. in Patelis (1997).

In Study I of this dissertation, I delve for the first time into the impact of monetary policy on the implied risk premia.[6] My results support the conclusion of Bernanke and Kuttner (2005) with respect to short-horizon premia. However, the effect on risk premia at longer horizons seem to have the opposite sign, i.e. the effect of monetary policy tightening on the average long-term equity premium is negative.

A bit similar result is found by Galí and Gambetti (2015), who use a theoretical framework with bubbles and in which investors are risk-neutral.[7]

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[6] See also Laine (in press) for some descriptive statistics of the risk premia.

[7] Galí and Gambetti (2015) use log-linearized form of the dividend discount model and realized dividends. However, one could apply the methodology of Study I of this dissertation to assess the effect of monetary policy on the bubble component directly, rather than indirectly by looking at the impulse response function of stock index as in Galí and Gambetti (2015). Their model, which has not been linearized, is represented by their equations (1) and (2). A direct way to model the bubble component would be to use equation (2) of Study I and assume that the risk premium is zero or constant. One could then numerically solve the fundamental component implied by the dividend forecasts and some proxy for the risk-free rate. The difference between this implied fundamental component and the observed market price would represent the bubble component of Galí and Gambetti (2015). In the language of Study I, this difference could be described as a "premium component." Thus, Galí and Gambetti (2015) and Study I essentially look at the same thing from different angles. Galí and Gambetti talk about bubbles and Study I about risk premia. Neither of these empirical studies disambiguates the part of the premium (or bubble) that represents the rational risk premium from mispricing.

Their work suggests that tightening monetary policy shock may raise stock prices persistently – even though the effect on the dividends is negative and the effect on the real interest rate is positive. They argue that the effect is likely not due to the rational risk premium but the bubble component. Study I notes that there are also reasons such as the inflation risk premium that could justify rational risk premium interpretation. While the theoretical reasons for the results are unclear, Galí and Gambetti (2015) and Study I show that the effect of monetary policy on the valuation of stocks may not necessarily comport with conventional wisdom.

A theoretical model that explains the effect on short-horizon premia documented in Bernanke and Kuttner (2005) and Study I is provided by Gust and López-Salido (2014). In their model, monetary expansion reduces the riskiness of the economy. The key feature of the model is that some households rebalance their portfolios infrequently. Monetary policy affects the share of households that are active in the market, thereby affecting risk. Monetary expansion increases the share of active households. Thus, aggregate risk becomes more broadly shared across households, lowering the risk premium.

The theory explaining the effect of opposite sign on the long-horizon premia is scarce. According to the explanation of Galí (2014) that is based on rational bubbles, a transitory rate hike does not alter the size of the bubble on impact, but has a positive effect on its subsequent growth rate that leads to a permanent enlargement. This effect could explain the results regarding the long-horizon implied premia of Study I. The experiment of Galí, Giusti and Noussair (2021) gives some support for this bubble-based explanation.

A potential rational-risk-premium-based explanation is provided by Caballero and Farhi (2013, 2018). Their idea is that central banks may increase the riskiness of risky assets by buying assets such as long-term government bonds that provide insurance against risks. One should note that this potential mechanism is related to specific type of asset purchases and unlikely to explain the entire effect documented in Study I. For example, the baseline local projections in the article tell mainly about the effects of conventional short-rate monetary policy.

Monetary policy may also affect the long-horizon risk premia via inflation. Modigliani and Cohn (1979) argue that, due to money illusion, higher inflation means too low stock prices (higher implied premia). Investors incorrectly discount expected real cash flows using nominal rates instead of real rates (or discount nominal cash flows using nominal rates and forget include inflation forecasts in their cash flow forecasts). While the results of Campbell and Vuolteenaho (2004) support this behavioural explanation, Bekaert and Engstrom (2010) argue that the phenomenon is also consistent with asset pricing theory that incorporates uncertainty about real growth prospects and habit-based risk aversion.[8] Whether rational or not, inflation that is influenced by monetary policy seems to affect stock market valuations.

2.2 Targeted longer-term refinancing operations

Funding for lending (FFL) is perhaps the most unconventional policy measure introduced during the past decade. It was adopted in some form by for example the ECB, the Fed, the Bank of Japan (BoJ) and the Bank of England (BoE). While FFL programmes differ across central banks, they share a common feature of channelling central bank funding to bank lending. In the euro area, these lending programmes are called targeted longer-term refinancing operations (TLTROs).[9]

As the TLTROs are rather new, there is only a thin body of work on FFL/TLTRO transmission. The majority of the existing literature focuses on *non-targeted* longer-term refinancing operations (e.g. Andrade, Cahn, Fraisse, and Mésonnier, 2019; Bednarek, Dinger, te Kaat and von Westernhagen, 2021; Carpinelli and Crosignani, 2021; Crosignani, Faria-e-Castro and Fonseca, 2020; Darracq-Paries and De Santis, 2015; Garcia-Posada and Marchetti, 2016). Without uncertainty, the maturity of refinancing operations plays no role as banks can just roll over their short-term loans. If there is uncertainty regarding future central bank refinancing, however, maturity becomes relevant (e.g. Carpinelli and Crosignani, 2021). The articles listed in this paragraph show that extending the maturity of central bank funding has increased bank lending to firms. On the other hand, there is also some evidence about possibly unintended effects on the government bond holdings of banks (Crosignani et al., 2020; Carpinelli and Crosignani, 2021).

TLTROs have built-in incentives for banks to increase their lending to the private sector. The incentives vary across operations, but recent operations (TLTRO II and III) have rewarded banks with low interest rates for lending to firms and households (excluding house purchase loans). Study II of this dissertation shows that TLTROs have boosted lending to the corporate sector without increasing the government bond holdings of banks. Broadly similar results are found in the contemporaneously published papers of Benetton and Fantino (2021) and Afonso and Sousa-Leite (2020). Andreeva and García-Posada (2021) assess the mechanism through which

[8] The "Fed model" discussed in Campbell and Vuolteenaho (2004) and Bekaert and Engstrom (2010) is a simplification of equation (2) in Study I. The Fed model assumes that the expected dividend growth rate is constant at all horizons. Equation (2) of Study I then simplifies to $P_0 = E_0[\frac{D_1}{r-g}]$, which is typically called "Gordon growth model" after Gordon (1962). We arrive at equation (1) of Campbell and Vuolteenaho (2004) by moving one period earlier and rearranging: $\frac{D_0}{P_{-1}} = r - g$.

[9] For the BoE's Funding for Lending Scheme, see the article by Churm, Joyce, Kapetanios and Theodoridis (2018). English and Liang (2020) discuss the Fed's Main Street Lending Program. See also Kandrac (2021).

TLTROs transmit to bank lending utilising data from the ECB's Bank Lending Survey (BLS). Because participating banks demand less market-based funding, that note that the bank bond yields of both participants and non-participants decline. The indirect effect on bank lending of non-participating banks is ambiguous as the participating banks can gain market shares at the expense of non-participants. Andreeva and García-Posada (2021) find evidence that this mechanism is not very significant economically and that TLTROs have increased lending on the aggregate level. They find that TLTROs have lowered the margins of participants on relatively safe borrowers and eased credit standards for non-participants.

Unlike Benetton and Fantino (2021) and Andreeva and García-Posada (2021), Study II analyses also the impact on loans to households for consumption. The article does not provide evidence in favour of positive effect on lending for consumption though this loan category was targeted by the ECB. This result is relevant for policy making, because households are more dependent on bank loans than firms because firms (especially large ones) have alternative financing sources as well. Study II also notes that the effects TLTROs vary across countries. This is hardly surprising as economic conditions vary across countries, but the observation is still important. Many papers that study targeted or non-targeted operations rely on data from a single country.[10]

At least two papers analyse the transmission of TLTROs to macroeconomic variables such as inflation or GDP (Balfoussia and Gibson, 2016; Nelimarkka and Laine, 2021). Both studies find that TLTROs have had non-negligible macroeconomic effects. Ambler and Rumler (2019) show that TLTRO announcements affect multiple interest rates, suggesting that the effects of TLTROs are not limited to bank lending.

2.3 Time-varying effects of conventional monetary policy

It is largely acknowledged that the effectiveness of monetary policy evolves over time (e.g. Primiceri, 2005; Boivin et al. 2010; Matějů, 2019). For example, credit markets evolve due to regulatory changes and technological updates, and central banks shift their monetary policy strategies. In addition to these long-term developments, the strength of transmission may also vary over the business cycle (e.g. Ciccarelli, Maddaloni and Peydró, 2013; Matějů, 2019; Paul, 2020).

Possibly the most significant change in the central banks' environment in recent decades has been the lowering of interest rates. Even so, it is not well understood how persistently low interest rates affect the transmission mechanism of conventional monetary policy. Brunnermeier and Koby (2018) show that the short-term interest rate policies may turn from

[10] García-Posada and Marchetti (2016) study the effects in Spain, Andrade et al. (2018) in France, Benetton and Fantino (2021) and Carpinelli and Crosignani (2021) in Italy.

accommodative to contractionary when interest rates are sufficiently low. Borio and Gambacorta (2017) provide empirical evidence that monetary policy is less effective in stimulating bank lending growth when interest rates reach a very low level. In contrast, Altavilla, Burlon, Giannetti and Holton (2021) find no evidence that monetary policy becomes ineffective when rates are negative.

Low interest rates may change the transmission mechanism of monetary policy via channels other than bank lending. Conventional monetary policy also transmits to the economy by affecting long-term real interest rates (Hanson and Stein, 2015). If long-term interest rates are already approaching their effective lower bound, cutting short-term policy rate may have a smaller effect on long-term rates than during normal times.

Empirical evidence regarding the time-variation of the effectiveness of conventional monetary policy on long-term or medium-term interest rates is mixed. Cenesizoglu, Larocque and Normandin (2018) provide some evidence that the effect of conventional monetary policy in the US on long-term interest rates declined after 2007. On the other hand, the results of Von Borstel, Eickmeier and Krippner (2016) suggest that the transmission of conventional monetary policy to lending rates and the term spread did not change in the euro area during or after the 2008 financial crisis.

Uncertainty and risk perceptions may also affect the effectiveness of monetary policy. Aastveit, Natvik and Sola (2017), Pellegrino (2018) and Hauzenberger, Pfarrhofer and Stelzer (2021) show that high uncertainty dampens the effects of monetary policy shocks. On the other hand, for example Mishkin (2009) argues that the effect of monetary policy could be stronger during crisis periods because then its effect on risk premia is stronger. Aysun, Brady and Honig (2013) provide empirical evidence about a positive relationship between financial frictions and the strength of monetary transmission. For example, Kashyap and Stein (2000) and Albertazzi, Nobili and Signoretti (2021) show that conventional monetary policy transmits more strongly via weak banks than strong banks.

Given this evidence, it is not unambiguous how the effectiveness of monetary policy has developed during the first decades of 21st century. If we ignore the transmission mechanism and look at the effects on the real economy and inflation the evidence is mixed. Ciccarelli et al. (2013) provide evidence in favour of stronger effects of monetary policy shocks on GDP during the crisis due to credit channel in the euro area. Lopez-Buenache (2019) finds evidence supporting increased effectiveness of monetary policy after 2008 in the US. The results of Salisu and Gupta (2020) suggest that monetary policy has been more effective during crisis periods than normal times in the UK. According to Bagzibagli (2014) and Finck (2019) monetary policy shocks have had almost identical impacts on macroeconomic variables before and during the financial crisis of 2008 in the eurozone. The results of Pellegrino (2018) suggest that conventional monetary policy was less effective during the financial crisis than before it in the euro area. Matějů (2019) argues that during the banking crises, monetary policy transmission has been weaker than otherwise. Study III provides some evidence in favour of weakened effect of conventional monetary policy after the financial crisis of 2008 in the euro area.

These wide-ranging empirical results highlight the fact that the time-variation in the effectiveness of monetary policy shocks is not fully understood. There are multiple issues that may explain the diversity of results, including the differences in empirical approaches, differences in shock identifications and different datasets. As is typical in empirical macroeconomics, the confidence bands around the impulse response functions are quite wide. Researchers may also draw differing conclusions from roughly similar results.

Looking ahead, it is important to understand the mechanisms driving the changing effectiveness of monetary policy. In this regard, the contribution of Matějů (2019) is significant. Another policy relevant issue is the possible asymmetry in the effectiveness of monetary policy: tight policy may be more effective than accommodative policy (Florio, 2004). This question is currently highly relevant as central banks consider raising their policy rates after a prolonged period of low interest rates.

3 Methodology and data

3.1 Econometric methods

The articles of this dissertation apply several methods, including vector autoregressions (VARs), local-projections and difference-in-differences.

VAR models have been widely used in empirical macroeconomics since the seminal work of Sims (1980). Their advantage over other models such as dynamic stochastic general equilibrium (DSGE) models is that they require fewer theoretical assumptions regarding the dynamic relationships between variables. Causal inference can be done by studying the responses of variables to an exogenous structural shock (impulse response functions, IRFs). This approach is used in the first and third articles of this dissertation.

A specific challenge with structural VARs is the identification of shocks from the error terms of a reduced form VAR. Ramey (2016), for example, discusses the different identification methods. Study I uses the sign restriction method of Rubio-Ramirez, Waggoner and Zha (2010).[11] Sign restrictions are widely used in the current literature. Faust (1998) and Uhlig (2005) note that researchers tend to accept VAR results that are “reasonable” in some sense. Sign restrictions make this hidden model selection transparent and explicit. Their key idea is that a shock of interest can be identified by

[11] I use the MATLAB code of Ferroni and Canova (2021) to implement the algorithm.

using a priori knowledge about the effects of the shock on those variables whose responses are not of interest in the study at hand. Study III uses the more traditional Cholesky decomposition for identification. It assumes that only some variables respond contemporaneously to a shock. The variables of VAR model are ordered based on their "level of endogeneity". Cholesky ordering is used in the third article because factor-augmented VARs are used (see Bernanke, Boivin and Eliasz, 2005). The factors in the model do not have economic interpretations, making it difficult to justify the use of sign restrictions.

Study I uses also local projections with external proxy for monetary policy shock as an alternative for VARs with sign restrictions. Local projections were introduced by Jordà (2005). One of their advantages over VAR models is that they are more robust to model misspecification. The idea here is to run a sequence of predictive regressions of a variable of interest on an exogenous shock for various prediction horizons. The estimated regression coefficients represent the IRF.

Study I applies the smooth local projection method of Barnichon and Brownlees (2019) that increases precision of estimates compared ordinary least squares. Different proxies for monetary policy shocks are obtained from the euro area monetary policy event-study database (see Altavilla, Brugnolini, Gürkaynak, Motto and Ragusa, 2019). The database contains intraday asset price changes around ECB policy announcements. These changes are interpreted as unpredictable exogenous policy surprises, thereby justifying causal inference. A kindred approach to study the effect of monetary policy on equities is used by Bernanke and Kuttner (2005).

Study II relies on a difference-in-differences approach to bank-level monthly data. The idea for this simple method has a long history. One well-known example of this method's application can be found in the article of Card and Krueger (1994). In recent years, the method has gained great popularity in empirical banking research (e.g. Rodnyansky and Darmouni, 2017). It compares the average change over time in the outcome variable in two groups. The policy change of interest is assumed to affect only one group ("treatment group"). The key assumption is parallel trends, i.e. the difference between the two groups should remain constant over time in the absence of "treatment." Because banks studied in the second article could choose whether to participate in TLTRO, the policy is not assigned randomly. This problem is tackled using instrumental variables. Propensity score matching is used as a robustness check (see Rosenbaum and Rubin, 1983).

3.2 Data

The dissertation uses data from many sources. Study I uses data about equity analysts' forecasts, interest rates, Eurostoxx 50 stock market index and prices of dividend future contracts from Bloomberg. The shadow rate estimate of Kortela (2016) is produced by the Bank of Finland. Alternative shadow rate estimates (Krippner, 2015; Wu and Xia, 2016) are obtained from the websites of Leo Krippner and Jing Cynthia Wu. Equity premium data from the US is downloaded from the website of Aswath Damodaran. Monetary policy surprises are from the euro area monetary policy event-study database.

In Study II, the main data are from the ECB's individual balance sheet items (IBSI) database. It is at bank level and its frequency is monthly. It covers the period from January 2015 to July 2018. The final dataset covers 187 banks from 18 countries. The ECB data about the bank level amounts of TLTRO are obtained courtesy of the Bank of Finland's Monetary Policy Implementation Division.

Study III uses almost 100 different monthly time series from January 1999 to July 2017. The data are mainly from Eurostat and the ECB. Some series are from MSCI, the Bank of Japan, OECD and the Bureau of Labor Statistics.

4 Summary of the articles

4.1 Study I: Monetary Policy and Stock Market Valuation

Study I considers the effect of monetary policy on stock market risk premia. The assessed risk premia are *implicit*, meaning that they are implied by a dividend discount formula, observed market price and dividend forecasts. The analysis focuses on the euro area.

The time-varying long-term average risk premium is first solved utilizing Eurostoxx 50 stock index and analysts' dividend expectations. This approach is quite similar to that of Claus and Thomas (2001), Gebhardt, Lee and Swaminathan (2001) and Damodaran (2020). The discount-horizon-specific premia are then examined. The implied horizon-specific premia are solved using a novel method that is based on the article of Binsbergen, Hueskes, Koijen, and Vrugt (2013). Study I combines the idea of implicit premia and the theoretical framework of pricing dividend future contracts presented by Binsbergen et al. (2013).

The article shows that even though (risk-free) interest rates have declined considerably after the financial crisis the average long-term implied discount rate has remained quite stable. This means that the implied average long-term risk premium has risen. When it comes to horizon-specific premia, short-horizon premia tend to rise during turbulent times. The behaviour of long-horizon premia seems to be much less volatile.

After documenting the behaviour of implied risk premia, the article studies the impact of monetary policy on the premia. The effect is considered using two empirical approaches. The article first uses local projections in which monetary policy shocks are proxied by the immediate market reactions of relevant interest rates to the ECB's monetary policy decisions.

Second, the article uses VAR models, where monetary policy is measured using shadow rate and the structural shock is identified using sign restrictions. Both approaches yield similar results. Monetary policy easing raises the average risk premium. The effect is explained by the positive effects on the long-horizon premia.

4.2 Study II: The Effect of Targeted Monetary Policy on Bank Lending

The second article studies the effect of the ECB's targeted longer-term refinancing operations (TLTROs) on bank lending.[12] The policy relevance of this work is high as TLTROs have become an important addition to the toolkits of many central banks. The article uses bank-level monthly data from multiple countries about bank balance sheets. A difference-in-differences method is used.

Because banks may choose whether to participate and how much they borrow from the central bank, the "treatment" is not random. Therefore, the article uses instrumental variables to reduce this problem. First, central bank borrowing prior the assessed operations is used as an instrument. Earlier borrowing is highly correlated to the borrowing in the assessed second series of TLTROs as banks used TLTRO II to replace their earlier TLTRO funding. While it is true that this instrument is maybe not perfectly valid as earlier central bank borrowing may have a direct effect on bank lending, the effect is likely to be minor because it is difficult for banks to forecast their lending opportunities multiple years ahead. Second, the article uses the amount of eligible loans (loans to non-financial corporations and households excluding mortgages) as an instrument. The predetermined maximum borrowing was based on this amount of eligible loans. Because many banks used their entire borrowing allowance, the weakness of instrumental variable is not a problem. The amount of eligible loans prior TLTRO II was predetermined, so the exclusion restriction assumption is likely hold. Propensity score matching is used as a robustness check.

Another issue in Study II is unobservable but important variable: credit demand. The article uses two ways to control for demand. First, country-time fixed effects are used. Second, loans to households for house purchase is used as a control variable. This credit category was not part of the ECB's targeting, so it is likely that changes in loans for house purchase reflect changes in credit demand.

The cumulative effect of TLTROs on participating banks' stock of corporate loans is estimated to be significant (about 20 percent). However, the effect on lending for consumption is found close to zero. Furthermore, the positive effects on corporate loans are found to be driven by crisis countries suggesting that the effectiveness of monetary policy depends on

[12] The title of section 5 should read "conclusions" not "conslusions."

the economic conditions. The paper also finds some evidence that the effect on government bond purchases is negative.

4.3 Study III: The effect of the ECB's conventional monetary policy on the real economy: FAVAR-approach

The third article studies the effects of conventional monetary policy. It applies the factor-augmented vector autoregressive (FAVAR) models proposed by Bernanke et al. (2005) to euro area data. The model is first estimated using the full observation sample from January 1999 to July 2017. Monetary policy shock is identified using Cholesky decomposition as in Bernanke et al. (2005). In the baseline identification the ECB's policy rate is ordered last, but other orderings are considered as robustness checks. The estimated impulse response functions are in line with previous research using euro area data (e.g. Soares, 2013).

The article provides evidence that impulse response functions have changed over time. In some subsamples after the 2008 financial crisis, the impulse response functions of industrial production and unemployment rate are close to zero and statistically insignificant. While theoretical explanation for the results is not assessed in the article, there are many reasons the monetary policy transmission mechanism might have changed (see subsection 2.3).

5 Conclusions

Modern central banks pursue their policy goals using a variety of monetary policy tools. The articles of this dissertation highlight the limited knowledge we have on their effects. The results of the first article suggest that the effect of monetary policy on stock market valuation is perhaps different than what policymakers might traditionally expect. The second article indicates that targeted refinancing operations (TLTROs) of central banks seem to affect lending to firms as desired, but more research on their design and incentives is needed. One puzzling finding is that the results offer no evidence of a positive effect on lending to households. The third article challenges earlier results suggesting no time-variation in the effectiveness of conventional monetary policy. Despite extensive research on the monetary policy transmission mechanism, clearly much work remains.

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Original Publications

Monetary Policy and Stock Market Valuation*

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Abstract

This paper estimates the effect of monetary policy on the term structure of stock market risk premia. The implied stock market risk premia are obtained using analysts' dividend forecasts and dividend future prices. The effect of monetary policy on risk premia is analysed using local projections and VAR models. According to the results, monetary policy easing raises the average risk premium. The effect is driven by a rise in long-horizon risk premia.

JEL codes: E52, G12

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1. Introduction

The notion that the value of a stock is the present value of its expected future dividends goes back at least to Williams (1938). Hence, the changes in stock prices must be explained by either changes in dividend expectations or changes in discount rates. The discount rate, or (approximately) expected rate of return, can be thought of as a sum of a risk-free rate and a risk premium. Hence, monetary policy should have an effect on stock prices through the risk-free rates which it partly controls. Monetary policy may also affect dividend expectations for example through the output of firms (e.g. Galí and Gambetti, 2015).

The effect of monetary policy on the risk premia is however less clear. There are reasons to believe that expansionary monetary policy lowers the risk premia as many papers provide empirical evidence that expansionary monetary policy generates an immediate rise in equity prices followed by a period of lower-than-normal excess returns (e.g. Bernanke and Kuttner, 2005). There are also some papers that provide potential explanations for this empirical observation (e.g. Gust and López-Salido, 2014). On the other hand, there are many reasons to believe that expansionary monetary policy does not necessarily lower the risk premia. For example, Caballero and Farhi (2013, 2018) argue that asset purchases may raise stock market risk premia by decreasing the supply of negative-beta assets. In addition, expansionary monetary policy may increase risk premia by affecting inflation (e.g. Modigliani and Cohn, 1979; Schotman and Schweitzer, 2000). Assuming that stock market bubbles exist, Galí (2014) shows that monetary policy easing may decrease the size of stock market bubble (which means increasing risk premia in the framework of this paper). Monetary policy may also affect differently on different parts of the equity risk premium curve that is used to discount the future dividends. For example, Binsbergen, Hueskes, Koijen and Vrugt (2013) show that empirically the slope of the curve changes over the business cycle.

In this paper I analyse the effect of monetary policy on the term structure of equity premia in order to better understand developments in stock market valuation after the global financial crisis and during the era of unconventional monetary policy measures. As a preliminary analysis, I solve for the implied risk premium using a dividend discount model and analysts' forecasts for the future dividends of the major eurozone stocks that are included in the Eurostoxx 50 index.[1] This approach yields an approximation of the time-varying average long-term equity premium. Then I move to the main topic of this paper: the term structure of equity premia. I combine the

[1] A kindred approach has been used by e.g. Claus and Thomas (2001), Gebhardt, Lee and Swaminathan (2001) and Damodaran (2020). None of these studies, however, provides empirical analysis of the impact of monetary policy on the risk premium.

prices of Eurostoxx 50 dividend futures with the analysts' dividend forecasts to calculate horizon-specific risk premia using the framework of Binsbergen et al. (2013).[2] I then analyse the impact of monetary policy on the term structure of equity premia. To the best of my knowledge, this has not been done in the earlier literature.

The effect of monetary policy on the risk premium is interesting as it may give information about the effect on investors' perception of risk. Thus, the issue is highly relevant for the risk-taking channel of monetary policy.[3] On the other hand, the effect on the implied risk premium may give information about the effect of monetary policy on stock market bubbles or mispricing as the premium captures also the potential bubble component.[4] Therefore, the paper also contributes to the literature studying the effect of monetary policy on stock market bubbles (e.g. Galí, 2014; Galí and Gambetti, 2015). However, in this paper, I do not attempt to distinguish whether the variation in the implied premium is due to mispricing or variation in the rational risk premium.

From the perspective of firms, equity premium represents a large share of the cost of capital. Therefore, the effect of monetary policy on the implied risk premium is crucial when it comes to the transmission of monetary policy to the investments of firms. For example, the results by Liu, Mian, and Sufi (2021) suggest that the aggregate risk premium may behave differently, when interest rates are low due to heterogeneity of firms. Finance literature has also shown that high expected profitability of firms is related to high excess returns (Hou, Mo, Xue and Zhang, 2021a, 2021b). Therefore, monetary policy may affect implied risk premia also by affecting the expected profitability of firms.

I analyse the effects of monetary policy with two methods. First, I study the effects using local projections, where changes in the overnight indexed swap (OIS) rates around ECB Governing Council announcements are used as a proxy for monetary policy shocks.[5] In the second approach, which serves as a robustness check, I study the effects in a VAR model in which monetary

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[2] Dividend futures have been used recently in many other applications as well. See for example the paper by Gormsen and Koijen (2020). In this paper, I focus on the euro area, because the available time series regarding the prices of dividend futures in the USA is very short. When it comes to estimation of horizon-specific premia, alternative approach that applies credit default swap spreads is proposed by Berg (2010) and Berg and Kaserer (2013).

[3] See, for example, the paper by Borio and Zhu (2012) about the risk-taking channel of monetary policy.

[4] The variation in the implied premium means that stocks are sometimes too expensive or too cheap for an investor whose required premium over the risk-free rate is constant over time (or whose required premium varies less than the implied premium).

[5] Intraday OIS-rate changes are obtained from the EA-MPD: https://www.ecb.europa.eu/pub/pdf/annex/Dataset_EA-MPD.xlsx?afecc88fe2e29c7abcdee5670b6d0f68. See the paper by Altavilla, Brugnolini, Gürkaynak, Motto and Ragusa (2019).

policy stance is measured using the shadow (policy) rate, taking account of the zero lower bound of the policy rate.[6] The monetary policy shock is identified using sign restrictions.

The results show that the average risk premium has increased considerably since the global financial crisis, and this change is driven by the change in long-horizon premia. The results show that monetary policy easing has had a positive and persistent impact on the average and long-horizon risk premia. At the same time, it is true that monetary policy easing decreases short-horizon risk premia (in line with e.g. Bernanke and Kuttner, 2005). This implies that expansionary monetary policy steepens the slope of the term structure of risk premia.

These results indicate that "leaning against the wind" policies are ineffective when it comes to the stock market. Contractionary monetary policy increases the short-term premia but decreases long-horizon premia. The effect on average risk premium is negative. Thus, monetary policy tightening makes stocks "expensive" in relation to the expected stream of dividends and the level of risk-free rates. The results provide no evidence that expansionary monetary policy causes stock market bubbles.

The underlying theoretical mechanism explaining the results is likely to be complex. The results do not seem to support inflation-based explanations. When it comes to mechanism described by Caballero and Farhi (2013, 2018), the results do not contradict this explanation, but more careful research is needed to draw any further conclusions. Another potential explanation that should be examined in the future research is "bubbles in dividend forecasts": the equity analysts' forecasts may be systematically too optimistic or pessimistic after the monetary policy shocks. All in all, the contribution of this paper is to provide stylized facts and raise questions rather than provide definitive answers.

The remainder of the paper is as follows. Section 2 explains the theoretical framework and empirical strategy. It is divided into three subsections. The first presents the general theoretical model, the second the empirical strategy and the third briefly reviews the main results of the previous literature. Section 3 presents the data and explains the econometric methods. It is also divided into three subsections: the first explains the construction of time-varying risk premia and the latter two explain the econometric methods to analyse how monetary policy affects the risk premia. Section 4 shows the local projection and VAR results. Section 5 provides some additional results that may guide the future research. Section 6 concludes.

[6] The concept of shadow rate was first introduced by Black (1995).

2. Monetary policy and equity premia

2.1 *Theoretical framework*

The value of a stock at time 0 can be expressed as:

$$P_0 = E_0\left[\sum_{t=1}^{n} m_{0,t} D_t + m_{0,n} P_n\right], \tag{1}$$

where $m_{0,t}$ is the stochastic discount factor and D_t is the expected dividend. The specific form of the $m_{0,t}$ depends on the assumptions regarding the investors' preferences. An alternative expression for the stochastic discount factor is $m_{0,t} = \prod_{i=1}^{t} \frac{1}{(1+r_i)}$, where r_i is the required rate of return.[7] Risk premium is the difference between the required rate of return and the risk-free rate: $r_i^{excess} = r_i - r_i^{risk-free}$.

2.2 *Solving for the implied risk premia*

I use two approaches to solve for the implied risk premia. The first one gives an approximation of the long-run average premium (see e.g. Fama and French, 2015). The second one allows me to solve for the horizon specific premia.

For the first approach, I assume that expected dividends, D_t, grow at constant rate, g, after the period n, and that the discount rate, r, is the same for all the horizons. This r approximates the (time-varying) long-run average required rate of return. With these assumptions, equation (1) simplifies to:

$$P_0 = E_0\left[\frac{D_1}{1+r} + \frac{D_2}{(1+r)^2} + \cdots + \frac{D_n}{(1+r)^n} + \frac{\frac{D_{n+1}}{r-g}}{(1+r)^n}\right]. \tag{2}$$

If the market price of the share and the dividend expectations are known, r can be solved numerically period after period. This gives an approximation how the expected long-run average return varies over time. Subtracting some proxy for the risk-free rate gives an estimate for the risk premium.

The second approach relates dividend expectations to the prices of dividend future contracts. Dividend futures are contracts that allow one to buy the dividends of a specific year.[8] The cash flows are paid at the end of

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[7] This expression follows from the definition of return: $r_1 = \frac{P_1 - P_0 + D_1}{P_0}$. This can be rearranged as: $P_0 = \frac{D_1 + P_1}{1+r_1}$. Using rational expectations one can solve this forward: $P_0 = E_0\left[\frac{D_1}{1+r_1} + \frac{D_2 + P_2}{(1+r_1)(1+r_2)}\right]$. Here, for example, $\frac{1}{(1+r_1)(1+r_2)} = m_{0,2}$.

[8] Wilkens and Wimschulte (2010) and Lamponi and Latto (2017) provide a good overview of these instruments and their pricing.

the year. Assuming no-arbitrage, the price of the dividend future at time 0 that matures after h years is given by[9]:

$$F_{0,h}^{DivFut} = e^{-\sum_{i=1}^{h} r_i^{excess}} E_0[D_h]. \tag{3}$$

The discount rate is equal to the premium, r_i^{excess}, because the cash flows are paid at the maturity.[10] This means that the future contract does not tie up money. Thus, the buyer of the future receives the risk-free return in excess to the return of the future contract. The price of the dividend future, considering the dividends of the next 12 months, is $F_{0,1}^{DivFut} = e^{-r_1^{excess}} E_0[D_1]$. Accordingly, the price of the future considering the dividends 12 to 24 months ahead is $F_{0,2}^{DivFut} = e^{-(r_1^{excess}+r_2^{excess})} E_0[D_2]$ and so forth. Therefore, knowing the prices and the dividend expectations, one can solve horizon-specific expected premia beginning from r_1^{excess}.

2.3 The role of monetary policy

The previous literature has focused on analysing the effect of monetary policy on excess returns rather than on risk premia implied by dividend expectations (e.g. Patelis, 1997; Bernanke and Kuttner, 2005).[11] Implied risk premia and realized excess returns are related, but the two may sometimes differ considerably. Figure 1 shows the development of implied 5-year average stock market risk premium based on equation (3) together with the realized stock market excess return.[12] For example, during the financial crisis of 2008 investors were pricing a roughly 20 percent risk premium. Yet, the actual realized excess returns were much lower. Thus, realized excess returns after the financial crisis tell little about stock market valuation during the financial crisis. For the purpose of this paper, it would probably be more interesting to see what happens after 2015. In January 2015, the ECB's large-scale asset purchase programme (APP) was extended to include purchases of sovereign bonds. After this policy, the implied stock market premium rose, and the realized excess returns were very low.

