Essays on Monetary Policy Transmission

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Henrike Michaelis

Referentin:	Prof. Dr. Dalia Marin
Korreferent:	Prof. Dr. Gerhard Illing
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Namen der Berichterstatter: Prof. Dr. Dalia Marin, Prof. Dr. Gerhard Illing, Prof. Gianni Amisano Ph.D.

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Preface

For a successful conduct of monetary policy, the monetary authority must have an accurate assessment of how its policy decisions are transmitted through the economy. This requires a sound knowledge of the underlying structural processes at work through which monetary policy unfolds. Moreover, a good understanding of the transmission channels of monetary policy and whether they are changing over time helps to evaluate the extent to which earlier policy decisions have influenced the economy. A great number of studies assess the monetary transmission mechanism. It represents one of the most largely studied areas of monetary economics. With regard to the empirical analysis of the monetary transmission mechanism, Vector Autoregression approaches (VARs) have become one of the most widespread tools by economists. In particular, VARs are used to evaluate the importance of monetary policy disturbances for business cycle fluctuations. Since the seminal work by Sims (1972), Sims (1980) and Sims (1986), VAR models have been used extensively on this subject.[1] The literature generally documents a temporary reduction in output, with a peak after about one year following a monetary tightening. Prices seem to respond with a delayed decline and then fall permanently.

However, the typical VAR maintains the assumption of constant coefficients over time. This assumption seems quite restrictive and should thus be tested. This is particularly relevant for the discussion of possible changes in the monetary policy transmission mechanism. As Canova (2007) describes, one can think of these changes in two ways: *first*, as abrupt switches which can be accounted for by structural breaks[2] or, *second*, as models with continuously evolving parameters which account for gradual changes over time. Studies incorporating continuously evolving parameters in a VAR include for example Cogley

[1] Beginning, e.g., with Bernanke and Blinder (1992), Sims (1992), Evans and Marshall (1998) and Christiano et al. (1999) for the US, and Van Aarle et al. (2003) and Peersman and Smets (2003) for the Euro Area.

[2] Structural breaks can be addressed by Markov switching or regime switching VARs (see e.g. Paap and Van Dijk (2003), Sims and Zha (2006) and Koop and Potter (2006).

(2005) and Primiceri (2005). Cogley (2005) employs a time-varying parameter VAR and additionally accounts for stochastic volatility in the variances. But the simultaneous relations between the variables are still assumed to be constant. Primiceri (2005) allows for a fully time-varying variance covariance matrix as well as for time-varying VAR coefficients. He develops the salient Bayesian time-varying parameter Vector Autoregression approach (TVP-VAR).

This line of research serves as the starting point of my thesis, which aims to provide further empirical evidence on a better understanding of how monetary policy shocks are transmitted through the economy. This thesis consists of three main chapters, all contributing to the literature on monetary policy transmission. In chapter I, I assess the time-varying impact of an unconventional monetary policy shock in Japan. Chapter II addresses how a conventional monetary policy and also an exchange rate shock are evolving across time in Poland. Chapter III focuses on the Euro Area by employing a standard VAR and a nonlinear VAR. It assesses the impact of a conventional monetary policy shock and its possibly different influence before and after the financial crisis.

Before the Lehman Brothers bankruptcy in 2008 which induced the financial crisis and led to the worst global recession since the 1930s, the intellectual and empirical basis for conducting monetary policy seemed well founded and robust. The main goal of monetary policy was to achieve low and stable inflation. For pursuing their price stability objective, the monetary authorities' policy framework was inflation targeting and its key policy instrument was the short-term interest rate. Nowadays, the conduct of monetary policy in advanced economies based on this framework generally is referred to as conventional monetary policy. This way of conducting monetary policy basically has not been challenged until the outbreak of the financial crisis. In the wake of this turmoil and its aftermath, inflation rates have been falling below the generally accepted and desired levels. Many central banks in advanced economies responded by cutting interest rates to historically low levels and by embarking on unconventional policies. In addition to inflation targeting, the monetary policy objective has shifted more and more to secure financial stability. However, according to Tinbergen's Law, a central bank needs at least as many instruments as it has policy targets. Therefore, monetary authorities started to extend their set of policy instruments by considering additional macroprudential measures. Furthermore, with short-term interest rates close to zero or even at the Zero Lower

Bound (ZLB), central banks have little room for further meaningful reductions of their policy interest rates to provide additional monetary stimuli. Therefore, an important tool for macroeconomic stabilisation and for escaping deflation might soon or already has become obsolete. Under such circumstances, the use of unconventional monetary policy tools to support economic and financial stability is increasingly taken into consideration by many central banks. This includes, for example, unorthodox measures like credit and quantitative easing. The nature and magnitude of the macroeconomic effects of such unconventional measures is of great interest to central banks, but historical evidence on unconventional measures is rare. However, for a successful conduct of monetary policy, it is crucial to learn from the experiences made within the past few years and to shed light on the efficiency of unconventional monetary policy.

A comprehensive summary of the growing literature which assesses the quantitative effects of unconventional policies on the economy is given by Joyce et al. (2012). It is generally a challenge to isolate the impact of an unconventional policy as there exist many other contributory factors driving the economy. For example, at the same time as central banks embarked on monetary easing, fiscal authorities were trying to stimulate demand. Furthermore, other countries implementing similar measures may have induced spillover effects. Given these difficulties, a wide range of methodologies has been applied in order to assess the impact of unconventional policy measures. With regard to the estimation framework, these methodologies can be generally divided into three branches.

The *first* strand is based on the so called 'plug-in' approach. Into standard macroeconomic models estimates of the effect of unconventional policies on asset prices are implemented (see Chung et al. (2012) among others).[3] For instance, Chung et al. (2012)) provide evidence that unconventional policies undertaken by the Federal Reserve possibly influenced output and inflation. However, as Kimura and Nakajima (2013) argue, (a) this approach rests on 'plugged-in' estimates coming from separate empirical studies on the response of financial markets and (b) it cannot differentiate between two different transmission channels, the portfolio rebalancing channel and the signaling channel, which may lead to

[3]Two of the first studies evaluating the Federal Reserves' large scale asset purchases (LSAPs) and its influence on asset prices are Gagnon et al. (2011) and Wright (2012). The Bank of Japan's policy is studied among others by Kimura and Small (2006) and Ueda (2012). Meier (2009) and Joyce et al. (2011) focus on the Bank of England's policy. These studies trace the effect of unconventional policies on asset prices and find that it has been successful in reducing medium and long-term interest rates. Their estimates can then be plugged into standard macroeconomic models, such as the FRB/US model (see Chung et al. (2012) among others).

a possible estimation bias.[4]

The *second* strand of literature is based on the structural model provided by the DSGE framework in order to assess the different influences of long- and short-term interest rates on the real economy. For example, Kiley (2012) and Chen et al. (2012) employ a structural model and account for financial market segmentation. They claim that the real economy is less affected by term and risk premiums than by short-term interest rates. Note that on the one hand, it appears useful to trace the influence of unconventional policies based on structural models that allow for a clear definition of transmission channels. On the other hand, there are numerous ways on how financial market segmentation can be incorporated.[5] Therefore, these approaches are crucially driven by the assumed underlying structure. So far however, it seems that we do not have enough information about all possible specifications and which one characterises the best structure of financial market segmentation (Kimura and Nakajima (2013)).

The *third* strand of research refers to the structural VAR. This constitutes a data-driven approach, which imposes very little theoretical structure on the data. It can be used to establish important stylised facts on the transmission mechanism of monetary policy. The transmission is evaluated by means of impulse responses. Within VAR models, several papers explicitly account for the possibility of a ZLB. As the historic evidence on economies close to or at the ZLB is rare, many studies focus on Japan. It is well known that the Japanese economy has been stuck in a liquidity trap since the mid-90s. Since then the Bank of Japan (BoJ) has implemented various different strategies to fight the ZLB and to provide economic stimuli. For example, Fujiwara (2006) employs a Markov switching VAR to analyse if there are structural breaks in the Japanese monetary transmission mechanism due to the ZLB. Iwata and Wu (2006) use a VAR model with constant parameters and model the policy interest rate as a censored variable. In contrast, Nakajima et al. (2011) refer to a TVP-VAR with stochastic volatility and the interest rate

[4] The portfolio rebalancing channel captures a compression of term premiums and the signaling channel a lowering of expected short-term interest rates. For further details on the incorrect assessment which may arise due to possible biases and uncertainties in the coefficients refer to Hamilton and Wu (2012) and Bauer et al. (2012)). Regarding the effects of the two channels refer to Stein (2012), Kiley (2012) and Chen et al. (2012).

[5] Consider for example Kiley (2012) and Chen et al. (2012). Both models are based on two types of agents. One agent is assumed to trade both long- and short-term bonds. The difference arises with respect to the second agent. Kiley (2012) assumes that the second agent has only access to short-term bonds and in contrast, Chen et al. (2012) assume that the second agent only trades long-term bonds.

as a censored variable.[6]

In the beginning of 2001, the BoJ changed the operating target from the call rate to the outstanding current account balances held by banks at the BoJ. Since then they have been using this measure for providing support to the financial markets. This measure expanded the overall size of the BoJ's balance sheet and, consequently, the monetary base. The respective increase in the monetary base is also reflected in an accumulation of bank reserves. This is also referred to as Quantitative Easing (QE) (see Lenza et al. (2010) for more details).[7] As these bank reserves have become the key monetary policy instrument in Japan, it is crucial for policy makers to learn about its effectiveness. Studies that explicitly trace the effect of a QE shock are, for example, Kimura et al. (2003) by means of a semi-TVP-VAR and Kimura and Nakajima (2013) by means of a TVP-VAR with stochastic volatility which is additionally combined with a latent threshold model.[8] However, these studies are based on recursive identification for identifying a QE shock. This identification can be quite restrictive especially in a model with quarterly data. The use of sign restrictions generally imposes fewer assumptions on the simultaneous relations between the variables. For example, Kamada and Sugo (2006) and Franta (2011) use sign restrictions for Japan. Franta (2011) appears to be the first work with sign restrictions in a TVP-VAR with stochastic volatility. He implements different sign restrictions to account for the differences in the conduct of monetary policy in Japan between 1981 and 2010.[9] Note that the monetary policy shock is the only identified shock in his analysis. This may not be sufficient as it is likely that other disturbances, e.g., business cycle fluctuations, enter the identified monetary policy shock. This highlights the importance of identifying further shocks. Moreover, his analysis is still silent on the recent influences of Japanese monetary policy as his estimation horizon lasts until 2010 only. The recent period of Japanese monetary policy is of particular interest to central banks around the

[6]Their estimation horizon ranges from 1980Q1 until 2008Q3. They find a visible and declining effect on CPI following a call rate increase until the end of the 80s and a price puzzle for the beginning of the 90s. Thereafter, the effect seems to be insignificant. The negative effect on output is significant until 2002.

[7]Note that the Fed, the BoE and the ECB also use their balance sheet for providing financial support. In the period before the financial crisis, financial support was also given by changing the composition of their balance sheets, but the overall size basically remained the same. Since the financial turmoil however, they have induced an expansion of the overall size of their balance sheets.

[8]Kimura et al. (2003) focus on a positive monetary base shock for the period 1985Q2 until 2002Q1. Kimura and Nakajima (2013) focus on a bank reserve shock for the period ranging from 1981Q2 until 2012Q3.

[9]His results indicate a difference in the transmission mechanism between 'normal' times and the QE periods. However, he does not seem to find visible differences during the ZLB period.

world, because the BoJ has adopted various strategies to fight the ZLB since the middle of the 90s. More specifically, they introduced a Zero-Interest Rate Policy (ZIRP) at the end of the 90s, from 2001 to 2006 they implemented the first Quantitative Easing Policy (QEP) and since 2013, they have been trying to combat the ZLB more aggressively with the so-called 'Abenomics' monetary policy easing strategy. Empirical evidence on the recent period is still rather scarce. Against this background, it is important to shed more light on the transmission of monetary policy.

In chapter I of my dissertation, I fill these gaps. It is joint work with Dr. Sebastian Watzka (LMU) and can be seen as the main chapter of my thesis. We use an identification scheme based on sign restrictions to study if the transmission of the QE shock has changed over time. Besides a QE shock, two business cycle disturbances, a demand and a supply shock, are identified. This allows us to avoid that business cycle disturbances enter the identified QE shock and to evaluate the quantitative importance of the QE shock relative to the other disturbances. We use relatively agnostic sign restrictions to incorporate the ZLB of short-term interest rates into a TVP-VAR. This identification scheme is based on the New Keynesian model of Eggertsson (2011). These restrictions were first implemented by Schenkelberg and Watzka (2013) in a Bayesian VAR (BVAR). We employ the TVP-VAR as it proved useful in previous studies analysing the monetary policy transmission in Japan. For example, among others, Nakajima (2011) shows that it is important to account for time variation. Our analysis also confirms the importance of using a TVP-VAR. Furthermore, as far as we are aware, this work is the first to trace a QE shock during the 'Abenomics' period based on a VAR setting. Our results show that both the effect on output and on inflation have become stronger and longer lasting over time. More specifically, the inflation response in 2013, a year likely to be strongly influenced by the current 'Abenomics' strategy shows a significant permanent increase in inflation following a QE shock. In contrast, the responses during the period of the ZIRP and the QEP only confirm an initially positive effect on prices. For output, we see a visible effect in 2013, but insignificant effects during the ZIRP and the QEP. Next to Japan, our findings are also of particular interest to other advanced economies with short term interest rates at the ZLB.

Chapter II of my dissertation is related to chapter I with respect to the methodological approach. It traces the effect of a conventional monetary policy shock in Poland one of the most important emerging countries in Eastern Europe. In addition to the monetary policy shock, it investigates the effect of an exchange rate shock. It is joined work with Dr. Olga Arratibel (ECB) and was published in January 2014 in the refereed Working Paper Series of the ECB. Poland's economy experienced several structural changes during the last decades. These include, e.g., the increasing trade openness, partly driven by the integration into the European Union, the shift from exchange rate targeting to an inflation targeting strategy and, more recently, the global financial crisis. These structural changes seem to strongly motivate a flexible estimation framework which allows for the possibility of time variation. Our approach follows the TVP-VAR with stochastic volatility and we provide empirical evidence that the impact of a monetary policy and an exchange rate shock has indeed varied over time from 1996Q1 to 2012Q3.

There are numerous studies employing VAR models for tracing the influence of both monetary policy and exchange rate shocks in Central and Eastern European countries (CEEs). A comprehensive summary is given by Égert and MacDonald (2009). Studies assessing the influence of monetary policy decisions are among others Darvas (2009) and Jarociński (2010). Darvas (2009) uses a semi TVP-VAR with recursive identification for the period from 1993Q1 to 2008Q2.[10] Jarociński (2010) uses a structural BVAR with a combination of sign and zero restrictions for four CEE countries (including Poland). However, his approach does not allow for assessing whether the shocks have time-varying effects. Studies focusing on the exchange rate pass-through in Poland by means of a standard VAR with recursive identification include, e.g., Bitans (2004) and Ca'Zorzi et al. (2007).[11]

This avenue of research serves as the starting point of my second chapter. It aims to provide empirical contributions to a better understanding of how monetary policy decisions and exchange rate changes have affected the Polish economy, which can be considered as a small open economy. Our work contributes to the literature with an explicit analysis whether there are time-varying effects following a monetary policy or an exchange rate shock in Poland. We employ a combination of sign and zero restrictions for identifying our two shocks. As far as we are aware, our study is the first to estimate a TVP-VAR with stochastic volatility for Poland and to provide evidence on which econometric frame-

[10]The approach allows for time-varying coefficients but assumes a constant variance. Following a monetary tightening, he finds visible effects on both GDP and CPI which become stronger over time.

[11]The approach by Bitans (2004) is based on two samples (1993-1999 and 2000-2003). Following an exchange rate shock, he estimates a larger pass-through on prices for the first sample.

work is the best approach for tracing the impact of monetary policy and exchange rate shocks in Poland. Four main findings stand out: (1) Our analysis confirms the importance of using a TVP-VAR. (2) Overall, the Polish economy seems to have become more resilient to monetary policy and exchange rate shocks over time. More specifically, (3) a monetary tightening has a visibly declining effect on GDP. Since 2000, absorbing such a shock has become less costly in terms of output, notwithstanding some reversal since the beginning of the financial crisis. With regard to prices, we estimate a stronger decline during the first half of our sample, when Poland experienced high inflation. (4) Following the exchange rate shock, defined as an appreciation, we see a changing effect on output over time. The price responses confirm the general finding of the literature of a slightly decreasing pass-through across time.

Chapter III also focuses on the effect of a conventional monetary policy shock. It investigates the Euro Area transmission mechanism. Especially against the background of the recent financial as well as the Euro crisis, which pose a great challenge to the conduct of monetary policy, it is crucial to understand the effect of monetary policy decisions. More specifically, most of the studies analysing the effect of a monetary policy shock in the Euro Area employ standard VAR approaches based on synthetic Euro Area data from 1980 onwards (see Van Aarle et al. (2003) and Peersman and Smets (2003)). During this time, a common central bank and hence a common monetary policy was not yet established. Studies tracing monetary policy effects for the Euro Area include, e.g., Weber et al. (2009) and Cecioni et al. (2011).[12] Cecioni et al. (2011) employ a BVAR with recursive identification or sign restrictions using Euro Area data from 1999M1 to 2007M7 or to 2009M8. They estimate a temporary decline in output and a permanent fall in prices following a monetary tightening. With a comparison between the two sample periods (until 2007 or 2009), they try to shed light on whether there are differences in the transmission of a monetary policy shock due to the financial crisis. Their results do not confirm a visible difference. However, as generally argued in the literature, the global financial crisis induced a rethinking of monetary policy frameworks. For the ECB, Gerlach and Lewis (2010) provide empirical evidence for a policy shift after the bankruptcy of Lehman Brothers. Against this background, it emphasises the importance of a sample split in September

[12] Weber et al. (2009) investigate whether the creation of the Euro Area influenced the monetary transmission mechanism by means of a standard VAR with recursive identification. Their Euro Area data covers the period from 1999Q1 until 2006Q4. Following a monetary tightening, GDP temporarily declines and prices show a delayed response, but then stay permanently negative.

2008. Therefore, the study by Cecioni et al. (2011) might not thoroughly account for the possible differences in the transmission mechanism.

Against the background of scarce evidence on the possible influence of the financial crisis following a monetary policy shock, it is important to shed more light on the Euro Area transmission mechanism. Therefore, I analyse whether there are possible differences in the influence of monetary policy before and after the global financial crisis by employing a VAR framework with recursive identification. I split the data in September 2008, the time of the outbreak of the financial turmoil. The pre financial crisis sample analyses the Euro Area transmission in 'normal' times from January 1999 until September 2008. The post financial crisis sample ranges from October 2008 until December 2014.[13] To my knowledge, this is the first approach using a standard VAR for the post financial crisis period. Furthermore, I provide new empirical evidence on the Euro Area monetary transmission mechanism by employing a nonlinear VAR model. More specifically, a standard linear VAR mixes the influence of an unexpected increase in the policy rate and an expected decrease which is not implemented (Hamilton and Jorda (2002)). Since these two expectations have a very different influence on the economy, it is important to differentiate between them. I address this point by following the empirical framework of Hamilton and Jorda (2002). Furthermore, I extend Hamilton and Jorda's (2002) approach by implementing exogenous variables into the nonlinear VAR specification. This controls for changes in world demand and inflation. My estimates show that (1) it seems important to differentiate between unexpected and expected policy changes which show very different effects concerning the monetary policy transmission. (2) Compared to the linear VAR, an unexpected increase in the policy rate has larger and longer lasting effects on output and inflation. This result is especially pronounced for the pre financial crisis sample. (3) As expected, the expected interest rate change reveals a rather minor and insignificant impact. (4) The influence of a monetary policy shock during the post financial crisis sample seems to be less strong. This suggests that monetary policy has become less effective in the post financial crisis period.

The three chapters analyse the monetary policy transmission in three very different economies. The results of each chapter deliver fresh and interesting insights into the

[13] Against the background of a rather short sample period from October 2008 until January 2014, it becomes necessary for the time series estimation to extend the post financial crisis sample until December 2014.

transmission mechanism of conventional as well as unconventional monetary policy decisions and help economists and policy makers to improve their understanding of the influence of their decisions on the economy.

All three chapters of this dissertation are self-contained and include their own introductions and appendices such that they can be read independently.

Chapter 1

Are there Differences in the Effectiveness of Quantitative Easing at the Zero-Lower-Bound in Japan over Time?*

1.1 Introduction

Using a time-varying parameter vector autoregression (TVP-VAR) framework, we study the changing effectiveness of the Bank of Japan's Quantitative Easing policies over time. It is well known that the Japanese economy has been stuck in a liquidity trap since the mid-90s. Since then the Bank of Japan (BoJ) has adopted various different strategies to combat the recession and stimulate the economy. We use a time-varying VAR framework with stochastic volatility to analyse how the effects of a Quantitative Easing (QE) shock have changed over time and when it was possibly effective.

We specifically analyse the so-called Zero-Interest Rate Policy (ZIRP) from 1999 to 2000, the Quantitative Easing Policy (QEP) from 2001 to 2006, and most recently the so-called 'Abenomics' monetary policy easing strategy implemented under current BoJ Governor Haruhiko Kuroda and carried out under the political leadership of Prime Minister Shinzo Abe.

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* This chapter is based on joint work with Dr. Sebastian Watzka (LMU).

To identify a QE shock, we follow Schenkelberg and Watzka (2013) and use a new sign restriction approach when the economy is stuck at the Zero Lower Bound (ZLB). To allow for time variation in the impulse responses, we embed this identification strategy in the TVP-VAR framework of Primiceri (2005). With this approach we are seeking to shed light on the changing nature of the monetary policy transmission mechanism in Japan during these different monetary policy stances.

We investigate whether the impact of a QE shock has varied over time in Japan through a marginal likelihood estimation which compares a constant coefficient VAR with our TVP-VAR. Our research confirms that the TVP-VAR is indeed a better fit for Japan and that a QE shock estimated for the Japanese economy does in fact have changing effects over time. In particular it seems that the effects on both real GDP and core CPI have become stronger and longer lasting over time. More specifically, the response for prices in 2013, a period probably highly influenced by the 'Abenomics' program, stays permanently significant. This is in contrast to the responses under the ZIRP and the first QE program. During these periods only an initial significant effect is observed. Regarding GDP, we estimate again a significant impact during the time of the 'Abenomics' strategy, whereas during the ZIRP and the first QE program no significant impact is reported. Generally, these findings are also supported by our variance decomposition analysis. Especially since 2013, the relative importance of QE shocks has increased. These effects are likely to be driven by some extent by the current 'Abenomics' program.

Our results are interesting not only for Japan, but also for other advanced economies with nominal interest rates close to zero or at the ZLB. The recent financial crisis has by now been going on for five years, by some already labeled as 'Great Recession'. It started with housing market bubbles bursting in the US, UK, and some Euro Area countries. Problems in highly leveraged banking sectors followed, and policy interest rates were subsequently lowered to historically low levels of virtually zero. A severe deleveraging of the private sector is currently hitting the real economy of most advanced countries. Inflationary pressure has generally been subdued. Hence, the current experiences of most advanced economies pretty closely mirror the Japanese experience. It is against this background that our study on the effectiveness of the QE policy in Japan sheds light on the potential effects of recently implemented QE policies in the US, the UK, and possibly the Euro Area.

The remainder of the paper is organised as follows: Section 1.2 gives an overview of related literature on Japan, Section 1.3 quickly summarises the Japanese monetary policy developments, Section 1.4 describes the setup of our empirical model, Section 1.5 briefly summarises the marginal likelihood results and Section 1.6 discusses our results. It is divided into the four following subsections. Subsection one presents our results on a QE shock. Subsection two briefly reports the effect of other business cycle disturbances. Subsection three outlines results on a forecast error variance decomposition (FEVD) and subsection four links our results to the underlying theoretical framework of Eggertsson (2011). Section 1.7 discusses our robustness checks and section 1.8 finally concludes.