This potentially surprising observation is consistent with the argument by Caballero and Farhi (2013, 2018) that buying sovereign debt may increase stock market risk premium. They note that "…long term public debt, by being a "bearish" asset that can be used to hedge risky private assets, has a safe asset multiplier effect that short term public debt lacks. That is, long

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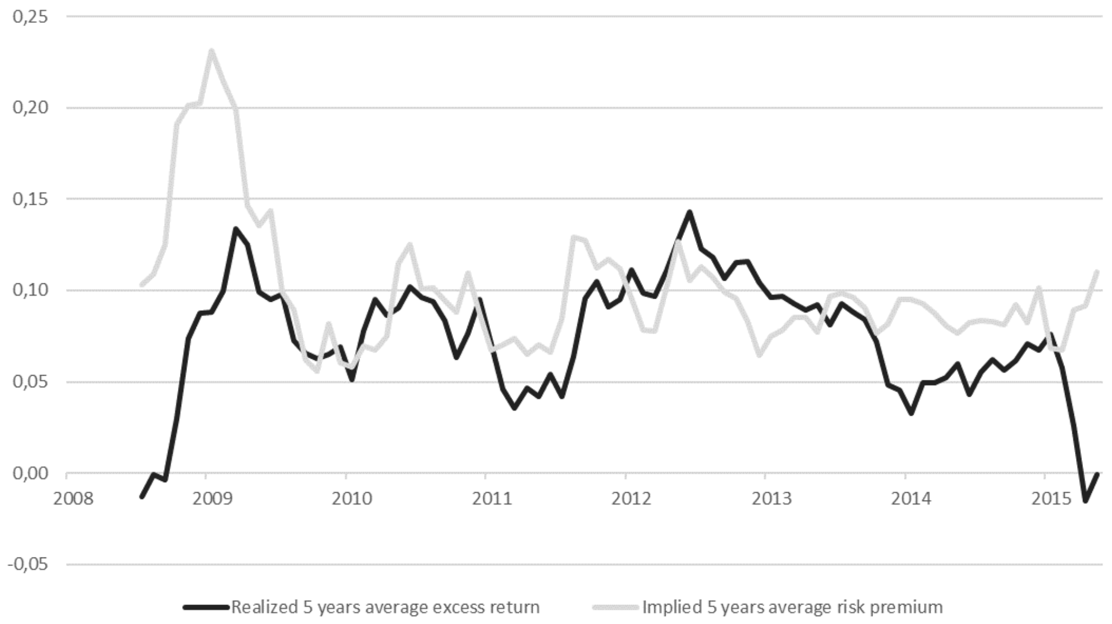
[9] See Binsbergen et al. (2013) for details.

[10] For example, the future or forward price (the price one observes) of a dividend future contract maturing after one year is: $F_{0,1}^{DivFut} = e^{-(r_1^{risk-free}+r_1^{excess})} E_0[D_1] e^{r_1^{risk-free}} = e^{-r_1^{excess}} E_0[D_1]$. See also Binsbergen et al. (2013, p. 504-505).

[11] For example, Bernanke and Kuttner (2005) consider expected future excess returns (risk premia) given by an econometric model.

[12] Details about empirical implementation are given in Section 3.

Figure 1. *Expected and realized stock market premia of Eurostoxx 50 index. The expected premia are based on analysts' dividend forecasts, Eurostoxx 50 dividend futures and application of equation (3). More details are given in Section 3. The realized return is calculated from log-differences of Eurostoxx 50 gross return index, which takes into account paid dividends. The German 1-year yield is used as the risk-free rate measure.*



term public debt is not only a safe asset in itself, but also makes risky private assets safer through portfolio effects".

Another potential mechanism through which monetary policy easing may increase stock market risk premia is inflation. Theoretically, one may argue that inflation is irrelevant for the premium, because it affects equally to expected nominal dividends and nominal risk-free rates. This is not true empirically. For example, Modigliani and Cohn (1979) and Schotman and Schweitzer (2000) have shown empirically that inflation matters for the stock market risk premium. Therefore, one mechanism could be that expansionary monetary policy raises the expected future inflation, which in turn, makes investors require higher risk premia for the future dividends.

On the other hand, if one assumes that stock market bubbles exist, it is possible that monetary policy easing decreases the size of stock market bubble (Galí 2014; Galí and Gambetti, 2015). In their papers, Galí (2014) and Galí and Gambetti (2015) use the risk-neutrality assumption and interpret all deviations from risk-free return as a bubble. In my paper, no assumptions regarding the risk attitude of investors are made. Therefore, my results can be interpreted also in the risk-neutral framework of Galí (2014): higher implied premium means smaller stock market bubble. One could argue that the monetary policy easing by the ECB decreased stock market bubble through the mechanism described by Galí (2014).

The conventional wisdoms that monetary policy easing lowers stock market risk premia and/or creates bubbles is motivated primarily by

Figure 2. *Solved expected rate of return and expected premium. The variables are solved applying equation (2) to Eurostoxx 50 stock market index and analysts' consensus forecasts in the period from 06/2006 to 04/2020.*

empirical results based of realized excess returns (e.g. Patelis, 1997; Bernanke and Kuttner, 2005). However, the theoretical mechanisms driving these results are not clear.[13] In addition, as discussed above, there are multiple reasons to presume that expansionary monetary policy raises rather than lowers stock market risk premia. In this paper, I challenge the conventional wisdoms that are based on empirical results by a set of new empirical results. Unlike in the previous literature, I do not analyse realized excess returns but implied risk premia. This approach allows me to focus on the whole term structure of risk premia that is used to discount the future dividends.

3. Data and empirical methods

3.1 Risk premia

To solve the implied risk premia, I use equations (2) and (3). First, I apply equation (2) to the Eurostoxx 50 stock market index and analysts' consensus forecasts for future dividends. Given the stock market index and aggregated dividend forecasts, I solve the expected annual rate of return, r. Further, the variation in expected return can be divided into variation in risk-free rate and

[13] See for example the discussion by Galí (2014) and Galí and Gambetti (2015).

stock market risk premium using some proxy for the risk-free rate. As a proxy for the risk-free rate, I use Germany's 10-year government bond yield following Claus and Thomas (2001), Gebhardt et al. (2001) and Damodaran (2020). The reason for using a long rate rather than a short one is that the rate should represent the average expected risk-free rate.[14] The results are about the same even though one uses some other proxy. All these data are from Bloomberg and cover the period from 06/2006 to 04/2020. The reason for using this sample is the availability of data about analysts' dividend forecasts.

I obtain monthly data about year-specific dividend forecasts from Bloomberg. The forecasts are at the index level. The index level forecasts are weighted averages across individual sell-side analysts' firm specific forecasts. The weights are the same as those that are used to form the stock index. One company is followed by approximately 30 analysts. The large number of analysts per company is one benefit from using Eurostoxx 50 rather than some broader index. I calculate the 12- and 24-month-forward dividend expectations (D_1 and D_2) as a weighted average of year-specific forecasts.[15] Additionally, analysts' forecast average earnings per share growth rate after 3–5 years. I assume that dividend payout ratio remains constant meaning that earnings per share growth rate can be used to calculate dividends: D_3, D_4, D_5 and D_{5+1}.[16] I assume that the expected long-run dividend growth rate, g, equals the historical average of the analysts' expected earnings per share growth rate in the sample (4.5 percent).[17]

With these assumptions, the only unknown variable in equation (2) is r. This variable can be solved for every month in the sample. Subtracting Germany's 10-year government bond yield, I get an estimate for the expected average annual stock market premium in each month. The developments of the expected rate of return and expected premium are presented in Figure 2.

[14] Alternatively, one could use different interest rates for different horizons. However, Claus and Thomas (2001) conclude that the results are almost identical. Therefore, for the sake of transparency, I use one interest rate.

[15] For example, I calculate the 12-months-ahead dividend forecast in April 2020 as follows: $(8 * Dividend\ forecast\ for\ the\ year\ 2020 + 4 * Dividend\ forecast\ for\ the\ year\ 2021)/12.$

[16] As the data about the average future earnings per share growth rate are rather noisy, I use a 12-month moving average to smooth the series. In addition, I remove clearly unrealistic extreme values (values below -30 percent and greater than 30 percent).

[17] This assumption affects mainly the level of risk premium, but not so much the variation of premium. It is also possible that there has been variation in the expected long-run growth. This issue is assessed in Appendix B. The results show that the long-run growth rate expectation should have declined more than 4 percentage points after the financial crisis to affect the conclusions of this paper.

***Figure 3**. Year-specific expected premia implied by Eurostoxx 50 dividend futures and analysts' dividend forecast from 7/2008 to 4/2020.*

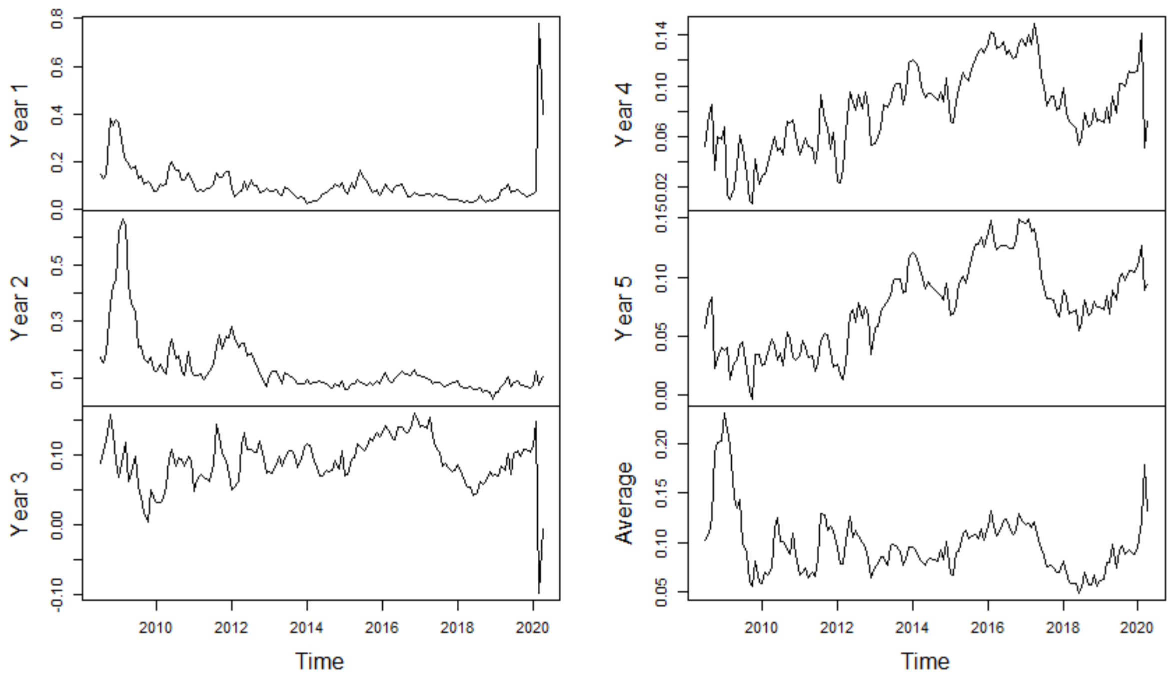


***Table 1.** Summary statistics of the year-specific expected premia implied by Eurostoxx 50 dividend futures and analysts' dividend forecast. The observation period runs from 7/2008 to 4/2020.*

	Year 1	Year 2	Year 3	Year 4	Year 5
Mean	0.10	0.14	0.09	0.08	0.08
Median	0.08	0.10	0.09	0.08	0.08
SD	0.09	0.11	0.04	0.03	0.04

In the sample the average expected premium, r^{excess}, is 7.3 percent and the average expected return, r, is 9.0 percent. These results are in line with the previous estimates based on a similar methodology.[18] Notably, the expected rate of return for equities has remained rather stable since the 2008 financial crisis, even as risk-free rates have declined substantially. This

[18] Claus and Thomas (2001) and Gebhardt et al. (2001) find that the average expected premium has been around 3 percent. However, recent estimates by Damodaran (2020) for the S&P 500 are quite similar to my results. The estimates for the S&P 500 suggest that the expected premium can deviate from its long-run average for many years. The expected premium in the United States was quite high during the 1970s and quite low during the 1990s (consistently with the results of Claus and Thomas (2001) and Gebhardt et al. (2001)). After the financial crisis, the expected premium has been very high in the United States and has exhibited an upward trend like in Europe.

means that the expected premium over the risk-free rates has increased considerably since the crisis.[19]

As was discussed earlier, expected returns differ across discounting horizons. To solve the horizon-specific risk premia, I apply equation (3) to the prices of Eurostoxx 50 dividend futures and analysts' dividend forecasts. A bit similar approach is used by Binsbergen et al. (2013).[20] I calculate the prices of 1, 2, 3, 4 and 5 years ahead dividend futures by taking the weighted average of the observed future prices.[21] Then, I calculate the implied expected premia for every horizon beginning from the 12-month discounting horizon. The data about the dividend futures are from 7/2008 to 4/2020 and have been collected from Bloomberg. The solved expected premia are shown in Figure 3. Key summary statistics appear in Table 1.

In agreement with Binsbergen, Brandt and Koijen (2012), Binsbergen et al. (2013) and Binsbergen and Koijen (2017), the variation in the expected premia is related mostly to the near future. The standard deviation of the year 1 premium is 0.09 and the standard deviation of the year 5 premium is 0.04. Year 1 premia are on average also higher than year 5 premia. The slope of the premium curve seems to be pro-cyclical as well. The premium for the year 2 has been the highest and the most volatile in this sample. This result is largely driven by the financial crisis, during which the year 2 premium reacted most strongly. This situation has been quite different during the covid-19 crisis. In March 2020, the year 1 premium climbed to about 80 percent, even as the year 2 premium remained at 8 percent.

One can also see that the premia of the years 4 and 5 exhibit similar upward trend as the premium implied by equation (2). Thus, these results are in line with the earlier conclusion that the average risk premium has increased after the financial crisis. Further, the results suggest that the rise in average premium is mainly driven by long-horizon expected premia.

3.2 Local projections

I use intraday interest rate changes around the time of ECB Governing Council announcements as a proxy for monetary policy shocks. These rate

[19] One potential explanation for this is that the long-run expected dividend growth rate, g, which is assumed constant, has declined. Appendix B studies this potential explanation. The results show that it is unlikely that the decline in very long-run expectations would explain the result. If one argues that the equity premium has not risen, it would mean that the long-run expected growth rate should have declined more than 4 percentage points.

[20] The key difference between this paper and the paper by Binsbergen et al. (2013) is that Binsbergen et al. (2013) use regression-based dividend forecasts instead of analysts' forecasts.

[21] For example, I calculate the price of the 1-year-ahead dividend future in April 2020 as follows: $(8 * Price\ of\ the\ future\ maturing\ in\ 2020 + 4 * Price\ of\ the\ future\ maturing\ in\ 2021)/12$.

changes are obtained from EA-MPD (see Altavilla et al., 2019).[22] I use change in the median quote from the 13:25-13:35 window before the press release to the median quote in the 15:40-15:50 window after the press conference. Assuming these changes represent exogenous variation in the ECB's monetary policy, one can assess the effect of monetary policy on different premia by estimating the following model for different horizons:

$$r_{t+h}^{excess} = \alpha_h + \gamma_h(L) r_{t-1}^{excess} + \beta_h Surprise_t + \sum_j^p \delta_{h,j}(L) Control_{j,t-1} + \epsilon_{t+h} \quad (4)$$

In the model, α_h is a constant, $\gamma_h(L)$ and $\delta_{h,j}(L)$ are lag-polynomials, ϵ_{t+h} is an error term and $Surprise_t$ is the rate change. The model includes p control variables. I use changes in $\log(Industrial\ production)$ and $\log(Harmonised\ Index\ of\ Consumer\ Prices)$ as controls. The data about industrial production and consumer prices are from Eurostat. EA-MPD covers surprise changes in many different maturities of OIS-rates and government bond yields. In the baseline analysis, I consider the unweighted average of 1-week, 1-month, 3-month, 6-month, 1-year, 2-year and 3-year OIS rate surprises. This average surprise is related to conventional and forward guidance policies. I use the average rather than some specific maturity, because the aim here is not to disentangle different policy tools but to assess monetary policy "in general".[23] The average is also less noisy than individual rate surprises. The rate surprises in longer maturities are considered in Section 5.3.

As an estimate for risk premium, I consider the premium implied by equation (2) and the horizon-specific premia implied by dividend futures and equation (3). The data cover the period from 06/2006 to 04/2020 except for horizon-specific premia that are not available before 7/2008.

The model is estimated using the smooth local projection method by Barnichon and Brownlees (2019).[24]

3.3 VAR model

My second approach that is used as a robustness check relies on VAR models. A VAR model can be written in reduced form as:

$$y_t = c + A(L) y_{t-1} + \mathrm{u}_t, \quad (5)$$

where y_t is $n\ x\ 1$ vector of endogenous variables, c is $n\ x\ 1$ vector of constants, $A(L)$ is $n\ x\ n$ matrix of lag polynomials and $\mathrm{u}_t = B\varepsilon_t$ is $n\ x\ 1$

[22] I assume that the value of the shock was zero if there was no announcement during the month. If there were two announcements, I calculate the average of the two.

[23] 3-month surprise is used in Appendix C.

[24] To estimate the model and to construct the confidence bands, I use the MATLAB code provided by Barnichon and Brownlees (2019) along their publication.

vector of error terms. B is $n\ x\ n$ matrix that relates error terms to structural shocks, ε_t, which are of interest.

In the baseline model, the stance of monetary policy is measured using the shadow rate proposed by Kortela (2016). The shadow rate permits analysis of the overall effect of monetary policy during a time when policy rates have been close to their effective lower bound. The use of shadow rates has been standard approach in recent macroeconomic literature (e.g. Wu and Xia, 2016). In the baseline model, y_t includes year-over-year changes in $\log(Industrial\ production)$ and $\log(Harmonised\ Index\ of\ Consumer\ Prices)$, the level of equity premium and the shadow rate. Industrial production and inflation are included as the monetary policy decisions are affected by the changes in real activity and price stability. The baseline model includes two lags. Many other model specifications (e.g. models with some other shadow rate estimates, a model in levels and with a different number of lags) are considered in Appendix A and in Online Appendix.

I identify monetary policy shock using sign restrictions and standard rejection method of Rubio-Ramirez, Waggoner and Zha (2010).[25] I assume that a positive monetary policy shock decreases the growth in industrial production and consumer prices.[26] These sign restrictions are assumed to hold the first nine months. The effect on shadow rate itself is also restricted positive for the first nine months. These restrictions are consistent with for example the results by Bernanke, Boivin, and Eliasz (2005) and Wu and Xia (2016). Alternative restrictions are considered in Appendix A and in Online Appendix.

4. Results

4.1 Local projection results

Figure 4 shows the estimated responses of different premia to the tightening of monetary policy. The responses of short-term and long-term premia have opposite signs. Contractionary monetary policy raises the year 1 expected premium, while the response of long-run average expected premium is negative. The negative effect on the average premium implies that a rate hike seems to make stocks expensive in comparison to the expected stream of dividends.

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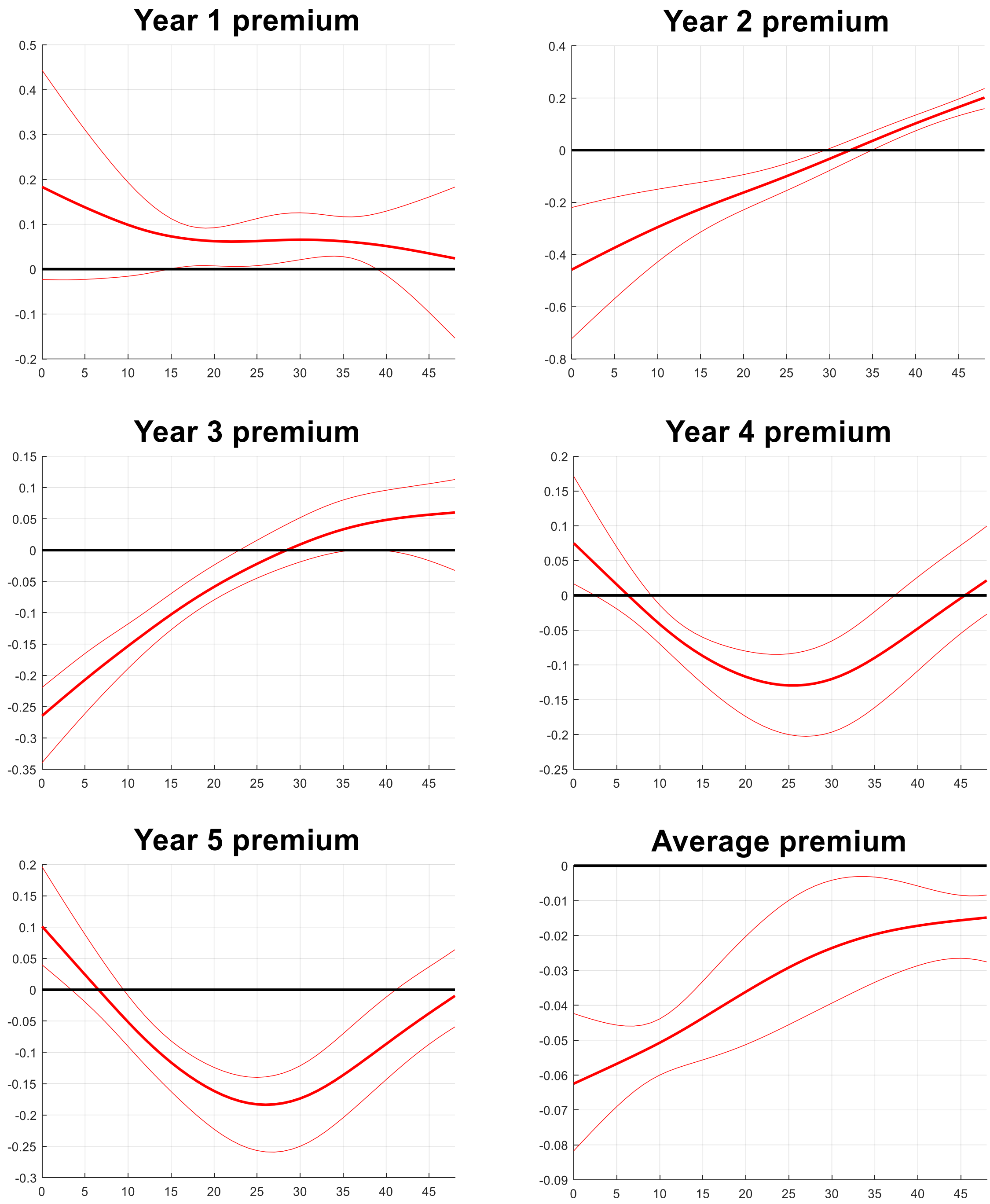
[25] The basic idea in sign restrictions is to randomly generate impulse response functions and store the impulse responses that are in line with sign restrictions. To implement the method, I use the MATLAB toolbox by Ferroni and Canova (2020). The documentation of the toolbox explains the practical implementation of the method in more detail.

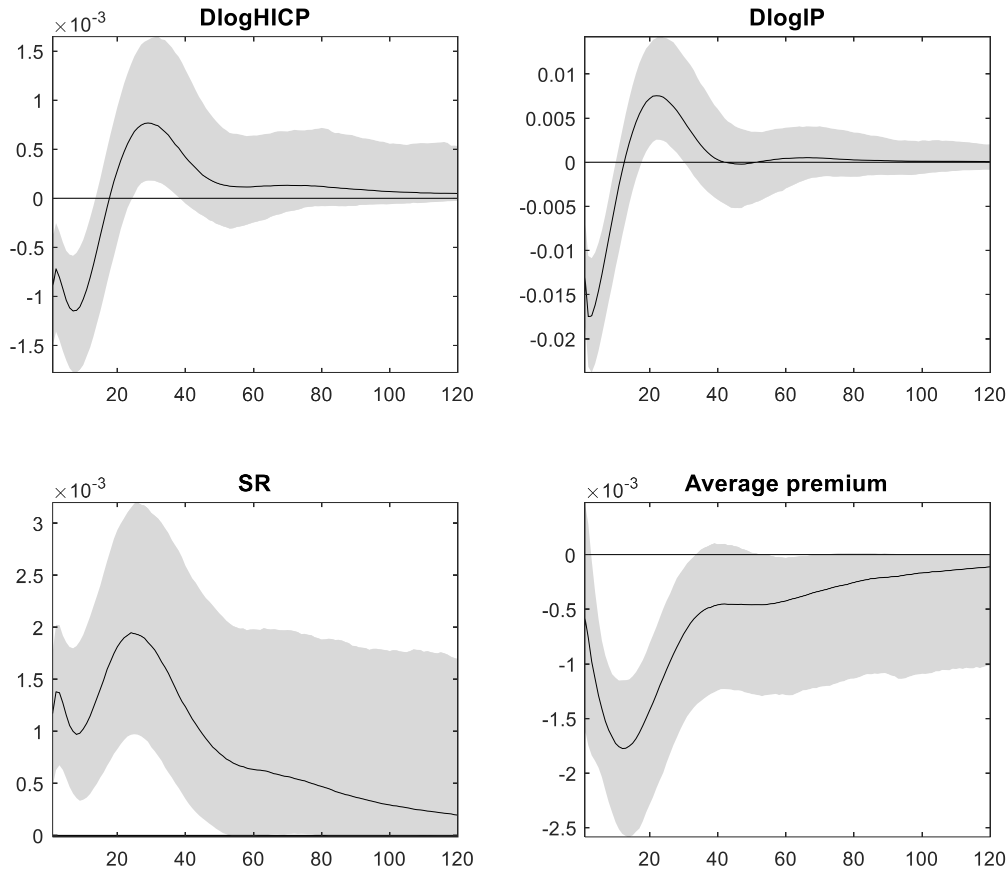
[26] The model is symmetric, and therefore the opposite applies to a negative shock.

Figure 4*. The local-projection-based impulse response function of the year 1, 2, 3, 4 and 5 premia implied by equation (3) and the long-run average expected premium implied by equation (2). Inflation and the growth rate of industrial production are used as control variables. The number of lags in the model is 3. The confidence band is 68 %. The impulse response functions and their Newey-West based confidence bands are estimated as in Barnichon and Brownlees (2019).*



The response of the year 1 premium is consistent with the evidence based on realized returns presented in Bernanke and Kuttner (2005). The signs of the responses of premia at horizons from 2 to 5 years vary. Notably, the effect on the long-term average premium suggests that expansionary monetary policy makes the long-run dividend stream riskier (or reduces the

Figure 5. *The impulse response functions to a one standard deviation positive shock to the shadow rate in the model in which the premium is the long-run average expected premium implied by equation (2). Sign restrictions are imposed on DlogIP, DlogHICP and SR for the periods 1-9. The shaded area represents the 68 % confidence interval. The number of lags is 2.*



size of the stock market bubble). This potentially surprising result supports the idea that mechanisms described in Section 2.3 may play a role.

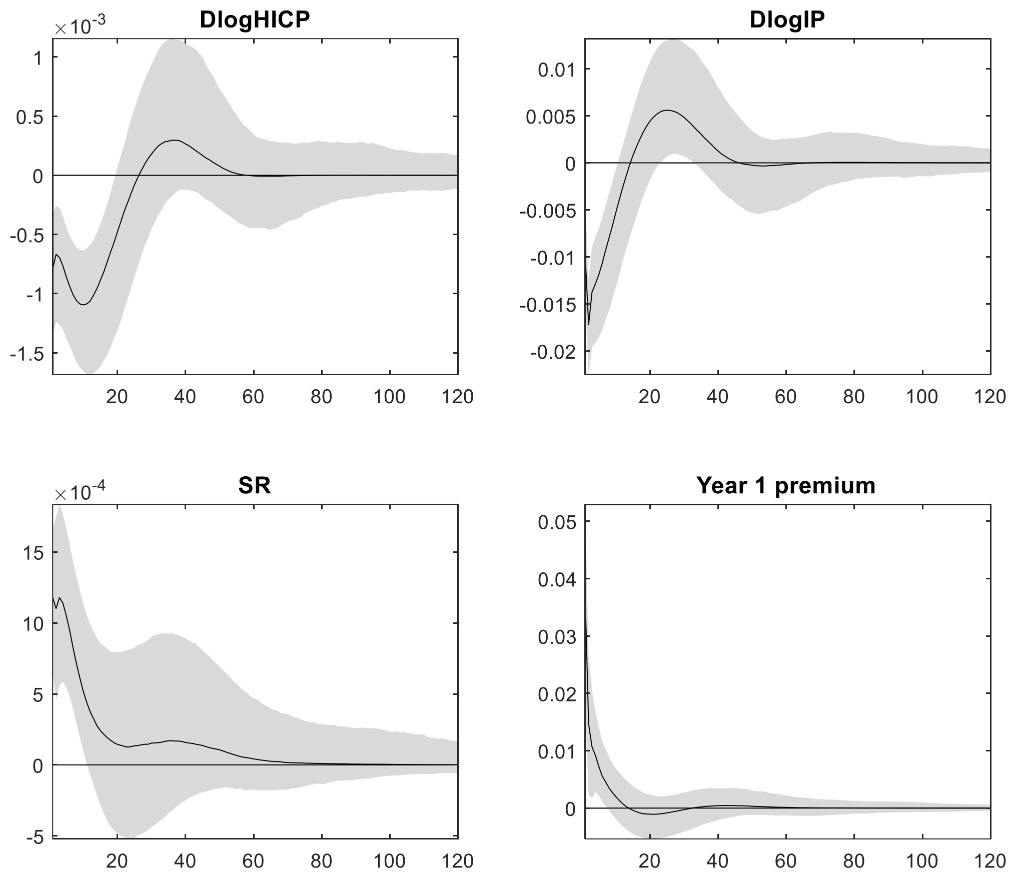
4.2 Robustness analysis using VAR models

Figures 5, 6 and 7 show the impulse responses to a positive shadow rate shock in three models with different implied premia.[27] In Figure 5, the premium is the one implied by equation (2). In Figure 6, the premium is the premium for the year 1 implied by dividend futures and equation (3). In Figure 7, the premium is for the year 5.

The responses of industrial production and consumer prices comport with the literature (e.g. Bernanke et al., 2005). The responses of short-term and long-term premia have opposite signs. Contractionary monetary policy raises the short-term expected premium temporarily (Figure 6), while the

[27] Multiple robustness checks are provided in Appendix A and in Online Appendix.

***Figure 6**. The impulse response functions to a one standard deviation positive shock to the shadow rate in the model in which the premium is the year 1 premium implied by equation (3) and dividend futures. Sign restrictions are imposed on DlogIP, DlogHICP and SR for the periods 1-9. The shaded area represents the 68 % confidence interval. The number of lags is 2.*



responses of long-run average expected premium and year 5 premia are negative and persistent (Figure 5, Figure 7).

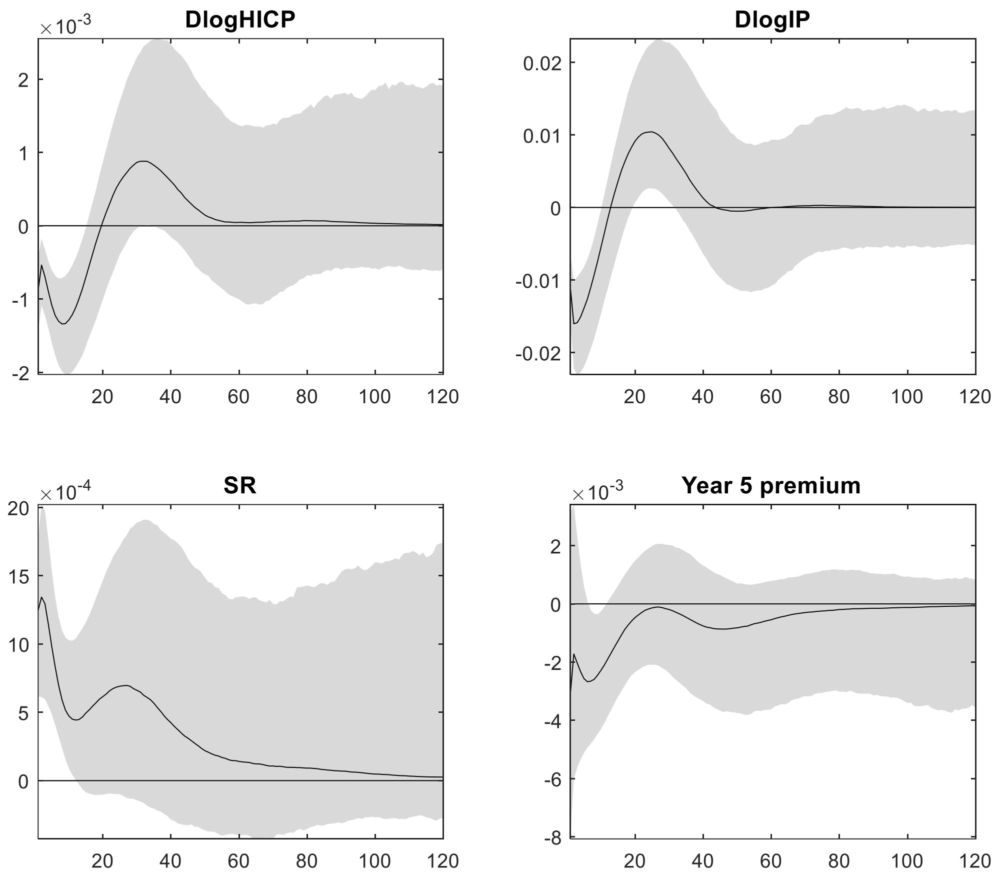
Whatever the theoretical explanation is, the empirical results are robust (see Appendix A and Online Appendix). The results do not depend on the model specification or on the used sign restrictions. The chosen shadow rate estimate is not crucial to the results. The results remain the same even though one uses the shadow rate by Krippner (2015) or Wu and Xia (2016). The effects seem not to vary over time.