1.2 Survey of Related Literature for Japan

Vector autoregression models (VAR) are a widely used tool for analysing the monetary policy transmission, also for Japan. These include for example Miyao (2002), who introduces the benchmark VAR model for estimating the impact of monetary policy during 1975 to 1998. Since it is likely that the transmission mechanism varies over time, more flexible models, accounting for time variation, are becoming increasingly a focus of research. For example, Kimura et al. (2003) estimate a VAR with time-varying coefficients for the period between 1971 and 2002. Still, their approach relies on a constant variance. In contrast, Nakajima (2011) employs a time-varying VAR (TVP-VAR) and allows for stochastic volatility in the variance covariance matrix.[1]

An important issue which is increasingly being discussed in the literature, is the monetary policy transmission when nominal interest rates are close to zero or even at ZLB. In these situations, central banks have only very little room for decreasing their short-term policy rates. Therefore, the impact of monetary policy is unlikely to operate through the conventional interest rate channel. Instead central banks then typically operate through what is now called ‘forward guidance’ or increases in the monetary base through some form of ‘Quantitative Easing’ (QE). Within the VAR framework, several papers deal explicitly with the ZLB for Japan and investigate the monetary policy transmission during these periods. Fujiwara (2006) uses a Markov switching VAR model for the period of 1985

[1] His sample is based on data ranging from 1977Q1 until 2007Q4. Following an increase in the call rate, he reports a fall in prices until the mid 90s and thereafter a small price puzzle. Further, he finds a visible effect on output.

to 2004 to investigate whether there are structural breaks in the impact of monetary policy due to the introduction of the zero nominal interest rate. Kamada and Sugo (2006) estimate a monetary policy proxy which can take on negative values to account for a whole range of different policy measures. Iwata and Wu (2006) model the nominal interest rate as a censored variable in a VAR model with constant coefficients. Nakajima et al. (2011) employ a TVP-VAR with stochastic volatility and in addition consider the nominal interest rate to be a censored variable.

It should be noted that since the beginning of 2001, the BoJ does not use the call rate as their operating target but the outstanding current account balances held by banks at the BoJ. Since this measure is the key monetary instrument in Japan, it is of importance to learn about its effectiveness. For example, Kimura et al. (2003), Kimura and Nakajima (2013) and Hayashi and Koeda (2014) trace the effect of such a QE shock. Hayashi and Koeda (2014) employ a regime switching SVAR for the period from 1988 until 2012 and find a visible effect on output and inflation following a QE shock. In contrast are the findings by Kimura et al. (2003) and Kimura and Nakajima (2013). Kimura et al. (2003) use a TVP-VAR with a constant variance from 1985Q2 until 2002Q1. For the ZLB periods they do not estimate a significant effect on output and prices. Kimura and Nakajima (2013) employ a TVP-VAR with stochastic volatility and combine it with a latent threshold model from 1981Q2 until 2012Q3. Following a QE shock, they also report a non-visible effect on both output and inflation. However, these studies use a recursive identification scheme for identifying a QE shock. As it is generally argued, this can be rather restrictive especially in a model with quarterly data. By referring to sign restrictions, usually less restrictions need to be imposed. For example, Kamada and Sugo (2006) and Franta (2011) incorporate sign restrictions to identify an unconventional monetary policy shock at the ZLB for Japan. Kamada and Sugo (2006) use Uhlig (2005)'s sign restricted VAR together with a special 'intermediate' monetary policy variable which is not a proper monetary policy instrument like the call rate or base money, but still closely related to the monetary policy instrument. Their sample spans from 1978 to April 2005. Franta (2011) uses the TVP-VAR with stochastic volatility but with altering sign restrictions to account for differences in the conduct of monetary policy in Japan between 1981 and 2010. He finds differences in the monetary policy transmission between the QE policy period and 'normal' times. However, he does not report visible differences to a QE shock during the ZLB periods. Note that he only identifies the monetary policy shock. This may not be sufficient since it does not ensure that other disturbances, such as e.g. business cycle fluctuations (demand or supply shocks), enter the identified monetary policy shock. Moreover, he does not capture the recent influence of the 'Abenomics' strategy. Against the background of still rather scarce empirical evidence on the recent period in Japan, it is crucial to provide more details on the monetary policy transmission.[2]

We hope to fill these gaps. More specifically, we use a relatively agnostic framework for analysing QE shocks when interest rates are close to zero. We employ a novel identification scheme based on sign restrictions to incorporate the ZLB of short-term nominal interest rates. These sign restrictions were introduced by Schenkelberg and Watzka (2013) in a BVAR to identify a QE shock when the economy is stuck at the ZLB.[3] Next to the QE shock, two business cycle disturbances, a demand and a supply shock, are identified for two reasons: (1) for avoiding that business cycle disturbances enter the identified QE shock and (2) for evaluating the quantitative importance of the QE shock relative to the business cycle shocks and how they have changed across time. Our estimation is based on the TVP-VAR with stochastic volatility, introduced by Primiceri (2005). Among others, Nakajima (2011) underlines the importance to allow for time variation. Also our analysis confirms the importance of a time-varying approach by means of a marginal likelihood comparison. Furthermore, as far as we are aware, this work is the first one to address the influence of a QE shock on key macro economic variables during the 'Abenomics' period. based on a VAR setting. Interestingly, the effectiveness of unconventional monetary policies[4] has been extensively studied on financial market effects, but there exists only a small body of empirical literature on real economic influences. It is against these backgrounds that our paper provides new interesting insights, not only for Japan, on the macro economic effects of QE shocks and whether they vary over time.

[2]See Hausman and Wieland (2014) for an exception

[3]This identification scheme is based on the New Keynesian model of Eggertsson (2011). See Appendix section D for a summary on the most important aspects of the Eggertsson model.

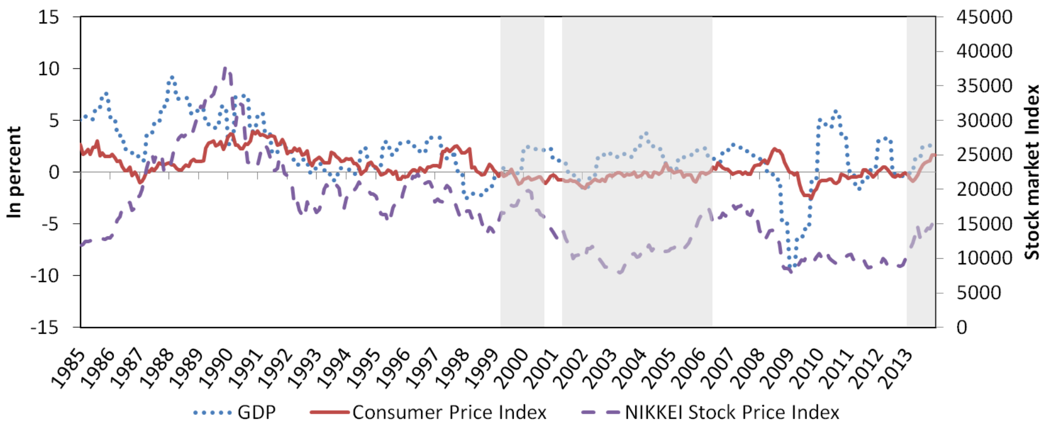
[4]Note, our focus is on the influence of QE as to other non-standard measures of unconventional monetary policy tools. The literature (see e.g. Bernanke and Reinhart (2004) and Bernanke et al. (2004) among others) generally divides unconventional monetary policy tools into three categories: (1) commitment to future policy stances, (2) QE, and (3) credit easing. A combination of these tools defines the term unconventional monetary policy.

1.3 Overview of Japanese Monetary Policy developments

This section briefly reviews the key developments of Japanese monetary policy over the last two decades. For a thorough discussion please refer to Mikitani and Posen (2000), Ugai (2007) and Ueda (2012). The bursting of the Japanese stock market bubble and the accompanying period of economic distress can be seen in Figure 1.1. The stock market was rising dramatically until around 1990. This went together with a rapid increase in industrial production under fairly low and constant rates of inflation. Realising that the elevated stock and land prices seemed out of touch with fundamentals the BoJ did in fact continuously increase the call rate.

Optimism turned into pessimism around 1990/1991 with both stock and land prices starting to fall rapidly. The Japanese economy was finally falling into deep recession. Japan had entered what is by now labeled 'Japan's lost decade'. Whilst Japanese GDP grew by an average rate of 3.9% per year in the pre-1991 period, growth slowed down to only 0.8% post-1991. Meanwhile the usually low Japanese unemployment rate more than doubled while the core inflation rate steadily trended below zero since 2000.

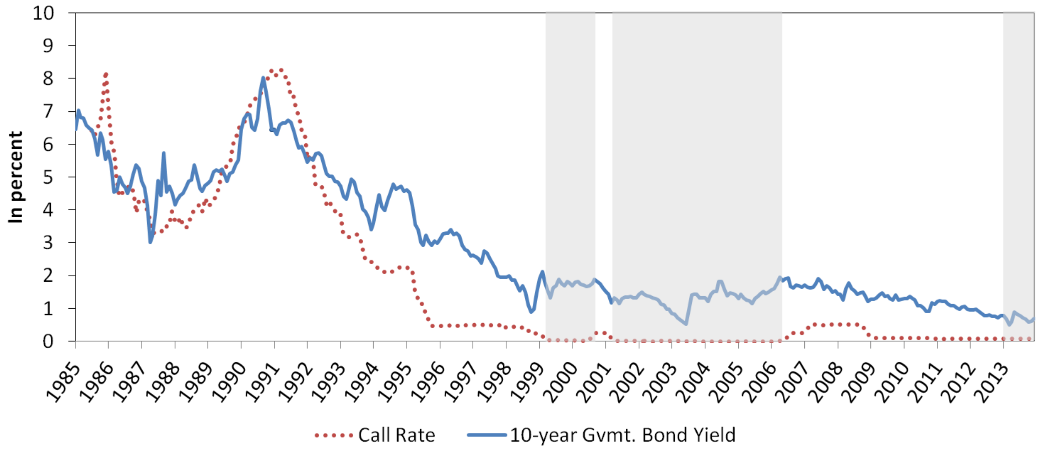
Figure 1.1: Industrial Production, Consumer Price Index and NIKKEI Stock Index



Shaded areas: 1999-2000 ZIRP, 2001-2006 1st QE period, 2013-end 'Abenomics' period

The initial response of the BoJ to the bursting of the asset price bubbles and the recession was rather slow and not very aggressive (Jinushi et al., 2000). In fact, Figure 1.2 shows that the call rate was high until 1992/1993 and decreased only very gradually until it reached 0.5% in the course of 1995.

Figure 1.2: Bank of Japan Short- and Long-Term Interest Rates



Shaded areas: 1999-2000 ZIRP, 2001-2006 1st QE period, 2013-end 'Abenomics' period

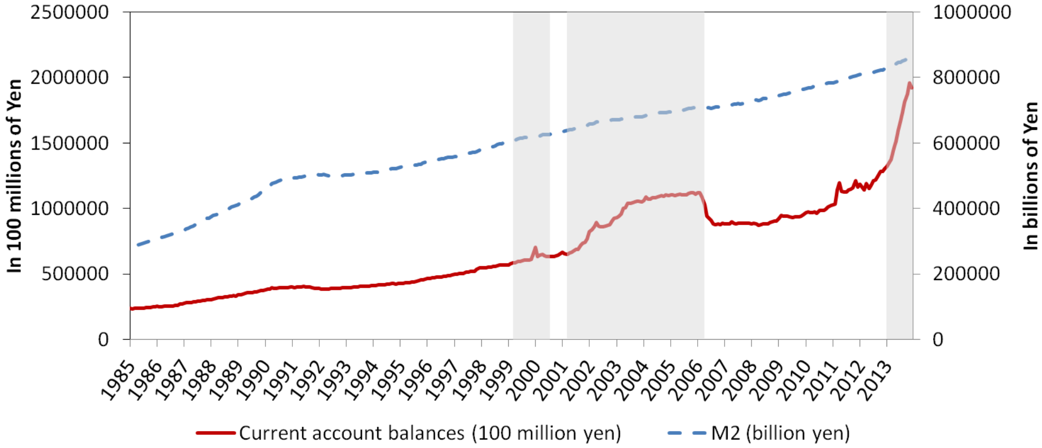
From February 1999 to August 2000, the BoJ officially introduced its so-called 'Zero Interest Rate Policy' (ZIRP) when it lowered the call rate to 0.03% (see Figure 1.2). It also tried to steer market expectations by adding commitments to its policy statements indicating that it would keep the call rate low for a longer time.

After a short-lived economic recovery and following the worldwide bursting of the IT-stock market bubbles, the BoJ introduced a more aggressive policy program. From March 2001 until March 2006 it implemented the so-called 'Quantitative Easing Policy' (QEP) which consisted of three main elements: (i) the operating target was changed from the call rate to the outstanding current account balances held by banks at the BoJ,[5] (ii) to commit itself to continue providing ample liquidity to banks until inflation stabilised at 0% or a slight increase, and (iii) to increase the amount of outright purchases of long-term Japanese government bonds.[6] The monetary development and the effect of the BoJ's QEP measures can be seen in Figure 1.3. We plot that part of the monetary base that relates to the current account holdings of banks at the BoJ. The figure shows the enormous increase in those reserves during the QEP period and later again when the recent financial crisis hit. At the same time the figure plots the evolution of the broader monetary aggregate M2 which can be seen not to be reacting in any obvious manner to the increases in bank reserves.

Most recently, the BoJ implemented its part of the so-called 'Abenomics' program. 'Abe-

[5]Current account holdings is the technical label for Japanese bank reserves being held at the BoJ.
[6]See the thorough survey by Ugai (2007) for more details.

Figure 1.3: Monetary Aggregates in Japan



Shaded areas: 1999-2000 ZIRP, 2001-2006 1st QE period, 2013-end 'Abenomics' period

nomics' essentially stands for a broader package of three different policy measures (also called the three arrows): monetary policy, fiscal policy and structural reforms of goods and labor markets. Concerning the arrow of monetary policy, Shinzo Abe, after having been elected prime minister in December 2012, appointed Haruhiko Kuroda as BoJ Governor in March 2013.

The most important monetary policy decision was then to raise the BoJ's inflation target from 1 to 2%.[7] The BoJ then basically committed itself to achieving this inflation target as soon as possible and not later than within two years. This was officially phrased in the BoJ's 'Quantitative and Qualitative Monetary Easing' policy statement which laid out the details of how the new inflation target is going to be implemented. Essentially, the BoJ will conduct money market operations so that the monetary base will increase at an annual pace of about 60-70 trillion yen. Further specifics are a maturity extension of the BoJ's holdings of Japanese Government Bonds and its commitment to continue expanding the monetary base until inflation stabilises at its 2% target. Although the preliminary evaluation of this much more expansionary monetary policy stance is widely regarded as supportive for the growth stimulus to the Japanese economy, it is unclear how these measures will play out in the medium and longer run (see also Hausman and Wieland (2014) for a preliminary evaluation of 'Abenomics').

Having these macroeconomic and monetary developments in mind we next present our identification strategy based on the reasonable assumption that the BoJ since 1995 did

[7]The BoJ previously was given an inflation target of 1% in February 2012.

not conduct its monetary policy through the call rate anymore - which was constrained by the ZLB - but by changing the reserve holdings of banks at the BoJ.

1.4 Empirical Model

We use the TVP-VAR model with stochastic volatility to trace the reaction of key economic variables to QE shocks over time. Our empirical approach closely follows Primiceri (2005) and Nakajima (2011).[8] The advantage of this framework is its flexibility to deal with the changing nature of the monetary transmission mechanism. This is particularly important for a country like Japan where monetary policy underwent significant changes in its stances.

The TVP-VAR model has both time-varying coefficient matrices as well as time-varying covariance matrices. The varying coefficients account for possible nonlinearities or time-variation in the lag structure of the model and the varying variance covariance matrices capture possible heterosecedasticity of the shocks and nonlinearities in the simultaneous relationships among the variables.

We estimate the following VAR model:

$$y_t = c_t + B_{1,t}y_{t-1} + \ldots + B_{l,t}y_{t-l} + u_t, \quad t = 1, \ldots, T \tag{1.1}$$

where y_t is a n x 1 vector of endogenous variables; c_t, is a n x 1 vector of time-varying intercepts; $B_{i,t}$, is a n x n matrix of time-varying coefficients with lag length $i = 1, \ldots, l$; and u_t, is a n x 1 vector of residuals. Ω_t describes the time-varying covariance matrix of u_t, which can be decomposed into:

$$VAR(u_t) \equiv \Omega_t = A_t^{-1}\Sigma_t\Sigma_t'(A_t^{-1})'. \tag{1.2}$$

Where A_t is a time-varying lower triangular matrix and Σ_t is a diagonal time-varying

[8]Further related studies are Franta (2011) and Canova and Ciccarelli (2009).

covariance matrix:

$$A_t = \begin{bmatrix} 1 & 0 & \dots & 0 \\ \alpha_{21,t} & 1 & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ \alpha_{n1,t} & \dots & \alpha_{n(n-1),t} & 1 \end{bmatrix} \quad \Sigma_t = \begin{bmatrix} \sigma_{1,t} & 0 & \dots & 0 \\ 0 & \sigma_{2,t} & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ 0 & \dots & 0 & \sigma_{n,t} \end{bmatrix} \tag{1.3}$$

The time-varying VAR can then be rewritten as:

$$y_t = X_t' \tilde{B}_t + A_t^{-1} \Sigma_t \varepsilon_t, \tag{1.4}$$

$$X_t' = I \bigotimes [1, y_{t-1}', \dots, y_{t-l}'],$$

where $\tilde{B}_t$ is a stacked vector containing all coefficients of the right hand side of equation 1.1. $VAR(\varepsilon_t) = I_n$ and the operator $\otimes$ denotes the Kronecker product.

The dynamics of the time-varying parameters (B_t and A_t) are following a driftless random walk, whereas the covariance matrix (Σ_t) evolves as a geometric driftless random walk:

$$B_t = B_{t-1} + \nu_t, \tag{1.5}$$

$$\alpha_t = \alpha_{t-1} + \xi_t, \tag{1.6}$$

$$log\sigma_t = log\sigma_{t-1} + \eta_t, \tag{1.7}$$

where α_t is a stacked vector of the lower triangular coefficients of the matrix A_t and the standard deviation σ_t is the vector of the diagonal elements of the matrix Σ_t. The vector of innovations $[\varepsilon_t', \nu_t', \xi_t', \eta_t']$ is assumed to be jointly normally distributed with variance-covariance matrix:

$$VAR(\varepsilon_t, \nu_t, \xi_t, \eta_t) = \begin{bmatrix} I_n & 0 & 0 & 0 \\ 0 & Q & 0 & 0 \\ 0 & 0 & S & 0 \\ 0 & 0 & 0 & W \end{bmatrix}, \tag{1.8}$$

where I_n is an n dimensional identity matrix and Q, S and W are positive definite matrices. S is assumed to be block diagonal and in order to reduce the dimensionality of the estimation, we restrict W to be a diagonal matrix.

1.4.1 Priors

For evaluating posteriors, prior distributions need to be specified. For the calibration of these priors, we use a training sample based on the period from 1980Q1 to 1995Q4 (see Appendix A) and run an OLS estimation on a fixed-coefficient VAR model.

The OLS point estimates ($\hat{B}_{OLS}$) and four times their variance specify the mean and the variance of B_0. We assume the same specification for the prior distribution of the simultaneous relation matrix A_0. The prior mean for the log standard errors is the log of the OLS point estimates ($\hat{\sigma}_{OLS}$), and the prior covariance matrix is specified to be $4 \cdot I_n$. The priors for the initial states of the time-varying VAR-parameters B_0, A_0 and $log\sigma_0$ follow a normal distribution. The hyperparameters Q, S and W are the covariance matrices of the innovations (see equations 1.5, 1.6 and 1.7). Matrices Q and S are distributed as an independent inverse-Wishart and W is assumed to follow an inverse-Gamma prior distribution. In summary:

$$
\begin{aligned}
B_0 &\sim N(\hat{B}_{OLS}, 4 \cdot V(\hat{B}_{OLS})), \\
A_0 &\sim N(\hat{A}_{OLS}, 4 \cdot V(\hat{A}_{OLS})), \\
log\sigma_0 &\sim N(\hat{\sigma}_{OLS}, 4 \cdot I_n), \\
Q &\sim IW(k_Q^2 \cdot \tau \cdot V(\hat{B}_{OLS}), \tau), \\
W &\sim IG(k_W^2 \cdot (1 + dim(W)) \cdot I_n, (1 + dim(W))), \\
S_b &\sim IW(k_S^2 \cdot (1 + dim(S_b)) \cdot V(\hat{A}_{b,OLS}), (1 + dim(S_b))),
\end{aligned}
$$

where τ has the size of the training sample, S_b with the index b refers to the corresponding blocks of a particular equation and $\hat{A}_{b,OLS}$ denotes the respective blocks of $\hat{A}_{OLS}$.[9] The degrees of freedom for W and S_b are specified as one plus its respective matrix dimension. The size of the training sample defines the degrees of freedom for Q. Finally, the parameters $k_Q = 0.01, k_W = 0.1$ and $k_S = 0.01$ define prior beliefs about the degree of time variation in the parameters, covariances and volatilities. For example, for

[9] The system consists of three blocks with the respective size: 2, 3 and 4.

the OLS estimation of the covariance matrix of the VAR coefficients, we allow for 1% ($k_Q = 0.01$) of uncertainty surrounding the $V(\hat{B}_{OLS})$ estimates to time variation (Kirchner et al. (2010)).[10]

We conduct a formal model selection since there are no economic reasons for choosing one ($k_Q, k_W,\ k_S$) combination over another. Posterior probabilities for a set of 18 models are estimated based on the reversible jump Markov chain Monte Carlo (RJMCMC) method (see Primiceri (2005)).[11] The selection of k_Q, k_W and k_S delivers a posterior probability for one combination which is almost one. Table 1.9 in the Appendix C reports the posterior probability estimates for the set of 18 models.

1.4.2 Identification and Estimation

So far, we have outlined the estimation strategy for a reduced form VAR which is estimated using Bayesian methods for the sample from 1996:Q1 to 2013:Q4. For maintaining the degrees of freedom, two lags are used. For approximating the posterior distribution, $40,000$ iterations of the Gibbs sampler are used and we drop the first $20,000$ iterations for convergence. For breaking the autocorrelation of the draws, only every $10th$ iteration is kept. Our final estimates are therefore based on $2,000$ iterations. The sample autocorrelation functions of the draws die out rather quickly. Furthermore, the convergence diagnostics reveal satisfactory results (a detailed overview is given in Appendix J).

To identify a QE shock, we follow Schenkelberg and Watzka (2013) and use their sign restriction approach when the economy is stuck at the ZLB. Essentially the idea is to let the variables of interest, in particular real GDP, unrestricted, whilst imposing relatively mild sign restrictions on the remaining variables. In addition to real GDP, the real effective exchange rate is left unrestricted as it might give us some indication of whether QE works through depreciation and the stimulating effects on exports. The sign restriction to identify the QE shock is imposed on bank reserves held at the BoJ, which have been the key monetary operating instrument of the central bank since 2001. More specifically, a QE shock is defined as a 1% increase in bank reserves. We restrict the price level to respond non-negatively to a positive QE shock. Since the price level is empirically known

[10] As a sensitivity check, we also experimented with other value combinations of these coefficients. The responses obtained are robust to those presented.

[11] The set of 18 models are constructed from all possible combinations of $k_Q = \{0.01; 0.05; 0.1\}$, $k_W = \{0.001; 0.01\}$ and $k_S = \{0.01; 0.025; 0.1\}$.

to move sluggishly, we also allow for a zero impact effect.[12]

In addition to the QE shock, we identify two business cycle disturbances: a positive demand and a positive supply shock. These are identified for two reasons: (1) to avoid that disturbances in business cycle fluctuations enter the identified QE shock and (2) to evaluate the explanatory power of the QE shock relative to the demand and supply shocks. The aggregate demand and supply shocks are identified according to the New Keynesian predictions of an economy at the ZLB (see Appendix D for a short overview of the Eggertsson (2011) model).[13] All sign restrictions are binding for three quarters after the shock. Table 1.1 summarises the restrictions.

Table 1.1: Sign Restrictions

		QE Shock	DE Shock	SP Shock
	Impact	0		
Prices	$K = 3$	$\geq$	$>$	$\leq$
GDP	$K = 3$	?	$\geq$	$<$
Reserves	$K = 3$	$>$	?	?
Exchange Rate	$K = 3$	?	?	?
Further restriction	$K = 3$		$\left\lvert\frac{\text{GDP}}{\text{CPI}}\right\rvert > 1$	$\left\lvert\frac{\text{GDP}}{\text{CPI}}\right\rvert < 1$

Note: ? denotes no restriction, $>$ defines a positive effect of the respective shock on the variable, vice versa for $<$. $K = 3$ indicates that the restriction horizon is three quarters. $|\text{GDP/CPI}|$ denotes the absolute value of the ratio between the GDP response and the CPI response.

For implementing the sign restrictions, we slightly modify the model specified in equations 1.4-1.7. So far, it is based on the recursive identification. We additionally specify an orthonormal rotation matrix G_t, i.e., $G_t'G_t = I_n$. The model in equation 1.4 can then be rewritten as

$$y_t = X_t'\tilde{B}_t + A_t^{-1}\Sigma_t G_t' G_t \varepsilon_t = X_t'\tilde{B}_t + A_t^{-1}\Sigma_t G_t' \tilde{\varepsilon}_t. \tag{1.9}$$

$\tilde{\varepsilon}_t = G_t'\varepsilon_t$ denotes the new shocks and the respective variance is $Var(\tilde{\varepsilon}_t) = G_t I_n G_t'$. We use the QR decomposition for finding G_t. Since we have a four variable VAR, G_t is a 4 x

[12]Ideally, we would include more variables to shed more light on the transmission mechanism, but are limited here by degrees of freedom problems from our estimation method.

[13]Because the main focus in this study is on the effects of QE on the macro economy, we will not go into a detailed discussion of any of the other shocks here. For a thorough overview of the demand and supply shocks, please refer to Eggertsson (2011) and Schenkelberg and Watzka (2013). For a brief summary on the identification strategy of the demand and supply shocks, refer to Appendix D.1.

4 matrix:

$$G_t = \begin{pmatrix} QR\left(\theta\left[1,1\right]\right) & \dots & QR\left(\theta\left[1,4\right]\right) \\ \vdots & \ddots & \vdots \\ QR\left(\theta\left[4,1\right]\right) & \dots & QR\left(\theta\left[4,4\right]\right) \end{pmatrix}. \tag{1.10}$$

In a first step we draw a 4 x 4 matrix, θ, from the $N(0,1)$ distribution. Step two: we take the QR decomposition of θ and construct the G_t matrix. This algorithm calculates a candidate structural impact matrix. Step three: it is checked whether this matrix is in line with the sign restrictions. Step four: if it satisfies the restrictions it is stored. Otherwise another θ is drawn from the standard normal distribution and we repeat the procedure from step two.

1.5 Empirical Evidence of Time Variation in Japan: is there any?

As a first step, we search for formal econometric evidence on whether the impact of a QE shock in Japan has changed across time. In particular, we calculate marginal likelihood estimates for a traditional constant-coefficient VAR model and our time-varying parameter (TVP-VAR) model with stochastic volatility.[14] The model that yields the largest marginal likelihood fits the given data best. We follow Nakajima et al. (2011) and use the modified harmonic mean estimator of the marginal likelihood due to Geweke (1999).[15] The log marginal likelihood value for the TVP-VAR, -557.5, is higher than the marginal likelihood estimate for the constant VAR, -700.062, suggesting that the TVP-VAR model with stochastic volatility is indeed a better model for Japan than the constant VAR.

[14] Prior for the constant parameter VAR: $B \sim N(0, 4 \times I)$, $\alpha \sim N(0, 4 \times I)$, $\sigma^{-1} \sim \text{Gamma}(2, 0.02)$

[15] For a detailed description of the harmonic mean estimator, please refer to Nakajima et al. (2011). The marginal likelihood calculation is based on the priors and number of lags as specified above. Additionally, we have to specify the parameter τ. We follow Nakajima et al. (2011) and set $\tau = 0.99$.

1.6 Results of the TVP-VAR

In what follows, section 1.6.1 presents the estimated median impulse responses of the QE shock (see also Appendix E). A key focus is on the different effects following the same QE shock over the different monetary policy stances of the BoJ during the entire ZLB period. It also includes an analysis on the posterior probability for the difference in the impulse responses. Section 1.6.2 shortly summarises the key findings of the demand and supply shocks. For evaluating the quantitative importance of the QE shock relative to the two business cycle disturbances, we present a forecast error variance decomposition (FEVD) in section 1.6.3. Section 1.6.4 briefly links our results to the theoretical framework of Eggertsson (2011). The time-varying posterior estimates of the covariance matrix are presented in Appendix G and Appendix J summarises the estimation on convergence diagnostics.

1.6.1 Impulse Responses to Quantitative Easing Shocks

Figure 1.4 presents the median impulse responses (over 17 quarters and the time period: 1996:1-2013:4) to a 1% increase in reserves in the given period across the sample.[16]

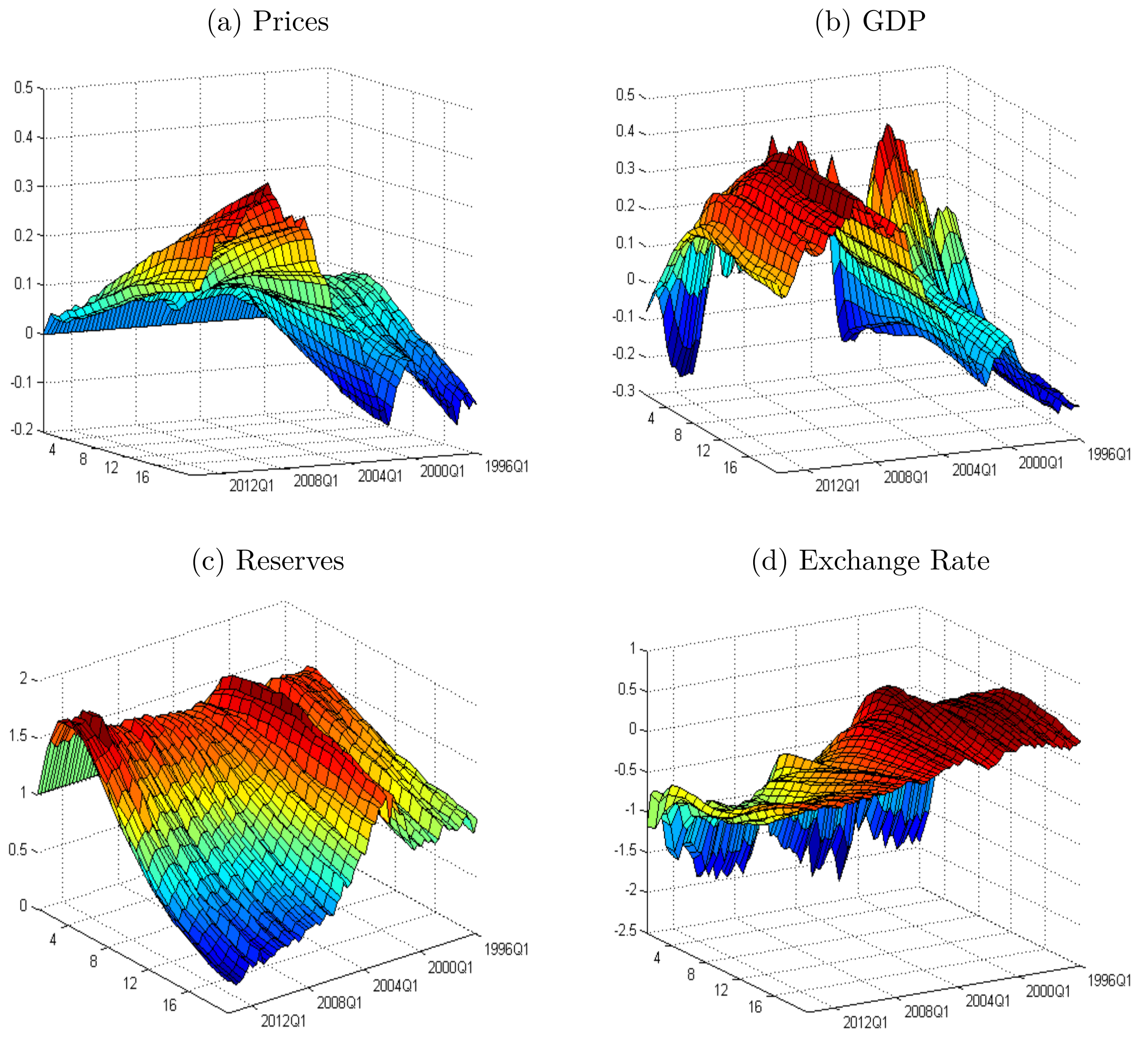
We clearly see that a QE shock has time-varying effects. Regarding the effect of a QE shock on *prices*, they exhibit time variation across our sample. The response of prices has been restricted to increase for three quarters following the QE shock, so the immediate rise is by construction (Figure 1.4a). The positive impact seems to be significant after four quarters and for eight quarters from 2006 onwards (Figures E.5a and E.5b in Appendix E plot the evolution of the responses at the 4th and 8th quarter with percentiles).[17] The long run impact on prices appears to change across our sample. More specifically, during the ZIRP from 1999 to 2000, the cumulative effect on prices seems to be close to 0.05% after three years. Since 2000, the long run impact seems to strongly decrease in size. Note that in August 2000 the ZIRP ended. From 2002 onwards, the long run impact appears to increase again. This effect may mirror the introduction of the first QE program in 2001. In 2003Q3 the effect on prices stand at about 0.15% after three years. Overall, our results

[16]The impulse responses are conditional on the current parameters from the relevant quarter. In the case where $t + T_{hor} > T$ the parameters from the last quarter are used for convenience.

[17]These figures give the 16th and 84th percentiles of the posterior distribution of the impulse responses as these confidence bands are standard in the literature. Based on the normality assumption, these percentiles refer to one-standard error bands (see also Uhlig (2005)).

suggest that the changes in the transmission mechanism of a QE shock may capture the end of the ZIRP and the beginning of the first QE program respectively. It is very likely that these monetary policy decisions influenced the economy.[18]

Figure 1.4: Time-Varying Impulse Responses to a Monetary Policy Shock



Median impulse responses to a 1% reserve shock.

Following a QE shock, *real GDP* initially increases for most of the periods and after 1999 the cumulative effect on output seems to be positive. More specifically, during the ZIRP from 1999 to 2000 the initial response of GDP seems to become stronger. With

[18] The BoJ implemented its ZIRP from February 1999 to August 2000 and the first QE program between March 2001 and March 2006. Note that we cannot explicitly control for the monetary policy stances in our estimation method. However, it is very likely that these episodes considerably influenced the transmission mechanism of QE shocks. The estimated changes in our impulse responses seem to occur exactly during these periods.

the end of the ZIRP, in the end of 2000, the initial response of GDP appears to weaken substantially. Since 2002, we see a reversal of this trend: output seems to become initially more responsive from about -0.2% in 2002 to 0.2% in 2003. Between 2004 and 2006, the initial impact of the QE shock has a positive effect on output of about 0.3% after four quarters. Especially the medium term and long run impact of the QE shock seems to be stronger from 2004 onwards (Figure 1.4b). The significance of this increase is also reported in Figures E.5a and E.5b (Appendix E). This positive significant impact on output also seems to hold for the year 2013. Interestingly, we can conclude that QE shocks appear to become visibly more effective as an output stimuli from 2006 onwards (Figures E.5a and E.5b). This is in contrast to the findings of Kimura et al. (2003) and Kimura and Nakajima (2013). However, also Franta (2011) finds an initial positive and significant impact on output. In line with our results, also Schenkelberg and Watzka (2013) estimate a significant impact on output after about two years.

Note that from about 2008 to 2011, the CPI and GDP responses seem to be larger compared to the other periods. A possible explanation might be that the global financial crisis in 2007/2008 led to a world-wide recession. This recession was to a large extend driven by a decline in aggregate demand. Under such conditions of insufficient aggregate demand - as opposed to structural problems of the economy - monetary policy tends to have a large effect. Our results possibly reflect this situation.

Finally, following a QE shock, the *real effective exchange rate* seems to initially depreciate (Figure 1.4d).[19] While the response of the exchange rate to a QE shock is insignificant thereafter, it is only since recently that the depreciation stays significant until about one year following the shock (Figures E.5a and E.5b). This finding suggests that a QE shock during the 'Abenomics' strategy leads to a somewhat longer depreciation of the Yen. The depreciation may in turn stimulate economic activity at the ZLB.

Comparison of impulse responses at different points in time

For a better illustration of the difference in the impulse responses to a QE shock across time, we also present Figures 1.5a and 1.5b. These allow for a comparison of impulse responses at specific points in time. Figure 1.5a plots the median impulse responses at 1999Q4, 2003Q3 and 2013Q3. The period around 1999Q4 reflects the environment under

[19] A negative value of the real effective exchange rate response reflects a depreciation of the Yen.

the ZIRP period, 2003Q3 under the influence of the first QE program and 2013Q3 under the influence of the current economic policies advocated by Shinzo Abe. The three different time periods for the comparison are chosen arbitrarily within each monetary policy stance. Figure 1.5b plots the impulse responses at 2003Q3 and 2013Q3 with percentiles.[20]

In terms of prices, as we described above, the positive and significant effect at the long run is visible in Figure 1.5b below. In contrast, the estimated impact of a monetary base shock in Kimura et al. (2003), Nakajima et al. (2011) and Kimura and Nakajima (2013) does not reveal a strong effect on prices. However, their results could be biased. Kimura et al. (2003) use a time-varying coefficient VAR but with a constant variance covariance matrix. Since the simultaneous relation matrix is time invariant in this case, it could lead to an underestimation of the shocks (see Primiceri (2005) for further details). In contrast, Nakajima et al. (2011) and Kimura and Nakajima (2013) use a fully time-varying VAR which allows for stochastic volatility. However, their results could be driven by the recursive identification system. In accordance with our findings are the results by Franta (2011) and Schenkelberg and Watzka (2013). Both approaches apply sign restrictions and estimate an initial positive and significant impact on prices. Contrary to Franta (2011), our results and Schenkelberg and Watzka (2013) also confirm a permanent significant increase in prices.[21]

Especially for the medium term and long run responses on both prices and GDP, there seems to be a significant difference between 2003Q3 and 2013Q3 (Figure 1.5a). The confidence bands in Figure 1.5b reveal a stronger significant impact of a QE shock in 2013Q3 compared to 2003Q3 and 1999Q4. More specifically, the effect on prices seems to be permanent in 2013Q3, whereas the impact in 2003Q3 becomes insignificant after seven quarters. Following a QE shock on GDP in 2013Q3, we estimate a positive and significant impact after seven quarters. Compared to 2013Q3, the effect on GDP in 2003Q3 is completely insignificant. Regarding the exchange rate, a QE shock leads to a somewhat longer depreciation in 2013Q3 than for example during earlier periods. This finding suggests that, in contrast to earlier policy programs by the BoJ, the 'Abenomics' strategy seems to have an impact on the exchange rate. These results are also supported

[20] Figure F.6 in Appendix F presents the average impulse response functions during the three respective monetary policy stances (ZIRP, first QE program, 'Abenomics'). They show that the arbitrarily taken periods (1999Q4, 2003Q3 and 2013Q3) reveal the same findings as the average response during the three monetary policy stances.

[21] See Table 1.6 to 1.8 in Appendix B for a more detailed overview of the results obtained in other studies focusing on the monetary policy transmission in Japan.

Table 1.2: Posterior probability for the difference in the impulse responses to a QE shock at different time periods

Horizon	1 Q (%)	4 Q (%)	8 Q (%)	12 Q (%)	16 Q (%)
HICP					
1999/2003	9.5	44.6	43.7	46.7	48.6
1999/2013	9.2	52.5	53.7	61.0	64.5
2003/2013	9.2	52.5	53.7	61.0	64.5
GDP					
1999/2003	36.8	40.6	46.4	49.4	50.5
1999/2013	9.4	41.7	66.1	67.4	66.6
2003/2013	9.4	41.7	66.1	67.4	66.6
Res					
1999/2003	46.0	47.1	46.4	47.6	48.0
1999/2013	67.2	70.1	60.5	51.4	47.6
2003/2013	67.2	70.1	60.5	51.4	47.6
ExR					
1999/2003	34.5	37.5	39.8	40.3	40.6
1999/2013	31.2	26.1	22.0	20.0	25.9
2003/2013	31.2	26.1	22.0	20.0	25.9

Note: Difference in impulse responses at the time periods 1999Q4, 2003Q3 and 2013Q3 for one, four, eight, 12 and 16 quarters ahead.

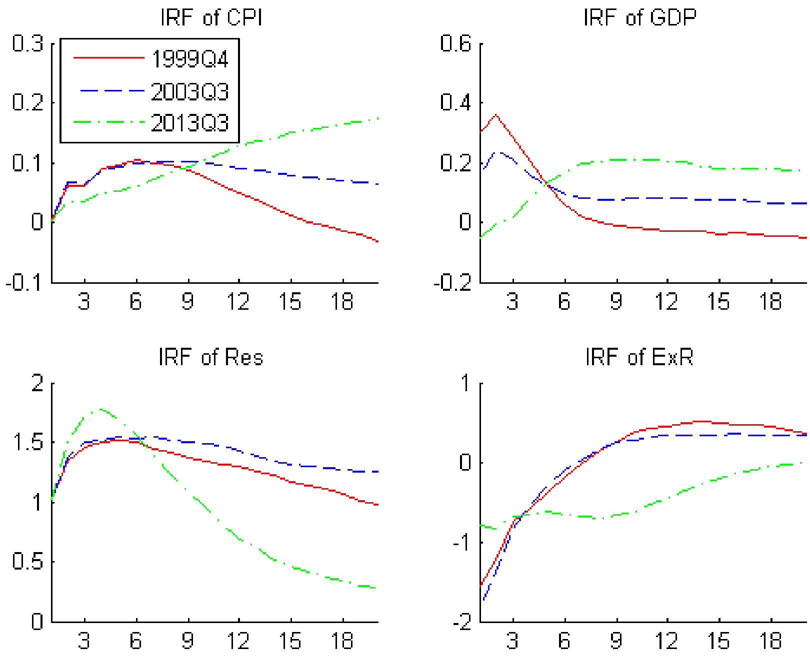
by our analysis on the posterior probability for the difference in the impulse responses.

We consider the statistical difference in the impulse responses between different time periods by calculating the ratio of the Markov Chain Monte Carlo draws (MCMC) of the responses between two time periods. More specifically, we estimate the posterior probability that the response at one given time period (first considered response) is smaller than at another given time period (second considered response). We consider again the three time periods referred to above and present the posterior differences in the impulse responses to the QE shock in Table 1.2. Posterior probability values close to 50% indicate a weak difference between the two periods. Values above (below) 50% imply that the first response is smaller (bigger) than the second response. Regarding prices, we estimate a strong difference in the initial responses of the three considered time periods as well as in the long run responses in 1999/2013 and 2003/2013, whereas the responses in 1999/2003 are quite similar in the long run. More specifically, values above 60% for 12 and 16 quarters ahead in 1999/2013 and 2003/2013 indicate a strong difference in the impulse responses. The estimated response in 2013Q3 is larger than in 2003Q3 or 1999Q4. Following a QE shock on GDP, a strong posterior difference is reported for all initial responses of the three considered time periods as well as for the time periods 1999/2013 and 2003/2013 from eight quarters ahead. The difference between 1999 and 2003 seems to be again very small in the long run. The evidence for time variation of the exchange rate responses is rather strong between all compared time periods. The responses in 2013 confirm a larger depreciation of the exchange rate than in 2003 and 1999.

Generally, our results suggest that a QE shock during the recent ‘Abenomics’ period seems to lead to a larger impact on output, prices and the exchange rate than during the ZIRP or the first QE program. A theoretical discussion of our results, based on the Eggertsson (2011) model, follows in section 1.6.4 below.

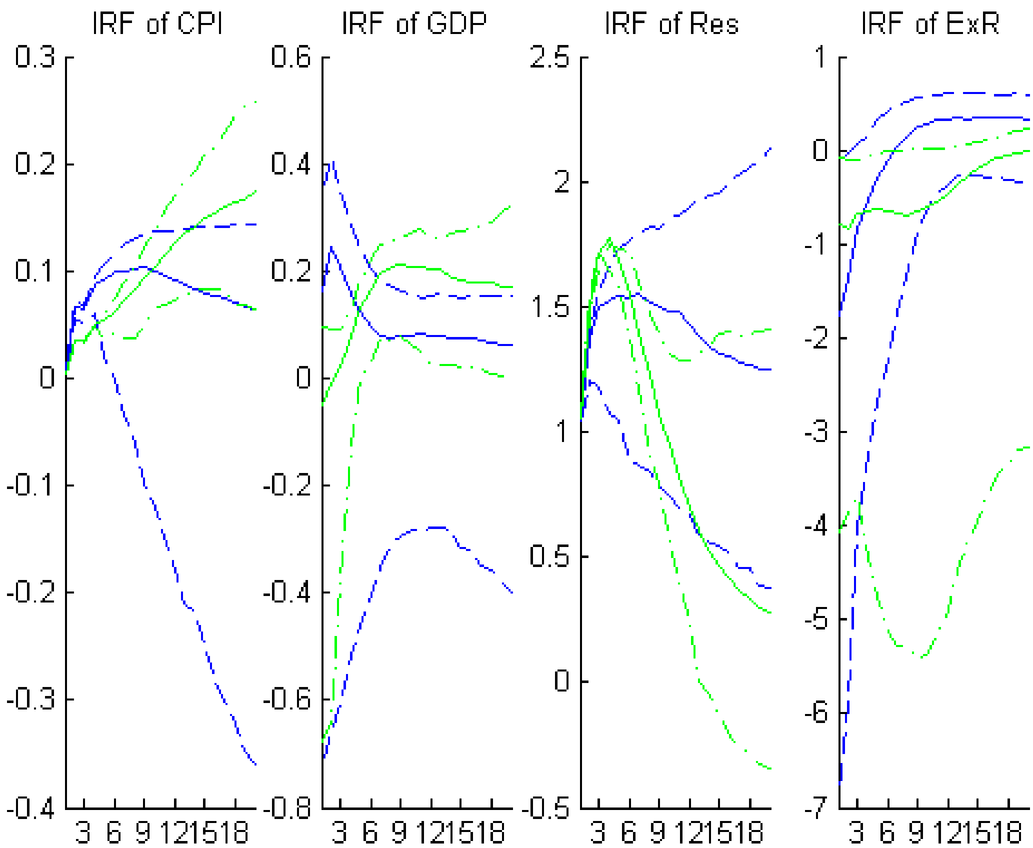
Figure 1.5: Responses at Different Time Periods to a QE Shock

(a) Responses without Percentiles



Median impulse responses to a 1% QE shock at 1999Q4, 2003Q3 and 2013Q3.

(b) Responses with Percentiles



Median impulse responses (solid line) to a 1% QE shock with 16-th and 84-th percentiles (dashed line) of the posterior distribution at 2003Q3 (blue) and 2013Q3 (green).

1.6.2 Impulse Responses to Demand and Supply Shocks

The impulse response functions for the demand and supply shocks are given in the Appendix H, Figures H.10 and H.11. Primarily, we identify these two shocks for (1) evaluating the explanatory power of the QE shock compared to the demand and supply shocks (see section 1.6.3 for details on the variance decomposition analysis) as well as (2) for ensuring that these business cycle fluctuations do not enter the identified QE shock. In the following, a brief overview of the main effects of the demand and supply shocks is given.

The responses of CPI and GDP is initially restricted to be positive following a demand shock, thus the immediate increase is by construction. More specifically, CPI rises by up to 2% in 1999Q4 and 2003Q3 and stays significantly above zero for a much longer period than restricted. Generally, this effect also holds for 2013Q3 except that it converges a bit faster back to the zero line. Turning to GDP, it becomes insignificant after the restriction horizon. Reserves and the exchange rate seem to respond insignificantly to a demand shock.

Figure H.11 plots the impulse response functions following a supply shock. By comparing the absolute size of the initial CPI and GDP responses to the demand and supply shocks in Figures H.10 and H.11, the determining sign restrictions can be seen. Following a demand shock, CPI is restricted to respond less strong than GDP and vice versa following a supply shock. The responses of CPI and GDP are as expected significantly negative. The impulse responses remain significant after the restriction horizon. Following a supply shock, we report a non visible effect on reserves and the exchange rate.

Overall, the transmission mechanism of the demand and supply shocks is similar to the findings of Schenkelberg and Watzka (2013).

1.6.3 Results on the Forecast Error Variance Decomposition

The forecast error variance decomposition (FEVD) allows us to analyse the explanatory power of the QE shock relative to the other structural shocks. We use the close-to-median impulse responses in this section for assessing the relative quantitative importance of our three structural shocks. This procedure generates the impulse responses which are closest to the median impulse response functions. By using the close-to-median impulses, we ensure that the variance shares add up to one as well as the orthogonality of the identified shocks. Table 1.3 summarises the estimated forecast error variance shares of all variables for each of the three identified shocks, for the horizons of 4, 8, 12 and 16 quarters as well as for the time periods 1999Q4, 2003Q3 and 2013Q3. The last column of each time period gives the sum of the variance shares of every endogenous variable over all identified shocks.