5. Some additional analyses, remarks and suggestions for the future research

5.1 *Long-run US data*

One may suspect that the results shown in the previous section depend on some specific choices made in the calculation of implied premia. The

Figure 7. *The impulse response functions to a one standard deviation positive shock to the shadow rate in the model in which the premium is the year 5 premium implied by equation (3) and dividend futures. Sign restrictions are imposed to DlogIP, DlogHICP and SR for the periods 1-9. The shaded area represents the 68 % confidence interval. The number of lags is 2.*

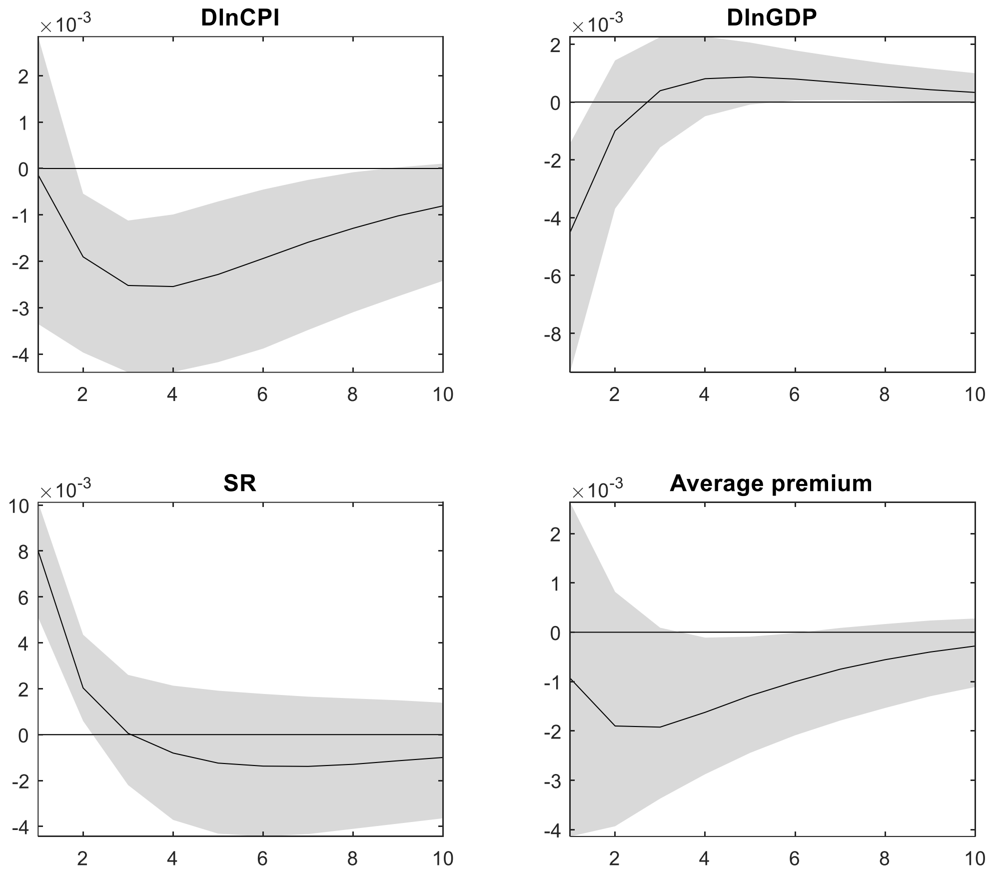


analysed time period is also limited, and the analysis has focused on the euro area. To show that these possible doubts are not justified, I run the following additional exercise.

Damodaran provides a dataset that includes equity premium estimate for the United States from 1961 to 2019 at the annual level.[28] The method used for the calculation of implied premium is described by Damodaran (2020). The method is very similar to the approach, where I apply equation (2). The premium can be interpreted as an expected long-run average premium. I include this variable into a VAR model together with the measure of monetary policy and the changes in logarithms of CPI and GDP. As a measure of monetary policy, I use the shadow rate constructed by Krippner (2015). In practice, this variable is equal to the effective fed funds rate most of the time. The model includes 1 lag and a constant. I identify the monetary policy shock using sign restrictions. I assume that a positive monetary policy

[28] http://pages.stern.nyu.edu/~adamodar/.

Figure 8. *The impulse response functions to a one standard deviation positive shock to the shadow rate in the model in which the premium is the long-run average expected premium by Damodaran (2020). Sign restrictions are imposed on DlogIP, DlogHICP and SR for the first period. The shaded area represents the 68 % confidence interval. The number of lags is 1. The data are from 1961 to 2019 and from the United States.*



shock has a negative effect on CPI and GDP one period after the shock. The sign restriction to the shadow rate is imposed only to the first period.

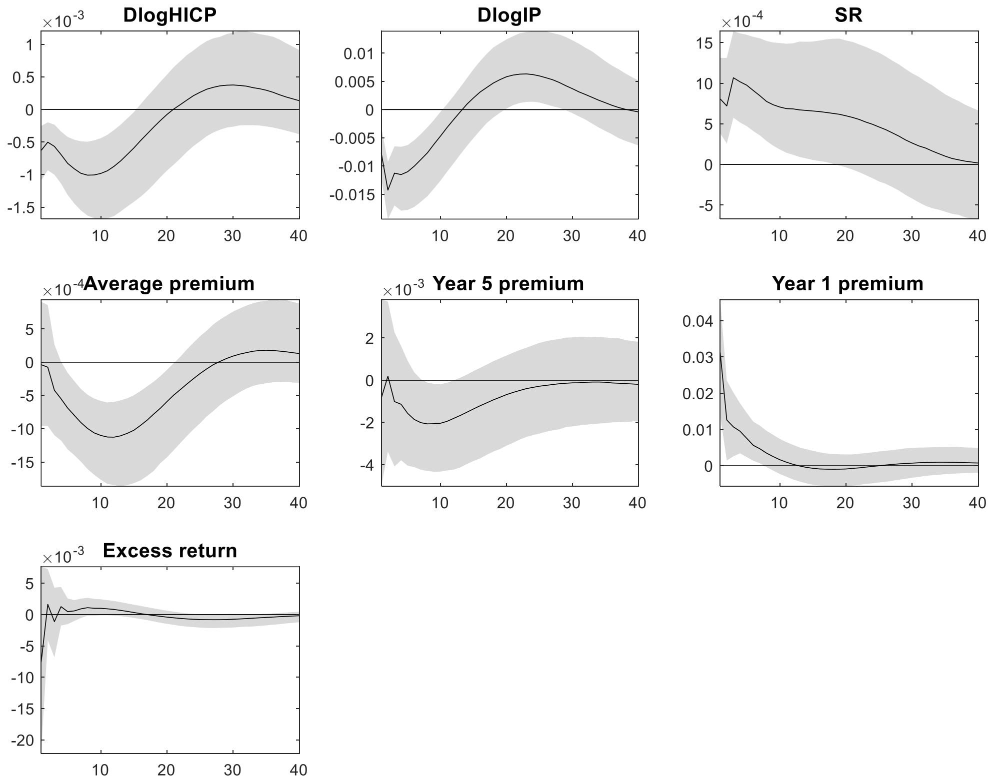
The results are shown in Figure 8. The impulse response function of the premium is very similar to the impulse response function in Figure 5. The response is negative, and it lasts about a decade.

5.2 *Results are in line with Bernanke and Kuttner (2005)*

The results may seem strange given the evidence by Bernanke and Kuttner (2005). After all, the results are in line with the previous literature. This subsection aims at clarifying this issue.

The literature has not assessed the effect of monetary policy on discounting of dividends at long horizons. Instead, it has analysed the effects on (rather short run) excess returns of stocks over the risk-free returns. In the paper by Bernanke and Kuttner (2005, p. 1250, Figure 6) the effect on excess return is assessed about two years meaning that the results by Bernanke and Kuttner (2005) should be compared to the impulse response functions of

Figure 9*. The impulse response functions to a one standard deviation positive shock to the shadow rate in the model, where, in addition to three different premia, realized monthly excess return is included. Sign restrictions are imposed on DlogIP, DlogHICP and SR for the periods 1-9. The shaded area represents the 68 % confidence interval. The number of lags is 2.*



short-horizon implied premia rather than long-horizon premia. The immediate response of the year 1 premium is sharp and positive meaning that the market price of dividends declines relative to the dividend expectations. The market price of dividends remains low relative to the dividend expectations about 12 months. In Bernanke and Kuttner (2005), the immediate excess return of stocks after a positive monetary policy shock is strongly negative (as in this paper, market price declines). Then the response remains positive about two years. This is consistent with the response of year 1 premium in this paper: the ex-ante one-year-ahead excess return remains positive about 12 months.

The relationship between the results of this paper and the results of Bernanke and Kuttner (2005) are illustrated in Figure 9. The figure shows the impulse responses to monetary policy shock in a model that is otherwise as the baseline VAR model, but includes long-horizon average, year 5 and year 1 premia at the same time, and the realized excess return. The excess return is calculated by subtracting monthly risk-free return from the monthly return of stocks.[29] The identification of the model is as in the baseline VAR model. The impulse response functions of the implied premia are as in the earlier analysis of this paper. The effect on excess return is as in Bernanke and Kuttner (2005). It is also worth noticing that the sign of the response changes after about two years, which is in line with the results regarding the long-horizon implied premia.

5.3 *Some tentative analysis and discussion about the mechanism*

So far, the paper has not said much about the underlying mechanism that could explain the results. This subsection considers some potential explanations briefly. At the same time, it hopefully raises some new ideas for the future research.

As mentioned earlier, the papers by Caballero and Farhi (2013, 2018) suggest that asset purchases of safe long-term government bonds may raise stock market risk premia. It matters whether the central bank buys risky or less risky assets. This mechanism receives some support as the effects of surprise in the spread between Italian and German government bond yields and the surprise in long-term risk-free rate seem to differ (see Appendix E).

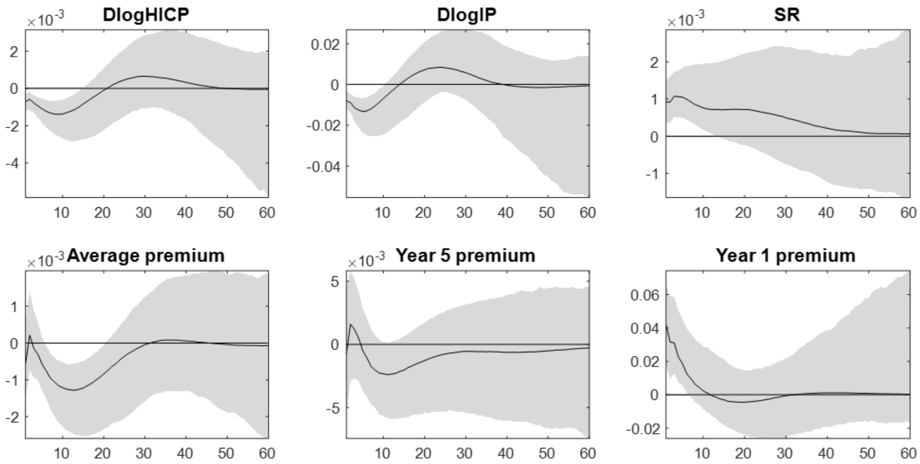
Another potential channel through which contractionary monetary policy could lower the long-horizon risk premia is expected inflation or inflation risk premium (Modigliani and Cohn, 1979; Schotman and Schweitzer, 2000). If this explanation was true, then it would seem natural that survey-based inflation expectations and the yields of inflation swaps would decline after a contractionary monetary policy shock. Appendix F augments the model with inflation expectations. The results do not support inflation-based explanations.

Third explanation considered here is the rationality of analysts' dividend forecasts. Easton and Sommers (2007) provide some evidence about the upward bias in the analysts' forecasts and show how this bias affects to their implied risk premium. Another article that provides evidence about the irrationality of survey-based expectations is by Greenwood and Shleifer (2014). Appendix G shows that these mechanisms may exist, but they seem to play economically insignificant role.

Another issue related to the analysts' forecasts is that equity analysts do not necessarily update their forecasts every month and the forecasts used in this paper may not reflect the analysts' actual expectations. To consider this issue, in the model of Figure 10, it is assumed that the analysts' expectations represent the expectations of the previous month. The lag of one month is

[29] The monthly risk-free return is obtained by using the Nelson-Siegel model to OIS rates. The maturities of the OIS rates vary between 3 months and 30 years.

***Figure 10**. The impulse response functions to a one standard deviation positive shock to the shadow rate in the model that includes three different premia. The analysts' dividend expectations are assumed to be updated with a lag of one month. Sign restrictions are imposed on DlogIP, DlogHICP and SR for the periods 1-9. The shaded area represents the 68 % confidence interval. The number of lags is 2.*



chosen because most dividend forecasts seem to be updated at least every second month. This does not seem to affect the results.[30]

The issues raised in this subsection are just examples about potential mechanisms that could explain the interesting responses of implied premia. There may, of course, be many other explanations for the results. One mechanism that was mentioned earlier, but has not been assessed in this subsection, is the one by Galí (2014). Another potential explanation, not addressed here, could be related to the results by Liu et al. (2021). In their paper they observe that low interest rates may have different implications for different firms (market leaders and followers) due to strategic behaviour. Eurostoxx 50 index includes large companies only, which may affect the results. Hou et al. (2021a, 2021b) in turn show that high the expected profitability of firms is related to high expected returns. Monetary policy, as it affects the profitability of firms, may have pricing implications through this channel. It is likely that the effect of monetary policy on the term structure of risk premia is not the same for all the companies given the evidence by, for example, Maio (2014). Therefore, it could be an interesting idea to apply the methodology of this article to individual stocks.

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[30] Appendix D shows that the results remain the same even though one uses the permanent component of Beveridge-Nelson (1981) decomposition as a measure of true dividend expectations. This method is proposed by Jokivuolle (1995) to measure the value of true stock index that is not directly observable due to infrequent trading
of stocks. In this paper, the "infrequent trades" are infrequent dividend forecast updates. One idea for the future research would be to use "flash estimates" of dividend forecast. Because analysts do not update their forecasts at the same time, one could use only the forecasts that have been updated during the past month. IBES provides these kind of flash estimates.

6. Conclusions

Interest rates have declined considerably since the global financial crisis, yet the expected average stock market return has remained quite stable at around 9 percent. This implies that expected average stock market premium has increased remarkably. This rise is mainly driven by the premia over a discounting horizon of four years.

These results may seem unintuitive as the prices of stocks have risen and ratios like price-to-earnings have been historically high. However, high price-to-earnings ratios or low dividend yields do not necessarily mean that stocks are expensive, because the value of a stock is the present value of its *future dividends*.

When it comes to the role of monetary policy, the results show that monetary policy easing decreases short-horizon required premia but increases longer-horizon premia. The effect on expected average premium is positive, i.e. expansionary monetary policy makes the prices of stocks cheap in comparison to the expected dividend stream and risk-free rates.

These results show that the conventional wisdom that monetary policy easing makes stock market valuation high is likely to be false. They underline the fact that the effect of monetary policy on stock market valuation is a puzzle that has not been solved yet.

One potential reason for the results is the channel proposed by Caballero and Farhi (2013, 2018) that was discussed in Section 2.3. However, the results seem to hold for conventional and unconventional policies, while the mechanism by Caballero and Farhi (2013, 2018) is related to a certain type of unconventional policies. These facts suggest that there are also some other mechanisms that explain the results. These mechanisms may be related to rational asset price bubbles as discussed by Galí (2014) and Galí and Gambetti (2015). Another explanation could be the fact that inflation seems to affect stock market risk premia even though the results of this paper do not support this idea (e.g. Modigliani and Cohn, 1979; Schotman and Schweitzer, 2000). The results may also be related to the empirical observation by Binsbergen et al. (2013) that the slope of equity risk premium curve varies over the business cycle. Analysts' potentially biased dividend forecasts may also be one part of the explanation. These open questions are left for the future research.

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Appendix A

This section shows a set of robustness checks for the baseline VAR model. The models in this section mainly consider long-run average premium. Similar robustness tests regarding the short-horizon premia are shown in Online Appendix.

Figure A1 shows the results after increasing the number of lags to 4. Otherwise, the model is as in the baseline VAR analysis. The results remain almost the same.

Figure A2 shows the results, when HICP and industrial production are in log-levels. The signs restrictions are still imposed for periods 1 to 9. The results are robust to using levels instead of differences.

There are many other shadow rate estimates than the one proposed by Kortela (2016). In Figures A3 and A4, the shadow rate by Krippner (2015) and Wu and Xia (2016) are used. The results remain about the same.

Figure A5 shows that the results do not change even though one imposes the sign restriction only for the first 3 periods. The results also remain the same even though one uses recursive identification instead of sign restrictions (Figures A6 and A7).

One could also ask, what happens if I include multiple premia in the model. The results remain the same as shown in Figure A8.

There are multiple empirical papers documenting that stock prices react positively to monetary policy easing. Assuming that this really is the case, Figure A9 reports the results when the change in log $(Eurostoxx\ 50)$ is included to the model and assumed to respond negatively the first 2 periods after the shock. The result regarding the effect on long-run risk premium is unaffected.

Finally, there is some evidence that the effects of monetary policy may vary over time (e.g. Laine, 2020). To study this issue, I conduct a simple rolling window analysis, in which the model includes only the shadow rate and the premium implied by equation (2) in this order. Monetary policy shock is identified recursively. The length of the window is 120 months, and the window is moved 6 months at the time. The first estimation window ends in July 2016. The results are shown in Figure A10. The effect on long-run average expected premium seems not to vary over time.

***Figure A1**. The impulse response functions to a one standard deviation positive shock to the shadow rate in the model in which the premium is the long-run average expected premium implied by equation (2). Sign restrictions are imposed on DlogIP, DlogHICP and SR for the periods 1-9. The shaded area represents the 68 % confidence interval. The number of lags is 4.*

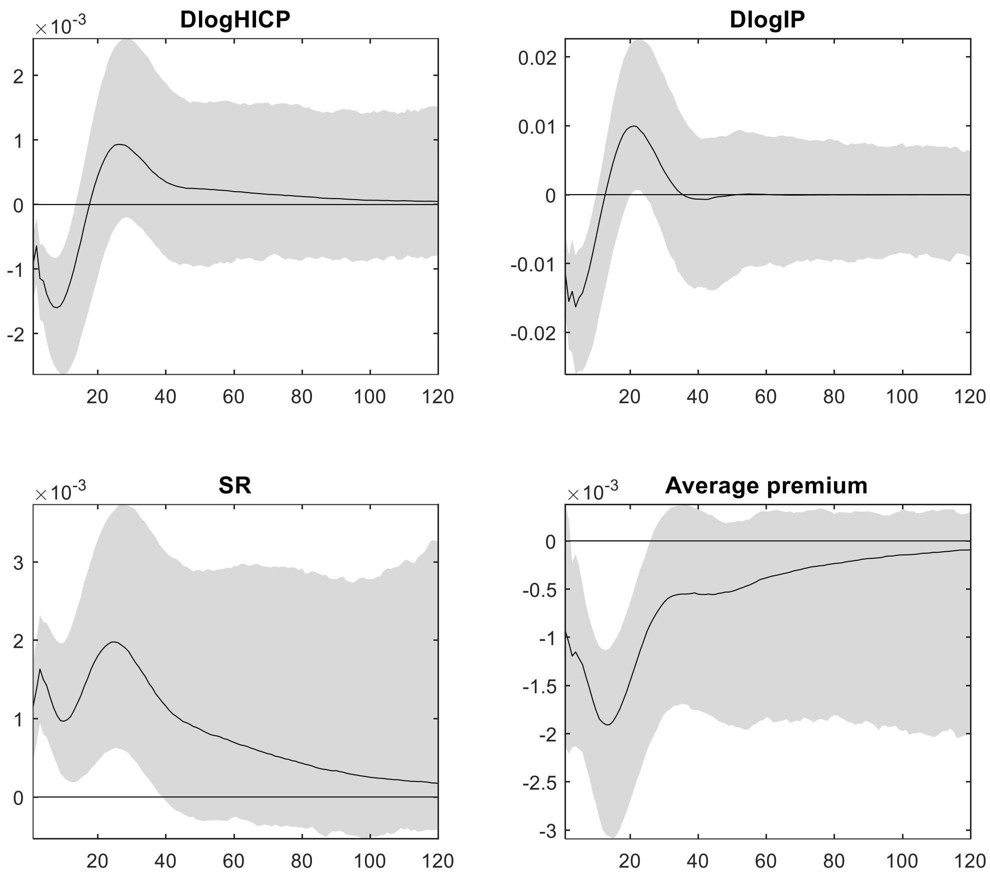


Figure A2. *The impulse response functions to a one standard deviation positive shock to the shadow rate in the model in which the premium is the long-run average expected premium implied by equation (2). Sign restrictions are imposed on logIP, logHICP and SR for the periods 1-9. The shaded area represents the 68 % confidence interval. The number of lags is 12.*

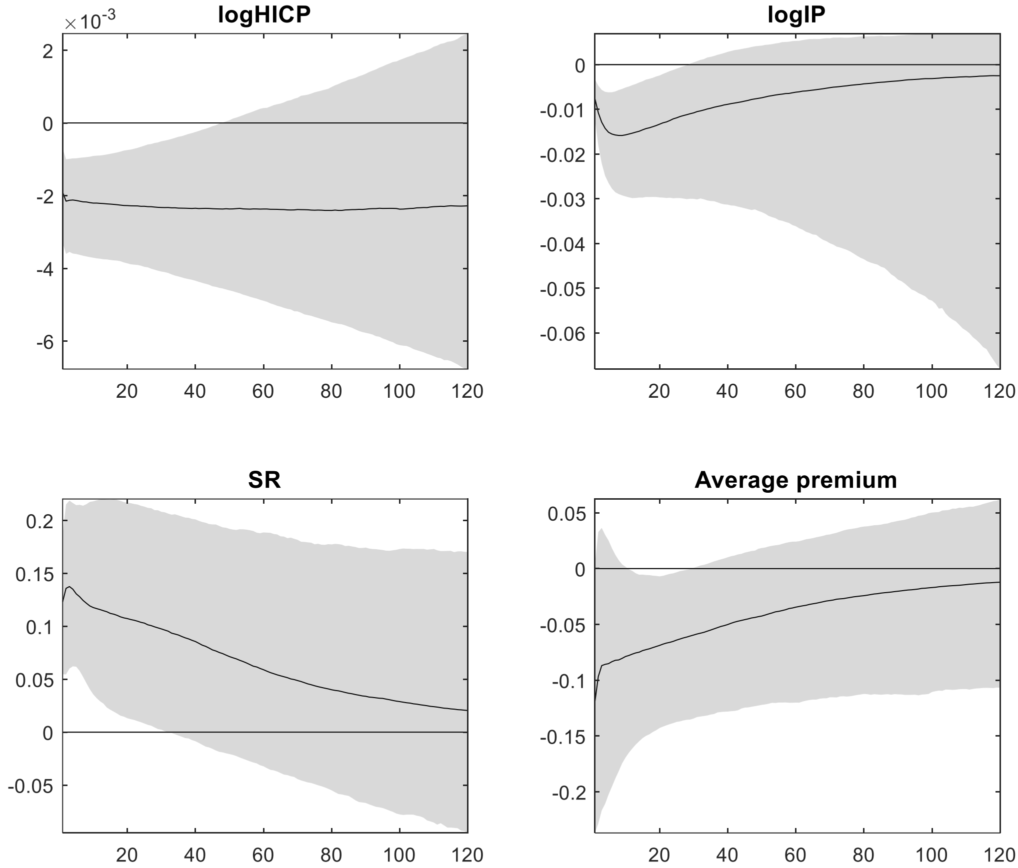


***Figure A3**. The impulse response functions to a one standard deviation positive shock to the shadow rate in the model in which the premium is the long-run average expected premium implied by equation (2). The shadow rate by Kortela (2016) is replaced by the shadow rate by Krippner (2015). Sign restrictions are imposed on DlogIP, DlogHICP and SR for the periods 1-9. The shaded area represents the 68 % confidence interval. The number of lags is 2.*

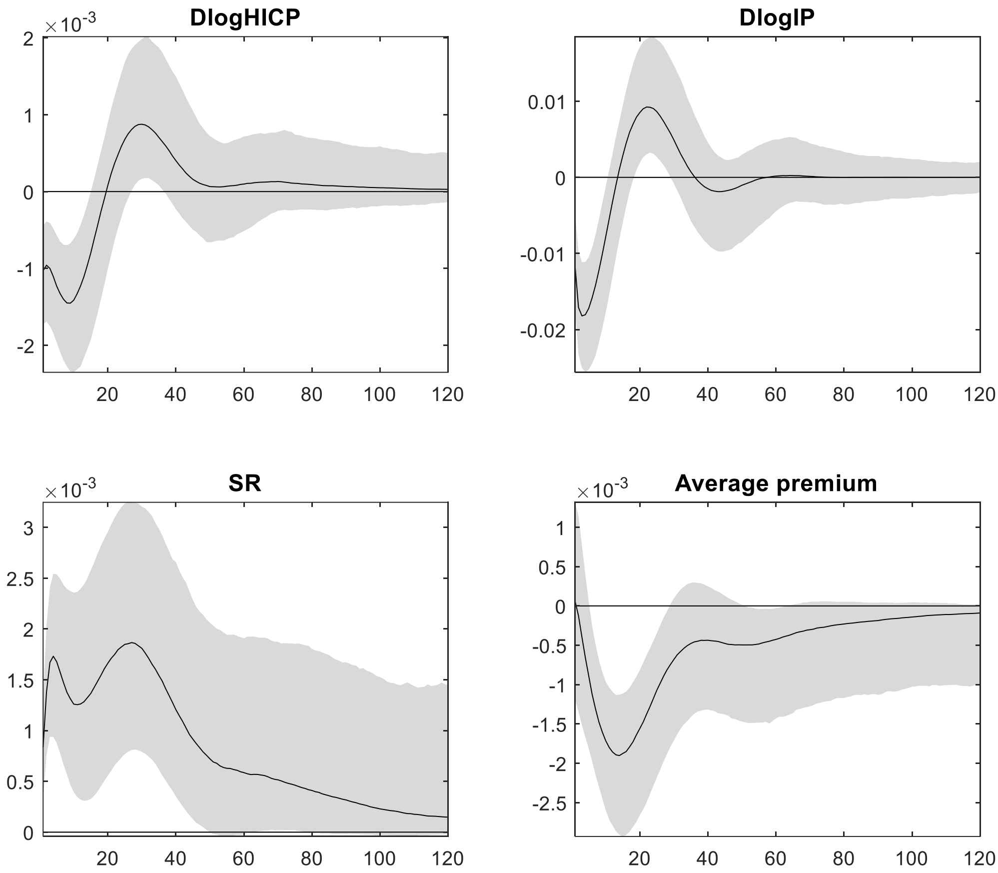


***Figure A4**. The impulse response functions to a one standard deviation positive shock to the shadow rate in the model in which the premium is the long-run average expected premium implied by equation (2). The shadow rate by Kortela (2016) is replaced by the shadow rate by Wu and Xia (2016). Sign restrictions are imposed on DlogIP, DlogHICP and SR for the periods 1-9. The shaded area represents the 68 % confidence interval. The number of lags is 2.*

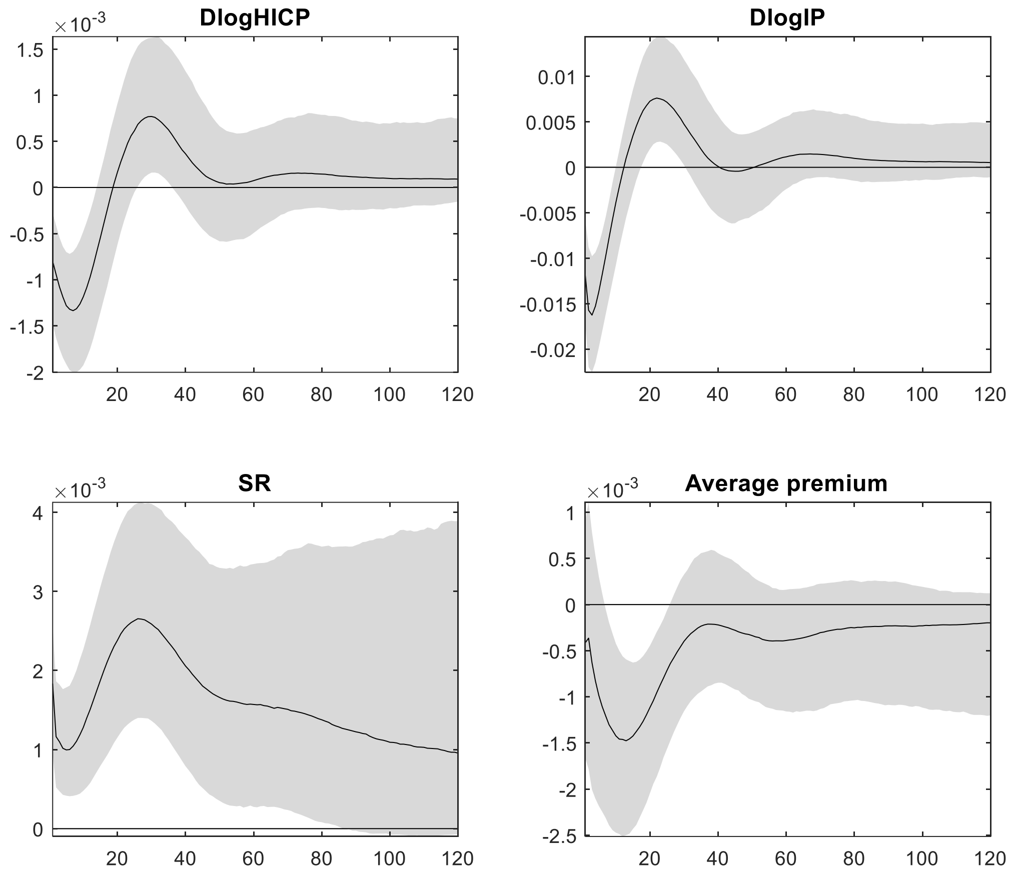


***Figure A5**. The impulse response functions to a one standard deviation positive shock to the shadow rate in the model in which the premium is the long-run average expected premium implied by equation (2). Sign restrictions are imposed on DlogIP, DlogHICP and SR. The sign restrictions are now set to hold only the first 3 periods. The shaded area represents the 68 % confidence interval. The number of lags is 2.*

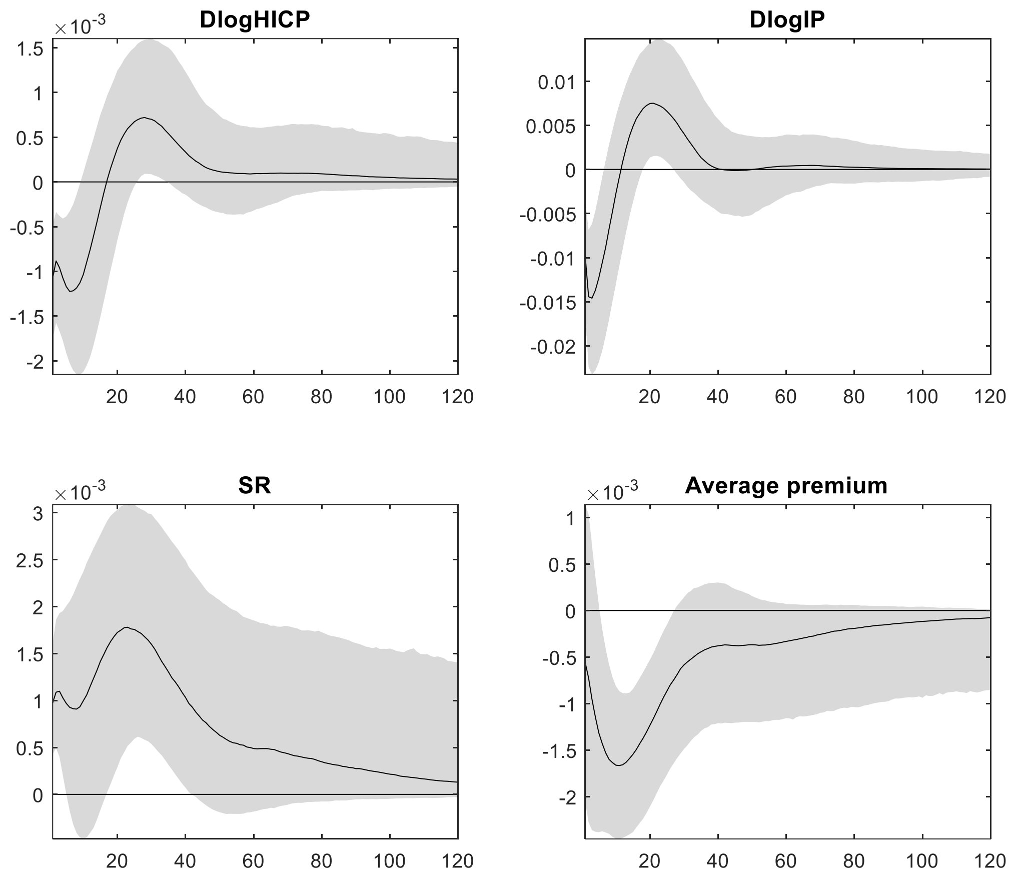


Figure A6. *The impulse response functions to a one standard deviation positive shock to the shadow rate in the model in which the premium is the long-run average expected premium implied by equation (2). The structural model is identified recursively. Average premium is ordered after the shadow rate. The shaded area represents the 68 % confidence interval. Model includes 2 lags.*

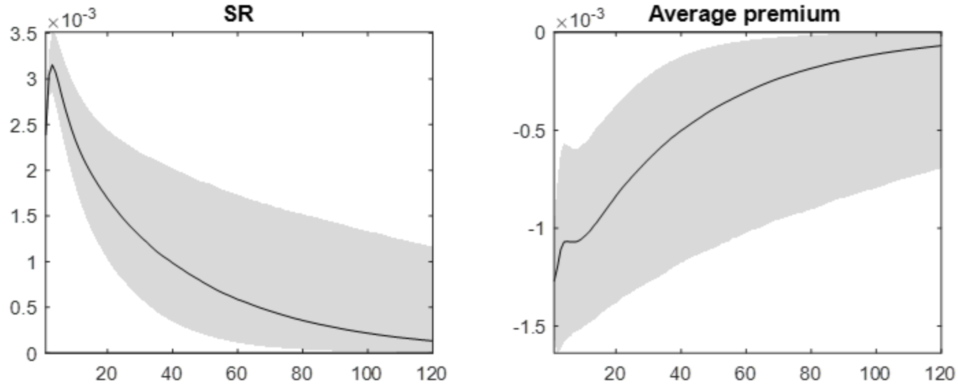


Figure A7. *The impulse response functions to a one standard deviation positive shock to the shadow rate in the model in which the premium is the long-run average expected premium implied by equation (2). The structural model is identified recursively. Average premium is ordered after the shadow rate. The shaded area represents the 68 % confidence interval. Model includes 12 lags.*

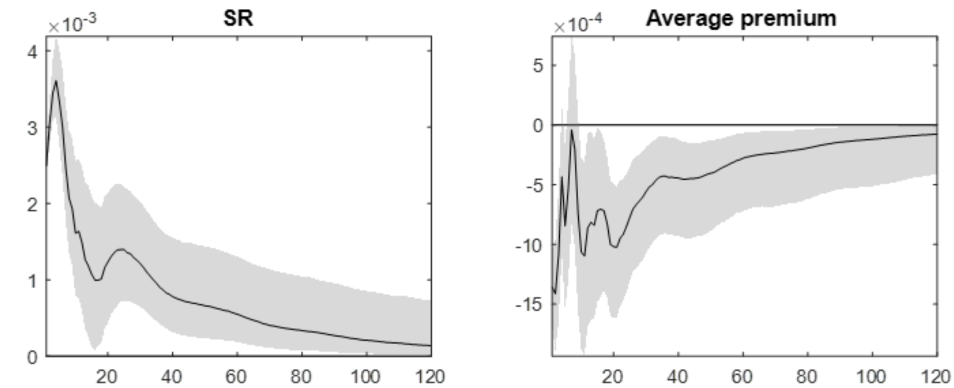


Figure A8. *The impulse response functions to a one standard deviation positive shock to the shadow rate in the model in which the premia are the long-run average expected premium implied by equation (2) and year 1, 2, 3, 4 and 5 premia implied by equation (3). Sign restrictions are imposed on DlogIP, DlogHICP and SR for the periods 1-9. The shaded area represents the 68 % confidence interval. The number of lags is 2.*

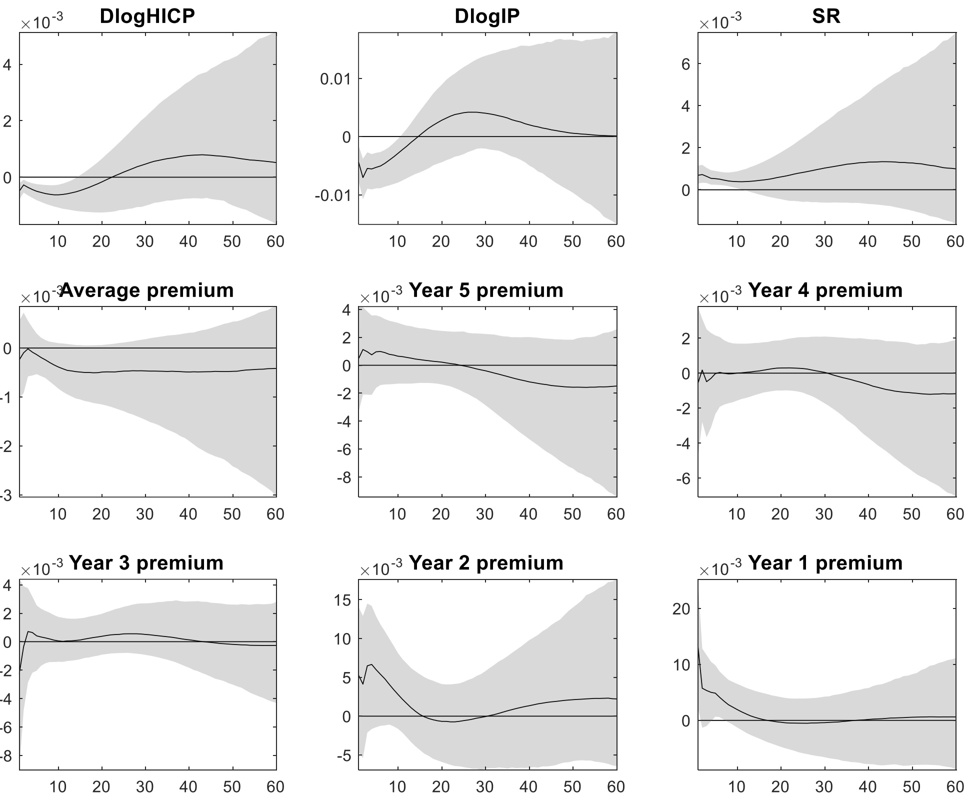


***Figure A9**. The impulse response functions to a one standard deviation positive shock to the shadow rate in the model in which the premium is the long-run average expected premium implied by equation (2). The change in log(Eurostoxx 50) is included into the model. Sign restrictions are imposed on DlogIP, DlogHICP, SR for the periods 1-9 and on DlogStockPrice for the periods 1-2. The shaded area represents the 68 % confidence interval. The number of lags is 2.*

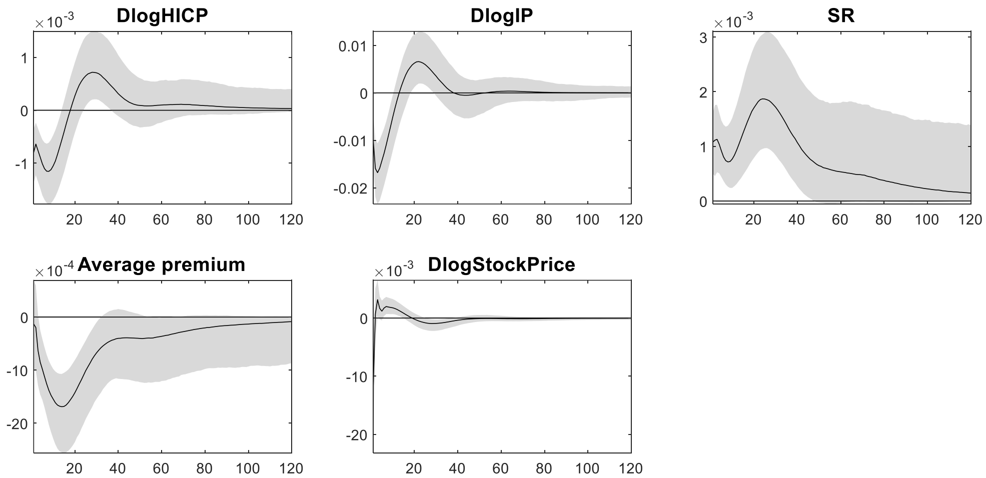
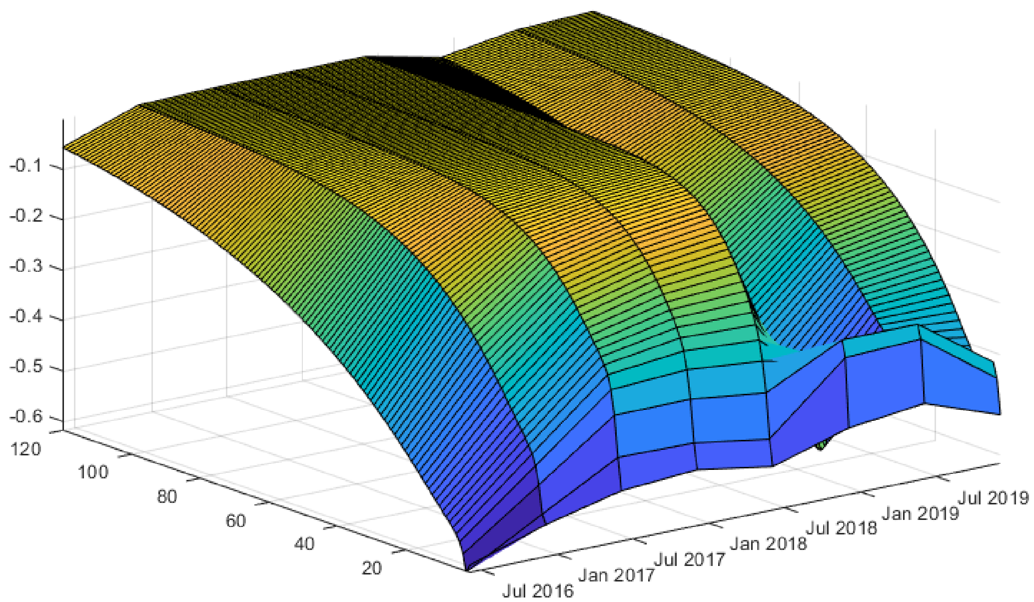


***Figure A10**. The time variation of the impulse response functions of the long-run average expected premium implied by equation (2) to one standard deviation positive shock to the shadow rate. The model includes SR and the premium in this order. The shock is identified recursively, and the size of the shock is normalized to 1 percent. The model has 2 lags. The dates on the axis tell the last period included in the estimation window.*

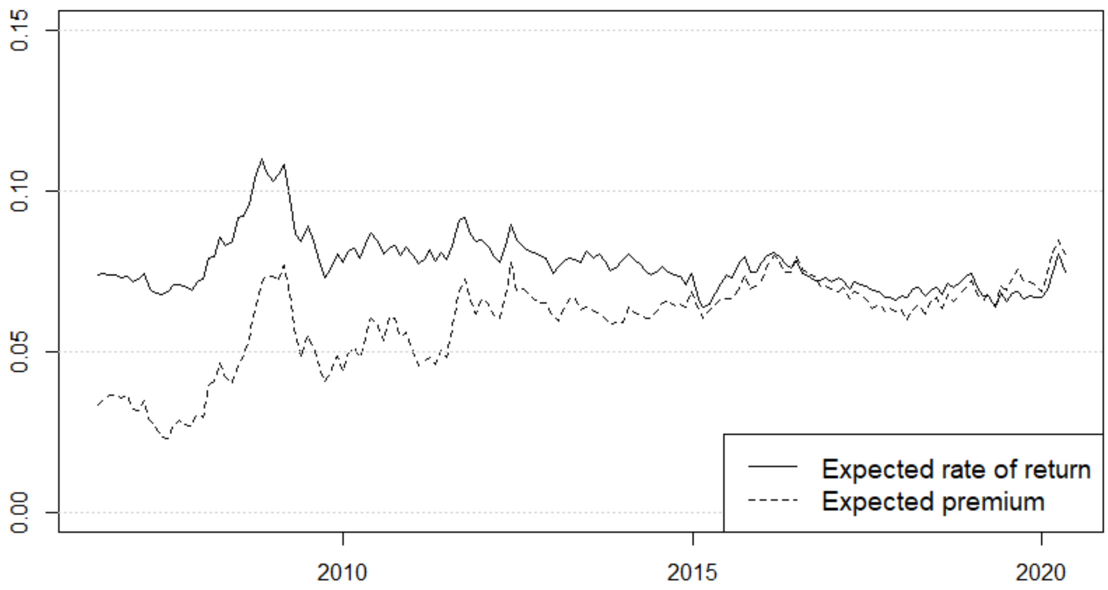


Appendix B

Figure B1 shows the development of risk premium and rate of return, when the long-run dividend growth rate, g, is assumed to have declined linearly from 4.5 percent to 0.5 percent during the period. Even in this case the risk premium seems to exhibit small upward trend. It does not seem very likely that the long-run expected growth rate would have declined more than 4 percentage points. Figure B2 shows the same development, if it was assumed that g is equal to 1 percent plus the average of (time-varying) rate of 10-year inflation swap and long-run expected inflation from the Survey of Professional Forecasters (SPF). Using inflation expectation as a proxy for the long-run nominal growth rate does not change the conclusions.

Figure B1. *Solved expected rate of return and expected premium, when g is assumed to have declined from 4.5 percent to 0.5 percent. The variables are solved applying equation (2) to Eurostoxx 50 stock market index and analysts' consensus forecasts in the period from 06/2006 to 04/2020.*

***Figure B2**. Solved expected rate of return and expected premium, when g is assumed to equal 1 % + (the rate of 10-year inflation swap + long-run inflation expectation from SPF)/2. The variables are solved applying equation (2) to Eurostoxx 50 stock market index and analysts' consensus forecasts in the period from 06/2006 to 04/2020.*



Appendix C

Figure C1 uses 3-month surprise instead of the average surprise. Figure C2 shows the results when the number of lags is increased to 4. The results are hardly affected.

***Figure C1**. The local-projection-based impulse response function of the year 1, 2, 3, 4 and 5 premia implied by equation (3) and the long-run average expected premium implied by equation (2). The number of lags in the model is 3. Inflation and the growth rate of industrial production are used as control variables. 3-month surprise is used as a shock. The confidence band is 68 %. The impulse response functions and their Newey-West based confidence bands are estimated as in Barnichon and Brownlees (2019).*

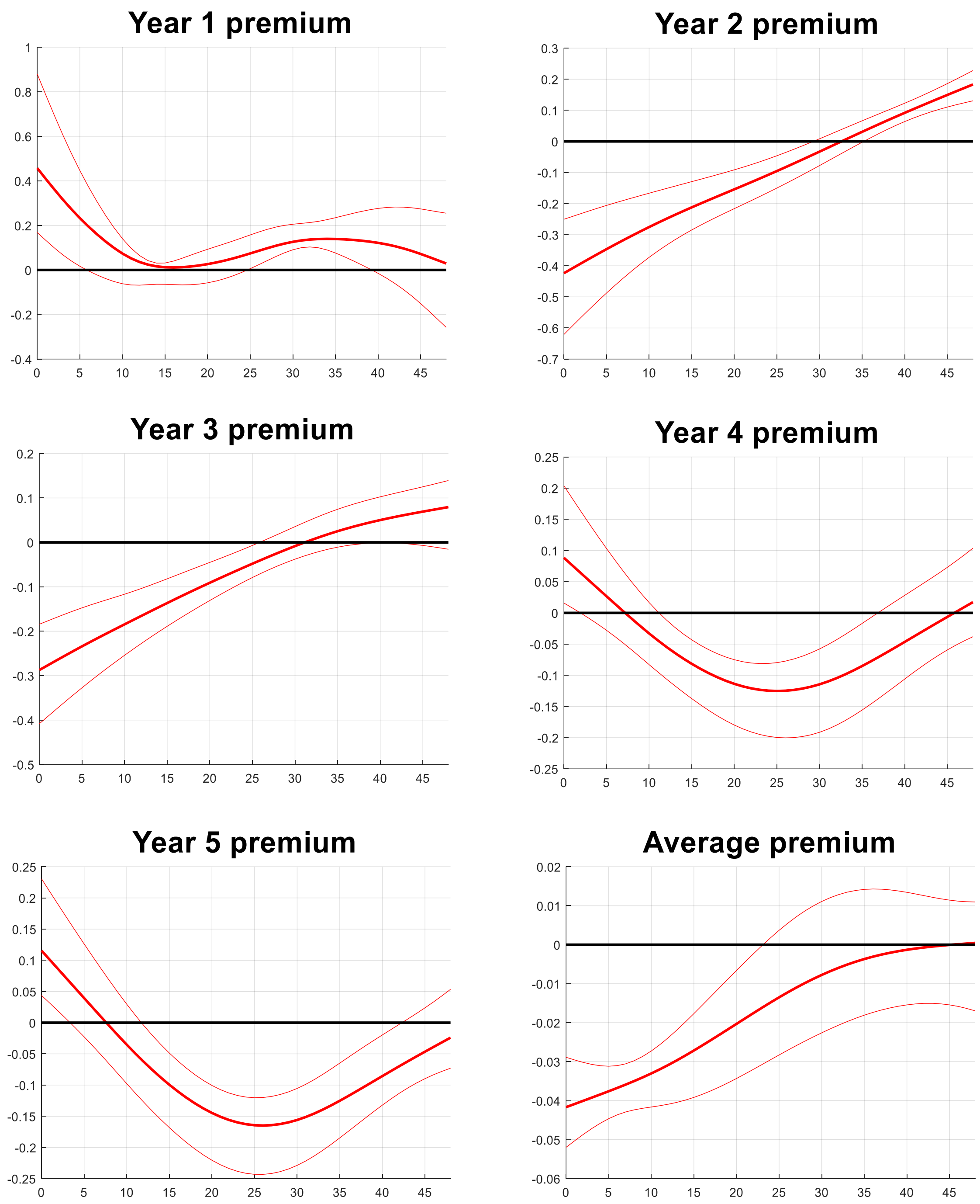
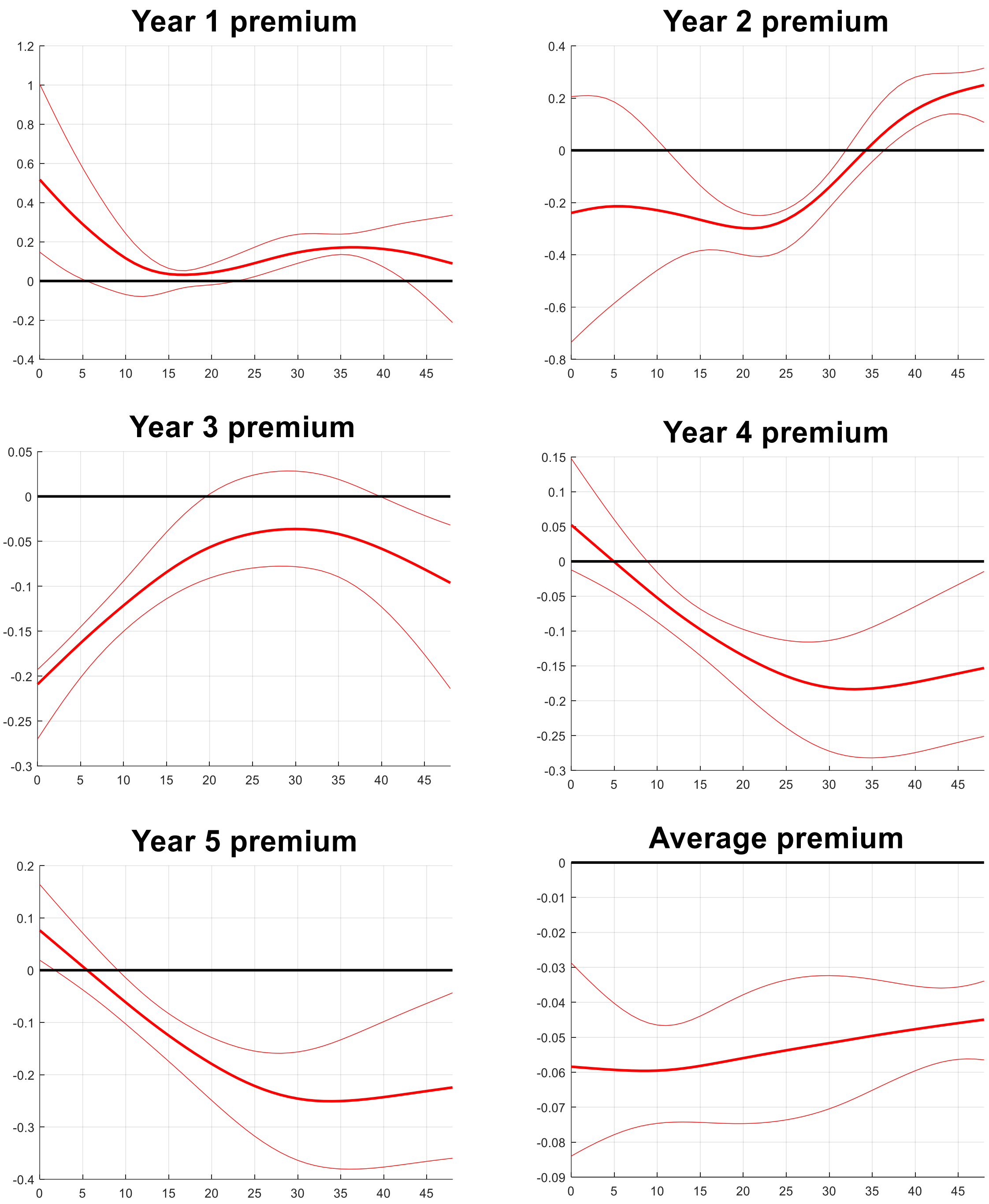


***Figure C2**. The local-projection-based impulse response function of the year 1, 2, 3, 4 and 5 premia implied by equation (3) and the long-run average expected premium implied by equation (2). The number of lags in the model is 4. Inflation and the growth rate of industrial production are used as control variables. The confidence band is 68 %. The impulse response functions and their Newey-West based confidence bands are estimated as in Barnichon and Brownlees (2019).*

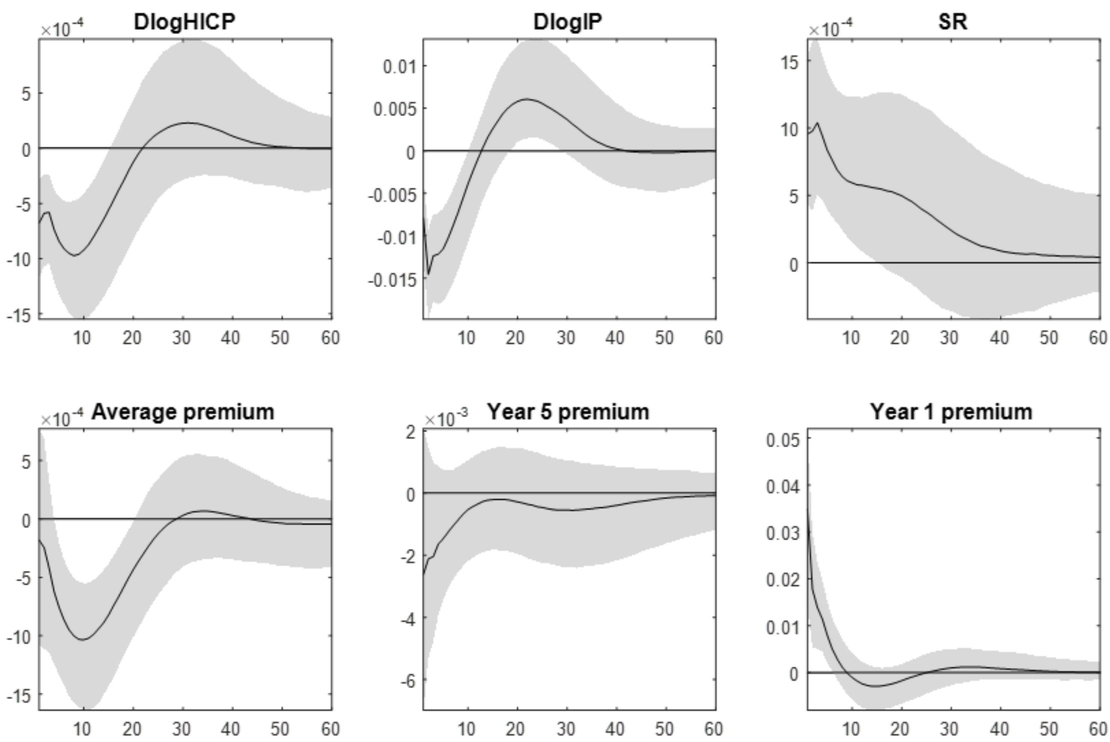


Appendix D

Jokivuolle (1995) shows that the permanent component of Beveridge-Nelson (1981) decomposition can be used to measure the true value of stock index that is not directly observable due to infrequent trading of stocks. I apply this method to analysts' dividend expectations: the "infrequent trades" are in my case infrequent dividend forecast updates. The method of Jokivuolle (1995) assumes that the log of true index follows a random walk with drift. The possible autocorrelation in the observed series should then reflect infrequent trading, and Beveridge-Nelson (1981) decomposition corrects this. Using the method of Jokivuolle (1995) is a conservative approach. For example, Bouchaud, Krueger, Landier and Thesmar, (2019) and Coibion and Gorodnichenko (2015) provide evidence that the expectations might be sticky. Therefore, the method of Jokivuolle (1995) probably overcorrects the expectations.

Figure D1 shows the impulse response functions, when the analysts' dividend forecasts are corrected using the method of Jokivuolle (1995). The number of lags assumed in the Beveridge-Nelson (1981) decomposition is 12. The method affects the implied risk premia, but the impulse response functions remain about the same.

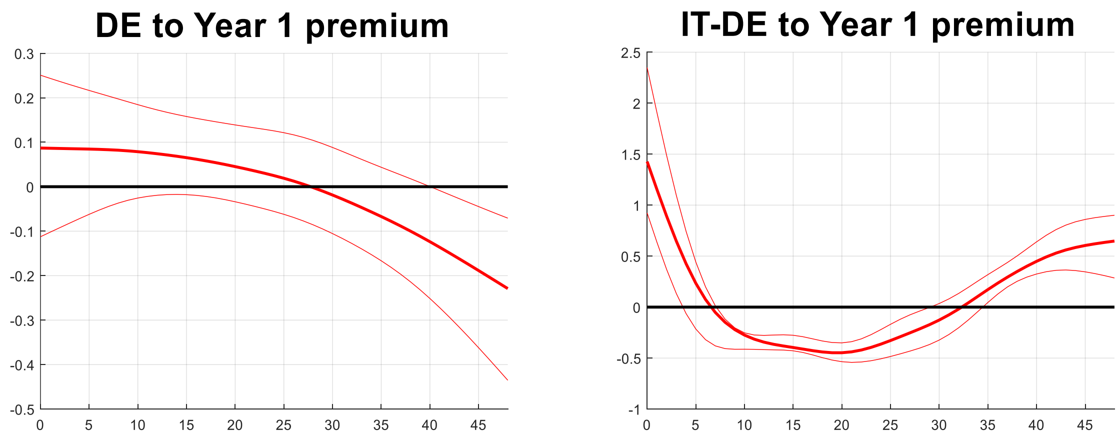
Figure D1. *The impulse response functions to a one standard deviation positive shock to the shadow rate in the model that includes three different premia. The analysts' dividend expectations are corrected using the method of Jokivuolle (1995). Sign restrictions are imposed on DlogIP, DlogHICP and SR for the periods 1-9. The shaded area represents the 68 % confidence interval. The number of lags is 2.*



Appendix E

Figure E1 shows the local projection impulse responses of the year 1 premium to rate surprises in a) the 10-year yield of Germany (DE) and b) the spread between the 10-year yields of Italy and Germany (IT-DE spread). The bond of Germany can be considered as a risk-free or negative-beta asset. The government bond of Italy, instead, is relatively risky asset. Surprises in the German yield are probably correlated to the announcements of policies related to asset purchases aimed at safe assets. Surprises in the spread are probably correlated to the announcements of asset purchases aimed at riskier asset. The immediate reaction to the surprise in the German yield is positive, but economically rather small compared to the response after the surprise in IT-DE spread. The positive response is also within the confidence bands unlike in the case of IT-DE spread. This simple analysis suggests that stock market risk premia respond differently to different kind of asset purchases. In the light of these results, it might be interesting to assess the mechanism of Caballero and Farhi (2013, 2018) more carefully in the future research. For example, the rate surprises are correlated with each other and announcements of different asset purchase programmes affect multiple yields and spreads. In the future research, it would be interesting to disentangle these different policies in a more careful way.[31]

Figure E1*. The local-projection-based impulse response function of the year 1 premium implied by equation (3). The changes in the German 10-year government bond yield and the spread between 10-year Italian government bond yield and the German one around the ECB's decisions are used as proxies for monetary policy shocks. The number of lags in the model is 3. The confidence band is 68 %. The impulse response functions and their Newey-West based confidence bands are estimated as in Barnichon and Brownlees (2019).*



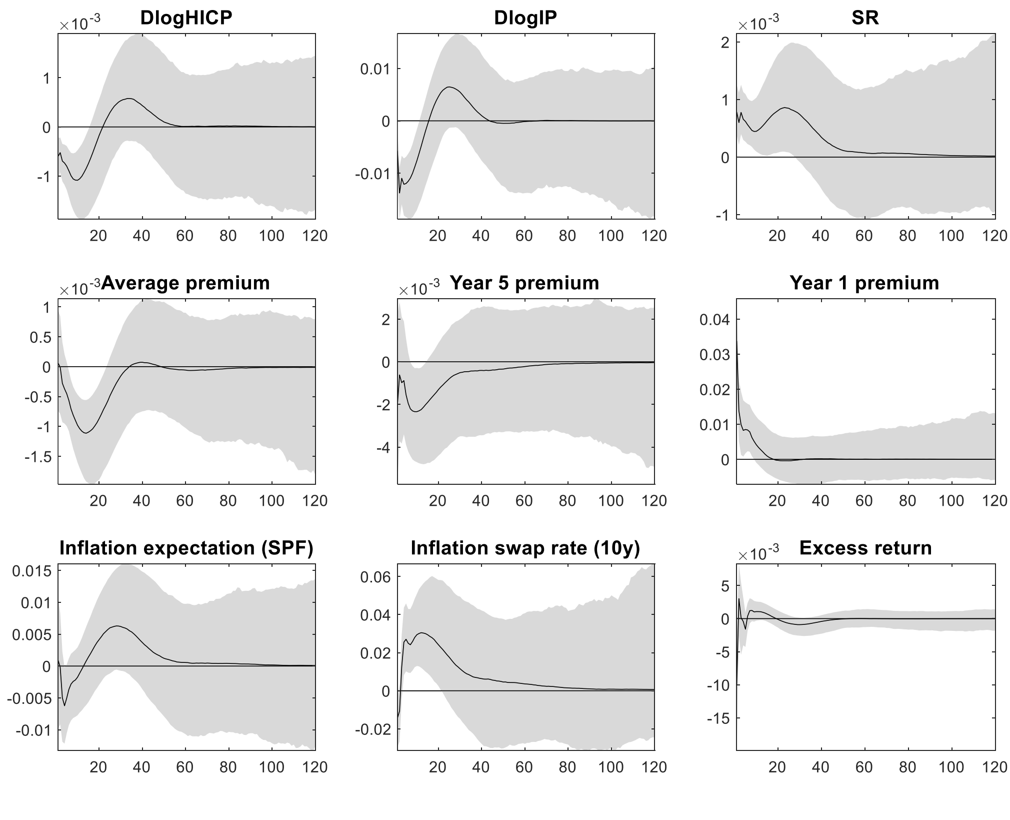
[31] One way to disentangle different monetary policies has been proposed recently by Kortela and Nelimarkka (2020). Nelimarkka and Laine (2021) use this methodological framework to assess the effectiveness of different monetary policy measures during the covid-19 pandemic.