To give an example, overall, our three shocks account for up to 98% of the GDP variability in 1999Q4.[22] More specifically, the QE shock explains up to 36% of the GDP variability in 1999Q4 and decreases substantially for 2003Q3. But the explanatory power on GDP in 2013Q3 is considerably larger than during the first QE program in 2003. Also 2013Q3 seems to have an overall larger explanatory power than during the ZIRP in 1999Q4. Regarding the quantitative importance of the QE shock on prices in 2003Q3, it does not appear to play a major role, explaining only 3% to 7%.[23] However, especially at longer horizons it seems that the variance shares for prices are higher in 2013Q3 and 1999Q4 compared to 2003Q3, accounting for up to 22% of the variability in 2013Q3.

The demand shock accounts for a non-negligible part of the variability in GDP with variance shares of up to 47% in 2003Q3. Similarily for the variance shares of CPI to a supply shock, these range from 48% to 10% in 2003Q3. It seems that the explanatory power of the demand shock on CPI and GDP decreases from 2003Q3 to 2013Q3 vice versa for the supply shock. As expected, the QE shock confirms the relatively important role on the variability of reserves; variance shares range from 49% to 47% in 2013Q3. In contrast, the other two identified shocks are of minor importance for the reserve variability. These findings are also in line with Schenkelberg and Watzka (2013).

Summing up, in 2013Q3, a QE shock seems to explain a larger variance share of both CPI and GDP. The relative importance of the QE shock in 2013Q3 on prices increases at longer horizons compared to 2003Q3. Also for GDP, the explanatory power of the QE shock is substantially higher in 2013Q3 than in 2003Q3. These measures possibly mirror the effects of the current 'Abenomics' program and support our previous findings.

[22]The sum would equal one when next to the three identified shocks also the unidentified disturbance, the exchange rate shock, is added.

[23]Generally, relatively low numbers where also estimated by Schenkelberg and Watzka (2013). However, as Peersman and Straub (2006) illustrate, sign restriction approaches often lead to rather small variance shares for some variables.

Table 1.3: Forecast Error Variance Decomposition

		1999Q4				2003Q3				2013Q3			
Variable	Horizon	QE	DE	SU	Sum	QE	DE	SU	Sum	QE	DE	SU	Sum
CPI	4 quarters	8	28	49	85	3	34	48	86	16	16	48	81
	8 quarters	17	51	25	93	6	45	22	74	16	36	35	86
	12 quarters	22	58	15	95	7	48	14	69	18	42	29	89
	16 quarters	25	59	12	96	7	51	10	69	22	41	26	89
GDP	4 quarters	36	46	16	98	2	42	11	55	44	29	12	85
	8 quarters	34	46	16	96	3	44	10	57	45	22	13	81
	12 quarters	35	46	15	96	3	46	10	59	47	17	16	79
	16 quarters	36	45	15	96	5	47	10	61	46	16	17	79
Res	4 quarters	90	9	1	100	1	0	1	2	49	15	10	74
	8 quarters	89	8	2	99	2	1	2	5	47	17	11	76
	12 quarters	88	7	3	99	5	1	2	8	47	17	12	76
	16 quarters	87	7	4	98	8	2	2	12	47	17	12	76
ExR	4 quarters	7	14	28	49	71	4	20	95	48	12	17	77
	8 quarters	11	14	35	59	54	3	26	83	47	12	18	78
	12 quarters	12	13	36	61	52	3	27	82	47	13	18	77
	16 quarters	12	13	36	62	50	4	26	80	46	14	18	78

Note: Variance shares of the QE shock, the demand and the supply shock. Entries are in percent.

1.6.4 Theoretical Discussion of our Results

Based on Eggertsson's (2011) model, we briefly discuss potential reasons why the observed effect of a QE shock in 2013 seems to have a significantly stronger impact on prices and GDP than in 2003 or 1999.

In the Eggertsson (2011) model, the economy at the ZLB has an upward sloping aggregate demand curve in inflation. That is, higher inflation expectations increase output because, for a given nominal interest rate, the real interest rate decreases due to a higher inflation rate. This stimulates consumption and thus output (refer to Appendix D for a brief summary of the Eggertsson (2011) model). As Krugman (2000) and Eggertsson (2011) illustrate, for fighting the liquidity trap, monetary policy needs to raise expected inflation through a credible commitment to expand the current and future money supply. This in turn will increase the current price level and thus current output. For these real effects to occur, two aspects have to hold: (1) the monetary expansion needs to be perceived by the markets as sustainable and (2) the central bank will not revert to normal practice of stabilising prices as soon as the recession is past. If the central bank cannot credibly promise to be irresponsible, the monetary expansion will be ineffective for fighting the liquidity trap (see subsection D.3 for the effect of a credible monetary policy expansion).

Figure 1.6 presents the development of Japanese inflation expectations. The development of the expectation of a higher inflation rate within the next six months and for a rise in inflation for one year forward suggest a respective increase during the three monetary policy stances of the BoJ (Figure 1.6a). As Figure 1.6b shows, since 2013 expected inflation increased substantially with a pronounced increase in expected inflation between 2 and 5% as well as more than 5%. This can also be confirmed by looking at the overall development of the expected inflation one year forward (Figure 1.6b). Moreover, also the medium and long run inflation expectations in Figure 1.6c reveal a considerable increase since the beginning of the 'Abenomics' strategy. Therefore, inflation expectations appear to be rising on the whole within the last year. Moreover, since the introduction of the 'Abenomics' strategy, also the overall index on consumer prices as well as the core price index already experienced a substantial increase (see Figure 1.6c). For example, the annual growth of CPI increased from -0.7% in April 2013 to 1.6% in December 2013.

Based on Eggertsson (2011) theory, one can argue that the current development in inflation expectations possibly mirrors that markets believe in a sustainable expansionary monetary policy under the 'Abenomics' program and thus to have considerable real effects.[24] This is also in line with the findings of Hausman and Wieland (2014). More specifically, they describe that, compared to the first QE program, the 'Abenomics' strategy seems to be perceived as non-temporary since the broad money supply experienced a remarkable growth since 2013 compared to the period under the first QE program. This is precisely what one would expect from a more credible monetary policy change. A perceived non-temporary QE program leads to expected lower future real interest rates since expected inflation rises. This in turn increases credit demand and thus induces money creation in the banking sector. Thus, the increase in the broad money supply (see Figure 1.7 for the differences in monetary base growth between the two programs). Therefore, a credible commitment to future money expansion may successfully fight the liquidity trap whereas temporary changes in the monetary base may fail to accomplish this.[25]

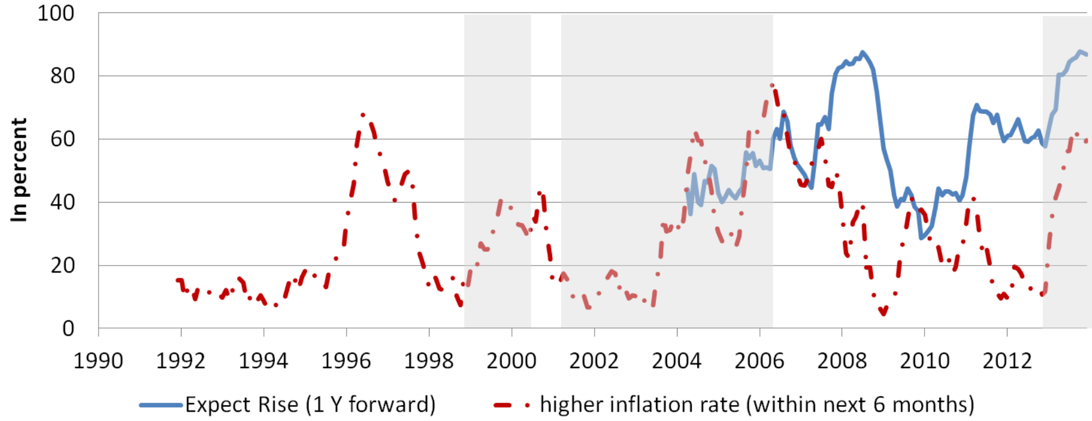
However, it should be stressed that it is still far too early to conclusively judge on the effectiveness of 'Abenomics'. Thus, our results should be seen as a tentative and preliminary evaluation of the short-term effects of 'Abenomics', much in the line and spirit of Hausman and Wieland (2014).

[24]Figure 1.6b also indicates a rise in consumers' inflation expectations for the period between 2007 and 2009. Based on Eggertsson's (2011) theory, this increase may reflect that markets believe in a credible monetary policy expansion during this period as well leading to a rise in prices and output. Note that following a QE shock, we estimate a significant impact on prices and GDP after two years since 2006 (compare Figure C.2b).

[25]As Hausman and Wieland (2014) outline, the first QE program was perceived as temporary. Although the broad money aggregate M3 increased after the beginning of the first QE program, the increase was rather small. As for example Krugman (1998) and Eggertsson (2011) (among others) argue, if QE is perceived to have temporary effects, it will only have small or no influence on expected real interest rates and the broad money supply.

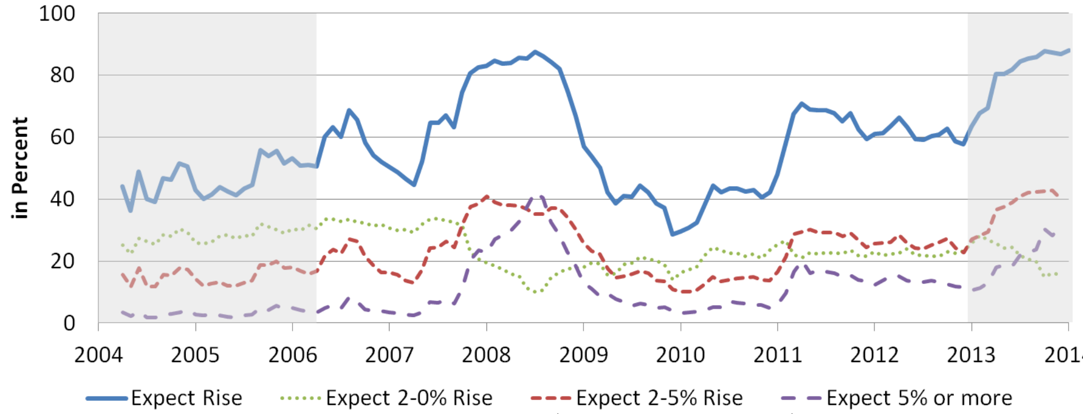
Figure 1.6: Japanese Inflation Expectations

(a) Expected Inflation within the next 6 months and 1 year forward



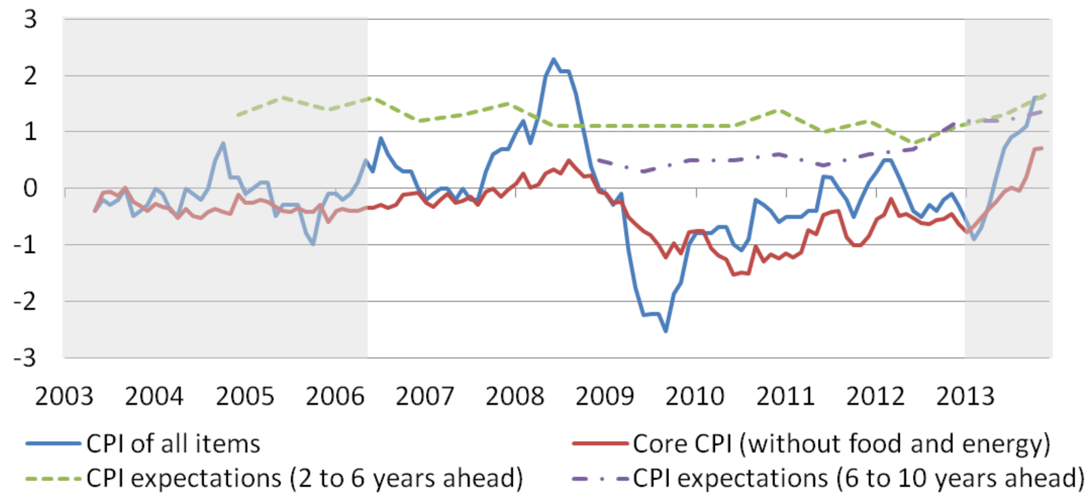
Medium term expectations of a higher inflation rate within the next 6 months (red dashed line), Source: ZEW Financial Market Survey; Expectations for rise in inflation 1 year forward (blue solid line), (incl. 1 person HH.), consumer confidence, not sa, Source: Cabinet Office Japan; Shaded areas: 1999-2000 ZIRP, 2001-2006 1st QE period, 2013-end 'Abenomics' period

(b) Details on Expected Inflation 1 year forward



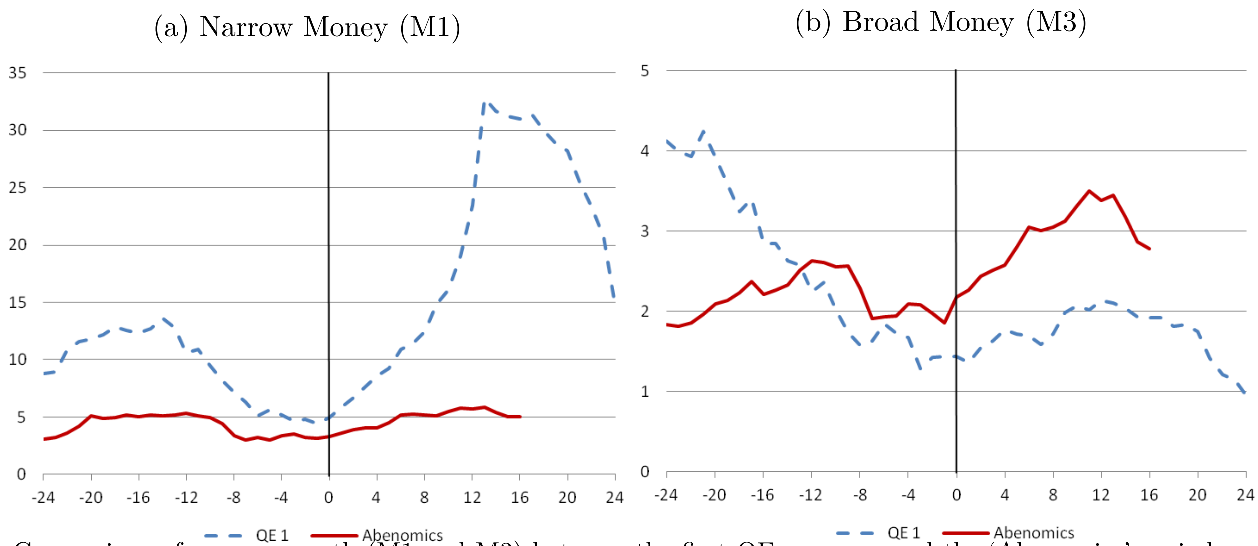
Expectations for rise in inflation 1 year forward, (incl. 1 person HH.), consumer confidence, not sa, Source: Cabinet Office Japan; Shaded areas: 2001-2006 1st QE period, 2013-end 'Abenomics' period

(c) Medium and Long Run Inflation Expectations and Inflation Development



Medium (Consensus forecast) and long run (ESP forecast) inflation expectations (annual average, % change). ESP forecast excludes the effects of the consumption tax hikes. *Source:* BoJ's Monthly Report of Recent Economic and Financial Developments, May 2014, BoJ. Percentage change from a year ago for CPI and core CPI. *Source:* St. Louis Fed. Shaded areas: 2001-2006 1st QE period, 2013-end 'Abenomics' period

Figure 1.7: Money Growth during the first QE- and 'Abenomics' program



Comparison of money growth (M1 and M3) between the first QE program and the 'Abenomics' period. Month 0 denotes March 2001 for the beginning of the first QE program and December 2012 for the 'Abenomics' period. *Source:* BoJ

1.7 Robustness Checks

The key focus of this paper is the investigation of the impact of a QE shock over time. Therefore, we concentrate in the following primarily on the QE shock. For reasons of clarity, we present the main figures of our robustness results in Appendix I. More detailed robustness results can be obtained upon request.

1.7.1 Close-to-median model

The median of the posterior of impulse responses determined by sign restrictions combines responses across different models. As a first robustness check, we investigate whether our results are sensitive to the close-to-median presentation. This procedure gives the impulse response that is the closest to the median impulse response. Figure I.12 in Appendix I.1 plots the median as well as the close-to-median impulse responses for the three time periods 1999Q4, 2003Q3 and 2013Q3 to a QE shock. It can be seen that the close-to-median responses are very similar to those based on the median.[26]

1.7.2 Robustness Checks on Priors

Strictly speaking, one can argue that the ZIRP from 1999 to 2000 is not an official QE policy. For a robustness check, we exclude this period from the estimation sample. Hence, the training sample is extended and based on data from 1980Q1 until 2000Q4. The results confirm those outlined in this paper. Since the TVP-prior could suffer from over-parameterisation, we use as an additional robustness check a hierarchical prior for B_0. It joins the Minnesota prior with the TVP-prior. The Minnesota prior allows for a shrinkage and thus reduces the risk of over-parameterisation. The results also support the findings presented in this paper.[27]

[26]Results on the close-to-median impulse responses of the demand and supply shocks resemble the findings based on the median responses as well. Results can be obtained upon request.

[27]The robustness results on the priors can be obtained upon request.

1.7.3 Robustness Checks on Identification

We attach special importance to robustness checks on the alteration of the identification of our structural shocks. In the following, we present three variations in the identification scheme.

Table 1.4: Sign Restrictions - alternative identification I

		QE Shock	DE Shock	SP Shock
	Impact	0		
Prices	$K = 3$	$\geq$	$>$	$\leq$
GDP	$K = 3$	?	$\geq$	$<$
Reserves	$K = 3$	$>$	$\leq$	?
Exchange Rate	$K = 3$	?	?	?
Further restriction	$K = 3$		$\left\lvert \frac{GDP}{CPI} \right\rvert > 1$	$\left\lvert \frac{GDP}{CPI} \right\rvert < 1$

Note: ? denotes no restriction, > defines a positive effect of the respective shock on the variable, vice versa for <. $K = 3$ indicates that the restriction horizon is three quarters. $\lvert GDP/CPI \rvert$ denotes the absolute value of the ratio between the GDP response and the CPI response.

First, since we use a rather agnostic identification in our benchmark model, we check whether our results to a QE shock are sensitive to an additional restriction on reserves following a demand shock. We assume that reserves respond negatively (≤ 0) to a demand shock for the first three quarters (Table 1.4). This additional restriction does not lead to changes in the transmission mechanisms of the QE and supply shocks (compare Figure I.13 for the impact of the QE shock).[28] Following a QE shock, prices increase significantly in the long and short run in 2013 whereas the impact in 2003 stays significant only until about two years after the shock. GDP starts to increase significantly after about two years. In contrast, the GDP response in 2003 is insignificant. Also the impact of the demand shock confirms our results presented, except for reserves. More specifically, we observe an insignificant response of reserves in our benchmark specification, while, the additionally imposed restriction in this identification scheme leads by definition to a slight significant decrease in reserves following a demand shock. The effect on the other endogenous variables is unaltered. In the light of this finding, we can assure to credibly distinguish the QE shock from the demand shock in our benchmark model.[29]

[28] Results on the demand and supply shocks can be obtained upon request.

[29] With the additional restriction on reserves in this section, we specify a stronger assumption on the differences between the transmission of the QE shock and the demand shock. This ensures to credibly distinguish between the QE and the demand shock. As this sensitivity check leaves the estimated effect on output, prices and the exchange rate compared to the benchmark model unchanged, it assures that

Second, again starting from our benchmark model we now assume a zero restriction on impact on GDP following the QE shock (see Table 1.5). Thus, CPI and GDP are both restricted to have a lagged response to a QE shock. Figure I.14 in Appendix I.3 shows that our results for the time periods 1999Q4 and 2003Q3 are basically insensitive to this change. For 2013Q3 we observe a slightly more visible effect on CPI and GDP. The responses seem to be stronger and longer lasting than under the benchmark identification scheme. Using a zero restriction on impact seems natural. However, against the background of quarterly data and the analysis of unconventional monetary policy shocks, imposing a more conservative identification scheme with a zero restriction on GDP might be too restrictive. Therefore, we recommend to stay agnostic on the initial impact on GDP following a QE shock.

Table 1.5: Sign Restrictions - alternative identification II

		QE Shock	DE Shock		SP Shock	
	Impact	0				
Prices	$K = 3$	$\geq$	$>$		$\leq$	
	Impact	**0**				
GDP	$K = 3$	?	$\geq$		$<$	
Reserves	$K = 3$	$>$	?		?	
Exchange Rate	$K = 3$	?	?		?	
Further restriction	$K = 3$		$\left\lvert\frac{\text{GDP}}{\text{CPI}}\right\rvert$	> 1	$\left\lvert\frac{\text{GDP}}{\text{CPI}}\right\rvert$	< 1

Note: ? denotes no restriction, $>$ defines a positive effect of the respective shock on the variable, vice versa for $<$. $K = 3$ indicates that the restriction horizon is three quarters. $\lvert$GDP/CPI$\rvert$ denotes the absolute value of the ratio between the GDP response and the CPI response.

Third, we employ a more traditional identification scheme and use a recursive identification based on monthly data with six lags. The ordering of our variables follows: $y_t = CPI_t, Output_t, Res_t, ExR_t$, where output refers to industrial production in the monthly version. These restrictions are implied by more traditional VAR studies such as Kimura et al. (2003) and Nakajima et al. (2011). Figure I.15 in Appendix I.4 summarises the respective results. They basically confirm those presented in this paper. The impact of the QE shock leads to a more pronounced increase in prices and output for 2013Q3 compared to the responses in 2003Q3 or 1999Q4. Note that the significance of the price response vanishes for 2003Q3 and it remains only slightly significant between about one until two years in 2013Q3. As before, the output response in 2003Q3 stays

we credibly distinguish the QE shock from the demand shock in our benchmark model.

insignificant and we observe a significant response in 2013Q3 between one and three years after the shock.

1.7.4 Robustness Checks on Data

Since many studies on the Japanese economy use output gap as an output measure, we reestimate our analysis and substitute GDP with output gap.[30] The results confirm those presented above. However, we abstain from using it due to its forward looking nature and the difficulty in general for estimating the Japanese output gap.[31] Results can be obtained upon request.

1.8 Conclusion

This paper uses a new kind of sign restriction in a TVP-VAR framework. It allows to analyse the impact of QE shocks over time when the economy is close to or at the Zero Lower Bound (ZLB). Our findings show that the reaction of macroeconomic variables in Japan to a QE shock has, indeed, varied over time. Especially GDP and prices reveal considerable time-varying effects across our sample from 1996 until 2013. Overall, our results suggest that the impact on GDP and prices, following a QE shock (1% increase in reserves), has become stronger and longer lasting over time.

More specifically, a comparison of the three considered time periods (1999Q4, 2003Q3 and 2013Q3)[32] shows that GDP responds significantly in 2013Q3 to a QE shock whereas in 2003Q3 and 1999Q4 the response is insignificant. Turning to prices, we generally report a visible response following a QE shock. Note that compared to 1999Q4 and 2003Q3, the price response in 2013Q3 stays permanently significant.

Regarding the relative explanatory power of a QE shock, we observe that a QE shock seems to explain a larger variance share of prices in 2013Q3 compared to 1999Q4 and 2003Q3. Also for GDP, the explanatory power of the QE shock is considerably higher in 2013Q3 than in 2003Q3. The generally more pronounced effects in 2013 possibly mirror

[30] Output gap of the total economy, seasonally adjusted, quarterly series, source: OECD.

[31] See Hausman and Wieland (2014) for a detailed discussion.

[32] The three time periods are chosen arbitrarily within each monetary policy stance of the BoJ. They refer to the ZIRP, the first QE program and the ‘Abenomics’ strategy respectively.

the influence of the current ‘Abenomics’ program. However, it should be noted that it may still be too early to make a definite evaluation on the effectiveness of ‘Abenomics’. Thus, our findings should be taken as a preliminary assessment on the short-term effects of ‘Abenomics’.

We would like to stress the robustness checks conducted for testing the consistency of our results. The use of the TVP-VAR with stochastic volatility is also supported by a marginal likelihood estimation based on the modified harmonic mean estimator that compares the TVP-VAR with a constant BVAR. Moreover, a sophisticated model selection algorithm is used to ensure the correct specification of the prior beliefs about the amount of time variation. Further, we implement several checks on the prior specifications, identifications and data. Additionally, we reestimate the impulse response functions based on the close-to-median responses. Our results presented in this paper seem to be largely insensitive to these alterations.