Appendix F

Figure F1 shows the impulse response functions of the model that has been augmented with the rate of 10-year inflation swap and long-term inflation expectation (from the Survey of Professional Forecasters, SPF). The excess return of stocks is also added. The identification is otherwise as in the baseline VAR model, but the immediate reactions of the inflation swap and the excess return are restricted to negative to further improve the identification.

The results do not support the idea that monetary policy easing would raise the long-horizon premia through the higher inflation risks. At least the identified policy shock does not seem to move survey- and market-based expectations to the right direction. Even though the immediate reaction of the inflation swap is restricted to negative, the data seem to think otherwise: the sign of the response changes rapidly.

Figure F1*. The impulse response functions to a one standard deviation positive shock to the shadow rate in the model, which has been augmented by survey-based long-term inflation expectation (SPF), 10-year inflation swap rate and the monthly excess return of stocks. Sign restrictions are imposed on DlogIP, DlogHICP and SR for the periods 1-9, and on the inflation swap rate and on the excess return for the first period. The shaded area represents the 68 % confidence interval. The number of lags is 2.*



Appendix G

Figure G1 shows the impulse response functions of the model that is augmented with the analysts' dividend forecasts 1-year-ahead (Div1) and 5-years-ahead (Div5). Div1 declines after contractionary monetary policy, but Div5 stays about the same or even rises. The reaction of Div5 may seem strange given that contractionary monetary policy can be expected to lower the profits of firms. To further analyse whether this reflects some systematic error in the analysts' expectations, I consider the forecast errors made by analysts.

Figure G2 shows the histogram of the analysts' forecast errors at the horizon of five years. The error has been calculated as a log-difference between the expected dividend and the realized dividend. The distribution is skewed to the right and the analysts have been on average about 10 percent too optimistic. The largest forecast errors happened during the financial crisis and the years 2015-2016.

The key question is whether the expectation errors are caused by monetary policy shocks. If this was the case, one could say that monetary policy causes positive or negative "bubbles in dividend expectations". Figure G3 shows the local projection impulse response that is followed by the surprise in the average OIS rate. The response is positive suggesting that contractionary monetary policy makes analysts too optimistic. Therefore, it is possible that some systematic errors made by analysts explain the results of this paper partly. However, it should be noted that the response in Figure G3 is not economically significant: 1 percentage point rate surprise causes only about 1 percent prediction error in 5-years-ahead dividend forecast.

***Figure G1**. The impulse response functions to a one standard deviation positive shock to the shadow rate in the model, which has been augmented by 1-year-ahead and 5-years-ahead dividend forecasts. Sign restrictions are imposed on DlogIP, DlogHICP and SR for the periods 1-9. The shaded area represents the 68 % confidence interval. The number of lags is 2.*

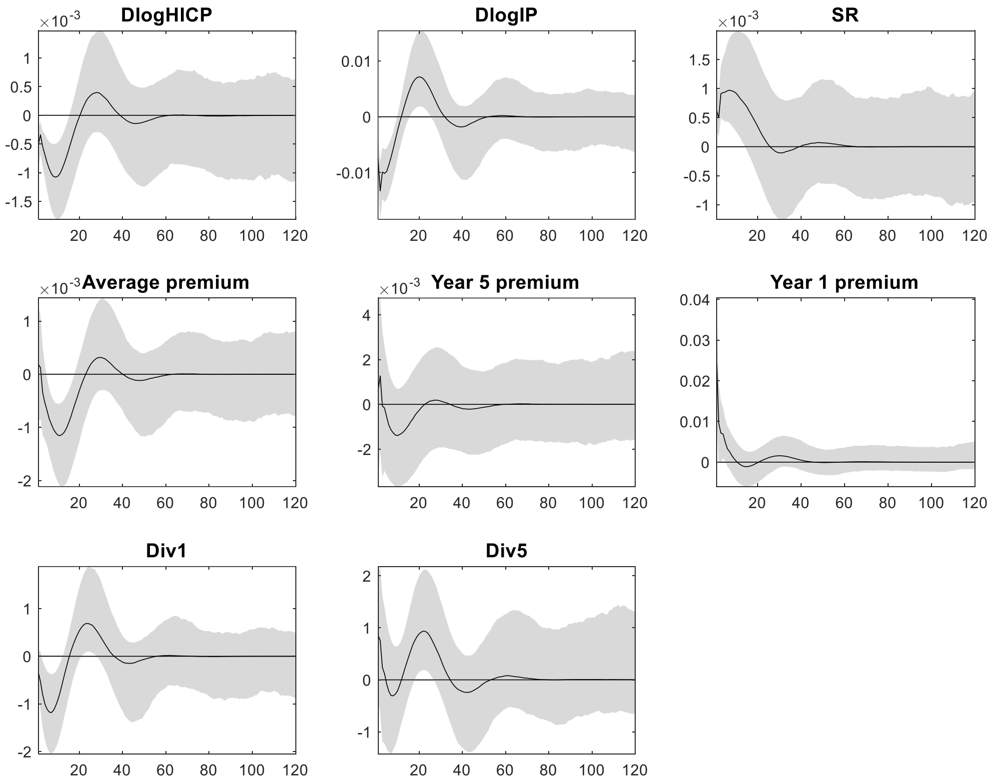


***Figure G2**. The histogram of 5-years-ahead forecast errors made by analysts. The error is defined as: log(dividend forecast)-log(realized dividends).*

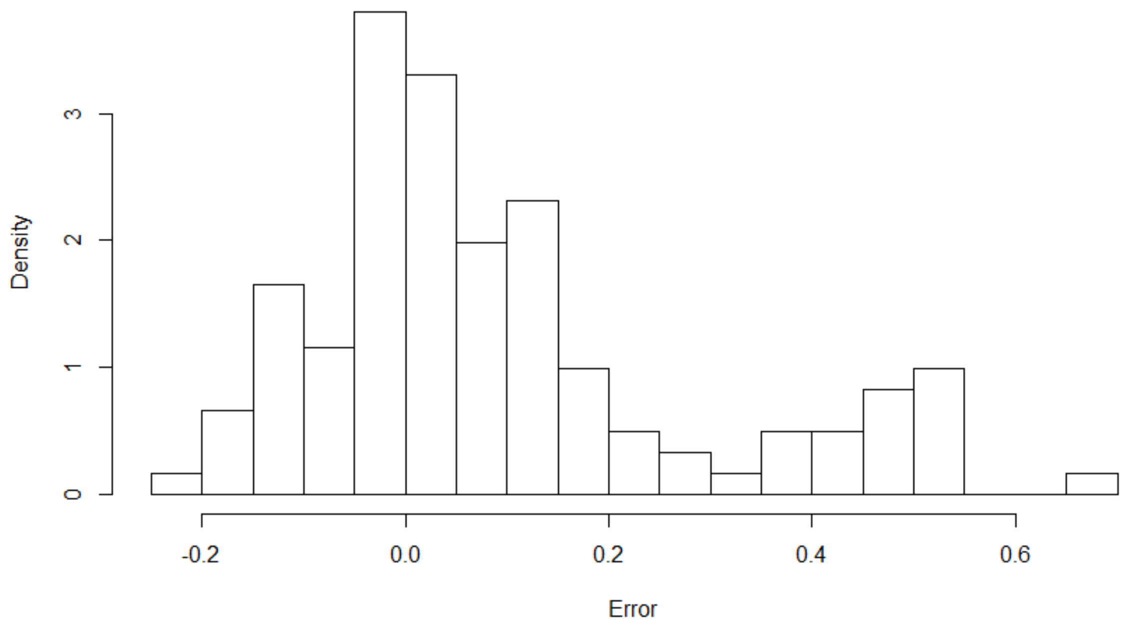
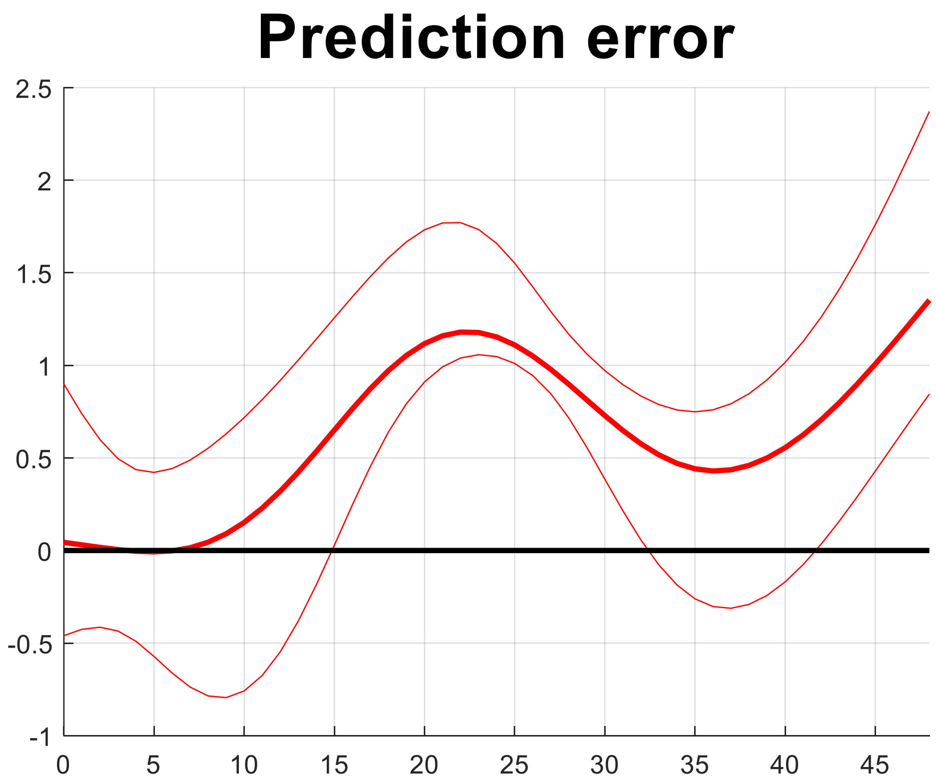


***Figure G3**. The local-projection-based impulse response function of the analysts' 5-years-ahead forecast error. The number of lags in the model is 3. The confidence band is 68 %. The impulse response function and its Newey-West based confidence bands are estimated as in Barnichon and Brownlees (2019).*



The Effect of Targeted Monetary Policy on Bank Lending

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ABSTRACT

This paper studies the effect of central banks' targeted refinancing operations on bank lending. It utilizes data from the European Central Bank's targeted longer-term refinancing operations (TLTROs) together with monthly bank level balance sheet data from multiple countries. The effect of targeted policy is identified utilizing the institutional setting that provides natural instrumental variables and a proxy for credit demand. Unlike previous papers, this paper studies the effects on corporate loans and loans for consumption separately. The cumulative effect of TLTROs on participating banks' stock of corporate loans is estimated to be significant (about 20 per cent). However, the effect on lending for consumption is found close to zero. Furthermore, the positive effects on corporate loans are found to be driven by crisis countries suggesting that the effectiveness of monetary policy depends on the economic conditions. The paper also finds some evidence that the effect on government bond purchases is negative. This result is very different from the earlier results regarding non-targeted liquidity operations.

JEL Classification: E44; E51; E52; G21

Keywords: unconventional monetary policy, credit supply, TLTRO, bank lending

1. INTRODUCTION

When policy rates have been close to the effective lower bound, central banks have adopted a range of unconventional tools to stimulate the economy. One channel through which these tools operate is bank lending.[2] The unconventional tools have included providing banks with cheap long-term credit. For example, the European Central Bank (ECB) has conducted several longer-term credit operations that have been geared to increasing bank lending to the non-financial private sector in order to stimulate activity in the real economy and accelerate euro area inflation. Andrade et al. (2018) find that these operations have increased bank lending to non-financial corporations. Though the earlier literature has provided some evidence that supports the effectiveness of these

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[2] See for example Jiménez et al. (2012), Rodnyansky and Darmouni (2017), Altavilla et al. (2020), Di Maggio et al. (2020).

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tools, many questions have remained unanswered. Especially, the literaure concerning so called targeted refinancing operations is scarce. These targeted operations are the focus of this paper.

The first shortage of the literature is that it has not studied the effects of these liquidity operations to other types of loans than loans to firms. Because the credit market is quite different for households and firms, it is likely that the liquidity operations have very different effects on lending to non-financial corporations and lending to households, though the banks‘ are given equal reward for lending to households and lending to firms. Second, the literature finds that liquidity operations have increased bank lending both on the extensive margin and on the intensive margin (e.g. Benetton and Fantino, 2021). In other words, both the participation to the operations and the borrowed ammount matters. However, central banks have launched several targeted and non-targeted operations that have rather different incentive structures. Therefore, the results regarding one operation cannot be necessarily generalised to another operation. Another issue regarding the generalisation of the previous results is that the earlier literature has focused on the effects in single countries, though the effects may be very different in different economic conditions.[3] Finally, one important reason for targeting the liquidity operations in the euro area was probably the observation that non-targeted longer-term refinancing operations seemed to be used for buying sovereign debt (see Crosignani et al., 2020). Therefore, it should be analysed whether the targeted operations have had an effect of sovereign bond holdings of the banks or not.

This paper applies a difference-in-differences estimation to bank level dataset from multiple countries to analyse the effects of the second series of the ECB‘s targeted longer-term refinancing operations (TLTRO-II). The paper shows that the ECB's liquidity operations have boosted lending to non-financial corporations, but not the lending to households for consumption. This finding is interesting as the ECB does not favour corporate loans over loans for consumption. Unlike the previous literature, the paper finds that the positive impact is mainly explained by the effect of participation (extensive margin). The allotted amount of TLTRO-II does not seem to have been very important. In addition, the paper shows that the positive effects on corporate lending are largely driven by crisis countries. This suggests that the effectiveness of longer-term refinancing operations depends on the economic conditions under which they are implemented. The results also show that TLTRO-II did not increase participating banks' sovereign bond purchases. Instead, the effect is found negative. Thus, the results suggest that the ECB's targeting strategy was effective in this respect.

The remainder of the paper is as follows. Section 2 reviews the earlier literature. Section 3 describes the data and the used methodology. It is dividied into three subsections. The first one describes the institutional setting, the second one represents the data and the third one explains the methods used in this study. Section 4 shows the results. It begins from the baseline results that focus on the participation effect (or the extensive margin). Then it shows that assuming continuous treatment (the amount of TLTRO) yields different results. After that the section analyses potential cross-country differences between the effectiveness of TLTROs and the issue related to sovereign bond purcahses. Finally, the section considers the robustnsess of the results. Section 5 concludes.

2. LITERATURE REVIEW

Conventionally, the maturity of refinancing operations provided by central banks has been very short. For example, the maturity of the ECB‘s main refinancing operations is one week. In recent years, central banks have began to refinance banking sector with loans that have maturity of multiple years. The rationale of this policy change is, as Carpinelli and Crosignani (2021) note, that ”In presence of uncertainty about the future role of the central bank as a liquidity provider,

[3] García-Posada and Marchetti (2016) study the effects in Spain, Andrade, Cahn, Fraisse, Mésonnier (2018) in France, Benetton and Fantino (2021) and Carpinelli and Crosignani (2021) in Italy.

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short-term liquidity is ineffective in stopping an ongoing credit contraction". Furthermore, central banks have began to incentivise banks to use this credit for lending to non-financial private sector (e.g. TLTROs in the euro area and Funding for Lending in the UK). In the euro area, the ECB launched the first series of TLTROs in the year 2014.

As these targeted tools are rather new, there are not many published papers that study their effectiveness. When it comes to non-targeted operations, Andrade et al. (2018), Carpinelli and Crosignani (2021) and García-Posada and Marchetti (2016) provide some evidence about their effectiveness using bank level data from single countries. VAR evidence is provided by Darracq-Paries and De Santis (2015).

When it comes to targeted operations, that is the main interest of this paper, Balfoussia and Gibson (2016) show that the first series of TLTROs (TLTRO-I) incresed lending to firms. In addition, contemporaneously with this paper, Benetton and Fantino (2021) show using data from Italy that TLTRO-I lowered the rates of corporate loans and incresed their amount. In addition, they find that the competition between banks matters for the effectiveness of targeted lending progmmes.

3. DATA AND METHODOLOGY

3.1. TLTRO-II

TLTRO-II was launched in June 2016 to ease private-sector credit conditions and stimulate credit creation. Four operations, one each quarter, were conducted, with the final operation taking place in March 2017. TLTRO-II loans carry a maturity of four years, so e.g. the first operation matured in June 2020. The borrower banks are also able to repay voluntarily the amounts borrowed at a quarterly frequency starting two years from the settlement of each operation.

Banks could borrow a total amount of up to 30 per cent of a specific eligible part of their loans in January 2016, less any amount previously borrowed and still outstanding under the first two TLTRO-I operations in 2014. Eligible loans included loans to non-financial corporations and households (excluding loans to households for house purchase).

The interest rate of the operations was fixed to match that of main refinancing operations (MROs) prevailing at the time of allotment. Nonetheless, the participating banks were given an incentive to increase their eligible lending by promising a lower rate if the eligible lending was increased enough in the period between February 2016 and January 2018 in comparison to bank specific benchmark. The lowered rate could be as low as the rate on the deposit facility (-0.40 per cent).

The bank-specific benchmark depended on eligible net lending as follows. For the banks with positive eligible net lending in the 12-month period before January 2016, benchmark net lending was set at zero. For the banks with negative eligible net lending, benchmark net lending was the same as eligible net lending in the 12-month period before January 2016.

The incentives in TLTRO-II to increase eligible lending differed from the incentives in TLTRO-I. In TLTRO-I, the banks were pushed to increase their lending by offering them more TLTRO-I credit when they increased their eligible lending. However, the banks were able to reduce their lending after they had borrowed their preferred amount of TLTRO-I credit. A key difference between TLTRO-I and TLTRO-II was also the maturity. TLTRO-I credit borrowed in September 2014 matured after four years, but the last operation of TLTRO-I matured after about two years. The key differences between VLTRO operations of 2011–2012 and TLTRO operations are summarised in Table 1.

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Table 1
Main features of the ECB's longer-term refinancing operations in recent years

	VLTRO	TLTRO-I	TLTRO-II
Implementation	2 operations (12/2011 and 2/2012)	8 operations between 9/2014 and 6/2016.	4 operations between 6/2016 and 3/2017.
Interest rate	Average MRO rate	First operations: MRO rate + 10bp at time of allotment. Subsequent operations: MRO rate only.	MRO rate at time of allotment. Possibility for lowered rate if eligible net lending increased sufficiently.
Maturity	Both operations carried maturities of 3 years.	All operations mature in 9/2018.	Every operation has a maturity of 4 years.
Amount	Full allotment	9/2014 and 12/2014: Max. 7% of eligible loans in 4/2014. 2015-2016: Max. 3 x eligible net lending relative to bank-specific benchmark.	Max. 30% of eligible loans in 1/2016, less any amount previously borrowed and still outstanding under the first two TLTRO operations in 2014.

Source: ECB's press releases.

Table 2
Descriptive statistics grouped by decision to participate in TLTRO-II

Variable	TLTRO-II participant (n = 97)		TLTRO-II non-participant (n = 90)	
	Mean	**Median**	**Mean**	**Median**
Balance sheet (million €)	106 989	40 043	72 290	14 203
Central bank credit to total liabilities	4.4%	2.5%	0.7%	0.0%
Household deposits to total liabilities	25.5%	24.3%	33.7%	36.5%
Equity ratio	10.4%	9.1%	10.0%	8.2%
Eligible credit to total assets	26.0%	24.9%	27.1%	27.0%

Note: The statistics are calculated from bank-level January 2015 to May 2016 averages, i.e. before TLTRO-II. Thus, statistics represent how the banks that participated in the credit operations and the other banks differed before treatment.

Source: Author's calculation

3.2. Data

The main data are taken from the ECB's individual balance sheet items (IBSI) database. The data are monthly and at bank level. The used data are from January 2015 to July 2018. The IBSI data are linked to confidential information about bank's total borrowing in TLTRO-II.

IBSI data offer several advantages. First, they make it possible to analyse TLTRO-II in multiple countries. Additionally, as the data are monthly and cover a sufficiently long time period after the treatment, it is possible to analyse how possible effects evolve over time. While IBSI does not cover all euro area banks the sample is quite large and includes about 300 large banks that are from all the euro area countries. The final dataset covers 187 banks from 18 countries due to missing data.[4] However, the data are still very representative as the interpreted bank

[4] All the banks that have missing data from necessary variables are excluded. Also, banks that experience periods during which they have not had any corporate credit, loans for consumption or loans for house-purchase are excluded because these variables are analysed in logs. This sample selection limits generalisation of the results, but makes the analysed banks more alike. All the banks from France are excluded because the data about central bank credit are missing.

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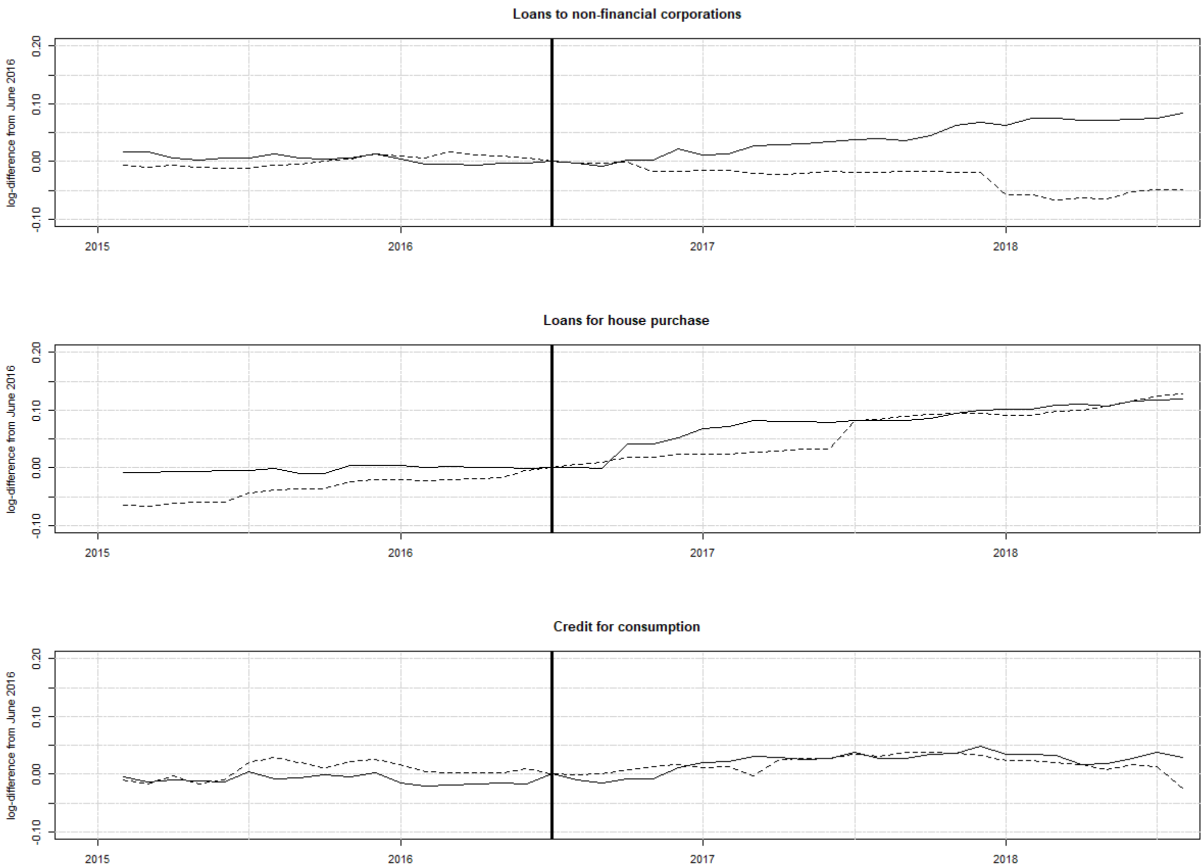
covered about 62 per cent of the total corporate loans in the euro area prior TLTRO-II. Some key descriptive statistics of the assessed banks, grouped by the decision to participate TLTRO-II, are shown in Table 2.

Figure 1 shows the average development of loans to non-financial corporations, loans for house purchase and loans for consumption by groups. The solid lines show the development of the TLTRO banks and dashed lines the developments of non-TLTRO banks. The TLTRO banks increased corporate lending compared to other banks after the beginning of TLTRO-II. Instead, it is rather difficult to observe significant diverging in other types of loans.

Figure 2 shows the average development of loans to non-financial corporations, loans for house purchase and loans for consumption among the banks that participated in TLTRO-II. Now, the grouping is based on the share of TLTRO-II in total liabilities. The size of balance sheet is from May 2016 (before TLTRO-II). The solid lines show the development of the banks that had the share of TLTRO-II above the median and dashed lines the developments of the banks that had a ratio below the median. The differences between groups remain rather constant. This suggests that the allotted amount of TLTRO-II was not essential.

This preliminary analysis has not taken into account the fact that banks could choose whether to participate in TLTRO-II or not. Additionally, this analysis has not considered the role of credit demand. These issues are assessed in the remaining sections.

Figure 1
The development of different types of credit in the treatment (solid line) and control (dashed line) groups in comparison to the situation as of June 2016



Note: The credit stocks are in logs. The treatment group includes 97 banks and the control group 90 banks.

Source: Author's calculation.

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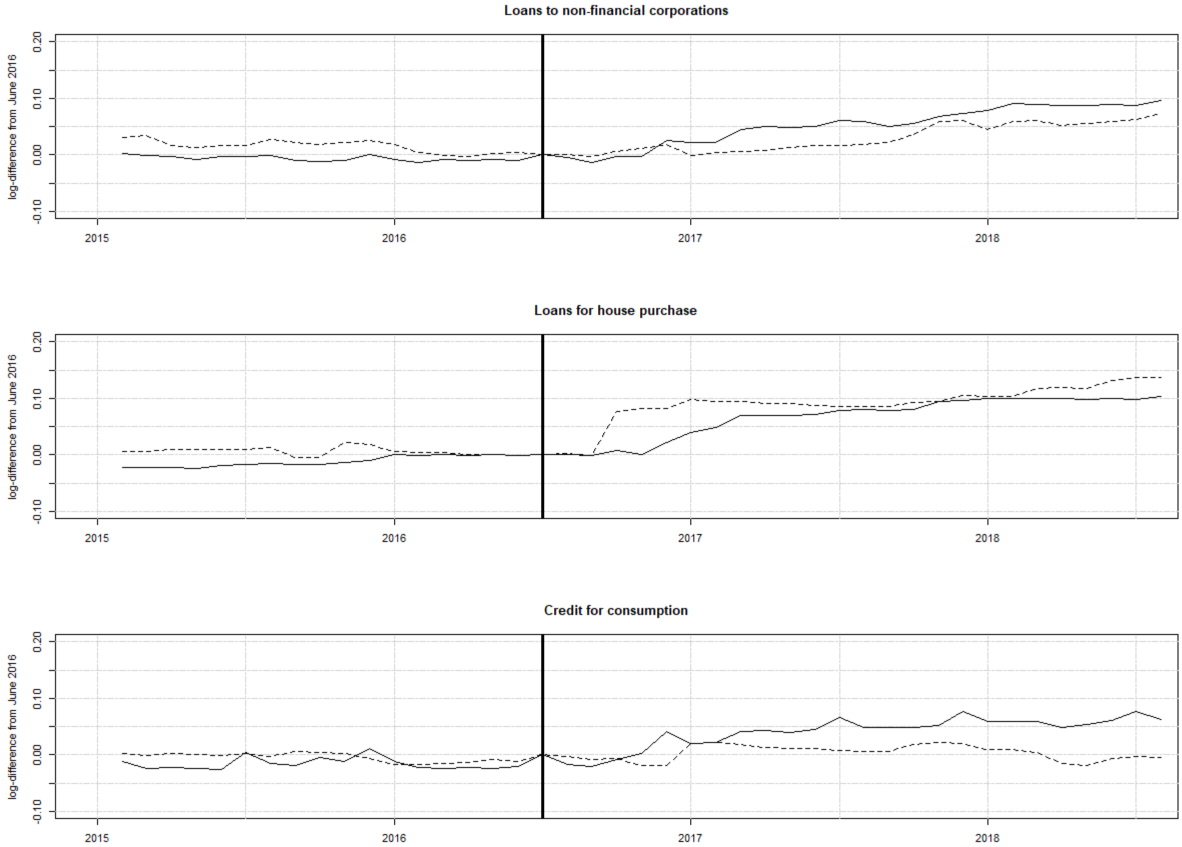
Figure 2
The development of different types of credit in the high-intensity participants (solid line) and low-intensity participants (dashed line) groups in comparison to the situation as of June 2016



Note: The credit stocks are in logs. The high-intensity group includes 49 banks and the low-intensity group 48 banks. High-intensity participants are those that had the ratio of TLTRO-II take-up to total liabilities (in May 2016) above the median. Low-intensity participants are the banks that borrowed in TLTRO-II, but had the ratio below the median.

Source: Author's calculation.

3.3. Methodology

The paper applies a difference-in-differences approach to study the effects of TLTRO-II on bank lending. It uses two types of specifications. First, in the baseline regression the treatment is assumed to be binary: participation in the TLTRO-II or not. This specification is used to assess the participation effect of TLTROs. Second, it is possible that the impact of TLTROs depends on the amount allotted to the banks. As was discussed earlier, some other studies find that the targeted operations have a positive effect on bank lending both on the extensive and on the intensive margin. The second specification is used to analyse this issue.

To be concrete, the baseline specification is:

$$\ln(Y_{ict}) = \alpha_{ic} + \tau_{ct} + \Sigma_h \beta_h (D_h \cdot TLTRO_{ci}) + \gamma Z_{ict} + e_{ict}, \tag{1}$$

where Y_{ict} is the stock of credit on the balance sheet of bank i in country c at time t, α_{ic} includes bank fixed effects, τ_{ct} includes country-time fixed effects, Z_{ict} includes time-varying bank-specific control variables that are the size of balance sheet in logs and equity ratio in the baseline analysis. $TLTRO_{ic}$ equals 1 if the bank participated in TLTRO-II and D_h, where $h \in \{2015Jan, \dots, 2018Jul\} \backslash \{2016Jun\}$, includes indicators for time periods. June 2016 is the reference month. This means that the regression coefficients β_h tell how the credit granted by TLTRO banks differed from other banks in a given month relative to the difference between the

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groups in June 2016. Standard errors are clustered at bank and month level to allow for serial correlation and heteroscedasticity in the error term e_{ict}.

A similar approach is used by Rodnyansky and Darmouni (2017) to investigate the effects of quantitative easing on bank lending behaviour in the United States. This specification is useful because it is not realistic to assume that the effect was the same in every month after treatment as is assumed in standard difference-in-differences models. If the effect was the same every month after treatment, it would mean that the stock of credit in TLTRO participant banks jumped immediately after June 2016 and remained the same thereafter. Additionally, the estimates for the interactions before the beginning of TLTRO-II should be zero. Otherwise, the assumption of common trends would not be credible. Adding these interactions in the regression allows testing the common trend assumption.

To assess whether the allotted amount of TLTRO-II was important, we use a specification slightly different from Eq. (1). The modified model is:

$$\ln(Y_{ict}) = \alpha_{ic} + \tau_{ct} + \Sigma_h \beta_h^* (D_h \cdot \log(TLTRO\ amount_{ci})) + \gamma Z_{ict} + e_{ict}, \tag{2}$$

where the binary treatment variable is replaced by the natural logarithm of the amount borrowed in TLTRO-II.

A central challenge in this study is justifying the assumption of common development of TLTRO banks and other banks if TLTRO-II had never been conducted. Banks were free to decide whether they wanted to borrow TLTRO-II credit or not, so banks that participated in TLTRO-II may have increased their lending anyway. The coefficients may also be biased downwards, if participating banks had strong deleveraging pressures.

To tackle this selection bias, we use instrumental variable estimation. We utilise two different novel properties of TLTRO-II. First, TLTRO-II, launched in June 2016, was mainly used to replace earlier TLTROs that were mainly borrowed in 2014 and in the beginning of 2015.[5] In May 2016, TLTRO-I covered about 83 per cent of the total credit from the ECB. Therefore, the amount of credit from the ECB prior TLTRO-II is highly correlated to the amount borrowed in TLTRO-II. The amount of earlier TLTROs is also a valid instrument as it is quite difficult for a bank to forecast its lending opportunities multiple years ahead. In addition, in the first series of TLTROs, the incentive structure was such that it motivated banks to increase their lending at very beginning of the operations.[6] Therefore, it is probable that participation in TLTRO-I was not affected by the expected lending opportunities during the years 2016–2018. Thereby, the amount of TLTRO-I is a valid instrument for the amount of TLTRO-II. In Eq. (1), where the treatment is binary, we use $\frac{Credit\ from\ the\ ECB\ in\ May\ 2016}{Balance\ sheet\ in\ May\ 2016}_{ci}$ as an instrument for the participation in TLTRO-II. In Eq. (2), the used instrument is $\log(Credit\ from\ the\ ECB\ in\ May\ 2016)_{ci}$.

Another novel property of TLTRO-II is the fact that the amount a bank could borrow was predetermined by the ECB. This property provides another potential instrumental variable. The maximum amount a bank could borrow in TLTRO-II was based on its amount of loans to non-financial corporations and loans for consumption (so called eligible loans) in January 2016. This constraint was predetermined by the ECB and hence exogenous. Thus, the amount of eligible loans in January is another potential instrument for the participation in TLTRO-II. A similar identification strategy is used by Benetton and Fantino (2021) to analyse the effects of TLTRO-I. Because all the banks in the sample had eligible loans in January 2016, the share of eligible

[5] In the initial operation of TLTRO-II in June 2016, banks borrowed 399 billion euros. Nevertheless, the total stock of TLTROs increased only by 38 billion euros.

[6] In TLTRO-I, the participating banks were motivated to increase their eligible lending by promising a possibility to borrow more TLTRO credit if they increased lending. Because all TLTRO-I credit had to be paid back in 2018, the incentive structure motivated banks to increase their lending in the beginning of TLTRO-I. The reason for this is that the last operations of TLTRO-I had only a maturity of about two years. Thus, it was reasonable to increase lending as early as possible, and then be able to borrow more TLTRO credit with a long maturity.

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loans in total assets is a weak instrument. Therefore, the amount of eligible loans is used as an instrument only in Eq. (2) where the treatment is continuous. Specifically, the used instrument is $\log(\textit{Eligibje loans in January } 2016)_{ci}$.

In addition to the instrumental variables, the paper considers propensity score matching as a robustness check and shows that the results are robust to controlling for many observable variables.

In addition to the selection bias, another problem is the role of credit demand which is difficult to control for. Many earlier studies have utilised the approach of Khwaja and Mian (2008) and controlled the demand at firm level. Because we have no data about firms or households that had loans from multiple banks, we use country-time fixed effects. The problem in the approach of Khwaja and Mian (2008) and country-time fixed effects is the possibility of capturing supply side effects as well. If TLTROs increased the lending of all the banks and not just the lending of participating banks, then country-time fixed effects (or firm-time fixed effects) would unintendedly capture these indirect effects as well. The problem with country-time fixed effects is also the assumption that all the banks within a country faced identical credit demand. To mitigate these concerns, we test the robustness of the results by replacing τ_{ct} by τ_t and adding $\log(\textit{Loans for house purchase})_{ict}$ into Z_{ict}. The idea behind this control variable is the following. Loans for house purchase were excluded from the eligible loans. Therefore, it is likely that changes in loans for house purchase reflect mainly changes in credit demand.

4. RESULTS

4.1. Baseline results

First, we estimate Eq. (1) using 2SLS. The instrument we use is the average share of central bank credit in total liabilities in May 2016. Specifically, we instrument the interactions $D_h \cdot TLTRO_{ci}$ by $D_h \cdot \left(\frac{\textit{Credit from the ECB in May } 2016}{\textit{Balance sheet in May } 2016}\right)_{ci}$. The banks that participated in the first series of TLTROs were likely to participate also in TLTRO-II. Therefore, it is not surprising that the F-statistics of the first-stage regressions are about 41. Thus, weak instruments are not an issue.

Figure 3 shows the estimated values of the vector β_h, i.e. the estimated effects of TLTRO-II in various months for different types of credit. The solid lines represent the point estimates, and the dashed lines 90 per cent confidence intervals. Appendix A provides some more information about the model. In every month before June 2016, the estimated effects do not differ from zero, which supports the common trend assumption. The effect on corporate loans is positive and statistically significant. F-statistic for the joint significance of interactions from July 2016 to July 2018 is 2.9 ($p = 0.001$). The cumulative effect of TLTRO-II on participating banks' corporate lending is estimated to exceed 20 per cent. Instead, the estimated effect on loans for consumption is actually negative, though not statistically significantly. F-statistic for the joint significance of interactions from July 2016 to July 2018 is 0.6 ($p = 0.935$). This is surprising as TLTROs were also targeted on loans for consumption.

In the sample, the banks that took up TLTRO-II had lent about 50 per cent of the outstanding corporate loans in June 2016. If this share could be generalised to the whole population and if TLTRO-II did not affect to the banks that did not participate, it would mean that TLTRO-II increased the total stock of corporate credit about 10 per cent cumulatively from June 2016 to July 2018.

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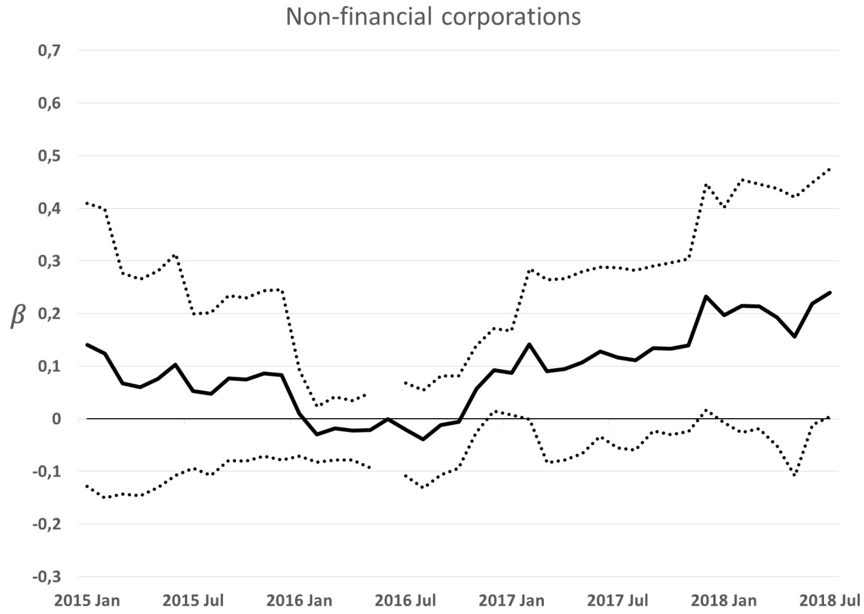
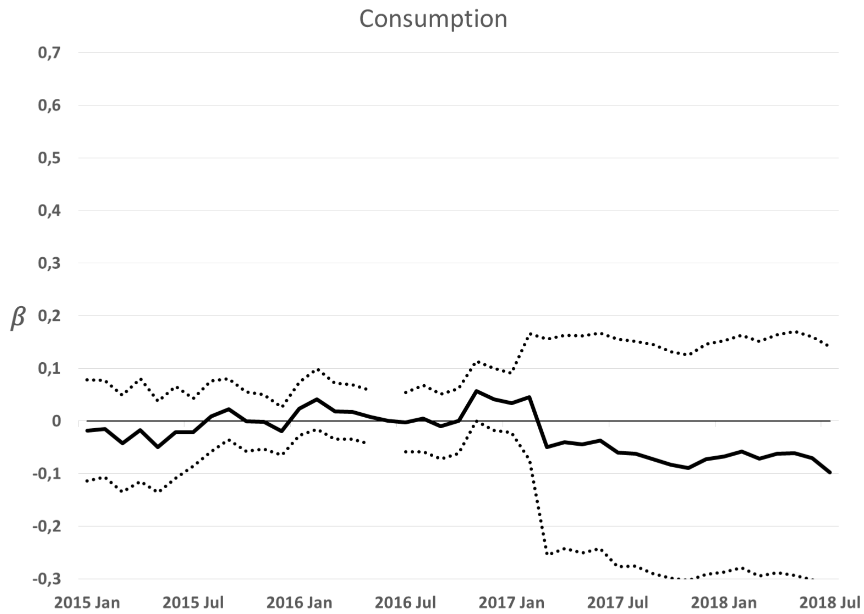
Figure 3
The estimated effects of TLTRO-II (parameters in vector β) on different types of credit





Note: The dashed lines represent 90 per cent confidence intervals. Standard errors are clustered at bank and month level. The share of central bank credit in total liabilities prior TLTRO-II is used as an instrument for participation in TLTRO-II.

Source: Author's calculation.

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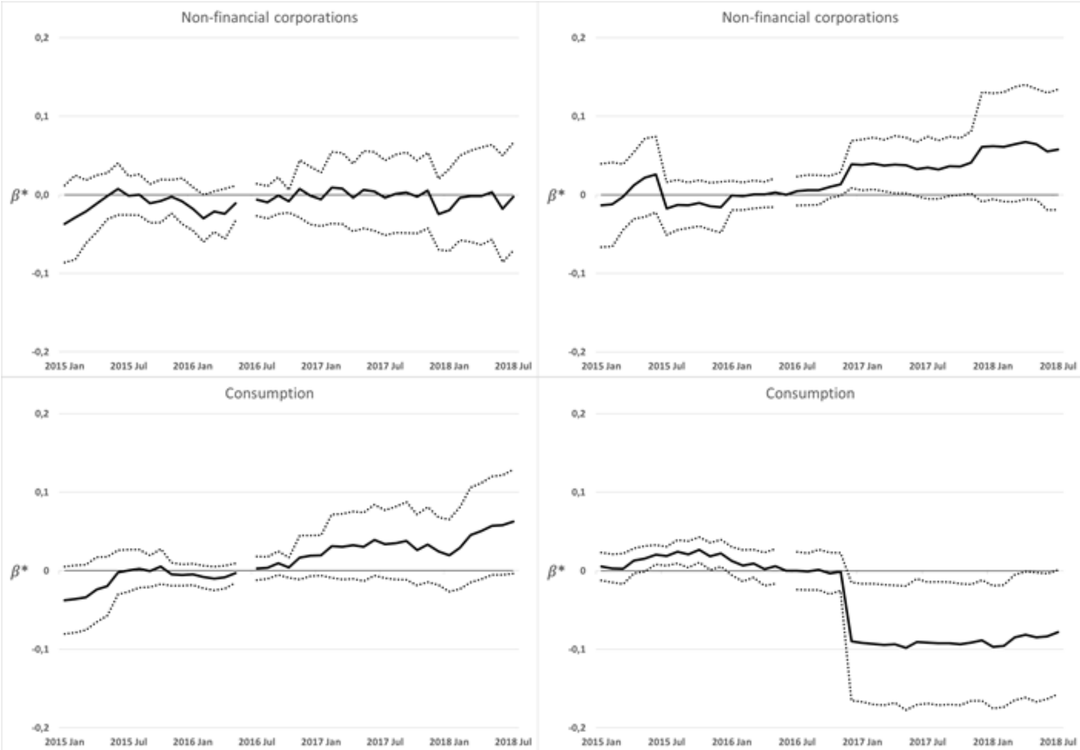
Figure 4
The estimated effects of the amount of TLTRO-II (parameters in vector β^*) on different types of credit



Note: The dashed lines represent 90 per cent confidence intervals. Standard errors are clustered at bank and month level. On left, the (log) amount of central bank credit prior TLTRO-II is used as an instrument for the (log) total borrowing in TLTRO-II. On right, the (log) amount of eligible loans in January 2016 is used as an instrument for the (log) total borrowing in TLTRO-II.

Source: Author's calculation.

4.2. The amount of TLTRO-II

So far, we have only considered the effects of a decision to participate in TLTRO-II. However, one might expect that the more a bank borrowed from the central bank, the more it increased its lending to non-financial corporations and to households for consumption. This kind of relationship is quite challenging to observe (see Figure 2). The correlation between TLTRO-II borrowing and growth in lending to non-financial corporations is practically zero (Pearson correlation is -0.02 and it is clearly insignificant).

To further asses this relationship, we drop all banks that did not participate in TLTRO-II from the baseline analysis (entire control group) and add the natural logarithm of total TLTRO-II into Eq. (2). In other words, we analyse only the banks that participated in TLTRO-II (97 banks) and group them by their TLTRO-II amounts. We instrument the (log) total take-up in TLTRO-II by the (log) amount of central bank credit in May 2016. Additionally, we use the (log) amount of eligible loans in January 2016 as an alternative instrumental variable.

Figure 4 shows the estimated effects. The estimates on the left-hand side are based on the amount of central bank credit in May 2016 and the estimates on the right-hand side are based on the amount of eligible loans in January 2016. The estimates based on eligible loans suggest that the allotted amount of TLTRO-II had an impact on bank lending. Instead, the estimates that are based on the amount of central bank credit are insignificant. The values of F-statistics for these two alternative instruments are about 14 and 269. Thus, assuming that both instrumental variables are valid, one should give more weight to the results based on the stronger instrument: amount of eligible loans in January 2016.

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The results are potentially unintuitive and puzzling, but there are also some good reasons for the conclusion that the amount of TLTRO did not matter so much. As was explained in Subsection 3.1, the banks were expected to achieve a certain threshold for their bank lending to receive lower interest rate. Therefore, banks with high TLTRO take-ups could use the part of TLTRO to something else than eligible lending without losing the low interest rate.

4.3. Cross-country differences

There are large cross-country differences when it comes to the state of banking sector or economic conditions. Therefore, it is likely that the effects of TLTRO-II were different in different countries. For example, Albertazzi, Nobili and Signoretti (2021) observe that the transmission of conventional monetary policy is stronger for weaker banks. However, their results suggest that when it comes to unconventional monetary policy, the transmission is stronger among strong banks. Boeckx, de Sola Perea and Peersman (2020) find some evidence in favour of the opposite conclusion. Thus, the literature regarding the bank lendig channel of uncoventional monetary policy tools is rather mixed. In addition there may some other reasons, why monetary policy may have different effects in different countries. More generally, the cross-country differences in the effects of monetary policy has been studied by Burriel and Galesi (2018). They find that countries with more fragile banking systems benefit the least from unconventional monetary policy measures.

To assess this question, we calculate a dummy variable that equals 1 if the bank's home country is Spain, Italy, Greece or Portugal. These countries form a group that we call "crisis countries". We replace the interactions $D_h \cdot TLTRO_{ci}$ in Eq. (1) by interactions $crisis_c \cdot D_h \cdot TLTRO_{ci}$. This means that the treatment group consists of the banks that participated in TLTRO-II and were located in the crisis countries. Otherwise, model specification and estimation are as in the baseline analysis.

The coefficient estimates are shown in Figure 5. The results hint that the effect on bank lending has been stronger in the crisis countries than elsewhere. This result indicates that it is problematic to generalise results obtained from a single country to euro area level. However, this issue requires more research. It is not clear, what is the underlying reason for heterogeneous effects. One potential reason is the state of the banking sector, but deeper analysis regarding this topic is left for the future research.

4.4. Effect on sovereign bond purchases

Crosignani et al. (2020) find that a large part of VLTROs went to buying sovereign bonds in Portugal. The fact that VLTROs was used to buy bonds in crisis countries was possibly one reason why the ECB chose to target its TLTROs. In principle, TLTROs create an incentive to replace government bonds by eligible loans. However, as discussed in the previous sections, the banks had to achieve a certain lending threshold, after which they were rewarded with a lower rate by the ECB. After achieving this threshold, the incentives in favour of eligible lending disappear. Therefore, the effect of TLTROs on sovereign bond purchases is ambiguous.

To investigate whether targeting worked as intended, we estimate the Eq. (1) as in the baseline analysis, but use the natural logarithm of sovereign bond holdings as a dependent variable and keep the treatment as in the previous section. The results are shown in Figure 6.

The results suggest that TLTRO-II worked as intended. TLTRO-II did not increase government bond holdings. Instead, the operations seem to have had a negative effect. However, the reason for this result is not necessarily the design of TLTRO-II. The different effect from Crosignani et al. (2020) might be driven, for example, by different macroeconomic conditions.

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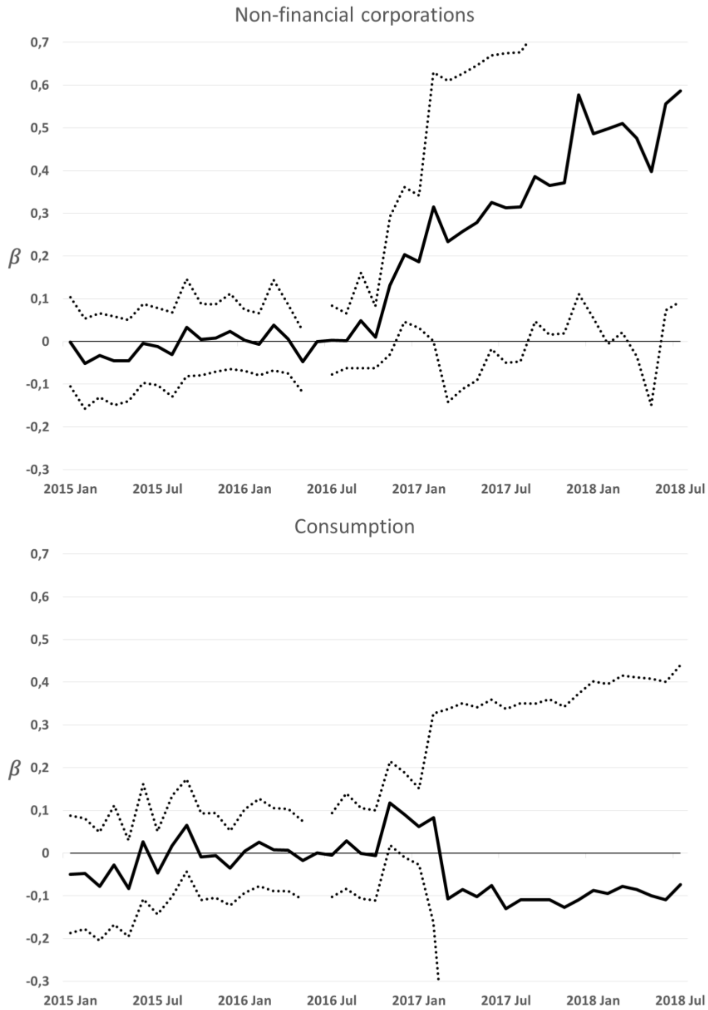
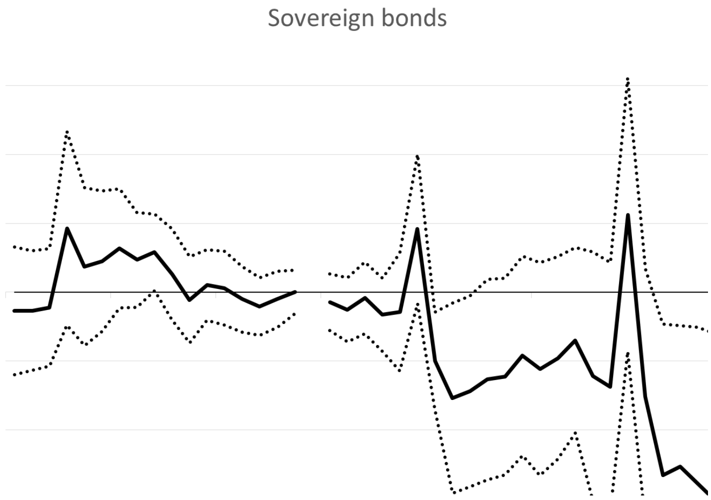
Figure 5
The estimated effects of TLTRO-II in crisis countries on different types of credit



Note: The treatment group consists of the banks that participated in TLTRO-II and are located in Spain, Italy, Greece or Portugal. The dashed lines represent 90 per cent confidence intervals. Standard errors are clustered at bank and month level. The share of central bank credit in total liabilities prior TLTRO-II is used as an instrument for participation in TLTRO-II.

Source: Author's calculation.

Figure 6
The effect of TLTRO-II on sovereign bond holdings



Note: The endogenous variable is (log) government bond holdings. The treatment group consists of the banks that participated in TLTRO-II and are located in Spain, Italy, Greece or Portugal. The dashed lines represent 90 per cent confidence intervals. Standard errors are clustered at bank and month level. The share of central bank credit in total liabilities prior TLTRO-II is used as an instrument for participation in TLTRO-II.

Source: Author's calculation.

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Table 3
Logit model used in the propensity score matching

	Participation in TLTRO-II
Predictors	**Log-Odds**
(Intercept)	-3.30*
Dlog(Loans to non-financial corporations)	-0.96
Dlog(loans for house purchase)	-16.58
Dlog(loans for consumption)	1.97
log(Balance sheet)	0.33**
Cash to total assets	-7.72
Household deposits to total liabilities	-0.91
Equity ratio	2.82
Observations	187
R^2 Tjur	0.100

* $p < 0.05$ ** $p < 0.01$ *** $p < 0.001$

Note: The used variables are calculated from bank-level January 2015 to May 2016 averages, i.e. before TLTRO-II.

Source: Author's calculation.

4.5. Robustness

As was shown earlier, the banks included in the sample were quite heterogeneous, for example, when it comes to their size (Table 2). With a perfect instrumental variable, this should not bias the results. However, one may always argue that the used instrumental variables are not valid, and there may be selection bias present. To analyse, if the results are driven by the differences in the treatment and control group, we use propensity score matching. Specifically, we estimate a logit model that predicts the participation in TLTRO-II based on banks' observable characteristics before TLTRO-II. Thereafter, the banks that borrowed in TLTRO-II are matched with other banks based on their estimated likelihood to participate using nearest-neighbour algorithm with replacement and calliper of 0.1.

In the logit model, we include such variables that could potentially affect the participation decision. Specifically, we include average growth rates of different types of lending before TLTRO-II. It is possible that such banks that were already increasing their lending self-selected into TLTRO-II because they believed that continuing increasing lending would be easy. On the other hand, it also possible that banks that were doing poorly self-selected into TLTRO-II, because they were unable to receive market-based funding. In addition, choosing loan growth variables makes the common trends assumption more reliable: we choose such banks that shared the common trend in loan growth. We also include the average size of the banks before TLTRO-II as the participating banks were much larger than the others. Additionally, we consider the share of cash, share of household deposits and equity ratio. The estimated logit model is reported in Table 3, and Figure 7 shows the results from the propensity score matching. The matching drops 7 banks from the treatment group (participants) and 40 from the control group (non-participants). The results show that it is rather difficult to find observable variables that could explain the participation decision. In other words, based on the observable variables, it is difficult to argue that selection bias plays a significant role.

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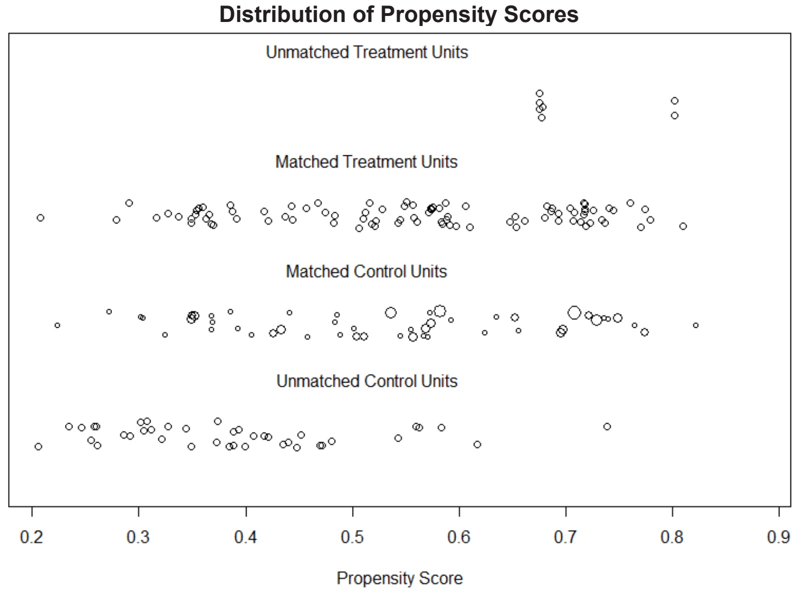
Figure 7
Results after using loans for house purchase as a proxy for credit demand



Note: The matching is done using nearest neighbour algorithm with replacement and 0.1 calliper. 7 banks are dropped from the treatment group and 40 from the control group. Thus, the final sample consists of 90 TLTRO banks and 50 other banks.

Source: Author's calculation.

The results from the baseline analysis with this subsample of banks is shown in Figure 8. The results remain roughly the same. Actually, the positive effect on corporate lending is now even more clearly statistically significant. The effect on lending for consumption is still close to zero and statistically insignificant. Therefore, it is difficult to argue that the results were biased downwards or upwards due to self-selection.

Another potential issue that may affect the results is demand. There are many ways to control for credit demand, and so far, we have used country-time fixed effects only. Figure 9 shows the results when country-time fixed effects are replaced by time fixed effects and $\log(\textit{Loans for house purchase}_{ict})$ is used as a control variable. Because TLTROs were targeted on loans to households excluding loans for house purchases, it is likely that the variation in the stock of mortgages reflects mainly variation in loan demand. The estimation is done using the full sample. This modification lowers the estimate for the effect on corporate lending a bit. The estimated effect on lending for consumption is still close to zero and statistically insignificant. Therefore, our results seem not to depend on the chosen way of controlling for credit demand.

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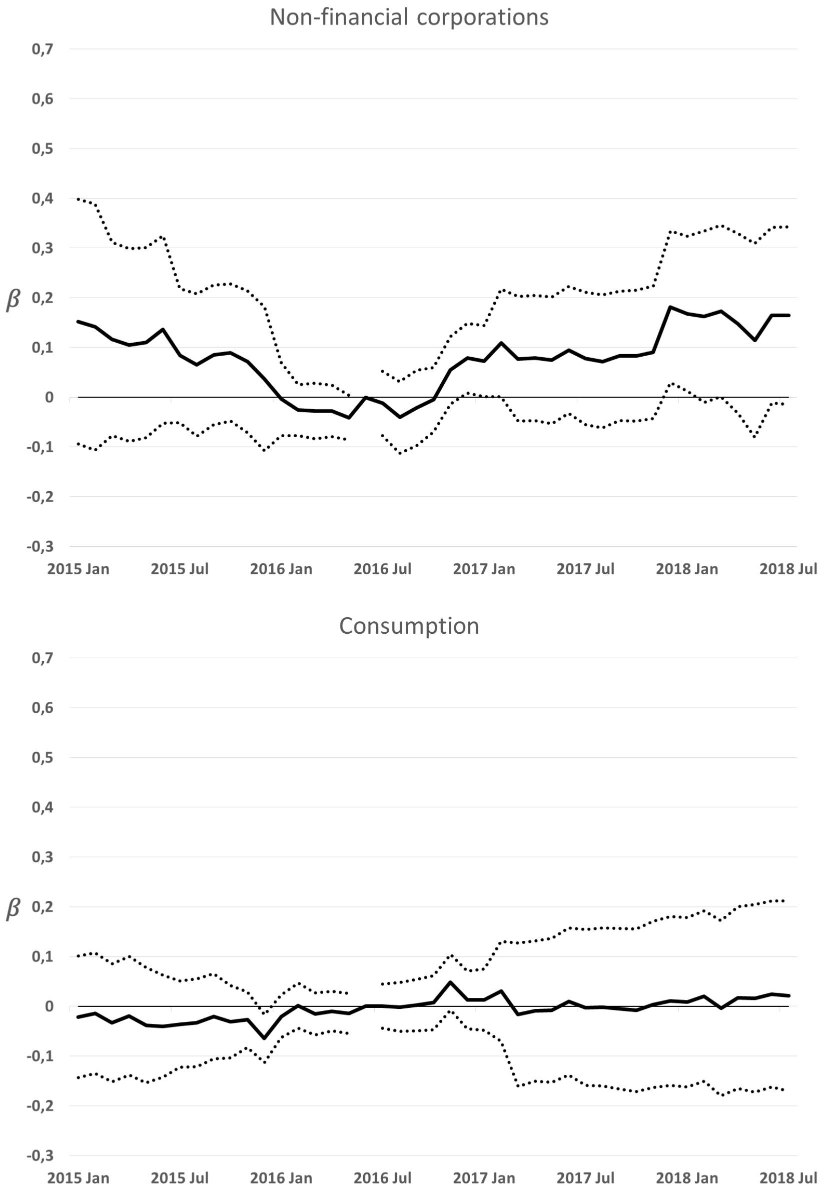
Figure 8
Results after propensity score matching



Note: The dashed lines represent 90 per cent confidence intervals. Standard errors are clustered at bank and month level. The share of central bank credit in total liabilities prior TLTRO-II is used as an instrument for participation in TLTRO-II.

Source: Author's calculation.

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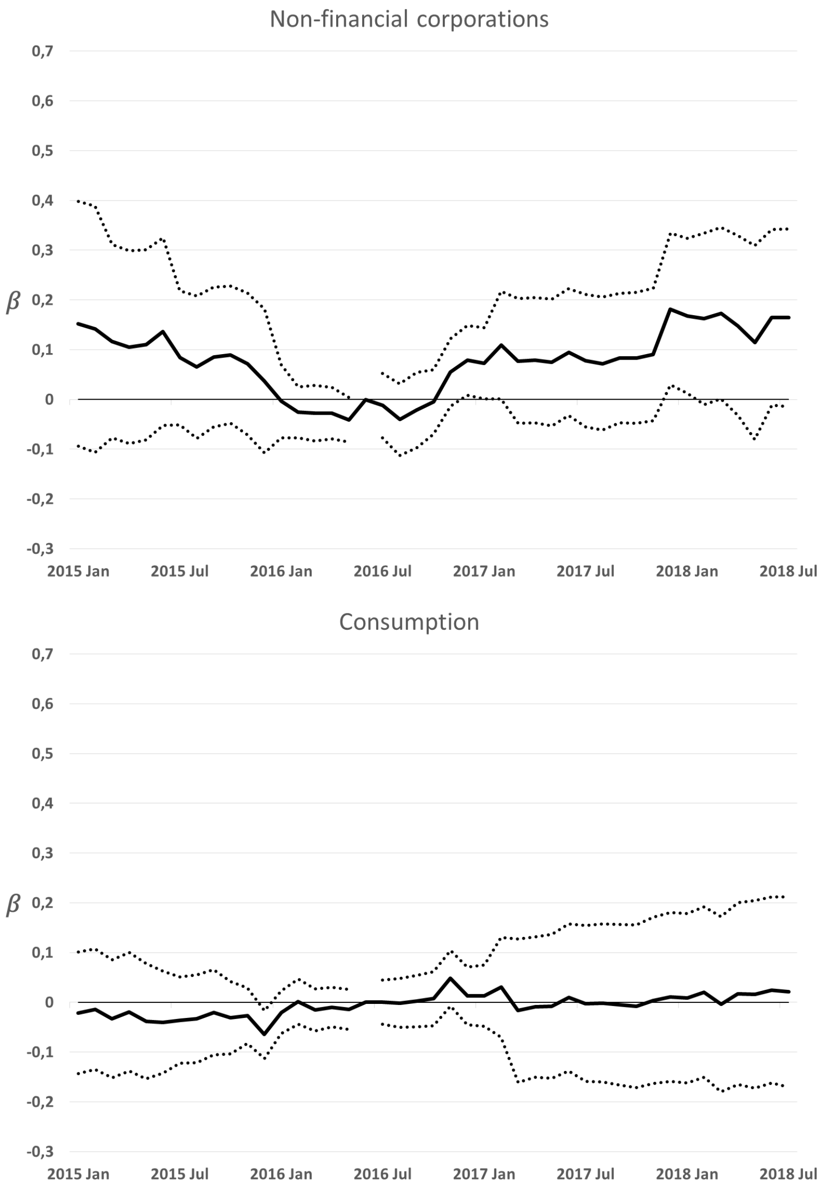
Figure 9
Results after using loans for house purchase as a proxy for credit demand



Note: The dashed lines represent 90 per cent confidence intervals. Standard errors are clustered at bank and month level. The share of central bank credit in total liabilities prior TLTRO-II is used as an instrument for participation in TLTRO-II.

Source: Author's calculation.

5. CONSLUSIONS

The results show that the effect of TLTRO-II on bank lending was positive. In particular, TLTRO-II boosted credit to non-financial corporations, while the effect on loans for consumption is estimated to be close to zero and statistically insignificant. This result is surprising as TLTRO-II was targeted equally at both consumption lending and corporate lending. Because the zero effect was unexpected, we do not have any obvious theoretical explanation for this in our mind. The explanation for the result might be related to, for example, differences in market power in different credit markets (Benetton and Fantino, 2021). One of the usual suspects for strange results in this field is the way loan demand is controlled for. In the baseline analysis, we use country-time fixed effects. This technique has its drawbacks, and therefore we assess the robustness of the results by controlling credit demand using loans to households for house purchase, which is

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excluded from the eligible lending and thus a good proxy for credit demand (especially regarding households). This alternative way of controlling for credit demand does not change the results. Another issue that might drive the results is the fact that the banks were rather heterogeneous before the treatment. If one had a perfect instrumental variable, this fact should not affect the results. Because the validity of instrumental variables is in the end a matter of argumentation, and one might maybe argue that central bank borrowing prior the treatment is not necessarily a perfectly valid instrument, we consider also propensity score matching as a supplementary technique for tackling potential selection bias. Controlling for potential variables that might explain the selection to the treatment does not change the results.