A Data sources

This paper uses quarterly data on Japan and covers a time horizon between 1980:1 and 2013:4. We estimate the model in detrended levels. Like Sims et al. (1990) state, this accounts for possible discrepancy which may arise in case of incorrectly assumed cointegration restrictions. Also, if there are unit roots in the data, it will not influence the likelihood function, since nonstationarity is of no concern in a Bayesian framework (see Sims and Uhlig (1991) for further discussions). In the following, the used time series are described:

Gross domestic product (GDP): Log of real gross domestic product (2010=100), seasonally adjusted, quarterly series. Source: Datastream.

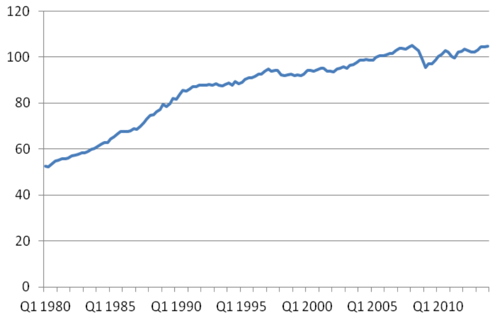
Core Consumer prices (CPI): Log of core consumer price index, all items less food (also less alcoholic beverages) and energy (2010 = 100), monthly index converted to a quarterly series (averaging over three respective months), not seasonally adjusted. Source: Datastream.

Reserves (Res): Log of monetary base (reserve requirement rate change unadj.), in hundreds of millions Japanese Yen, average outstanding amounts, current prices, seasonally adjusted. Monthly series converted to a quarterly series (averaging over three respective months). Source: Bank of Japan.

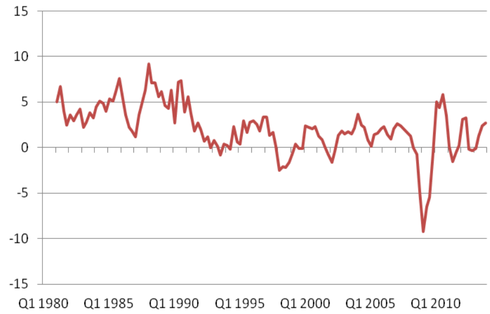
Exchange rate (ExR): Log of Japanese Yen real effective exchange rate index (2010=100), trade weighted exchange rate, CPI deflated, not seasonally adjusted. Monthly index converted to a quarterly series (averaging over three respective months). Source: Bank of Japan.

Figure A.1: Quarterly Data, Japan

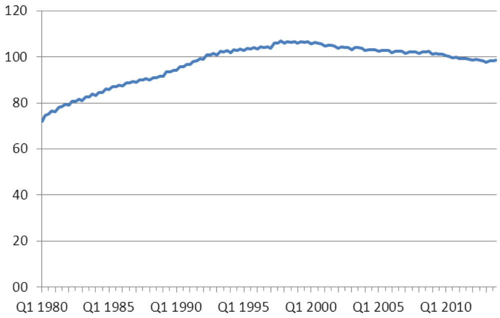
(a) GDP in levels



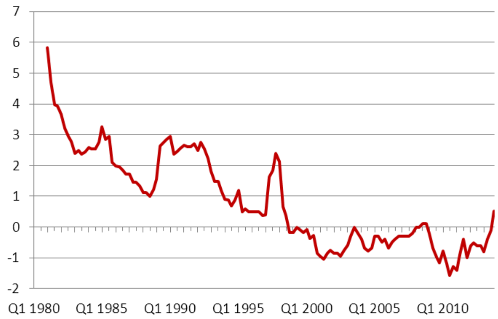
(b) GDP, yearly growth rate



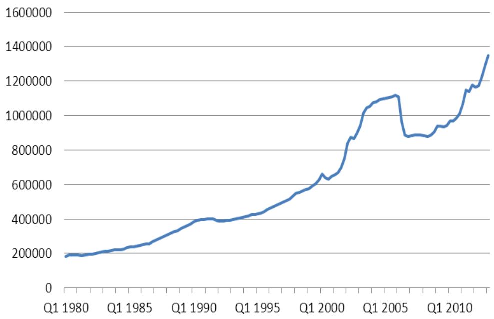
(c) Core CPI in levels



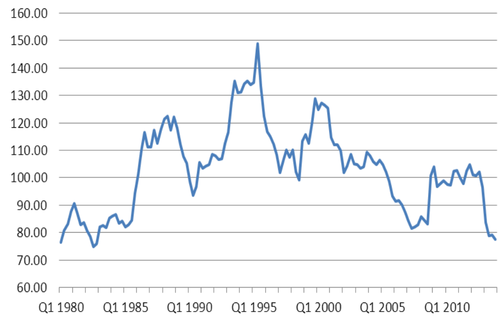
(d) Core CPI, annual rate of change



(e) Res in levels



(f) ExR in levels



B Literature Overview of VAR-studies on Japan

Table 1.6: VAR-Literature Overview of Monetary Transmission in Japan

Author(s)	Methodology	Identification	Data	Results
Schenkelberg and Watzka (2013)	Bayesian VAR	Sign restrictions on QE shock	$1995M3$ – $2010M9$; CPI, Industrial production (IP), reserves, government bond (10 years), exchange rate	- Benchmark VAR, **QE shock** (positive): CPI positive significant; IP during first 12 month negative insignificant thereafter positive and slightly significant; exchange rate initially depreciates, insignificant - Pre 1995 period VAR with call rate, MP shock (negative): CPI and IP increase significantly; exchange rate initially appreciates significantly, then depreciates significantly - also estimate price and output shock at the ZLB
Franta (2011)	TVP-VAR with stochastic volatility	Sign restrictions on monetary policy shock (specific to monetary policy regime)	$1971Q1$ – $2010Q3$; Industrial production (IP), CPI, call rate or current outstanding amounts, monetary base	- Comparision of pre ZIRP with QE: IP positive and significant for two quarters, similar response for both periods; CPI initially positive for QE, pre ZIRP no effect - Comparision of ZIRP and QE: no difference between two periods for IP and CPI responses - Comparision Financial Crisis: no difference between crisis periods for IP and CPI responses
Kimura et al. (2003)	VAR and TV-VAR with time-varying coefficients but constant variance matrix	Recursive identification	$1980Q$ – $2002Q1$; CPI, GDP gap, call rate, monetary base	- VAR, **monetary base shock** (positive): CPI and GDP gap increase but insignificant - VAR, interest rate shock (positive): CPI increase significantly; GDP gap decreases insignificantly - TV-VAR, **monetary base shock** (positive): CPI increases stronger in 1985Q2; in 2002 no impact visible; GDP gap no impact visible for 1985 and 2002

Table 1.7: VAR-Literature Overview of Monetary Transmission in Japan

Author(s)	Methodology	Identification	Data	Results
Kimura and Nakajima (2013)	Latent threshold model by Nakajima and West (2013) combined with a TVP-VAR with stochastic volatility	Recursive identification switching with a time-varying overidentifcation for the interest rate ZLB	1981Q2 – 2012Q3; CPI, GDP gap, call rate, outstanding balance of current accounts, goverment bond (10 years)	- Call rate shock (negative) if not QE period: CPI until mid 90s positive (slightly significant at some periods), thereafter insignificant; GDP until end of 90s positive significant, thereafter insignificant - **Bank reserves shock** (positive) during QE period: CPI and GDP positive but insignificant - Long term IR shock (negative) during QE and not QE period: CPI and GDP positive, significant until beginning/mid 90s - also estimate price and output shock
Hayashi and Koeda (2014)	Regime switching SVAR	Recursive identification	1988-2012 (monthly); CPI, GDP gap, call rate, excess reserves	**QE shock** (positive), base period Feb 2004: inflation increases significantly between 8-12 months after shock; GDP increases significantly and lasts for two years
Nakajima et al. (2011)	TVP-VAR with stochastic volatility and short rate as censored variable	Recursive identification	1981Q1 – 2008Q3; CPI, Industrial production (IP), call rate, monetary base (average outstanding amounts)	Call rate shock (positive): CPI decreases until end of 80s, gets significant after 1 year, increases in beginning of 90s significantly (price puzzle), decreases again in the beginning of 00s insignificantly; IP decreases, for 2002 insignificant, before significant after 5 months

Table 1.8: VAR-Literature Overview of Monetary Transmission in Japan

Author(s)	**Methodology**	**Identification**	**Data**	**Results**
Nakajima and Ginkō (2011)	TVP-VAR with stochastic volatility and censored interest rate with latent variable	Recursive identifcation, 4 lags	1977Q1 – 2010Q2; CPI, GDP gap, call rate, government bond (5 years)	- Call rate shock (positive): CPI (price puzzle at 1 year horizon) negative until 2000, thereafter no impact; GDP gap strongly negative until 1995 - Bond shock (positive): CPI initially since 1995 no impact, at 1 year horizon negative impact across sample; GDP gap since 1995 no impact, at 1 year horizone slightly negative and after 2 years positive - also estimate price and output shock
Nakajima (2011)	TVP-VAR with stochastic volatlity	Recursive identification	1977Q1-2007Q4; CPI, GDP gap, call rate or government bond (5 years)	- MP shock with bond (positive): CPI until mid 90s initially positive (price puzzle) then negative, since mid 90s in medium-term/long run no impact; GDP gap until mid 90s strongly negative, thereafter no impact - MP shock with call rate (positive): CPI, price puzzle less evident, decline until mid 90s, thereafter slightly positive; GDP gap initially negative across whole sample, medium term/long run negative until 2000 - also estimate price and output shock

C Posterior Probability Estimates for k_Q, k_W and k_S

Table 1.9: Posterior Probability Estimates for k_Q, k_W and k_S based on the RJMCMC Method

Model	k_Q	k_W	k_S	Posterior probability
1	0.0100	0.0100	0.0010	0
2	0.0500	0.0100	0.0010	0
3	0.1000	0.0100	0.0010	0
4	0.0100	0.0250	0.0010	0
5	0.0100	0.0250	0.0100	0.02
6	0.0100	0.1000	0.0010	0
7	**0.0100**	**0.1000**	**0.0100**	**0.994**
8	0.0100	0.0100	0.0100	0
9	0.0500	0.0250	0.0010	0.0007
10	0.0500	0.0250	0.0100	0.0003
11	0.0500	0.0100	0.0100	0
12	0.0500	0.1000	0.0010	0.0027
13	0.0500	0.1000	0.0100	0
14	0.1000	0.0250	0.0010	0
15	0.1000	0.1000	0.0010	0
16	0.1000	0.1000	0.0100	0
17	0.1000	0.0250	0.0100	0.0003
18	0.1000	0.0100	0.0100	0

Note: Posterior probability estimates are based on the reversible jump Markov chain Monte Carlo method for the set of 18 models. These are constructed from all possible combinations of $k_Q = \{0.01; 0.05; 0.1\}$, $k_W = \{0.001; 0.01\}$ and $k_S = \{0.01; 0.025; 0.1\}$.

D Eggertsson (2011) Model

In the following, we briefly summarise the New Keynesian model of Eggertsson (2011).[33]

A negative preference shock moves the economy to the ZLB and induces a fall in output. After this shock, the economy reverts back to its steady state with probability $1-\mu$. It stays at the ZLB with probability μ. Monetary policy is assumed to follow a Taylor rule and is approximated by:

$$i_t = max\left\{0, r_t^e + \phi_\pi \pi_t + \phi_Y \hat{Y}_t\right\} \tag{1.11}$$

with r_t^e denoting the exogenous shock, π_t the inflation rate and $\hat{Y}_t$ the output gap. ϕ_π and ϕ_Y are the respective Taylor rule coefficients. The 0 in equation 1.11 accounts for the ZLB state. Monetary policy can then be written to have the following form:

$$i_t = r_H^e \text{ for } t \geq T^e$$

$$i_t = 0 \text{ for } 0 < t < T^e,$$

where r_H^e refers to the negative preference shock in the non-recession state and T^e depicts some stochastic date when the economy returns back to its steady state. In the case of the binding ZLB ($t < T^e$), the AD and AS equations[34] are given by:

$$AD \quad \hat{Y}_L = \mu\hat{Y}_L + \sigma\mu\pi_L + \sigma r_L^e \tag{1.12}$$

$$AS \quad \pi_L = \kappa\hat{Y}_L + \beta\mu\pi_L, \tag{1.13}$$

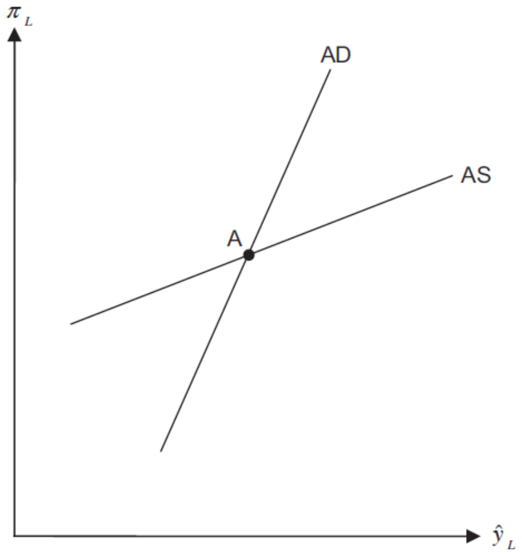
where L denotes the recession state, r_L^e the negative preference shock in the recession state and $\sigma, \kappa > 0$ and $0 < \beta < 1$.

Note that for the case of a multiperiod recession, where the ZLB is binding for more than one period ($\mu > 0$), both the aggregate demand and supply curve are upward sloping in inflation. The peculiar case of an upward sloping demand curve occurs since, for a given nominal interest rate, the real interest rate increases due to a lower inflation rate today which implies lower expected inflation ($\mu\pi_L < 0$).[35] Higher real interest rates

[33] The notation follows Schenkelberg and Watzka (2013).
[34] AD and AS are respectively referring to aggregate demand and aggregate supply
[35] At the ZLB, the central bank cannot reduce its nominal interest rate to offset the deflationary

are contractionary and induce aggregate demand to fall because current consumption is relatively more expensive than future consumption. Thus, households increase their current savings for future consumption. This in turn reduces current output. Therefore, inflation and output are both upward sloping (compare Figure D.2).

Figure D.2: The Effect of Multiperiod Recession at the ZLB in Eggertsson (2011)



D.1 Why the AD curve is steeper than the AS curve

Rewriting the AD and AS equations in 1.12 and 1.13 and solving for π_L gives:

$$AD' \quad \pi_L = \frac{(1-\mu)\hat{Y}_L}{\sigma\mu} - \frac{r_L^e}{\mu} \tag{1.14}$$

$$AS' \quad \pi_L = \frac{\kappa\hat{Y}_L}{(1-\beta\mu)} \tag{1.15}$$

If the slope coefficients of the AD and AS curves are the same, both curves are parallel. Eggertsson calls this a *deflationary black hole*:

$$\frac{1-\mu}{\sigma\mu} = \frac{\kappa}{1-\beta\mu} \text{ or } (1-\beta\mu)(1-\mu) = \sigma\mu\kappa.$$

In this case no solution exists. For a liquidity trap equilibrium to exist, the following has to hold:

$$(1-\beta\mu)(1-\mu) > \sigma\mu\kappa.$$

pressure.

This is satisfied for a sufficiently low probability μ of staying at the ZLB. When μ increases, the AD curve becomes flatter and the AS curve steeper. Eggertsson (2011) assumes μ to be small enough so that the *deflationary black hole* does not occur.

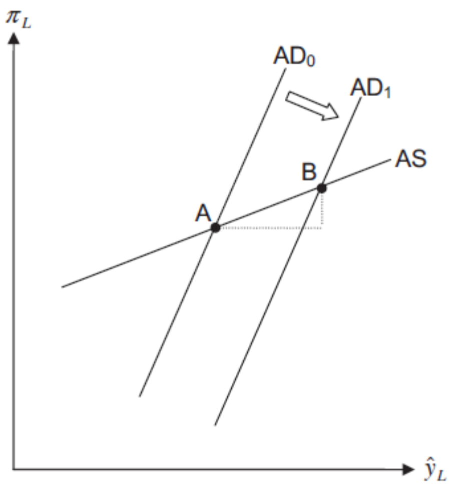
D.2 Identification of the demand and supply shocks

The usual identification strategy for demand and supply shocks does not hold if the nominal interest rate is at the ZLB. We briefly outline important aspects of how aggregate demand and supply shocks are identified as well as passed through in the Eggertsson (2011) model.

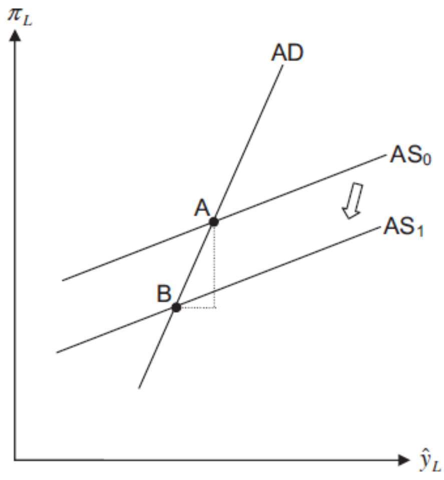
For identifying the demand from the supply shock at the ZLB, we use the Eggertsson assumption that the AD curve is steeper than the AS curve. Therefore, a positive demand shock increases output gap more than inflation. Vice versa for a positive supply shock. The sign restrictions in Table 1.1 capture this aspect by restricting the ratio of the output and inflation variables to be larger than one in absolute terms for the demand shock. The opposite holds for the supply shock. Furthermore, a positive demand shock is restricted to increase output and inflation (Figure D.3a), whereas a positive supply shock is defined to decrease both variables (Figure D.3b).

Figure D.3: Aggregate Business Cycle Shocks

(a) Demand Shock



(b) Supply Shock



D.3 Effects of a credible expansion of monetary policy at the ZLB

Eggertsson (2011) models a monetary expansion as an increase in future money supply. Monetary policy is assumed to follow:

$$i_t = max\left\{0, r_t^e + \pi^* + \phi_\pi(\pi_t - \pi^*) + \phi_Y(\hat{Y}_t - \hat{Y}^*)\right\}, \tag{1.16}$$

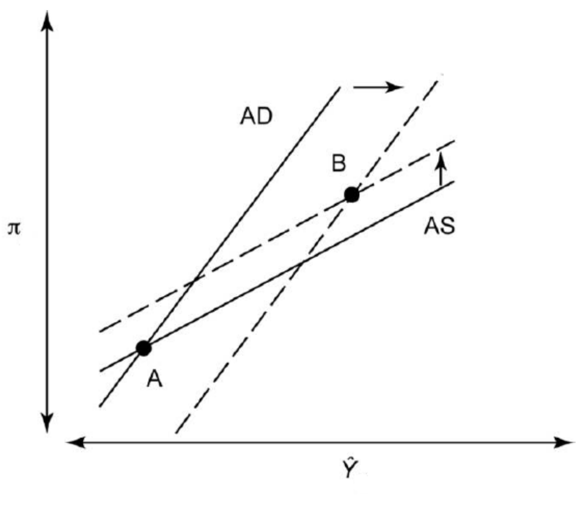
where π^* denotes the inflation target and the long run output target is given by $\hat{Y}^* = (1-\beta)\kappa^{-1}\pi^*$. A large π^* corresponds to a larger future growth rate in money supply. To see the effect of a higher π^*, we rewrite the AD and AS equations for the recession state when the ZLB is binding:

$$AD'' \quad \hat{Y}_L = \mu\hat{Y}_L + (1-\mu)\hat{Y}^* + \sigma\mu\pi_L + \sigma(1-\mu)\pi^* + \sigma r_L^e \tag{1.17}$$

$$AS'' \quad \pi_L = \kappa\hat{Y}_L + \beta\mu\pi_L + \beta(1-\mu)\pi^*, \tag{1.18}$$

If $\pi^* = 0$ is raised to $\pi^* > 0$, the AD curve moves to the right and the AS curve to the left (see Figure D.4). An increased inflation target for $t \geq T^e$ leads to a fall in the real interest rate in $t < T^e$. This in turn increases current consumption and thus stimulates current output. Note that next to higher expected inflation in $t \geq T^e$, a higher π^* also induces a higher inflation rate at all periods in which the ZLB is binding (Figure D.4).

Figure D.4: Effect of a credible inflation commitment at the ZLB

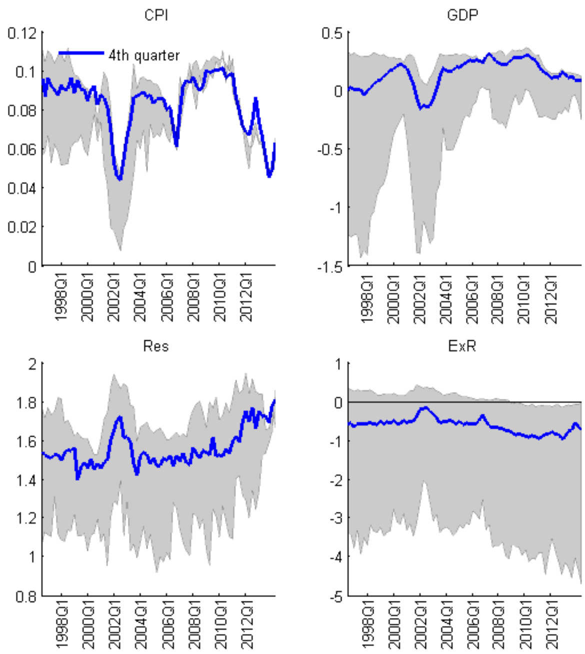


E Impulse Responses to a QE Shock at the 4th and 8th Quarter

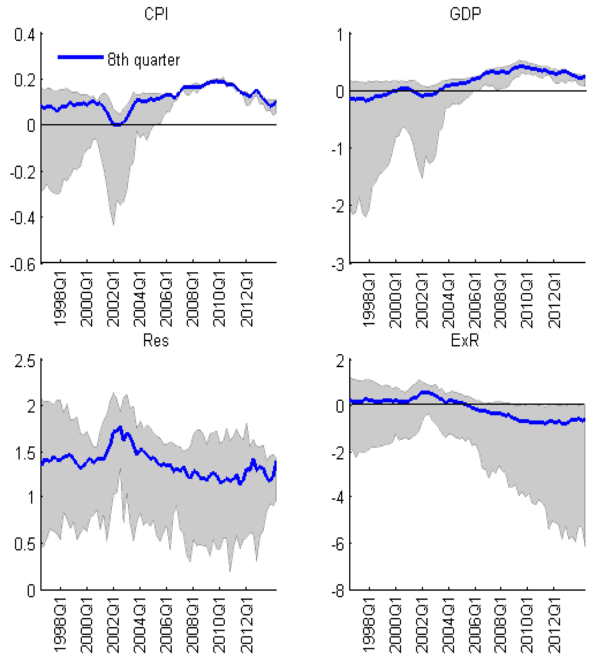
E.1 Estimation at 4th and 8th Quarter

Figure E.5: Impulse Responses at Different Horizons

(a) Responses at the 4th quarter to a QE Shock



(b) Responses at the 8th quarter to a QE Shock

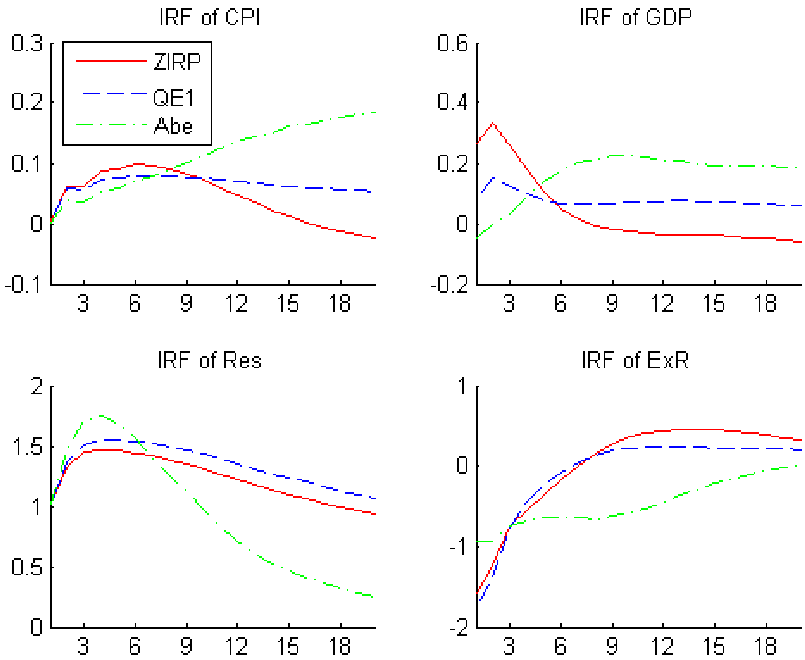


Median impulse responses (blue solid line) to a 1% QE shock with 16-th and 84-th percentiles (grey area) of the posterior distribution of the responses at the 4th and 8th quarter, respectively.

F Average of the QE Shock for ZIRP, QE- and 'Abenomics' program

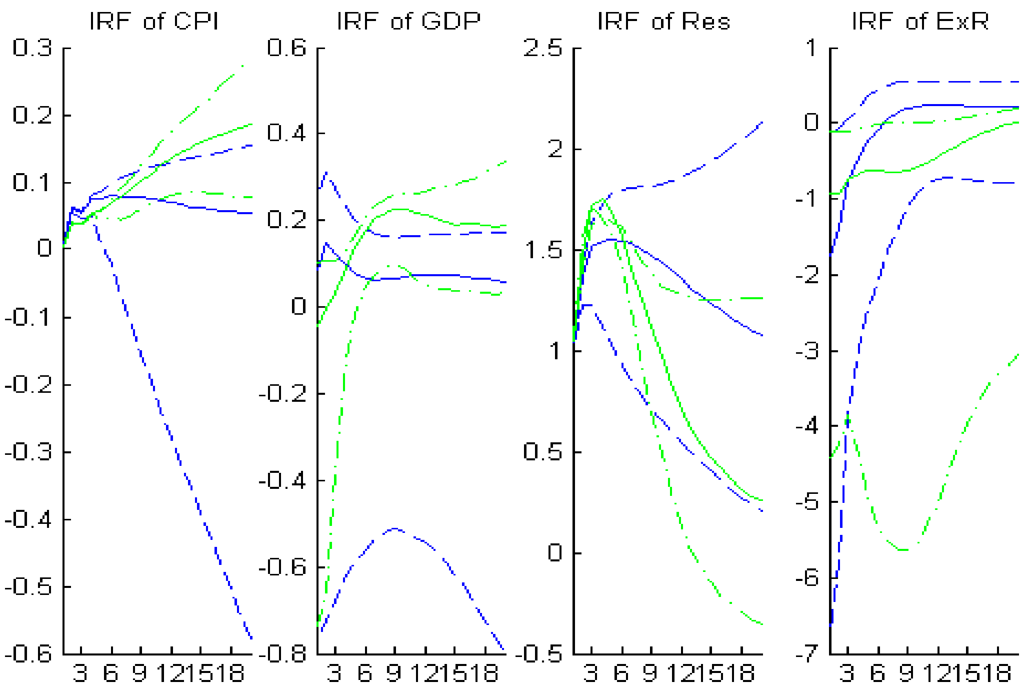
Figure F.6: Responses at Different Time Periods to a QE Shock

(a) Responses without Percentiles



Average of the median impulse responses to a 1% QE shock during ZIRP, first QE- and 'Abenomics' program.

(b) Responses with Percentiles



Average of the median impulse responses (solid line) to a 1% QE shock with 16-th and 84-th percentiles (dashed line) of the posterior distribution during the first QE- and 'Abenomics' program.

G Time-Varying Posterior Estimates of the Stochastic Covariance

The time-varying variance-covariance matrix of the residuals is decomposed as $A_t^{-1}\Sigma_t\Sigma_t'(A_t^{-1})'$ and comprises two matrices: (1), the time-varying matrix Σ_t which denotes the diaogonal matrix of the variances of the structural shocks ε_t and (2), the time-varying lower triangular matrix A_t which defines the simultaneous relations. The latter is structured as:

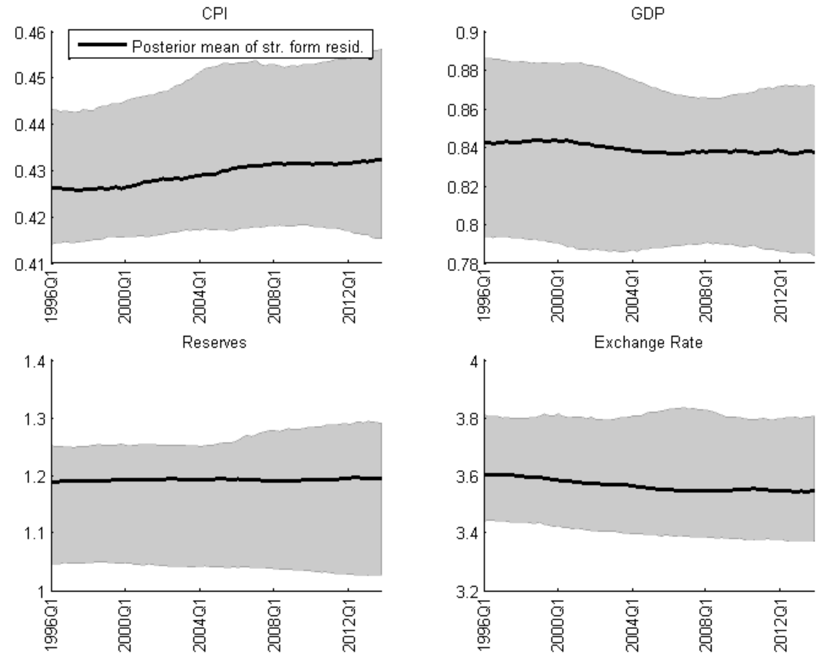
$$A_t = \begin{bmatrix} 1 & 0 & 0 & 0 \\ \alpha_{\pi y,t} & 1 & 0 & 0 \\ \alpha_{\pi r,t} & \alpha_{yr,t} & 1 & 0 \\ \alpha_{\pi e,t} & \alpha_{ye,t} & \alpha_{re,t} & 1 \end{bmatrix}$$

where π_t denotes inflation, y_t GDP, r_t reserves and e_t the exchange rate. More specifically, $\alpha_{re,t}$ captures the simultaneous impact of a reserves shock on the exchange rate.

Concerning Σ_t, not much time variation is visible. Figure E.6 below shows the estimated stochastic volatility of the structural shock on GDP, prices, reserves and the exchange rate. It plots the posterior mean and the 16th and 84th percentile of the standard deviation of the shock. The second matrix, the time-varying simultaneous relations are plotted in Figure G.9. The simultaneous effects are clearly time varying for almost every element in this matrix.

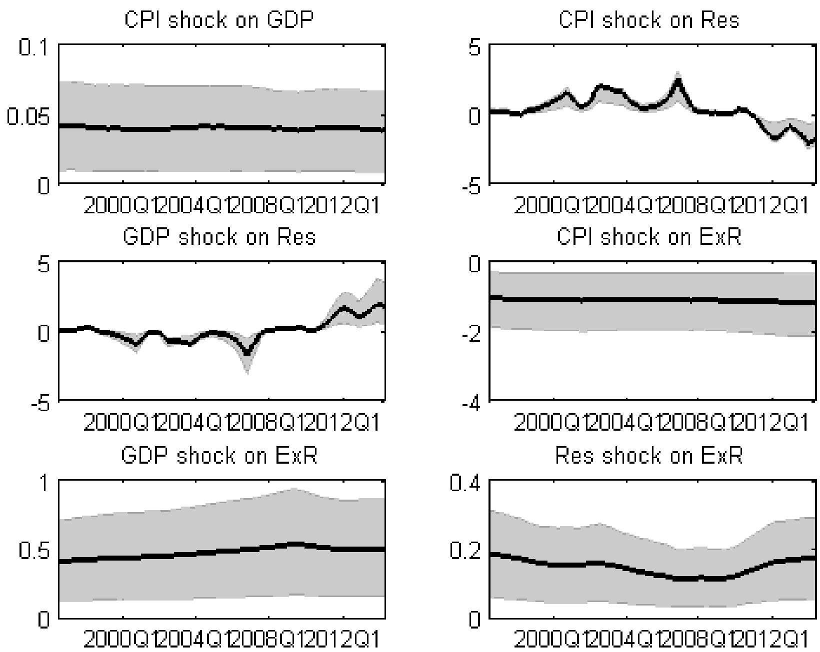
Figure G.7: Estimated Stochastic Volatility of the Structural Shocks

Figure G.8: Volatility of the Structural Shocks



Posterior mean (solid line), 16-th and 84-th percentiles (in grey) of the standard deviation of residuals of the CPI, GDP, reserves and exchange rate equation.

Figure G.9: Posterior Estimates for the Simultaneous Relation $\tilde{\alpha_{it}}$



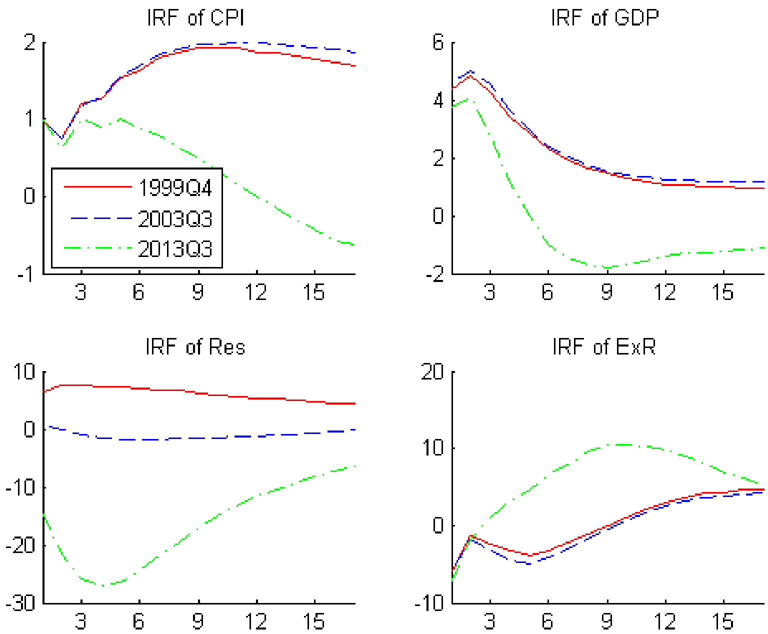
Posterior estimates for the simultaneous relations. Posterior mean (solid line), 16-th and 84-th percentiles (in grey).

H Impulse Responses to a Demand and Supply Shock

H.1 Demand Shock

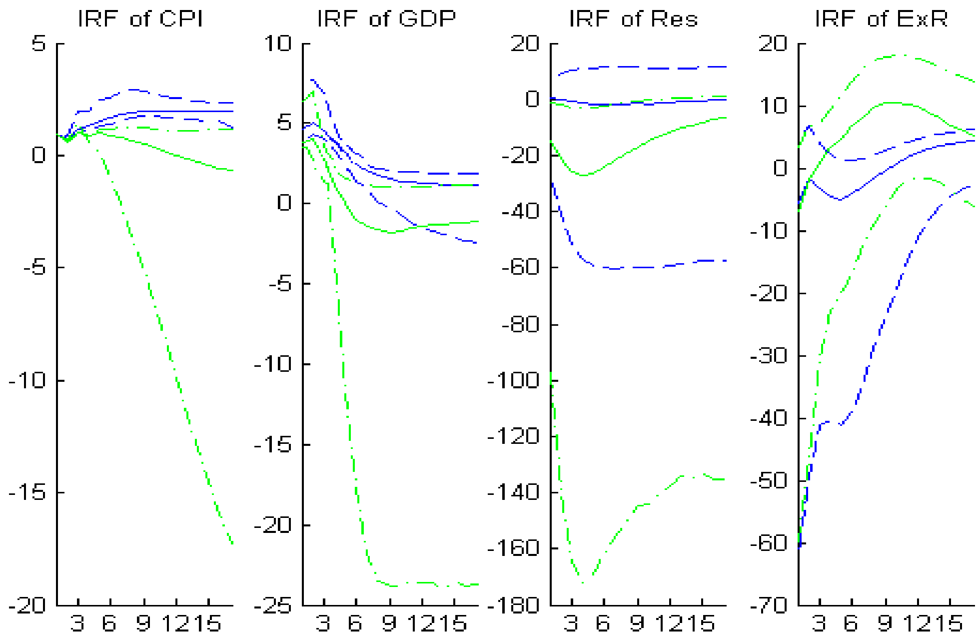
Figure H.10: Responses at Different Time Periods to a DE Shock

(a) Responses without Percentiles



Median impulse responses to a 1% DE shock at 1999Q4, 2003Q3 and 2013Q3.

(b) Responses with Percentiles

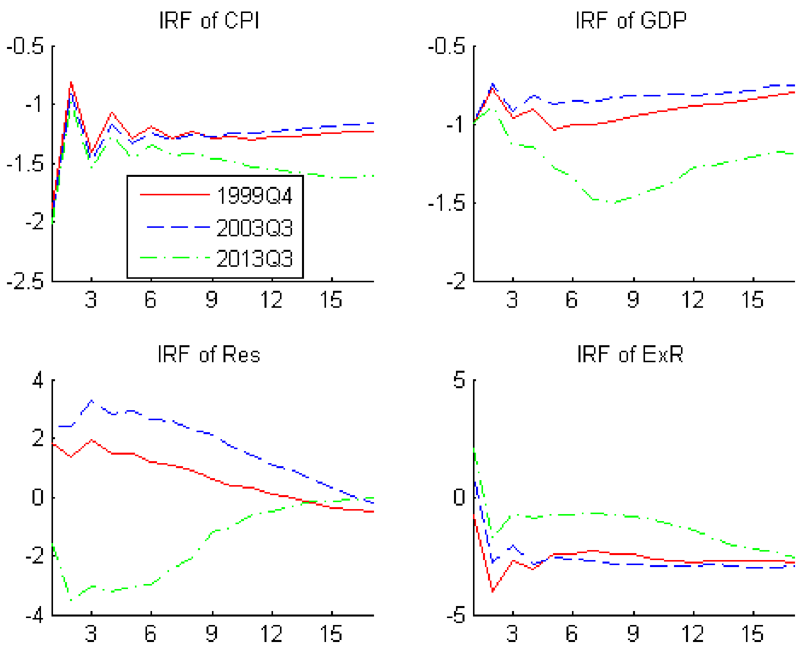


Median impulse responses (solid line) to a 1% QE shock with 16-th and 84-th percentiles (dashed line) of the posterior distribution at 2003Q3 (blue) and 2013Q3 (green).

H.2 Supply Shock

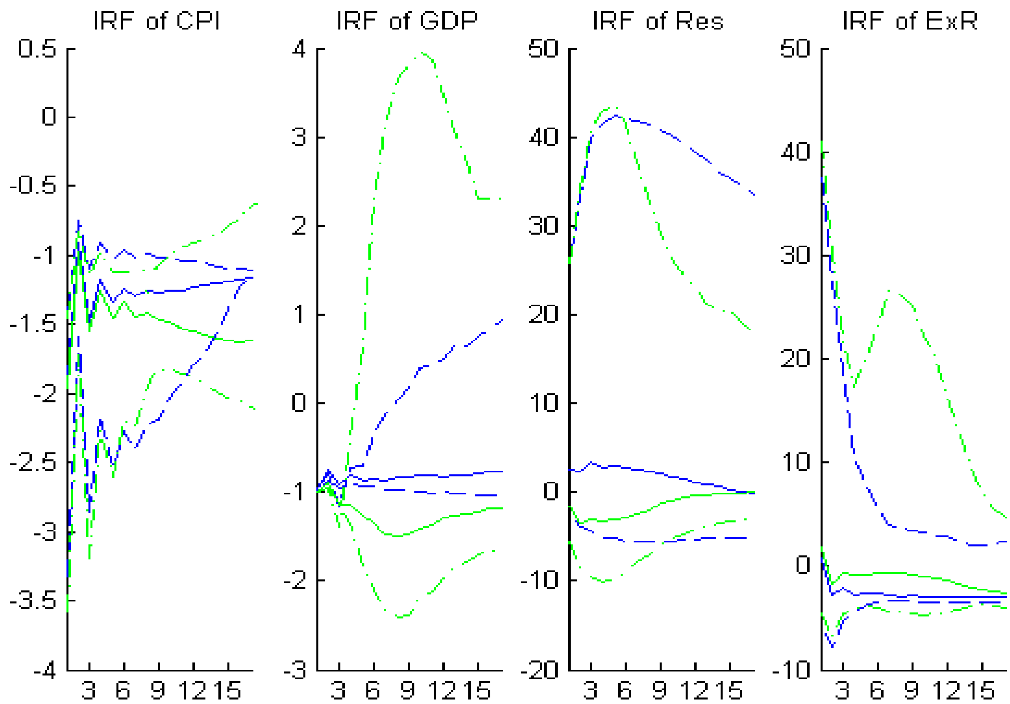
Figure H.11: Responses at Different Time Periods to a SP Shock

(a) Responses without Percentiles



Median impulse responses to a 1% SP shock at 1999Q4, 2003Q3 and 2013Q3.

(b) Responses with Percentiles

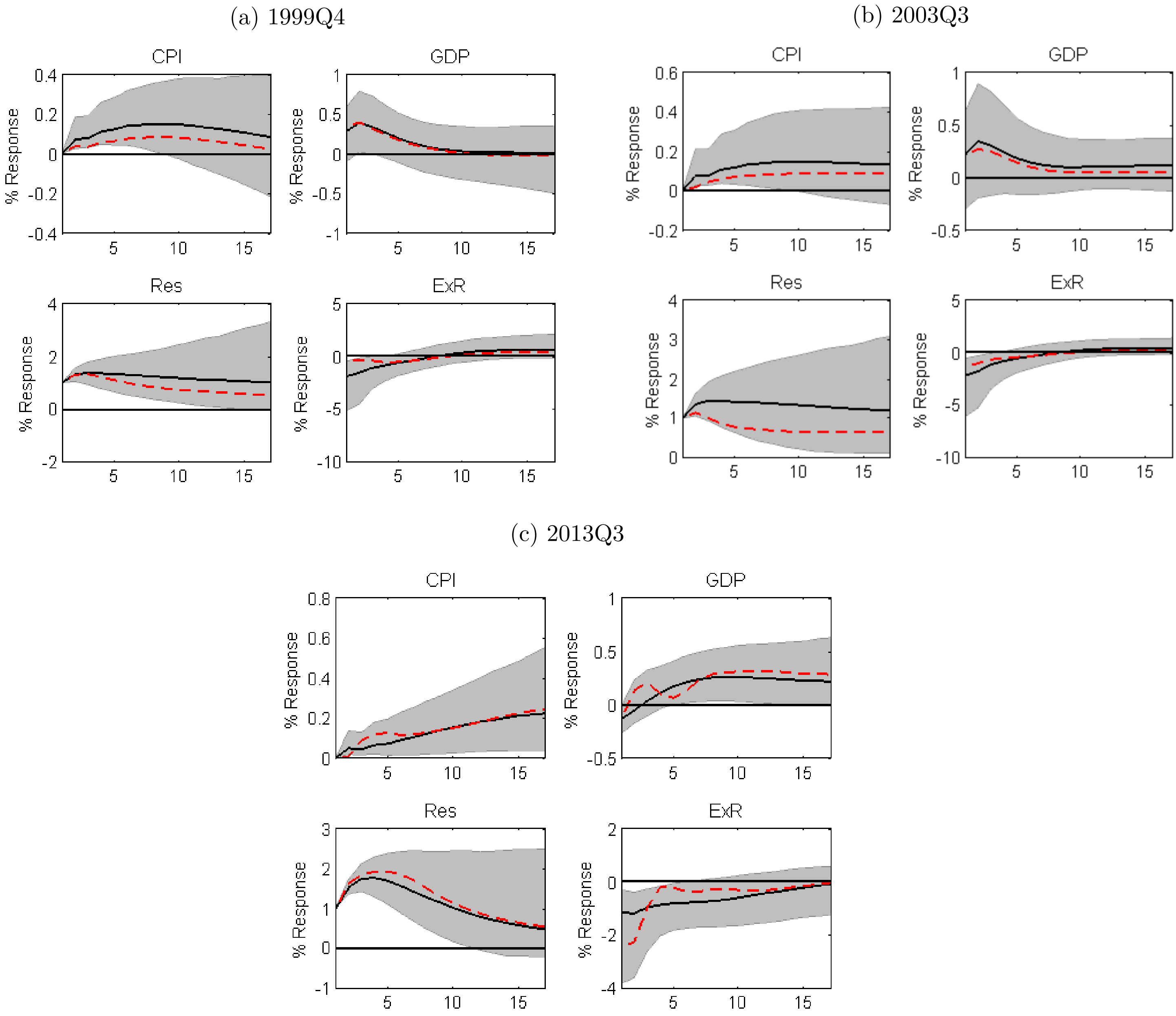


Median impulse responses (solid line) to a 1% SP shock with 16-th and 84-th percentiles (dashed line) of the posterior distribution at 2003Q3 (blue) and 2013Q3 (green).

I Robustness of the Results

I.1 QE shock based on Close-to-Median Impulse Responses

Figure I.12: Impulse responses to a QE shock

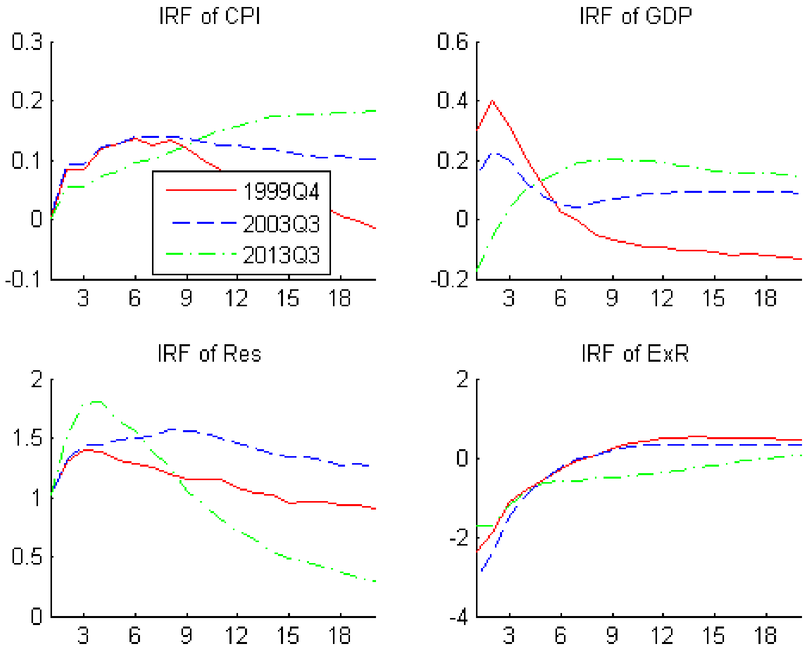


Note: Solid lines denotes the median impulse responses to a 1% QE shock, grey area the 16-th and 84-th percentiles of the posterior distribution of the responses. The red dashed lines plot the responses based on the close-to-median model.

I.2 QE shock based on Alternative Identification I

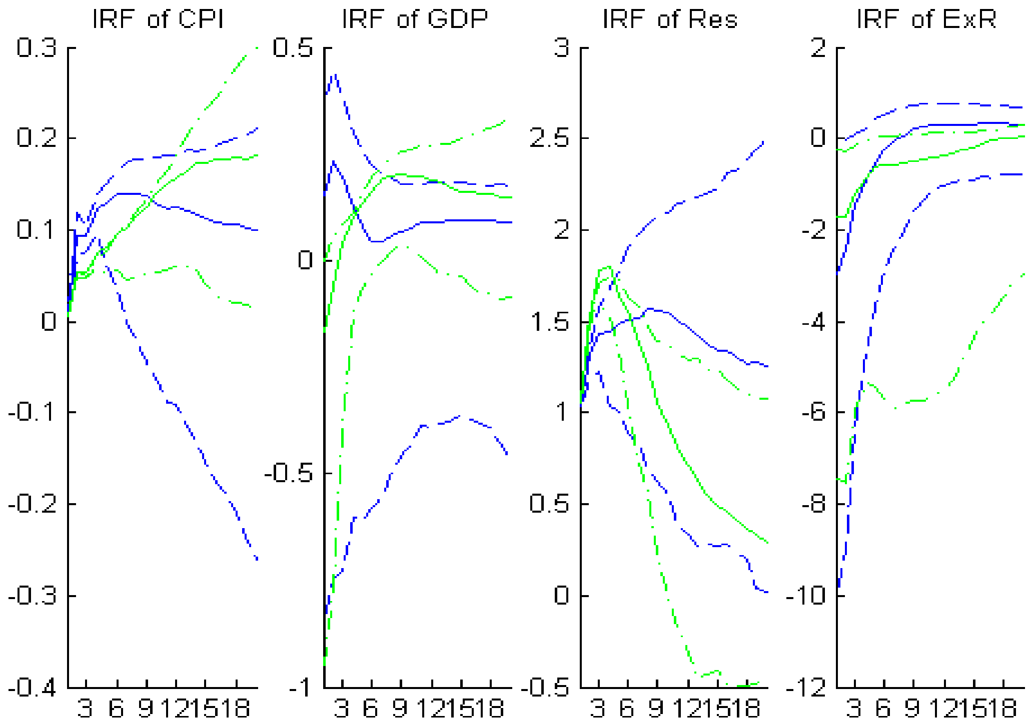
Figure I.13: Responses at Different Time Periods to a QE Shock

(a) Responses without Percentiles



Median impulse responses to a 1% QE shock at 1999Q4, 2003Q3 and 2013Q3.