The results also suggest that the effects of TLTROs have not been the same in all the countries. This is not surprising as there are many papers that show that the effects of monetary policy are different in different countries. However, the result is important, because the earlier studies that analyse the effects of longer-term refinancing operations with microdata focus on single countries. Thus, these results are difficult to generalise to other countries. According to the results, the effects have been strongest in countries most affected by the crisis.

The results show as well that TLTRO-II did not increase the government bond purchases of the participating banks in crisis countries. Thus, the effect of TLTRO-II was quite different from the effect of the VLTROs (see Crosignani et al. 2020) and suggests that the targeting of credit operations mattered.

Though this paper has covered many open questions related to the targeted monetary policy, there are certainly many questions that should be answered in the future research. One shortage in the current literature is that it is mainly empirical. As the targeted tools are becoming more and more “conventional” in the central banks’ toolboxes, it would be necessary to understand better how and why these tools work.

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APPENDIX

Table A1
Some key information about the baseline regressions

	Dependent variable:	
	Log(Loans to NFCs)	**Log(Loans for consumption)**
	(1)	(2)
log(Total Assets)	0.915**	0.914**
	(0.202)	(0.242)
Equity ratio	0.085	2.368*
	(0.886)	(1.186)
Observations	8,041	8,041
R^2	0.990	0.987
Adjusted R^2	0.989	0.986
Residual Std. Error (df = 7054)	0.191	0.213

* $p < 0.05$ ** $p < 0.01$

Note: The interactions are reported in Figure 3. Standard errors are clustered at bank and month level.

Source: Author's calculation.

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The effect of the ECB's conventional monetary policy on the real economy: FAVAR-approach

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Abstract

This study applies factor-augmented vector autoregressive models to investigate the effect of the European Central Bank's (ECB) conventional monetary policy on the real economy. More specifically, the study examines how unanticipated changes in the ECB's policy rate have affected unemployment rate and industrial production. The effect of monetary policy on unemployment rate and industrial production is estimated to be strong and statistically significant using the data from January 1999 to July 2017 or from the pre-crisis period. However, after the beginning of the crisis the responses weaken drastically and become sometimes statistically insignificant, indicating that the effect of the ECB's conventional monetary policy became weaker after the financial crisis. This finding is extremely interesting because one could presume either weaker or stronger effect based on economic theory. Additionally, the previous studies that have analysed the possible changes in the monetary policy effectiveness in the euro area have not found any changes (e.g. Bagzibagli in Empir Econ 47(3):781–823, 2014; Von Borstel et al. in Int J Money Finance 68:386–402, 2016).

Keywords Monetary policy · Real economy · FAVAR · Low interest rates · Financial crisis

JEL Classification E52 · E58 · E6

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1 Introduction

The financial crisis of 2008 was followed by a remarkable decline in nominal interest rates globally. In the euro area, the European Central Bank (ECB) lowered its policy rate first from 4.25 to 1.00. After dawning economic recovery and accelerating inflation, the rate was raised to 1.50 where it stayed only a little while before it was declined to zero after the escalation of the European debt crisis.

There is a great amount of literature concerning, for example the optimal monetary policy in zero lower bound (ZLB) or the effects of unconventional monetary policy in ZLB. However, there are surprisingly few papers that investigate the effect of conventional monetary policy when nominal rates are close to zero. According to Keynes (1936), the effectiveness of monetary policy diminishes as nominal interest rates approach to zero. That is, there are nonlinearities present.

In this research, I investigate the effect of the ECB's conventional monetary policy on the real economy. Specifically, I examine whether the effect has been weaker during the low interest rate period. The euro area is particularly interesting subject of research because the policy rates have been both raised and declined during the low rates period. To analyse the possible change in the effectiveness of monetary policy, I apply factor-augmented vector autoregressive models (FAVAR models) proposed by Bernanke et al. (2005). FAVAR models have many advantages compared to traditional VAR models. The main advantage is that a large amount of information can be included in the model. Traditional VAR models typically include no more than six to eight variables because the number of parameters to be estimated increases rapidly too high. FAVAR models include typically dozens or even hundreds of variables. It is, therefore, possible to estimate the effect of monetary policy on a large number of macroeconomic variables. In addition, the large information set makes identification of monetary policy shock more reliable as central banks observe literally hundreds of time series in reality.

The results of this study can be summarised as follows. The effect of conventional monetary policy on the real economy is found to be in line with the previous studies. Yet, the effect of conventional monetary policy weakened drastically or came even impotent after the ECB's policy rate was lowered to 1.00 in May 2009. The results contradict the earlier literature concerning the effects of the ECB's conventional monetary policy (Bagzibagli 2014; Von Borstel et al. 2016). The results are consistent, for example, with the results by Cenesizoglu et al. (2018) and Liu et al. (2019). They find similar kind of results in the USA.

Thus, the most important contribution of this article is to find evidence of the impact of the ECB's conventional monetary policy, which contradicts the previous literature. On the other hand, the results support evidence from the USA. The paper investigates, in addition, the effects during the pre-crisis period. The effects before the crisis are found to be similar to those of Bagzibagli (2014). This further strengthens the argument that the effects probably changed after the crisis. More broadly, the paper contributes to the literature concerning the time variation in the effects of macroeconomic shocks (e.g. Cogley and Sargent 2005; Primiceri 2005; Boivin et al. 2010; Korobilis 2013; Mumtaz and Zanetti 2015).

In the light of economic theory, the results can be seen either expected or surprising. There are at least three reasons to presume that the effect of conventional monetary policy has been weaker in the euro area after the financial crisis. However, the same three matters can be used as arguments for a stronger impact.

First, the effect of monetary policy may be weaker when nominal interest rates are low. One reason for this is the speculative demand for money as was proposed by Keynes (1936). In addition, low interest rates may have a negative impact on banks' profits (e.g. Borio et al. 2017). This in turn may reduce loan supply and weaken the effectiveness of expansionary monetary policy (Borio and Gambacorta 2017). Additionally, lowering policy rates to unforeseen levels may be seen as "Delphic", meaning that market participants believe that the central bank has lowered the rates because it expects economic situation to worsen in the future (see Campbell et al. 2012). Nevertheless, there are reasons to believe that monetary policy could have been very effective when nominal rates have been low. There is evidence that the natural rate of interest has declined considerably (e.g. Holston et al. 2017). If the natural rate was very low as the ECB raised its policy rate from 1.00 to 1.50 in 2011, one might expect that this hike would have had a more negative effect on the real economy than during previous years when the natural rate was probably higher.

Second, the problems of asymmetric information typically worsen during crisis periods (e.g. Mishkin 1990). Bernanke (1983), for example, proposes that increasing uncertainty makes people await more information and postpone investment decisions. The real economy, therefore, does not respond to monetary policy as in normal times. This proposition is supported by the results of Aastveit et al. (2013). On the other hand, Mishkin (2009) argues that the effect of monetary policy could actually be stronger during crisis periods because then its effect on risk premia is stronger.

Third, the financial intermediation was impaired in the euro area after the financial crisis. As the financial intermediaries play a crucial role in the transmission of monetary policy, one could think that broken banking system would weaken the effect of monetary policy (e.g. Diamond 1984). However, the bank lending channel of monetary policy might be especially strong when banking sector is weak because an increase in asymmetric information may increase the sensitivity of lending supply to changes in monetary policy (e.g. Albertazzi et al. 2016; Holton and McCann 2017).

When it comes to the empirical research concerning the euro area, there are few papers that investigate the possible change in the effectiveness of conventional monetary policy. Bagzibagli (2014) applies FAVAR models to examine whether the transmission of conventional monetary policy changed during the financial crisis. He concludes that the transmission has probably not changed as the impulse response functions are very similar before and after the beginning of the crisis. The problem in this study is the short data. Bagzibagli's (2014) last observation is in the end of 2011. Thus, the study does not concern the period during which the ECB's policy rate has been low for a long period of time.

Another interesting research is made by Von Borstel et al. (2016). They investigate whether the monetary policy transmission to nominal interest rates changed after the financial crisis using FAVAR models. According to their results, the effect of conventional monetary policy on interest rates remained roughly the same. However, they do not analyse the possible change in the responses of real variables such as unemployment and industrial production.

The remainder of the paper is as follows. Section 2 represents the FAVAR model. Section 3 describes the data. Section 4 analyses the results. Section 5 concludes.

2 Model

The model closely follows Bernanke et al. (2005). Let Y_t denote $Mx1$ vector containing observable variables. Typically, Y_t contains the policy instrument of the central bank and possibly some other economic variables that are assumed to be observable. Let F_t denote $Kx1$ vector that contains unobservable factors that represent abstract phenomena such as economic activity or confidence. These phenomena are impossible to observe through some single indicator. Together vectors Y_t and F_t form the following model:

$$\begin{bmatrix} F_t \\ Y_t \end{bmatrix} = \Theta^*(L)\begin{bmatrix} F_{t-1} \\ Y_{t-1} \end{bmatrix} + v_t, \tag{1}$$

where $\Theta^*(L)$ is a matrix of finite lag polynomials. The number of lags in the model is d, so the lag polynomials are order $d-1$. The symbol v_t denotes a vector containing error terms that are assumed to have mean zero and covariance matrix Q. Equation (1) is referred as a FAVAR model. The model cannot be estimated directly because the factors F_t are unobservable. However, these factors can be estimated from a large number of relevant time series. These time series are denoted by the $Nx1$ vector X_t. The time series X_t also contain the variables in Y_t. The relation between these time series, factors F_t and the observable variables Y_t is summarised by the equation:

$$X_t = \lambda^f F_t + \lambda^y Y_t + e_t, \tag{2}$$

where the matrix λ^f is NxK and the matrix λ^y is NxM. The matrix λ^f contains so-called factor loadings. In factor analysis, it is typical to use some rotation to make it easier to interpret the results. Here, the factor loadings are just unrestricted regression coefficients that are estimated after the estimation of factors. Similar method is used by Von Borstel et al. (2016). The vector e_t is $Nx1$ that contains error terms that are assumed to be mean zero but may display some small degree of cross-correlation.

When it comes to the estimation of the FAVAR model, there are basically two different methods. The first one is one-step Bayesian method and the second one is two-step method that applies principal component analysis. Bernanke et al. (2005) find the methods equally good. Thus, I apply computationally easier two-step method. In the first step, the factors F_t are estimated using principal component analysis. In the second step, F_t in Eq. (1) is replaced by the estimate $\hat{F}_t$. Thereafter, Eq. (1) is estimated using OLS.

It is assumed that the time series X_t can be divided into fast-moving and slow-moving variables. The fast-moving variables are assumed to respond contemporaneously to unanticipated changes in monetary policy. The slow-moving variables are assumed not to respond to monetary policy shocks during the same period. In practice, fast-moving variables are assumed to be, for example, asset prices and the slow-moving variables are mainly real variables like industrial production and unemployment rate.

The first step has two stages. In the first stage, principal components are estimated both from the slow-moving variables and from all of the variables. Principal component analysis is applied to correlation matrix as the variables have different scales. Another possibility would be covariance matrix. The first K principal components estimated from all the time series are denoted by the $Kx1$ vector $\hat{C}(F_t, Y_t)$, and the first K principal components of the slow-moving variables are denoted by the $Kx1$ vector $\hat{C}^*(F_t)$. In the second stage of the first step, the effect of the observable variables Y_t is purged from the principal components $\hat{C}(F_t, Y_t)$. This is carried out by estimating the equation:

$$\widehat{C_k}(F_t, Y_t) = a_k \widehat{C_k^*}(F_t) + b_k' Y_t + u_{kt}, \tag{3}$$

where a_k is the regression coefficient of the kth slow-moving principal component, b_k is a vector containing the regression coefficients of the observable variables Y_t and u_{kt} is an error term. That is to say, each principal component estimated from all the time series X_t is explained by the corresponding slow-moving principal component and by all the observable variables Y_t. The equation is estimated for all the K principal components using OLS. Thereafter, it is straightforward to calculate the estimate for the vector of factors F_t:

$$\hat{F}_{kt} = \widehat{C_k}(F_t, Y_t) - b_k' Y_t = a_k \widehat{C_k^*}(F_t) + u_{kt}. \tag{4}$$

In the second step, F_t in Eq. (1) is replaced by the estimate $\hat{F}_t$ and the equation is estimated using OLS like a standard VAR model.

In further analysis, the effect of monetary policy is investigated by examining impulse response functions. The impulse response functions can be calculated for FAVAR models like for VAR models. The monetary policy shock is identified using Cholesky decomposition. Cholesky decomposition is chosen as many other identification strategies require some of the factors to be identified as specific economic concepts like output gap. In the baseline model, the ECB's policy rate (MRO) is ordered last which means that the ECB's total assets/liabilities, inflation and all the factors are assumed to have a contemporaneous effect on the MRO. The total assets/liabilities is ordered second last and inflation third last. The FAVAR model can be written in structural form:

$$A \begin{bmatrix} F_t \\ Y_t \end{bmatrix} = \Psi^*(L) \begin{bmatrix} F_{t-1} \\ Y_{t-1} \end{bmatrix} + \varepsilon_t, \tag{5}$$

where A is a matrix of coefficients, $\Psi^*(L)$ is a matrix of finite lag polynomials and ε_t is the vector of structural shocks. The equation can also be represented in a vector moving average form:

$$\begin{bmatrix} F_t \\ Y_t \end{bmatrix} = \Psi(L)^{-1} \varepsilon_t \tag{6}$$

where $\Psi(L)^{-1}$ is a matrix of infinite lag polynomials. The impulse response functions for all the time series X_t can be calculated as:

$$X_t^{irf} = \lambda^f F_t + \lambda^y Y_t = \left[\lambda^f \lambda^y\right] \Psi(L)^{-1} \varepsilon_t. \quad (7)$$

The estimates for λ^f and λ^y are obtained by estimating Eq. (2) using OLS. To demonstrate the uncertainty of the estimates, confidence intervals are estimated following the method proposed by Yamamoto (2012). The method takes into account the uncertainty related to the estimation of factors.

3 Data

The data are mainly from Eurostat and the ECB. Other sources are MSCI, the Bank of Japan, OECD and the Bureau of Labor Statistics. The data include 90 monthly time series from January 1999 to July 2017. All the variables, their source and possible transformation are listed in "Appendix A". Most of the variables are seasonally adjusted.

Some studies, for example Soares (2013), use disaggregated quarterly data to increase the information set. However, disaggregation is always somewhat uncertain. In addition, many quarterly series are published with a considerable lag. Thus, it is not very realistic to assume that these data are always part of the ECB's governing council's information set as the council conducts monetary policy.

When it comes to the euro area as an entity, one needs to consider what the euro area actually is. In 1999, the euro area consisted of 11 countries, but the number of countries has increased to 19. It would be best to consider only the original countries. Unfortunately, the data are rarely available to this set of countries. The majority of the variables are, therefore, calculated for the current euro area (see "Appendix A"). However, this is hardly a problem in this analysis as the eight countries that joined the euro after 1999 joined quite early (Greece 2001, Slovenia 2007, Cyprus 2008, Malta 2008, Slovakia 2009, Estonia 2011, Latvia 2014, Lithuania 2015).

4 Results

4.1 The estimation of factors and model specification

Figure 1 shows the total variance explained by the first 10 principal components that are estimated from all the 90 variables. The first principal component explains 24 per cent of the total variance. Together all the 10 principal components explain 75 per cent of the total variation in the 90 variables. It is not unambiguous how many principal components should be used in the FAVAR model. Every principal component adds more information to the model, but on the other hand the idea of principal component analysis is to reduce the dimensions of the data.

There are some techniques to evaluate the optimal number of principal components. I apply two information criteria (IC1 and IC2) proposed by Bai and Ng (2002).

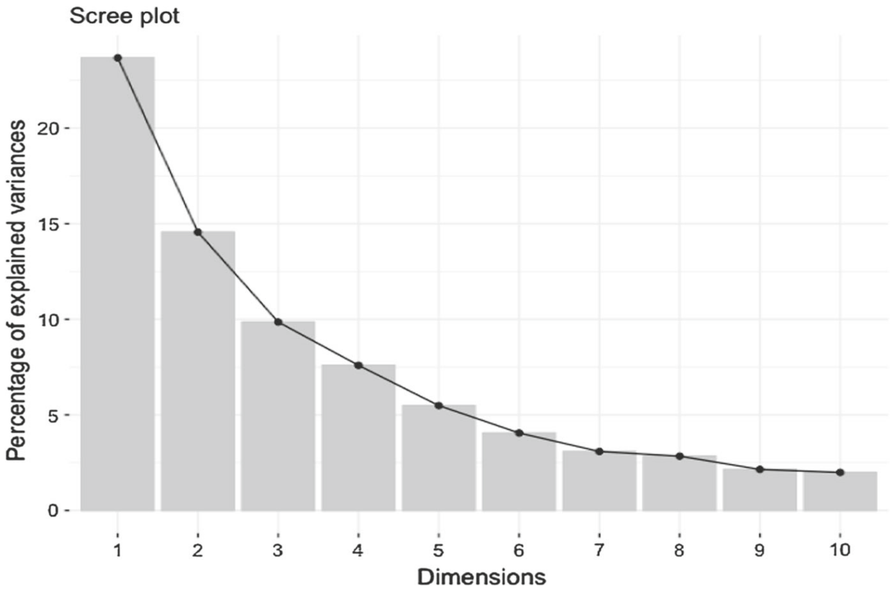

Fig. 1 Variance explained by the first 10 principal components

Table 1 The values of different information criteria in different models

Model	Total variance explained (%)	AIC	FPE	SC	HQ	IC1	IC2
FAVAR (5 factors, 3 lags)	64	– 25.21	1.14e–11	– 21.89	– 23.87	13.54	13.91
FAVAR (3 factors, 12 lags)	58	– 22.79	1.53e–10	– 15.71	– 19.93	13.58	13.80
FAVAR (8 factors, 2 lags)	73	– 25.54	8.16e–12	– 21.33	– 23.84	13.67	14.28
FAVAR (3 factors, 3 lags)	58	– 23.42	6.76e–11	– 21.51	– 22.65	13.58	13.80
FAVAR (1 factor, 3 lags)	37	– 23.47	6.43e–11	– 22.58	– 23.11	14.12	14.20

AIC Aikake information criterion, *FPE* final prediction error, *SC* Schwarz criterion, *HQ* Hannan–Quinn criterion. *IC1* and *IC2* two information criteria proposed by Bai and Ng (2002)

In addition, I estimate the FAVAR model using many different specifications and evaluate the goodness of these models using traditional information criteria used in the VAR literature (AIC, FPE, SC and HQ). Some examples of the results are shown in Table 1. In all the models, I assume that the observable variables, Y_t, are inflation (HICP, YoY, %), the change in the natural logarithm of the total assets/liabilities of the Eurosystem and the MRO. The Eurosystem's total assets/liabilities are included to control unconventional monetary policy. Inflation is included as it is the main objective variable of the ECB and a key determinant of the stance of monetary policy. The models are estimated using the whole data from January 1999 to July 2017. All the models include constant and deterministic trend.

Based on these results, I use the FAVAR model with 5 factors and 3 lags as a baseline model when evaluating the effect of monetary policy using the whole data. As I analyse

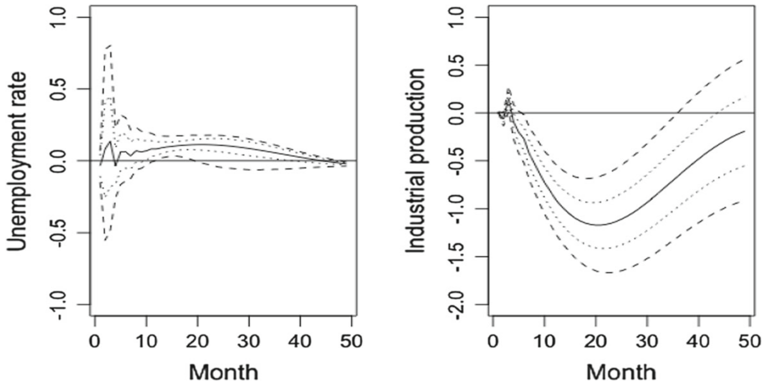

Fig. 2 The impulse responses of unemployment rate and industrial production to a 0.25 percentage points shock to the MRO estimated using the data from January 1999 to July 2017. The dashed lines around the impulse response functions represent 95% and dotted lines 68% CI

the possible time variation using only 99–120 observations, I use the model with 3 factors and 3 lags (baseline 2). The number of parameters might otherwise be too large for such a small sample. Nevertheless, I test the robustness of the results using different number of factors and lags.

4.2 The effect of conventional monetary policy on real variables

Figure 2 shows the impulse response functions of unemployment rate and industrial production to a 0.25 percentage points shock to the MRO (some more responses are shown in "Appendix B"). The estimated model includes the MRO, the ECB's total assets/liabilities, inflation, 5 factors, 3 lags, constant and linear trend. The impulse response functions are in line with previous research (e.g. Soares 2013; Bagzibagli 2014). The 0.25 percentage points shock to the MRO increases unemployment 0.11 percentage points. The reaction peaks after about 2 years. The shock has a negative impact on industrial production. The reaction is at its deepest 1.2 per cent after nearly 2 years.

The effects are quite robust to changes in the number of lags or factors (see "Appendix G"). The inclusion of trend term is not important either (see "Appendix E"). The assumed order of the observable variables is not the key driver of the results either (see "Appendices D and F"). The results are also robust to exclusion of the ECB's total assets/liabilities (see "Appendix C"). The results suggest that the model produces reasonable outcomes that are in line with previous findings. The model is, therefore, a good starting point for analysing whether the effects of monetary policy have changed after the drastic decline in nominal interest rates.

4.3 The effect might have changed

Figure 3 shows the impulse response functions that are estimated using pre-crisis and post-crisis data. The beginning of the crisis is assumed to be in July 2007. The same

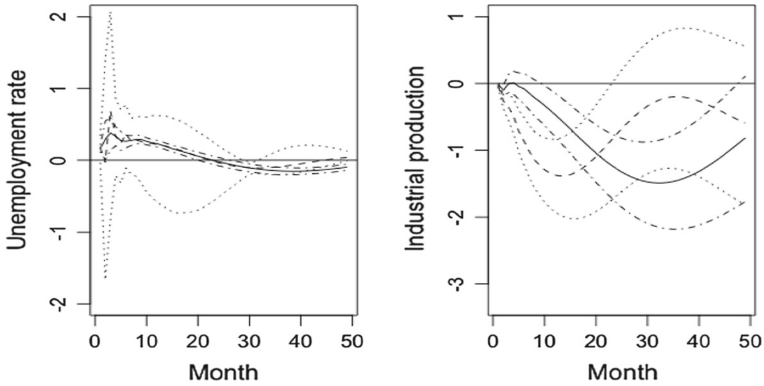

Fig. 3 The impulse responses of unemployment rate and industrial production to a 0.25 percentage points shock to the MRO. The solid line represents the response estimated using data from January 1999 to July 2007. The dashed line represents the response estimated using the data from August 2007 to July 2017. The dotted and dotdashed lines around the impulse response functions represent 95% CI

definition was used by Bagzibagli (2014) who notes that stock market peaked then. Now, the FAVAR model includes only 3 factors and 3 lags. The shock is again 0.25 percentage points.

Industrial production shows no signs of weakened reaction. The magnitudes of the pre- and post-crisis reactions are roughly the same, but after the crisis, industrial production has reacted somewhat faster. Instead, the reaction of unemployment rate becomes statistically insignificant after the crisis.

In July 2007, the MRO was still as high as 4.00 and was even raised to 4.25 in July 2008. Therefore, the period after July 2007 does not represent a period of low interest rates. To examine how the real economy reacted to monetary policy shocks when the policy rate and rates in general were low, I estimate the impulse response functions for unemployment rate and industrial production using the data from May 2009 to July 2017. In July 2009, the sharp decline of the MRO from 4.25 to 1.00 was over. Thereafter, the MRO was both raised and lowered, and it varied between 0.00 and 1.50. Thus, the time interval can be defined as a period of low interest rates.

The impulse response functions estimated from the period of low interest rates are shown in Fig. 4. Now, the reaction of unemployment rate is statistically significant but still very uncertain.

The impulse response function of industrial production is instead statistically insignificant.

The insignificant response of unemployment rate (Fig. 3) and industrial production (Fig. 4) is interesting. The FAVAR model was estimated using many different time intervals, and the responses of both variables were robustly statistically significant when the whole data or the pre-crisis period data are used (see "Appendices C, D, E, F, G"). The argument that the responses of industrial production and unemployment rate changed during or after the crisis is supported by multiple robustness tests. The responses are hardly affected when the number or the order of the observed variables is changed or the trend term is excluded (see "Appendices C, D, E, F"). Instead, the

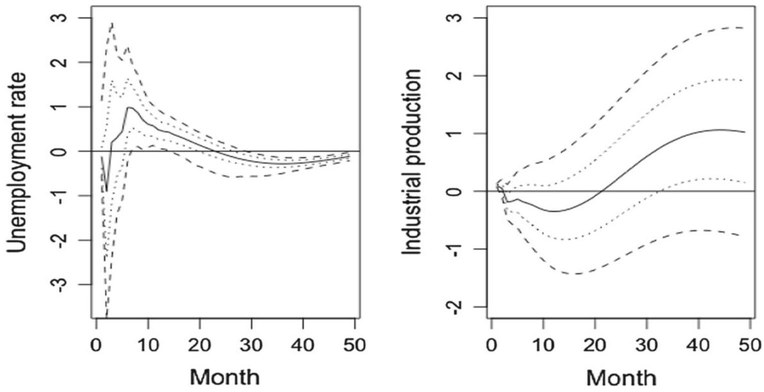

Fig. 4 The impulse responses of unemployment rate and industrial production to a 0.25 percentage points shock to the MRO estimated using the data from May 2009 to July 2017. The dashed lines around the impulse response functions represent 95% and dotted lines 68% CI

reactions vary as the number of lags or factors is changed (see "Appendix H"). For example, reducing the number of factors to 2 makes the response positive. Increasing the number of factors to 4 makes the response negative.

Another issue is the time frame. The last analysis considered only two different periods of time (from July 2007 to July 2017 and from May 2009 to July 2017). What happens if the beginning of the time period was somewhere between July 2007 and May 2009? This is considered in "Appendix I". The alternative starting periods are January 2008, January 2009 and February 2009. Including the whole year 2008 means that the time span covers also maybe the most dramatic months of the financial crisis. During those months, the MRO was also considerably high. In January 2009, the rate was lowered from 2.5 to 2.0. Using the data from February 2009 onwards excludes this rate cut that potentially dominates the results. The results show that the impulse responses of industrial production remained roughly the same before the low interest rate period that began in May 2009. The impulse response functions of unemployment rate remain about the same in every period. However, the confidence intervals are considerably wider than before the crisis in all the chosen time spans.[1] The differing behaviour of industrial production and unemployment rate is interesting and difficult to explain theoretically. Nevertheless, the results clearly show that the effect of conventional monetary policy remained hardly the same after the crisis. This conclusion is opposite to the conclusion made by Bagzibagli (2014, p. 798–799): "First of all, there is little sign of any variation in the real activity measurements such as industrial production, investment and employment. The same conclusion applies to real ULC, nominal wages, producer prices, trade, interest rates, stock market and consumer confidence. That is to say, the monetary policy shocks hitting the economy either before or after the crisis periods have almost identical impacts on these macroeconomic and financial indicators."

[1] Those confidence intervals are not drawn in "Appendix I" as the figure would be too messy. However, the confidence intervals are close to the confidence intervals in Figs. 3 and 4.

5 Conclusions

The results suggest that the transmission of conventional monetary policy to the real economy was weakened after the financial crisis of 2008 in the euro area. The reason for that might be, for example, the low level of nominal interest rates, increased uncertainty or broken banking system. The finding is interesting and policy relevant as the ECB is about to raise its policy rate from zero in some point of time. Conventional monetary policy during low interest rate periods is a surprisingly unknown area which should be examined more—both empirically and theoretically.

The results also support the inclusion of time-varying parameters in FAVAR models (TVP-FAVAR) (e.g. Cogley and Sargent 2005; Primiceri 2005; Korobilis 2013). As the responses of economic variables to shocks vary over time, it is problematic to apply a model with constant parameters.

Appendix A

In the following table, the description (EA) means the changing euro area and (EA19) the current euro area of 19 countries. The description (SCA) means that the time series is both seasonally and working-day adjusted, the description (SA) means seasonal adjustment, only and the description (NA) means that the series is not seasonally nor working-day adjusted. The description (S) means that the variable is assumed to be slow-moving.

Variable	Transformation	Source
Production (volume)		
1. Consumer goods (EA19) (SCA) (S)	Log-difference	Eurostat
2. Durable consumer goods (EA19) (SCA) (S)	Log-difference	Eurostat
3. Non-durable consumer goods (EA19) (SCA) (S)	Log-difference	Eurostat
4. Intermediate goods (EA19) (SCA) (S)	Log-difference	Eurostat
5. Energy (EA19) (SCA) (S)	Log-difference	Eurostat
6. Capital goods (EA19) (SCA) (S)	Log-difference	Eurostat
7. Total excluding construction (EA19) (SCA) (S)	Log-difference	Eurostat
8. Manufacturing (EA19) (SCA) (S)	Log-difference	Eurostat
9. Construction (EA19) (SCA) (S)	Log-difference	Eurostat
Price changes (percentage change year over year)		
10. Manufacturing (EA19) (NA) (S)	No transformation	Eurostat
11. Industry (except construction, sewerage, waste management and remediation activities) (EA19) (NA) (S)	No transformation	Eurostat

Variable	Transformation	Source
12. Capital goods (EA19) (NA) (S)	No transformation	Eurostat
13. Intermediate goods (EA19) (NA) (S)	No transformation	Eurostat
14. All-items HICP (YKHI) (EA) (NA) (S)	No transformation	Eurostat
15. Food and non-alcoholic beverages (EA) (NA) (S)	No transformation	Eurostat
16. Alcoholic beverages, tobacco and narcotics (EA) (NA) (S)	No transformation	Eurostat
17. Clothing and footwear (EA) (NA) (S)	No transformation	Eurostat
18. Housing, water, electricity, gas and other fuels (EA) (NA) (S)	No transformation	Eurostat
19. Furnishings, household equipment and routine household maintenance (EA) (NA) (S)	No transformation	Eurostat
20. Health (EA) (NA) (S)	No transformation	Eurostat
21. Transport (EA) (NA) (S)	No transformation	Eurostat
22. Energy and unprocessed food (EA) (NA) (S)	No transformation	Eurostat
23. Overall index excluding housing, water, electricity, gas and other fuels (EA) (NA) (S)	No transformation	Eurostat
24. ECB Commodity Price index (EA19) (NA) (S)	No transformation	ECB SDW
Unemployment		
25. Unemployment rate (EA19) (SA) (S)	No transformation	Eurostat
Exchange rates		
26. USD (NA)	Log-difference	Eurostat
27. JPY (NA)	Log-difference	Eurostat
28. GBP (NA)	Log-difference	Eurostat
29. CHF (NA)	Log-difference	Eurostat
30. RUB (NA)	Log-difference	Eurostat
31. ECB nominal effective exch. rate of the Euro against euro area-19 countries and the EER-19 group of trading partners (AU, CA, DK, HK, JP, NO, SG, KR, SE, CH, GB, US, BG, CZ, HU, PL, RO, HR and CN) excluding the Euro (EA19) (NA)	Log-difference	ECB SDW
Confidence		
32. Evolution of the current overall order books in retail (EA19) (SA)	No transformation	Eurostat
33. Employment expectations over the next 3 months in retail (EA19) (SA)	No transformation	Eurostat
34. Price expectations over the next 3 months in retail (EA19) (SA)	No transformation	Eurostat
35. Retail confidence indicator (EA19) (SA)	No transformation	Eurostat
36. Own financial situation over the next 12 months (EA19) (SA)	No transformation	Eurostat
37. General economic situation over the next 12 months (EA19) (SA)	No transformation	Eurostat
38. Price trends over the next 12 months (EA19) (SA)	No transformation	Eurostat

Variable	Transformation	Source
39. Unemployment expectations over the next 12 months (EA19) (SA)	No transformation	Eurostat
40. Expectation of the demand over the next 3 months in services (EA19) (SA)	No transformation	Eurostat
41. Expectation of the employment over the next 3 months in services (EA19) (SA)	No transformation	Eurostat
42. Services confidence indicator (EA19) (SA)	No transformation	Eurostat
43. Evolution of the current overall order books in construction (EA19) (SA)	No transformation	Eurostat
44. Employment expectations over the next 3 months in construction (EA19) (SA)	No transformation	Eurostat
45. Price expectations over the next 3 months in construction (EA19) (SA)	No transformation	Eurostat
46. Construction confidence indicator (EA19) (SA)	No transformation	Eurostat
47. Employment expectations over the next 3 months in manufacturing (EA19) (SA)	No transformation	Eurostat
48. Production expectations over the next 3 months in manufacturing (EA19) (SA)	No transformation	Eurostat
49. Selling price expectations over the next 3 months in manufacturing (EA19) (SA)	No transformation	Eurostat
50. Industrial confidence indicator (EA19) (SA)	No transformation	Eurostat
Foreign trade		
51. Imports (EA19) (SCA) (S)	Log-difference	ECB SDW
52. Exports (EA19) (SCA) (S)	Log-difference	ECB SDW
53. Capital account (EA19) (NA) (S)	No transformation	ECB SDW
54. Financial account (EA19) (NA) (S)	No transformation	ECB SDW
55. Current account (EA19) (NA) (S)	No transformation	ECB SDW
Money		
56. Total assets/liabilities of the Eurosystem (EA) (NA)	Log-difference	ECB SDW
57. Monetary aggregate M1 (EA) (SCA)	Log-difference	ECB SDW
58. Monetary aggregate M2 (EA) (SCA)	Log-difference	ECB SDW
59. Monetary aggregate M3 (EA) (SCA)	Log-difference	ECB SDW
Stocks		
60. Dow Jones Euro Stoxx log-difference 0 Price index (NA)	Log-difference	ECB SDW
61. Dow Jones Euro Stoxx Price index (NA)	Log-difference	ECB SDW
62. Dow Jones Euro Stoxx Basic Materials E index (NA)	Log-difference	ECB SDW
63. Dow Jones Euro Stoxx Consumer Goods index (NA)	Log-difference	ECB SDW
64. Dow Jones Euro Stoxx Consumer Services index (NA)	Log-difference	ECB SDW
65. Dow Jones Euro Stoxx Financials index (NA)	Log-difference	ECB SDW
66. Dow Jones Euro Stoxx Technology E index (NA)	Log-difference	ECB SDW
67. Dow Jones Euro Stoxx Healthcare index (NA)	Log-difference	ECB SDW
68. Dow Jones Euro Stoxx Industrials index (NA)	Log-difference	ECB SDW

Variable	Transformation	Source
69. Dow Jones Euro Stoxx Oil and Gas Energy index (NA)	Log-difference	ECB SDW
70. Dow Jones Euro Stoxx Telecommunications index (NA)	Log-difference	ECB SDW
71. Dow Jones Euro Stoxx Utilities E index (NA)	Log-difference	ECB SDW
72. MSCI gross index of large and middle cap enterprises in Europe (NA)	Log-difference	MSCI
73. Annual real return of stocks (MSCI), taxation not taken into account. Formula: e^[12*Dln(72. variable)]/[1 + (14. variable)] − 1. (NA)	No transformation	MSCI, Eurostat
Interest rates		
74. Euro area 10-year Government Benchmark bond yield (EA) (NA)	No transformation	ECB SDW
75. Euro area 3-year Government Benchmark bond yield (EA) (NA)	No transformation	ECB SDW
76. Euro area log-difference-year Government Benchmark bond yield (EA) (NA)	No transformation	ECB SDW
77. Real 3-month Euribor (EA) (NA)	No transformation	ECB SDW
78. Euribor 1-month (EA) (NA)	No transformation	ECB SDW
79. Euribor 1-year (EA) (NA)	No transformation	ECB SDW
80. Euribor 6-month (EA) (NA)	No transformation	ECB SDW
81. Main refinancing operations rate (EA) (NA)	No transformation	ECB SDW
82. Spread between real 3-month Euribor and the main refinancing operations rate (EA) (NA)	No transformation	ECB SDW
83. Spread between Euro area 10-year Government Benchmark bond yield and the main refinancing operations rate (EA) (NA)	No transformation	ECB SDW
84. Real Euribor 1-year. Formula: 79. variable − 14. variable. (EA) (NA)	No transformation	ECB SDW, Eurostat
Foreign variables		
85. CPI-All Urban Consumers (NA) (S)	No transformation	BLS
86. Federal funds rate (NA)	No transformation	FED
87. Monetary aggregate M1 in OECD countries (SA)	Log-difference	OECD
88. Monetary aggregate M3 in OECD countries (SA)	Log-difference	OECD
89. Bank of Japan interest rate (NA)	No transformation	BoJ
90. Industrial production in the USA (SCA) (S)	Log-difference	OECD

Appendix B

The following figure shows some additional impulse response functions. The estimated model includes the MRO, the ECB's total assets/liabilities, inflation, 5 factors, 3 lags, constant and linear trend. The shock to the MRO is 0.25 percentage points. The response of production in construction is cumulative.

See Fig. 5.

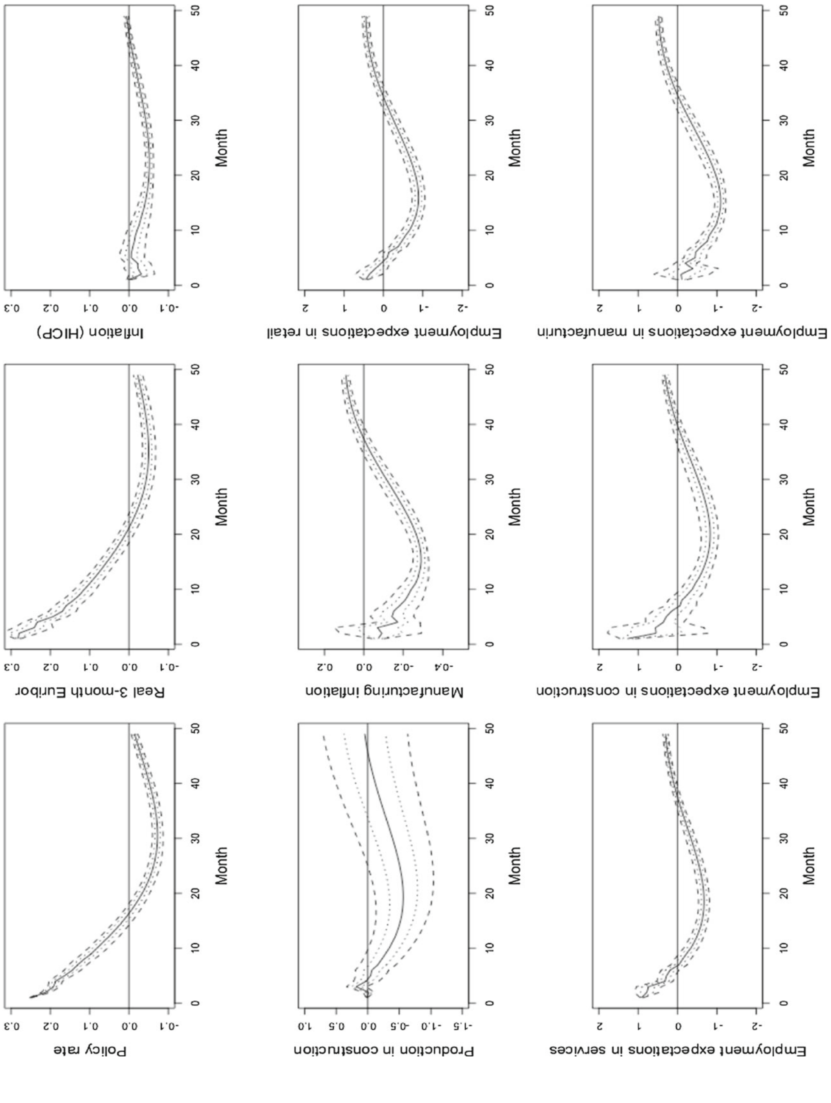

Fig. 5 Some additional impulse response functions estimated using the data from January 1999 to July 2017. The dashed lines around the impulse response functions represent 95% and dotted lines 68% CI

Appendix C

The following figures show the impulse responses when the ECB's total assets/liabilities are excluded.

See Figs. 6, 7 and 8.

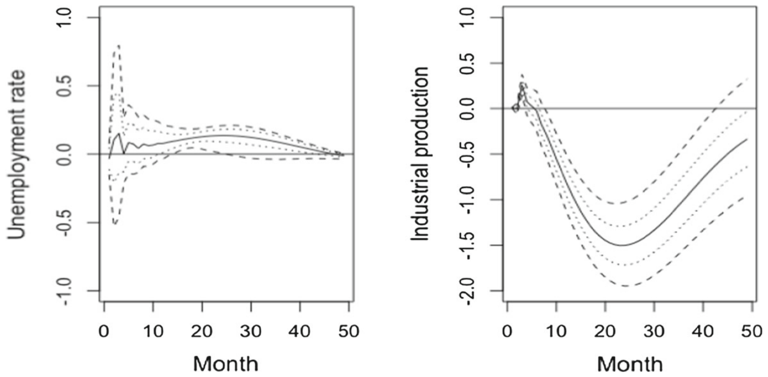

Fig. 6 The impulse response functions of unemployment rate and industrial production estimated using the data from January 1999 to July 2017. The dashed lines around the impulse response functions represent 95% and dotted lines 68% CI

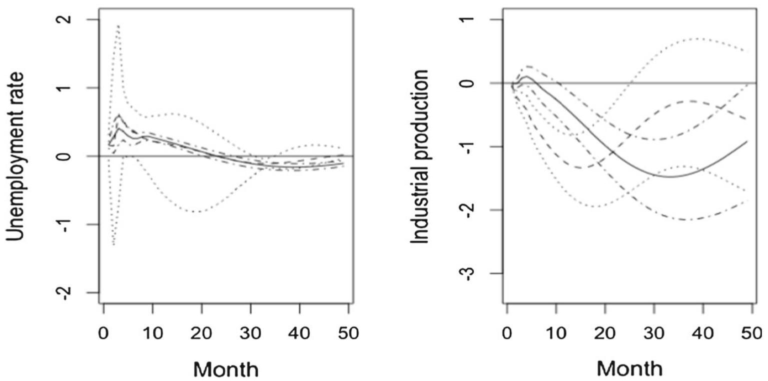

Fig. 7 The impulse response functions of unemployment rate and industrial production. The solid line represents the response estimated using data from January 1999 to July 2007. The dashed line represents the response estimated using the data from August 2007 to July 2017. The dotted and dotdashed lines around the impulse response functions represent 95% CI

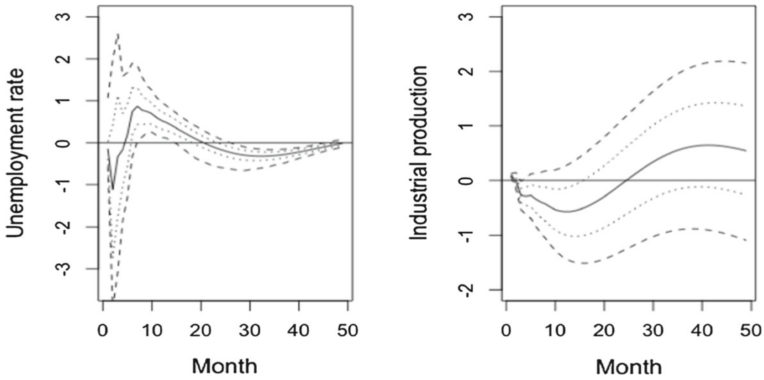

Fig. 8 The impulse response functions of unemployment rate and industrial production estimated using the data from May 2009 to July 2017. The dashed lines around the impulse response functions represent 95% and dotted lines 68% CI

Appendix D

The following figures show the impulse response functions when the order of the observed variables is inflation, MRO, total assets/liabilities.

See Figs. 9, 10 and 11.

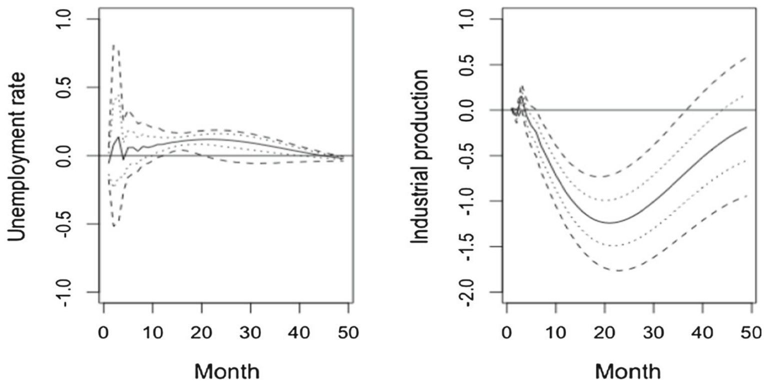

Fig. 9 The impulse response functions of unemployment rate and industrial production estimated using the data from January 1999 to July 2017. The dashed lines around the impulse response functions represent 95% and dotted lines 68% CI

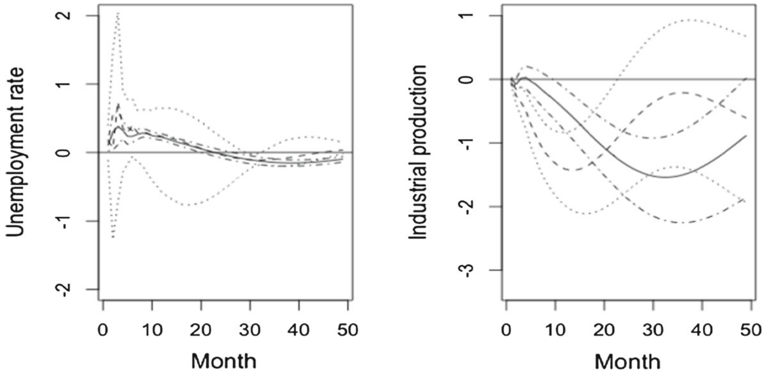

Fig. 10 The impulse response functions of unemployment rate and industrial production. The solid line represents the response estimated using data from January 1999 to July 2007. The dashed line represents the response estimated using the data from August 2007 to July 2017. The dotted and dotdashed lines around the impulse response functions represent 95% CI

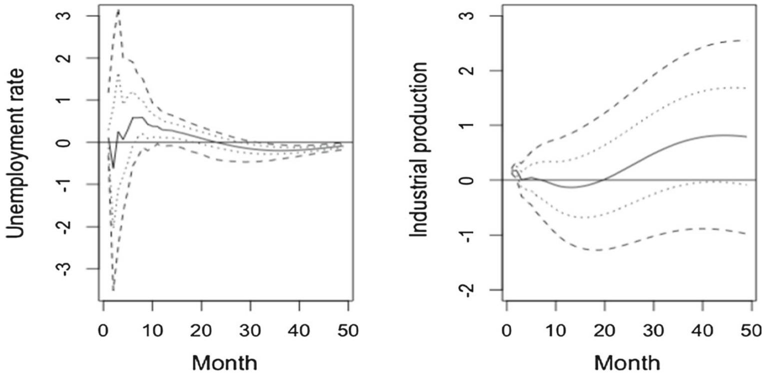

Fig. 11 The impulse response functions of unemployment rate and industrial production estimated using the data from May 2009 to July 2017. The dashed lines around the impulse response functions represent 95% and dotted lines 68% CI

Appendix E

The following figures show the impulse response functions when trend is left out.

See Figs. 12, 13 and 14.

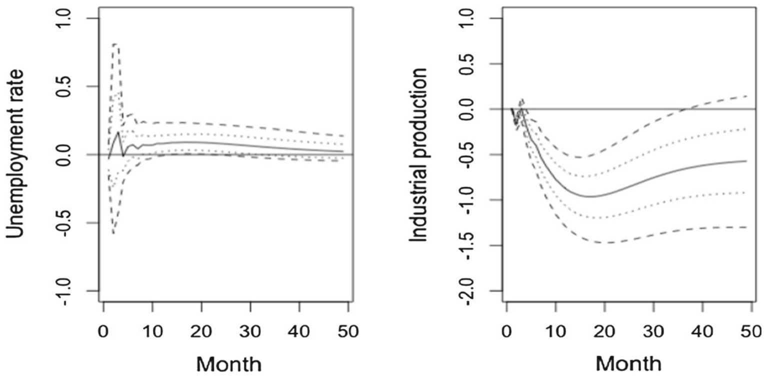

Fig. 12 The impulse response functions of unemployment rate and industrial production estimated using the data from January 1999 to July 2017. The dashed lines around the impulse response functions represent 95% and dotted lines 68% CI

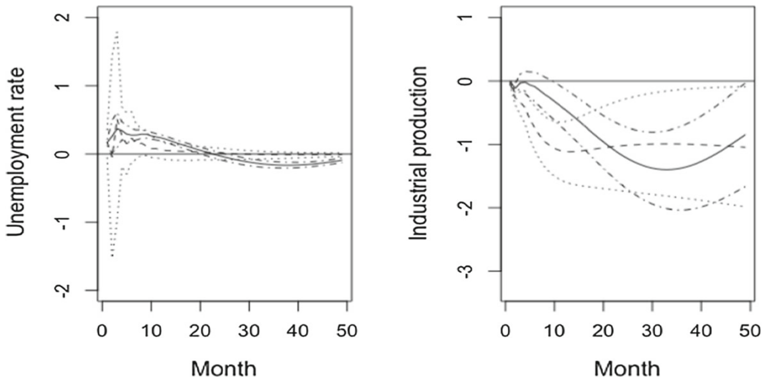

Fig. 13 The impulse response functions of unemployment rate and industrial production. The solid line represents the response estimated using data from January 1999 to July 2007. The dashed line represents the response estimated using the data from August 2007 to July 2017. The dotted and dotdashed lines around the impulse response functions represent 95% CI

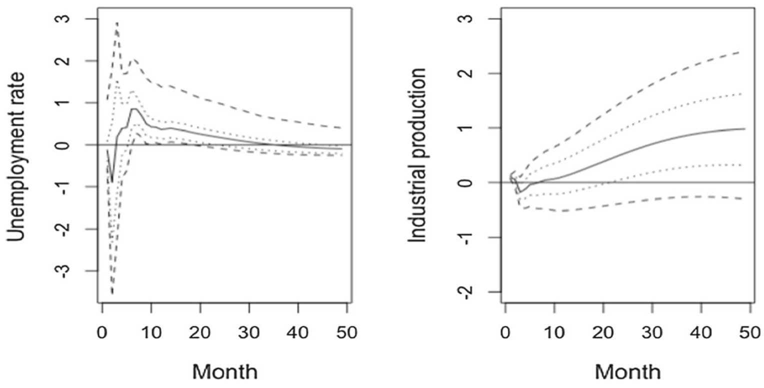

Fig. 14 The impulse response functions of unemployment rate and industrial production estimated using the data from May 2009 to July 2017. The dashed lines around the impulse response functions represent 95% and dotted lines 68% CI

Appendix F

The following figures show the impulse response functions when the order of the observed variables is total assets/liabilities, inflation, MRO.

See Figs. 15, 16 and 17.

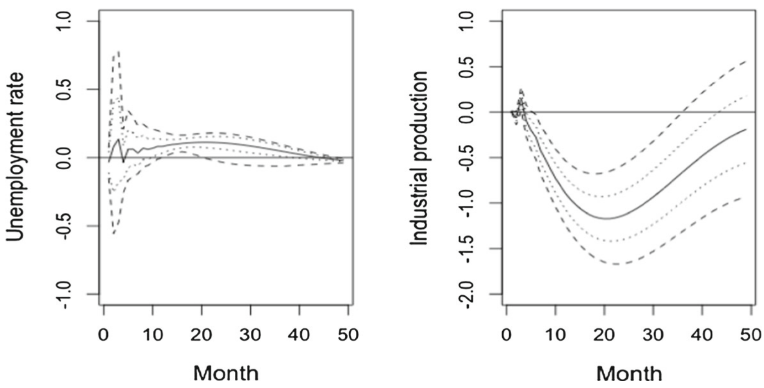

Fig. 15 The impulse response functions of unemployment rate and industrial production estimated using the data from January 1999 to July 2017. The dashed lines around the impulse response functions represent 95% and dotted lines 68% CI

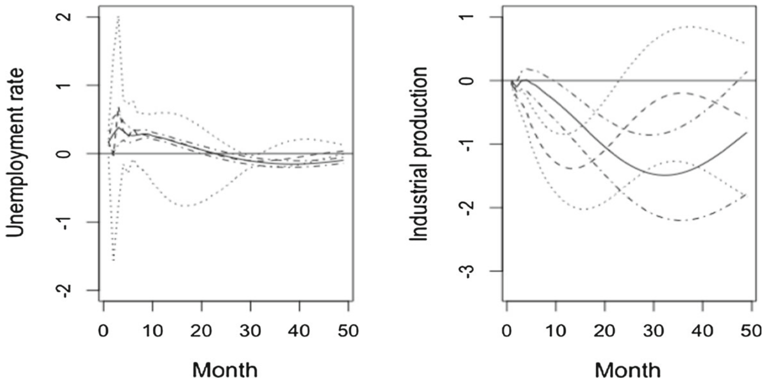

Fig. 16 The impulse response functions of unemployment rate and industrial production. The solid line represents the response estimated using data from January 1999 to July 2007. The dashed line represents the response estimated using the data from August 2007 to July 2017. The dotted and dotdashed lines around the impulse response functions represent 95% CI

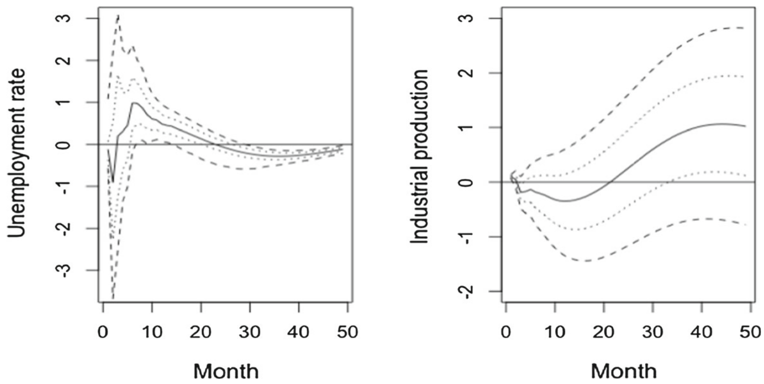

Fig. 17 The impulse response functions of unemployment rate and industrial production estimated using the data from May 2009 to July 2017. The dashed lines around the impulse response functions represent 95% and dotted lines 68% CI

Appendix G

The following figures show how the impulse responses estimated from the whole sample vary when the number of factors and lags is changed.

See Figs. 18, 19, 20 and 21.

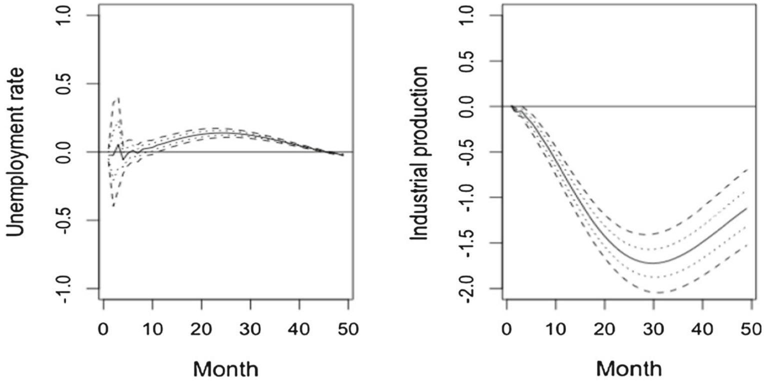

Fig. 18 The number of factors is 3. Everything else is as in the baseline model. The data are from January 1999 to July 2017. The dashed lines around the impulse response functions represent 95% and dotted lines 68% CI

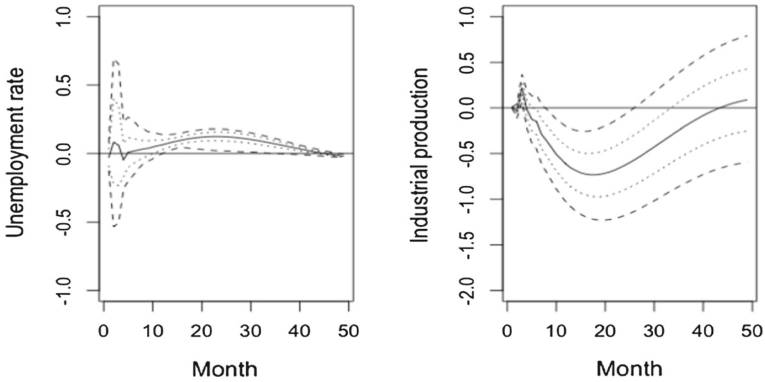

Fig. 19 The number of factors is 7. Everything else is as in the baseline model. The data are from January 1999 to July 2017. The dashed lines around the impulse response functions represent 95% and dotted lines 68% CI

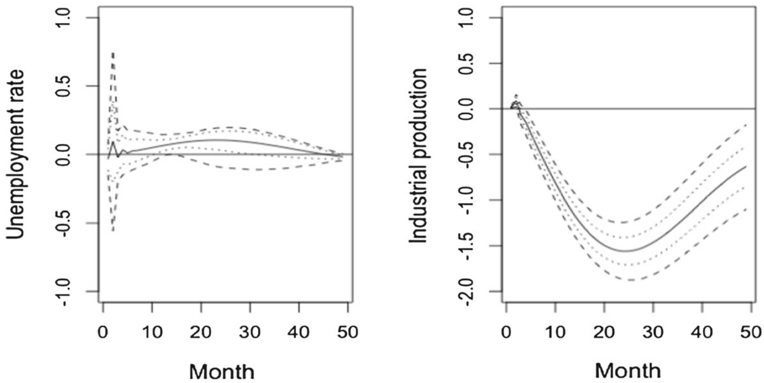

Fig. 20 The number of lags is 2. Everything else is as in the baseline model. The data are from January 1999 to July 2017. The dashed lines around the impulse response functions represent 95% and dotted lines 68% CI

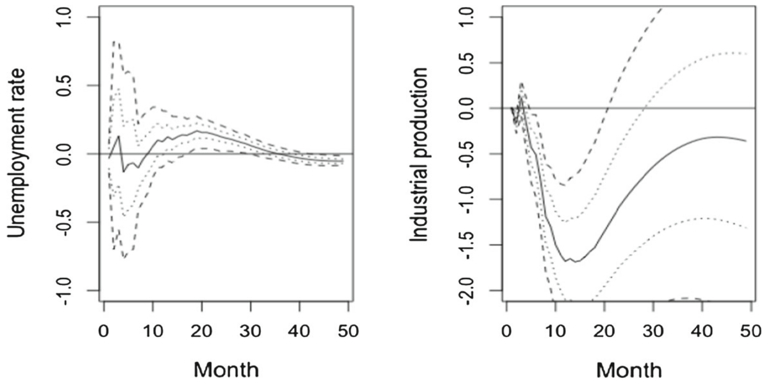

Fig. 21 The number of lags is 6. Everything else is as in the baseline model. The data are from January 1999 to July 2017. The dashed lines around the impulse response functions represent 95% and dotted lines 68% CI

Appendix H

The following figures show how the impulse responses estimated using the data are from May 2009 to July 2017 vary when the number of factors and lags is changed.

See Figs. 22, 23, 24 and 25.

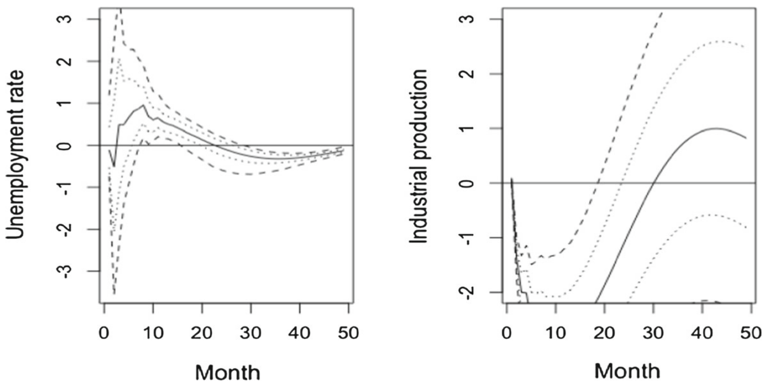

Fig. 22 The number of factors is 4. The data are from May 2009 to July 2017. Everything else is as in the baseline 2 model. The dashed lines around the impulse response functions represent 95% and dotted lines 68% CI

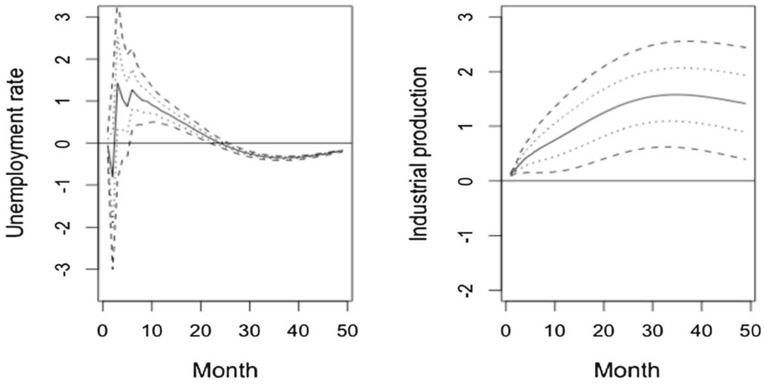

Fig. 23 The number of factors is 2. The data are from May 2009 to July 2017. Everything else is as in the baseline 2 model. The dashed lines around the impulse response functions represent 95% and dotted lines 68% CI

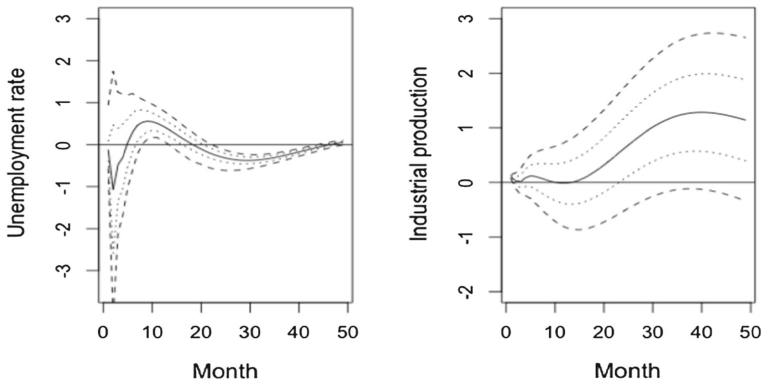

Fig. 24 The number of lags is 2. The data are from May 2009 to July 2017. Everything else is as in the baseline 2 model. The dashed lines around the impulse response functions represent 95% and dotted lines 68% CI

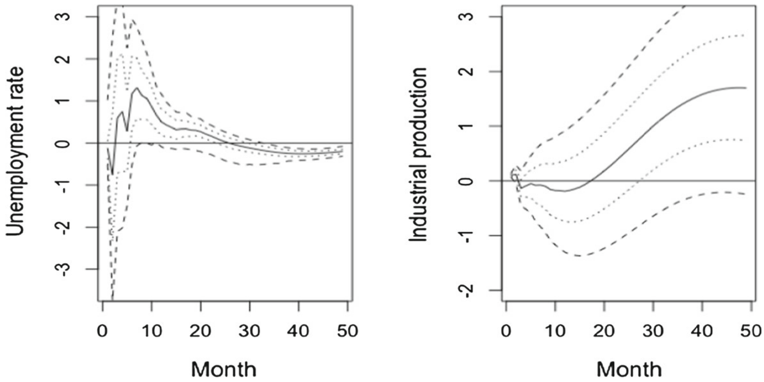

Fig. 25 The number of lags is 4. The data are from May 2009 to July 2017. Everything else is as in the baseline 2 model. The dashed lines around the impulse response functions represent 95% and dotted lines 68% confidence intervals

Appendix I

The following figure shows several impulse responses that are estimated using data from post-crisis period, and the beginnings of the time spans are varied.

See Fig. 26.

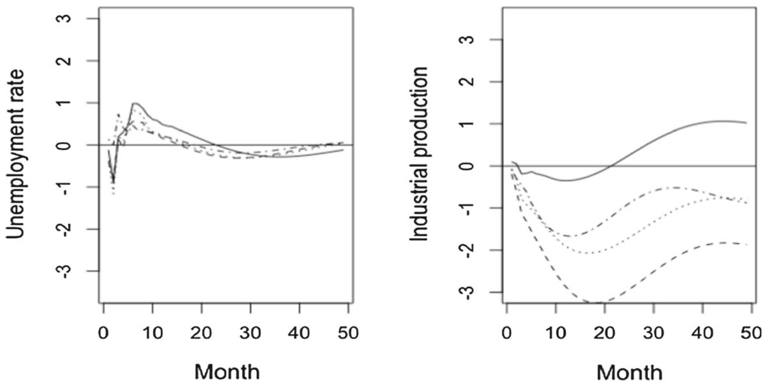

Fig. 26 The impulse response functions of unemployment rate and industrial production estimated using different time windows. The dotdashed lines represent impulse response functions estimated using the data from January 2008 to July 2017, dashed lines from January 2009 to July 2017, dotted lines from February 2009 to July 2017 and solid lines from May 2009 to July 2017

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