(b) Responses with Percentiles

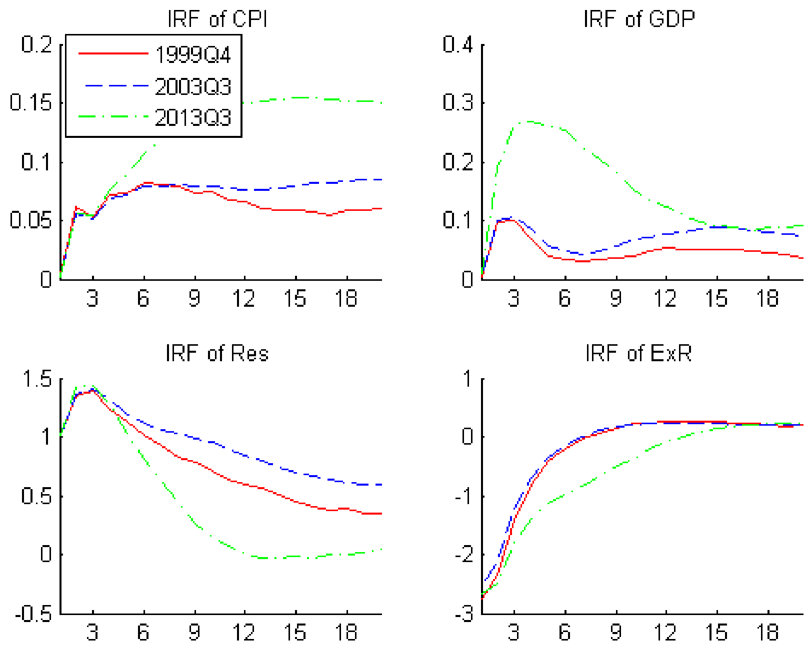


Median impulse responses (solid line) to a 1% QE shock with 16-th and 84-th percentiles (dashed line) of the posterior distribution at 2003Q3 (blue) and 2013Q3 (green).

I.3 QE shock based on Alternative Identification II

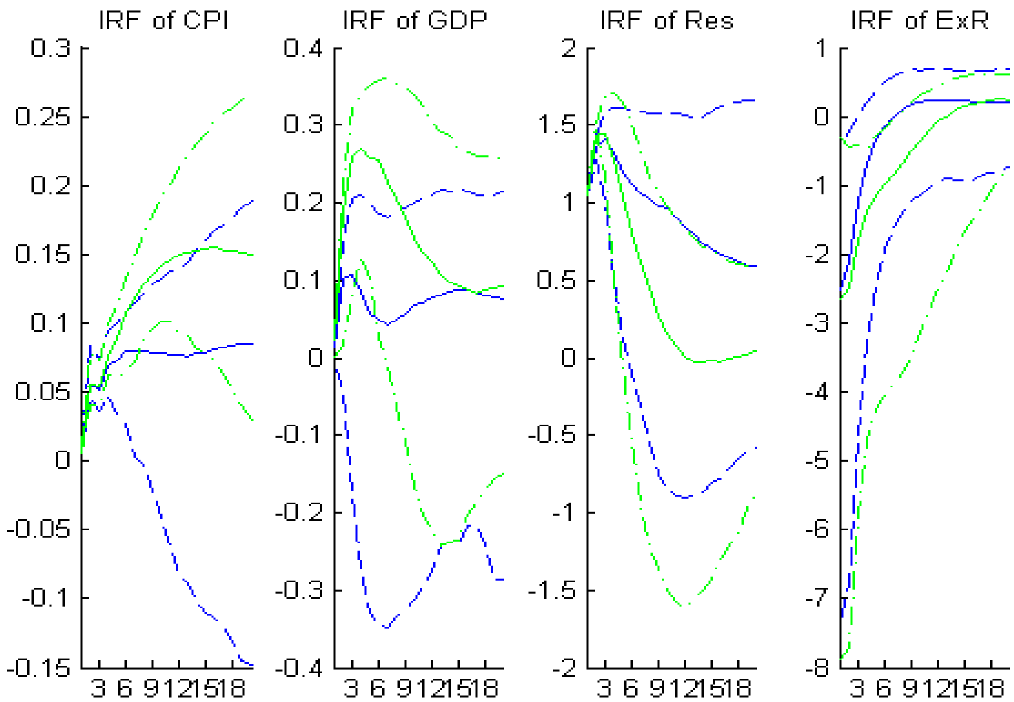
Figure I.14: Responses at Different Time Periods to a QE Shock

(a) Responses without Percentiles



Median impulse responses to a 1% QE shock at 1999Q4, 2003Q3 and 2013Q3.

(b) Responses with Percentiles

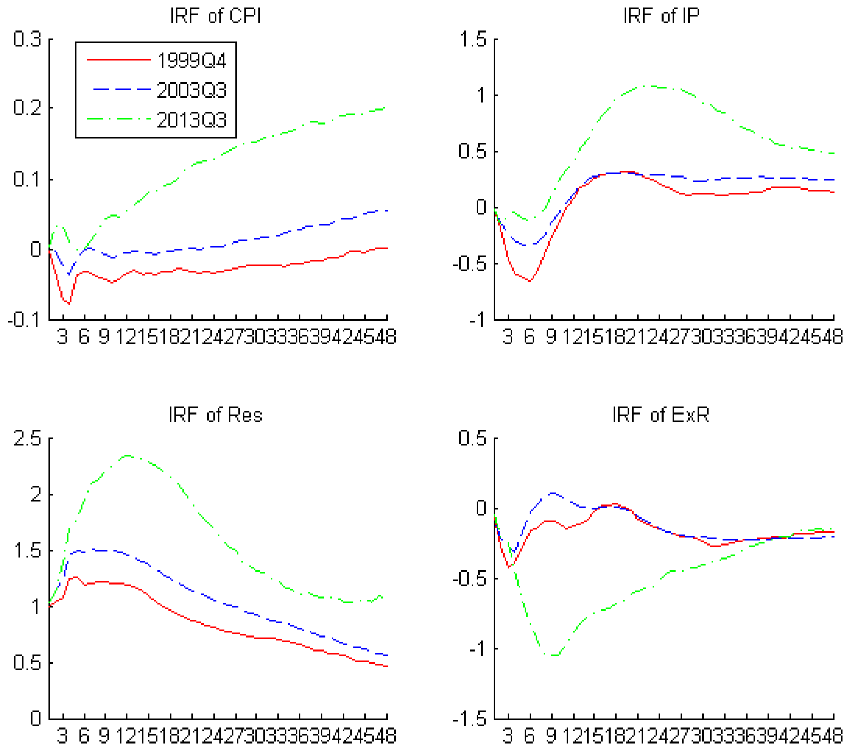


Median impulse responses (solid line) to a 1% QE shock with 16-th and 84-th percentiles (dashed line) of the posterior distribution at 2003Q3 (blue) and 2013Q3 (green).

I.4 QE shock based on Alternative Identification III

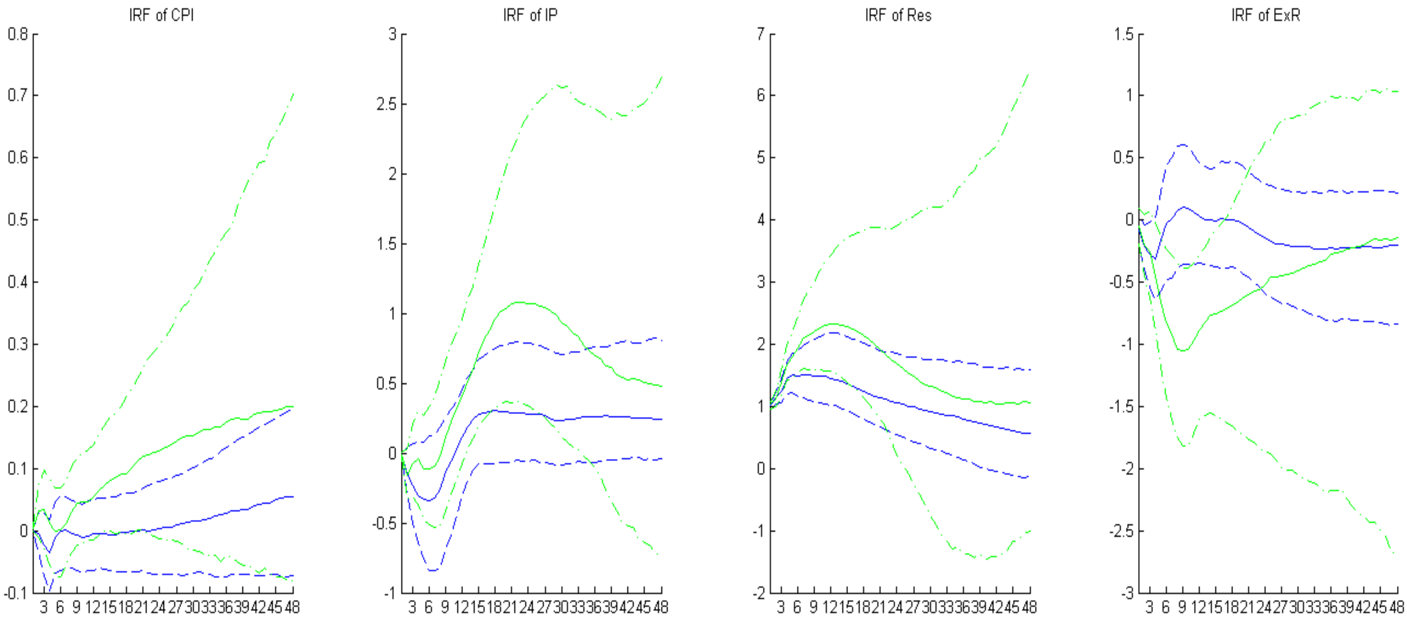
Figure I.15: Responses at Different Time Periods to a QE Shock (Cholesky)

(a) Responses without Percentiles



Median impulse responses to a 1% QE shock at 1999M11, 2003M8 and 2013M8.

(b) Responses with Percentiles



Median impulse responses (solid line) to a 1% QE shock with 16-th and 84-th percentiles (dashed line) of the posterior distribution at 2003Q3 (blue) and 2013Q3 (green).

J Convergence Diagnostics

This section gives convergence diagnostics of the Markov chain Monte Carlo algorithm. We follow Primiceri (2005) to calculate the convergence diagnostics. These autocorrelations measures are based on the Econometric Toolbox illustrated by LeSage (1999). For space reasons, the convergence diagnostics are only given for estimates of the point 2013Q3.[36]

We refer to three measures of convergence diagnostics: (i) 10-th-order sample autocorrelation of the draws; (ii) inefficiency factors (IFs) for the posterior estimates of the parameters, it is an estimate of $(1 + 2\sum_{k=1}^{\infty} \rho_k)$, with ρ_k as the k-th-order autocorrelation of the chain, adequate estimates are below or above the value of 20; (iii) and the Raftery and Lewis (1992) diagnostics, calculating the necessary number of runs to obtain a certain precision (the desired precision = 0.025, necessary probability for obtaining this precision = 0.95, calculated for the 0.025 quantile of the marginal posterior distribution).

J.1 Convergence Diagnostics of the Markov Chain Monte Carlo Algorithm

The (a) panel of Figure J.16 refers on the horizontal axis throughout the points 1-36 to B (time varying coefficients), points 37-42 correspond to A (time varying simultaneous relations), and points 43-46 refer to Σ (time varying volatilities). Respectively, the hyperparameter panels (b), (c) and (d) of Figure J.16, relate throughout the points 1-1296 to Q, points 1297-1332 to S and points 1333-1348 to W.

We start with a short summary of the 10-th-order autocorrelation. It is useful to scrutinise the autocorrelation function of the draws, to evaluate how well the randomly selected chain mixes. For an efficient algorithm, the draws need to be independent from each other. This is verified by low values of the autocorrelation function (see Figure J.16a and J.16b). The autocorrelation estimates for Σ exhibit some correlation indicating inefficiency (see below for discussion).

The diagnostics concerning the inefficiency factors (IFs) calculates values very much below 20, thus suggesting efficiency. An overview is also given in Table 1.10 below. Concerning

[36] Compared to other points in time, the respective estimates are very similar.

the IFs of A and B, the statistics show very low estimates. However, the IFs referring to Σ indicate some inefficiency. Considering the higher dimensionality of our problem, however, these results seem satisfactory (Kirchner et al. (2010)). Also Franta et al. (2011) illustrate that some inefficiency should be of a minor concern when the total number of runs required by the Raftery and Lewis (1992) statistics is well below the actual number used in this study. As can be seen in Figures J.16a and J.16d, the suggested number of iterations is below the actual number used. Furthermore, the impulse responses are calculated with respect to normalised shocks, hence, the inefficiency problem should not matter (Franta et al. (2011)).

To sum up, the total number of suggested iterations is far below the number used in this paper and, on average, we obtain satisfying IFs as well as autocorrelation estimates. Hence, the convergence diagnostics are sufficient.

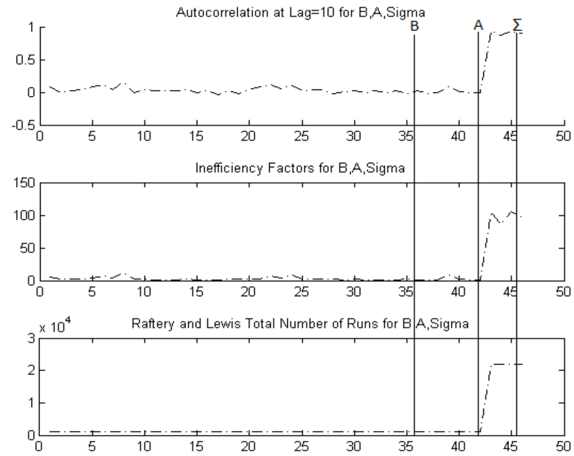
Table 1.10: Distribution of the Inefficiency Factors

	Median	Mean	Min	Max	10-th Percentile	90-th Percentile
A	1.1341	2.173	0.9536	6.7792	0.9812	4.5915
B	2.335	3.1415	0.7967	10.4348	1.4623	4.9349
Σ	136.1581	132.7426	106.3865	152.2678	104.035	151.5319

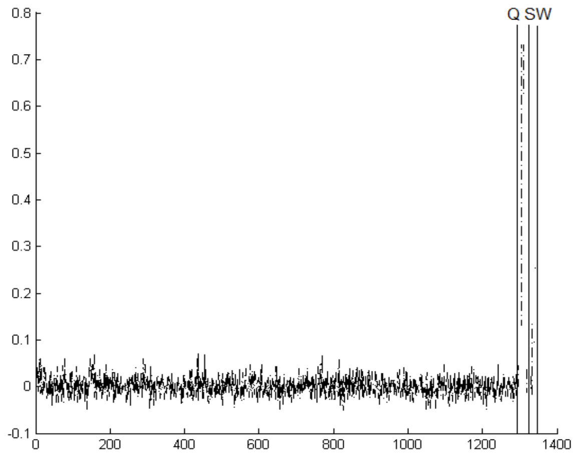
Overview of the inefficiency factors (IFs) for the posterior estimates of different sets of time varying parameters. A: time varying simultaneous relations; B: time varying coefficients; Σ: time varying volatilities.

Figure J.16: Convergence Diagnostics

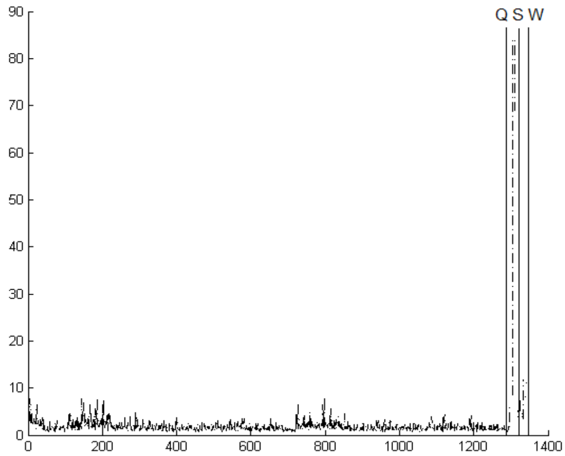
(a) Convergence Diagnostics for A, B and Σ



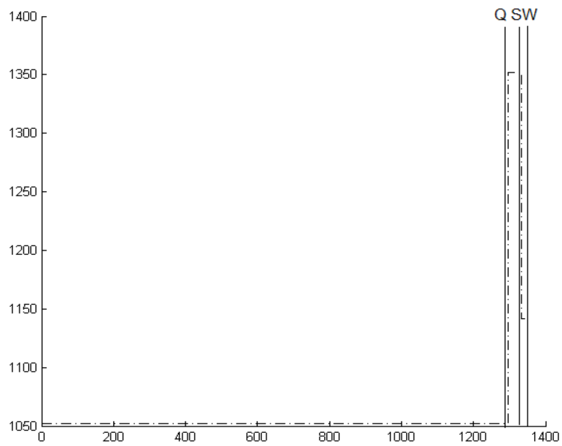
(b) 10-th-order Autocorrelation for Q, S and W



(c) IFs for Q, S and W



(d) Raftery & Lewis tot. no of runs for Q, S and W



Panel (a) refers on the horizontal axis throughout the points 1-36 to B (time varying coefficients), points 37-42 to A (time varying simultaneous relations), and points 43-46 to Σ (time varying volatilities). The hyperparameters in panels (b), (c) and (d) relate throughout the points 1-1296 to Q, points 1297-1332 to S and points 1233-1348 to W.

Chapter 2

The Impact of Monetary Policy and Exchange Rate Shocks in Poland: Evidence from a Time-Varying VAR*

2.1 Introduction

Over the past few decades, Poland has experienced significant structural changes in its economy. For instance, increasing trade openness, partly stimulated by integration into the European Union, the shift from exchange rate targeting to an inflation targeting strategy[1] and, more recently, the influence of the financial crisis may have led to changes in the transmission of monetary policy and exchange rate shocks. This highlights the importance of a flexible estimation framework that accounts for the possibility of time variation.

A large number of papers have analysed the impact of monetary policy and exchange rate shocks on key macroeconomic variables with standard techniques also in the case of Poland. More recently, however, a flexible estimation framework that accounts for the possibility of time variation has received attention. Taking this into account, this paper

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*This chapter is based on joint work with Dr. Olga Arratibel (ECB).

[1]The introduction of inflation targeting in 1998 helped to curb inflation, which was much higher in the 90s than in the last 10 years (Figure A.1d). For more information on Poland's monetary policy strategy, see *Medium-term strategy of monetary policy (1999-2003)* and *Monetary policy strategy beyond 2003* published by the National Bank of Poland. http://www.nbp.pl/homen.aspx?f=/en/publikacje/o_polityce_pienieznej/strategia_po_2003.html

follows the Bayesian time-varying VAR (TVP-VAR) approach with stochastic volatility developed by Primiceri (2005). We investigate whether the impact of monetary policy and exchange rate shocks has varied over time in Poland through a marginal likelihood estimation which compares a constant coefficient VAR with our TVP-VAR. Our research confirms that the TVP-VAR is indeed a better fit for Poland and, hence, that there is time variation in the impact of the shocks. To the best of our knowledge, this paper is the first attempt to estimate a Bayesian TVP-VAR with stochastic volatility for Poland and to provide formal evidence on which modeling approach is the best suited tool for analysing the impact of monetary policy and exchange rate shocks in Poland.

The main empirical findings are: (1) output seems more responsive to an interest rate shock (100 basis points interest rate increase) at the beginning of our sample, when this translates into a cumulative output cost of about 1% after two years in 1996. Notwithstanding some reversal since the beginning of the global financial crisis, from 2000 to 2007, a monetary policy shock is less costly to absorb with the output loss declining to about 0.4% after two years in 2004. The exchange rate shock (1% appreciation of the zloty) has also a time-varying effect on output. From 1996 to 2000, output seems to decline, whereas between 2000 and 2008 it has a positive significant effect. Thereafter, this effect on output mitigates. (2) Prices appear more responsive to an interest rate shock during the first half of our sample, when Poland experienced high inflation. Prices decline by about 1.4% after two years between 1996 and 1998, when the impact of the shock is the largest. In 2012, prices decline by 0.2%. During this period, the shock has the smallest impact on prices. The pass-through to consumer prices of an exchange rate shock seems to decrease slightly across time. The same exchange rate shock is also estimated for import and producer prices. It seems that the magnitude of the exchange rate shock on import prices is larger than on consumer or producer prices, confirming a decrease along the pricing chain.[2]

Overall, the findings confirm the importance of using a time-varying framework and suggest that the Polish economy has become more resilient over time to monetary policy and exchange rate shocks.

The rest of the paper is organised as follows. A brief literature overview is given in

[2]The decrease along the pricing chain, that is a stronger exchange rate pass-through on import prices than on consumer or producer prices, has been found in other studies using similar methodologies. Factors affecting the exchange rate pass-through are macroeconomic factors (such as the inflation rate and inflation persistence) and microeconomic factors (like menu costs, the size of the non-tradable sector, or the structure of imports), see e.g. Bitans (2004) for further details.

section 2.2. Section 2.3 describes the econometric model and estimation strategy. Section 2.4 briefly summarises the marginal likelihood results. The results of the TVP-VAR are presented in section 2.5. Section 2.6 summarises our robustness checks and section 2.7 concludes.

2.2 General Survey of Related Literature

Initial applications of VAR models revealed counter-intuitive results, such as the price puzzle[3] and other anomalies (Sims (1992)). A number of proposals have been made to tackle these issues.[4] In particular, identification schemes have been widely applied. In a small open economy like Poland, the price puzzle may arise when estimating monetary policy shocks. In that case, the sign restriction approach, as used by Faust (1998), Canova and Pires Pina (2000), Canova and de Nicoló (2002) and Uhlig (2005), is relevant. For our work on Poland, we follow Franta et al. (2011) (see Section 1.4.2).

Although relevant for our research, the above mentioned literature maintains the assumption of constant coefficients over time (Koop and Korobilis (2010)). This is a strong assumption because economic time series are driven by evolving features. As laid down in Canova (2007), one can think of these changes in two ways. First, as abrupt switches that can be addressed by structural breaks[5] and, second, as models with continuously evolving coefficients which capture gradual changes over time.

Allowing for stochastic volatility, but still assuming constant VAR coefficients, Uhlig (1997) introduced time variation into the VAR model. Alternatively, Cogley and Sargent (2001) developed a VAR model with drifting coefficients and a constant variance. Cogley (2005) accounted for stochastic volatility in the variance covariance matrix, but simultaneous relations among variables were nevertheless non-time-varying in his model. The salient approach by Primiceri (2005) allows the entire variance covariance matrix of the shocks as well as the coefficients to be time-varying.[6]

[3]The price puzzle denotes the counter intuitive response of a rise in inflation after a monetary policy tightening.

[4]For instance, Sims (1992) and Christiano et al. (1999) suggest to include further price variables for overcoming the price puzzle. Bernanke et al. (2005) account for an even richer data set (FAVAR).

[5]In this case, two possible models which could be applied are Markov switching or regime-switching VARs (Paap and Van Dijk (2003), Sims and Zha (2006), Teräsvirta (1994) and Koop and Potter (2006)).

[6]Following Primiceri (2005), who estimates the impact of monetary policy shocks for the US, Benati and Mumtaz (2005) apply the TVP-VAR with sign restrictions for the U.K. Other examples of this TVP-

Regarding empirical studies on Poland, there are a number of papers based on VAR methods that estimate monetary policy and/or exchange rate shocks in Central and Eastern European countries (CEEs). An excellent summary is given by Égert and MacDonald (2009). Examples of a standard VAR to examine the impact of monetary policy shocks are Creel and Levasseur (2005) and Lyziak et al. (2012). An analysis based on time-varying coefficients and contemporaneous restrictions via the standard recursive ordering is done by Darvas (2009). However, he does not account for changes in the variance covariance matrix of the shocks and, instead of a Bayesian approach, he applies a maximum likelihood framework. Jarociński (2010) estimates a structural Bayesian VAR with a combination of sign and zero restrictions. He compares the monetary policy transmission of four CEE countries (including Poland) to that of five Euro Area countries. However, his approach is based on constant coefficients and does not allow for conclusions on the evolution of the impact of the shocks across time. Concerning studies on the exchange rate pass-through in Poland, Coricelli et al. (2006) make use of a cointegrated VAR while Ca'Zorzi et al. (2007) use a standard VAR with recursive identification. Bitans (2004) also estimates a recursive VAR but on two different subsamples for Poland (1993-1999 and 2000-2003). Finally, Darvas (2001) uses an error correction model which accounts for time variation in the parameters but not in the variance matrix.

Our paper contributes to this literature with an examination of whether the impact of monetary policy and exchange rate shocks has varied across time in Poland. By allowing for time variation in the parameters and in the variance covariance matrix, we are able to analyse changes in the impact of monetary policy and exchange rate shocks across time. Given the significant structural and institutional changes experienced by the Polish economy over the last few decades, it is particularly important to take the possibility of such time variation into account. As far as we are aware, this work is the first one to address this matter and to apply a Bayesian TVP-VAR with stochastic volatility to monetary policy and exchange rate shocks as well as to provide formal evidence on which modeling approach is the preferred tool for analysing the effect of such shocks in Poland.

VAR literature are Baumeister et al. (2008) for the Euro Area and Nakajima et al. (2011) for Japan. A growing number of papers also estimate TVP-VARs to analyse dynamics in, for example, fiscal policy (Kirchner et al. (2010), Pereira and Lopes (2010)), oil prices (Baumeister and Peersman (2008)) and exchange rates (Mumtaz and Sunder-Plassmann (2010)).

2.3 Empirical Model

Our empirical approach closely follows Primiceri (2005). It is a multivariate time series framework with time-variation in the coefficients as well as in the covariances of the residuals. Varying coefficients capture possible nonlinearities or time-variation in the lag structure of the model. Furthermore, the varying variance covariance matrix accounts for possible heteroscedasticity of the shocks as well as nonlinearities in the simultaneous relationships between the variables.

We estimate the following VAR model:

$$y_t = c_t + B_{1,t}y_{t-1} + \ldots + B_{l,t}y_{t-l} + u_t, \tag{2.1}$$

where $t = 1, \ldots, T$; the vector of endogenous variables y_t is of the size n x 1; c_t, the vector of time-varying coefficients which multiply constant terms is of the size n x 1; the time-varying coefficients $B_{i,t}$, with the lag length $i = 1, \ldots, l$, have the size n x n; and u_t, size n x 1, are unknown heteroscedastic shocks with time-variation in the covariance matrix of the residuals Ω. The stochastic covariance matrix of the residuals u_t is factored as

$$VAR(u_t) \equiv \Omega_t = A_t^{-1}H_t(A_t^{-1})', \text{with } H_t = \Sigma_t\Sigma_t'. \tag{2.2}$$

The time-varying diagonal matrix Σ_t and the time-varying lower triangular matrix A_t are denoted as:

$$\Sigma_t = \begin{bmatrix} \sigma_{1,t} & 0 & \ldots & 0 \\ 0 & \sigma_{2,t} & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ 0 & \ldots & 0 & \sigma_{n,t} \end{bmatrix} \quad A_t = \begin{bmatrix} 1 & 0 & \ldots & 0 \\ \alpha_{21,t} & 1 & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ \alpha_{n1,t} & \ldots & \alpha_{n(n-1),t} & 1 \end{bmatrix} \tag{2.3}$$

The time-varying VAR can then be summarised as:

$$y_t = X_t'\tilde{B}_t + A_t^{-1}\Sigma_t\varepsilon_t, \tag{2.4}$$

where $X_t = I \bigotimes [1, y_{t-1}', \ldots, y_{t-l}']$, $\tilde{B} = vec([c_t, B_{1,t}, \ldots, B_{l,t}])$ and $VAR(\varepsilon) = I_n$.

The possibility of time-variation in A_t in equation 2.4 permits the shock to one endogenous variable to have a time-varying effect on the other variables in the system. This is a crucial aspect for modeling simultaneous relations among variables. It provides a flexible approach for estimating the transmission mechanism of structural innovations, particular important for transition economies like Poland.

The dynamics of the time-varying parameters (B_t and A_t) are following a driftless random walk, whereas the covariance matrix (Σ_t) evolves as a geometric driftless random walk:

$$B_t = B_{t-1} + \nu_t, \tag{2.5}$$

$$\alpha_t = \alpha_{t-1} + \xi_t, \tag{2.6}$$

$$log\sigma_t = log\sigma_{t-1} + \eta, \tag{2.7}$$

where α_t is a vector, stacked by rows, of only non-zero and non-one elements of the matrix A_t and the standard deviation σ_t is a vector containing the diagonal elements of the matrix Σ_t. The vector of innovations $[\varepsilon_t', \nu_t', \xi_t', \eta_t']$ is distributed according to the following assumption:

$$\begin{bmatrix} \varepsilon_t \\ \nu_t \\ \xi_t \\ \eta_t \end{bmatrix} \sim N(0, V), \ \text{ with } V = \begin{bmatrix} I_n & 0 & 0 & 0 \\ 0 & Q & 0 & 0 \\ 0 & 0 & S & 0 \\ 0 & 0 & 0 & W \end{bmatrix}, \tag{2.8}$$

where I_n is an n dimensional identity matrix and Q, S and W are positive definite matrices. S is assumed to be block diagonal, implying that the parameters of the simultaneous relations among variables are restricted to be independent. The respective $n-1$ blocks of S relate each to separate equations.

Specifying the underlying dynamics on the basis of the random walk provides a flexible framework. It allows to capture the evolution of different parameters coming from policy and structural changes in the economy.

2.3.1 Priors

VARs are not parsimonious models. Usually the estimation of VAR models require a large amount of parameters which can easily add up to a few hundred. Without prior information, it is almost impossible to obtain precise estimates.[7]

To specify the priors, we use a training sample based on the whole sample 1996Q1-2012Q3 (see Appendix A). We follow Canova (2007) and Canova and Ciccarelli (2009), who motivate this approach when a separate training sample is not available.[8] Therefore, we run an OLS estimation on a fixed-coefficient VAR model for calibrating our priors.

The mean and the variance of B_0 are, respectively, the OLS point estimates ($\hat{B}_{OLS}$) and four times their variance. The same holds for the prior distribution of the simultaneous relation matrix A_0. For the log standard errors, the prior mean is specified as the log of the respective OLS point estimates, whereas the prior covariance matrix is restricted to be I_n. The hyperparameters Q, S and W are the covariance matrices of the innovations (see equations 2.5, 2.6 and 2.7). Matrices Q and S follow the inverse-Wishart prior distribution and we follow Cogley and Sargent (2005) for defining W, which is based on the inverse-Gamma prior distribution. Furthermore, we restrict the matrix W to be diagonal for reducing the dimensionality of the estimation.

$$\begin{aligned}
B_0 &\sim N(\hat{B}_{OLS}, 4 \cdot V(\hat{B}_{OLS})),\\
A_0 &\sim N(\hat{A}_{OLS}, 4 \cdot V(\hat{A}_{OLS})),\\
log\sigma_0 &\sim N(\hat{\sigma}_{OLS}, 4 \cdot I_n),\\
Q &\sim IW(k_Q^2 \cdot \tau \cdot V(\hat{B}_{OLS}), \tau),\\
W &\sim IG(k_W^2 \cdot (1 + dim(W)) \cdot I_n, (1 + dim(W))),\\
S_b &\sim IW(k_S^2 \cdot (1 + dim(S_b)) \cdot V(\hat{A}_{b,OLS}), (1 + dim(S_b))),
\end{aligned}$$

where τ has the size of the training sample, S_b refers to the respective blocks of S and $\hat{A}_{b,OLS}$ denotes the respective blocks of $\hat{A}_{OLS}$. The parameters $k_Q = 0.05, k_W = 0.1$ and $k_S = 0.01$ specify prior beliefs about the amount of time variation in the estimates of the coefficients, covariances and volatilities. For example, for the OLS estimation of the covariance matrix of the VAR coefficients, we allow for 5% ($k_Q = 0.05$) of uncertainty surrounding the $V(\hat{B}_{OLS})$ estimates to time variation.

In order to justify our selection of k_Q, k_W and k_S, we do a formal model selection. Posterior

[7] In a VAR model, the number of free parameters increases substantially with the number of endogenous variables and lags (e.g., for a VAR with four variables and two lags, Q comprises 666 free parameters).

[8] For robustness, we also estimate the priors on a subset of the sample (1996Q1-2007Q4). Please refer to Section 2.6 for further details.

probabilities for a set of 18 models are estimated[9] based on the reversible jump Markov chain Monte Carlo (RJMCMC) method (see Primiceri (2005)). The selection of k_Q, k_W and k_S delivers a posterior probability for one model which is almost one. Table 2.4 in the Appendix B reports the posterior probability estimates for the set of 18 models.

Regarding the degrees of freedom for W and S_b, they are defined as one plus the dimension of each matrix. For Q they are set equal to the size of the training sample.

2.3.2 Estimation

So far, we have outlined the estimation strategy for a reduced form VAR which is estimated using Bayesian methods for the sample from 1996:Q1 to 2012:Q3. For maintaining the degrees of freedom, two lags are used. For approximating the posterior distribution, $40,000$ iterations of the Gibbs sampler are used and we drop the first $20,000$ iterations for convergence. For breaking the autocorrelation of the draws, only every $10th$ iteration is kept. Our final estimates are therefore based on $2,000$ iterations. The sample autocorrelation functions of the draws die out rather quickly. Furthermore, the convergence diagnostics reveal satisfactory results (a detailed overview is given in Appendix F).

To identify monetary and exchange rate shocks[10] we follow Jarociński (2010), Franta et al. (2011), Farrant and Peersman (2006) and An and Wang (2011). We assume an open economy with a flexible exchange rate and allow for simultaneous responses among monetary policy and exchange rate shocks.[11] Furthermore, our exchange rate shock restrictions are consistent with the uncovered interest rate parity condition.[12]

In order to identify the shocks, some restrictions are assumed and imposed on the impulse responses, both at the time of the impact as well as in the first and second period (see Table 2.1). We use zero and sign restrictions as follows[13]:

[9] The set of 18 models are constructed from all possible combinations of $k_Q = \{0.01; 0.05; 0.1\}, k_W = \{0.001; 0.01\}$ and $k_S = \{0.01; 0.025; 0.1\}$.

[10] An extensive amount of literature focuses on the identification of monetary policy shocks. For a review refer to Christiano et al. (1999)

[11] This applies also to the beginning of our sample, before Poland adopted a free floating exchange rate regime.

[12] The uncovered interest rate parity condition states that interest rate differentials account for expected changes in the exchange rate.

[13] Faust (1998), Canova and de Nicoló (2002) and early versions of Uhlig (2005) were quite influential for the application of sign restrictions. For instance, Artis and Ehrmann (2006) use a SVAR and identify monetary and exchange rate shocks applying short-run zero restrictions.

- No simultaneous response of GDP and prices either to a monetary policy or exchange rate shock.
- A monetary policy shock (100 basis points (BPs) rise in the policy interest rate) leads to an appreciation of the exchange rate.
- An exchange rate shock (1% rise in the exchange rate) is associated with a decrease in the interest rate and an exchange rate appreciation.

Table 2.1: Sign Restrictions

		MP Shock	ExR Shock
GDP	Lag 0	0	0
	Lag 1	?	?
	Lag 2	?	?
Prices	Lag 0	0	0
	Lag 1	?	?
	Lag 2	?	?
Interest Rate	Lag 0	$\geq$	$\leq$
	Lag 1	$\geq$	$\leq$
	Lag 2	$\geq$	$\leq$
Exchange Rate	Lag 0	$\geq$	$\geq$
	Lag 1	$\geq$	$\geq$
	Lag 2	$\geq$	$\geq$

Note: ? denotes no restriction, $\geq$ defines a positive effect of the respective shock on the variable, vice versa for $\leq$.

For implementing the sign restrictions, we need to slightly modify the model specified in equations 2.4 - 2.7. So far, it is based on the recursive identification scheme. We additionally specify an orthonormal rotation matrix G_t, i.e. $G_t'G_t = I_n$ The model in equation 2.4 can then be rewritten as

$$y_t = X_t'\tilde{B}_t + A_t^{-1}\Sigma_t G_t' G_t \varepsilon_t = X_t'\tilde{B}_t + A_t^{-1}\Sigma_t G_t' \tilde{\varepsilon}_t. \tag{2.9}$$

$\tilde{\varepsilon}_t = G_t'\varepsilon_t$ denotes the new shocks and the respective variance is $Var(\tilde{\varepsilon}_t) = G_t I_n G_t'$. Technically, the sign restrictions are implemented using the QR-decomposition method for finding G_t. We have a four variable VAR, implying a 4 x 4 G_t matrix. Due to zeros in the first two rows of the sign restriction matrix, the decomposition is restricted to the last two columns. Thus, the G_t matrix has the following form:

$$G_t = \begin{pmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & \multicolumn{2}{c}{\multirow{2}{*}{$\begin{bmatrix} QR\left(\theta\left[1,1\right]\right) & QR\left(\theta\left[1,2\right]\right) \\ QR\left(\theta\left[2,1\right]\right) & QR\left(\theta\left[2,2\right]\right) \end{bmatrix}$}} \\ 0 & 0 & & \end{pmatrix}. \tag{2.10}$$

In a *first* step, we draw a 2 x 2 matrix, θ, from the $N(0,1)$ distribution. In a *second* step, we take the QR decomposition of θ and construct the G_t matrix. This algorithm calculates a candidate structural impact matrix. In a *third* step, we check whether this matrix is in line with the sign restrictions. Finally, in a *fourth* step, if the matrix satisfies the restrictions it is stored; otherwise, another θ is drawn from the standard normal distribution and we repeat the procedure from the second step.[14]

This form ensures the respective zero restrictions on GDP and prices, so that the structural shocks to monetary policy and the exchange rate do not simultaneously influence GDP and prices.

2.4 Empirical Evidence of Time Variation in Poland: is there any?

As a first step, we search for formal econometric evidence on whether the impact of monetary and exchange rate shocks in Poland has changed across time. In particular, we calculate marginal likelihood estimates for a traditional constant-coefficient vector autoregressive (VAR)[15] model and our time-varying parameter (TVP-VAR) model with stochastic volatility. The model that yields the largest marginal likelihood fits the given data the best. We follow Nakajima et al. (2011) and use the modified harmonic mean estimator of the marginal likelihood due to Geweke (1999).[16] The log marginal likelihood

[14]Maximum number of possible draws for θ is 100. In case a candidate structural impact matrix is not obtained, we move to the next iteration of the Gibbs sampler. On average, 19 values of θ have to be drawn to generate the structural impact matrix that satisfies all sign restrictions. The fraction in which the structural impact matrix does not satisfy the sign restrictions is only 4.49%.

[15]Prior for the constant parameter VAR: $B \sim N(0, 4 \times I)$, $\alpha \sim N(0, 4 \times I)$, $\sigma^{-1} \sim \text{Gamma}(2, 0.02)$.

[16]For a detailed description of the harmonic mean estimator, please refer to Nakajima et al. (2011). The marginal likelihood calculation is based on the priors and number of lags as specified above. Additionally, we have to specify the parameter τ. We follow Nakajima et al. (2011) and set $\tau = 0.99$.

value for the TVP-VAR, 356.852, is higher than the marginal likelihood estimate for the constant VAR, 173.675, suggesting that the TVP-VAR model with stochastic volatility is indeed a better model for Poland than the constant VAR.

2.5 Results of the TVP-VAR

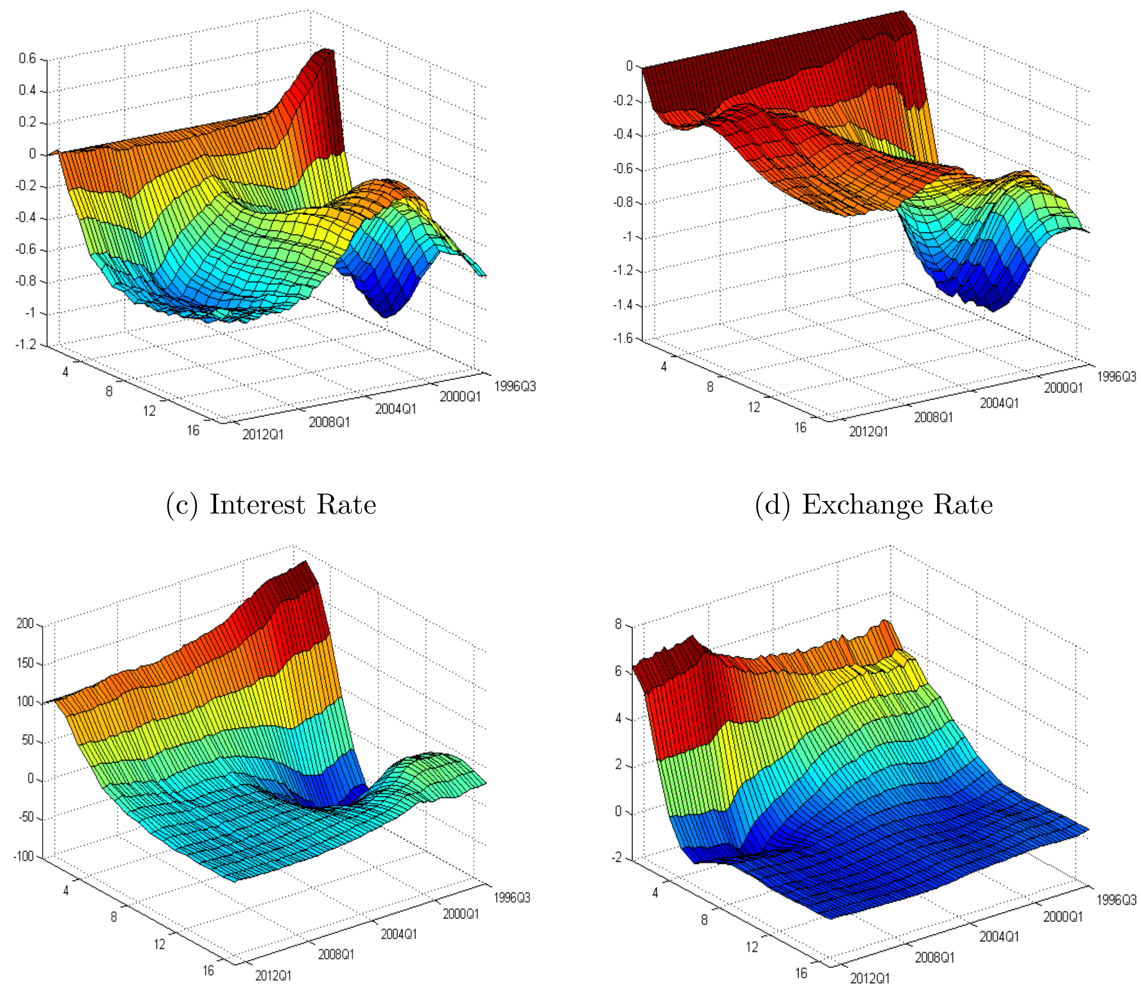
In what follows, sections 2.5.1 and 2.5.2 present, respectively, the estimated median impulse responses of the monetary policy and exchange rate shocks (see also Appendix C) and include an analysis on the posterior probability for the difference in the impulse responses. In section 2.5.2, we also substitute the HICP index with an index of import prices or producer prices to analyse the pass-through of an exchange rate shock on these price levels (the respective Figures are given in Appendix D). The time-varying posterior estimates of the stochastic covariance matrix are presented in Appendix E and Appendix F summarises the estimation on convergence diagnostics.

2.5.1 Impulse Responses to Monetary Policy Shocks

Figure 2.1 presents the median impulse responses (over 17 quarters and the time period: 1996:1-2012:3) to a 100 basis points (BPs) interest rate increase in the given period across the sample. This monetary policy shock has the expected impact on GDP (↓), prices (↓), interest rate (↑) and exchange rate (initially ↑).

We clearly see that a monetary policy shock has time-varying effects. Specifically, the decline in *real GDP* after a monetary policy shock is stronger in the beginning of the sample, while since 2000 until 2008 it is weaker. More specifically, between 1996 and 1998 the cumulative output loss stands at about 1% after eight quarters compared to only 0.4% in 2004Q3 (Figure 2.1a). These results are similar to those found by Lyziak et al. (2012) in a structural VAR accounting for boom/bust cycles. Our results may partly reflect the adoption of an inflation targeting framework by the Polish central bank in 1998 and the fact that a more credible central bank is generally able to achieve its inflation objective at lower output costs, see also Darvas (2009). Since the beginning of the financial crisis in 2008, real GDP seems to react somewhat stronger again, but this effect is nevertheless insignificant (Figures C.2a, C.2b).

Figure 2.1: Time-Varying Impulse Responses to a Monetary Policy Shock



(a) GDP

(b) Prices

(c) Interest Rate

(d) Exchange Rate

Median impulse responses to a 100 BPs monetary policy shock.

Regarding the effect of a monetary policy shock on *prices*, they exhibit a very large degree of time variation across our sample. In line with the theory, prices decrease after a monetary policy shock (Figure 2.1b). The impact on prices seems to be strongest between 1996 and 2001, a period during which Poland experienced high inflation (Figure A.1d). The largest accumulative effect is estimated at about 1.4% in 1996Q4 after eight quarters. A possible explanation for this time variation could be that, at the beginning of our sample, the central bank managed to curb inflation significantly and bring it down to a more moderate rate. This may have contributed to enhancing the central bank's credibility and explain the weaker impulse responses from 2004 onwards. At the end of the sample in 2012Q1 the median impulse response decreases to about 0.3% after four quarters.

The effect of the monetary policy shock on the *interest rate* is particularly stable since 2002 (Figure 2.1c). Interestingly, this effect has not changed since the beginning of the global financial crisis.

Finally, the impact of a monetary policy shock on the *nominal effective exchange rate* is, as expected, initially positive (Figure 2.1d). Furthermore, this shock seems to be absorbed much more quickly since 2004. This is in contrast to Darvas (2009) who, in a setting that accounts for time-varying coefficients in a VAR with recursive identification and a constant variance covariance matrix, estimates rather stable impulse responses of the exchange rate over time. This leads him to conclude that there is time variation mainly in real GDP. In contrast, our results reveal time variation next to GDP, also in prices and in the exchange rate.

Comparison of impulse responses at different horizons and points in time

The evolution of the responses at the 4th and 8th quarter with their percentiles is given in Figures C.2a and C.2b in Appendix C. In terms of real output, a monetary tightening has a negative significant effect at the 8th quarter horizon at the end of the 90s, while it does not have any significant impact afterwards. The same holds for the effect on prices for the 4th and 8th quarter. Concerning the interest rate, the impact of the monetary policy shock converges to zero after two years (see Figures C.2a and C.2b). The influence on the nominal effective exchange rate seems to be different across time (Figures C.2a and C.2b), converging more quickly towards zero from 2004 onwards.

For a better illustration of the difference in the impulse responses to a monetary policy shock across time, we also present Figures 2.2a and 2.2b. These allow for a comparison of impulse responses at specific points in time. Figure 2.2a plots the median impulse responses at 1996Q3, 2000Q1 and 2012Q1. The three different time periods for the comparison are chosen arbitrarily. The period around 1996Q3 reflects the environment under the exchange rate targeting regime, 2000Q1 under inflation targeting and 2012Q1 under the influence of the financial crisis. Figure 2.2b plots the impulse responses at 2000Q1 and 2012Q1 with percentiles. Especially for the price impulse responses, there seems to be a strong difference between 2000Q1 and 2012Q1. This result is also supported by our

Table 2.2: Posterior probability for the difference in the impulse responses to a monetary policy shock at different time periods

Horizon	1 Q (%)	4 Q (%)	8 Q (%)	12 Q (%)
GDP				
1996/2000	12.9	39.8	64.9	65.2
1996/2012	12.4	43.3	64.1	52.3
2000/2012	12.8	53.9	50.6	36.6
HICP				
1996/2000	43.7	55.9	52.4	53.3
1996/2012	43.5	71.3	82.7	73.1
2000/2012	44.5	68.7	90.6	77.2
IR				
1996/2000	52.6	43.4	42.9	43.4
1996/2012	49.0	40.8	62.9	41.0
2000/2012	45.6	47.1	74.4	46.8
ExR				
1996/2000	54.9	55.5	47.5	46.0
1996/2012	58.4	34.4	33.7	41.1
2000/2012	54.5	26.8	35.4	45.8

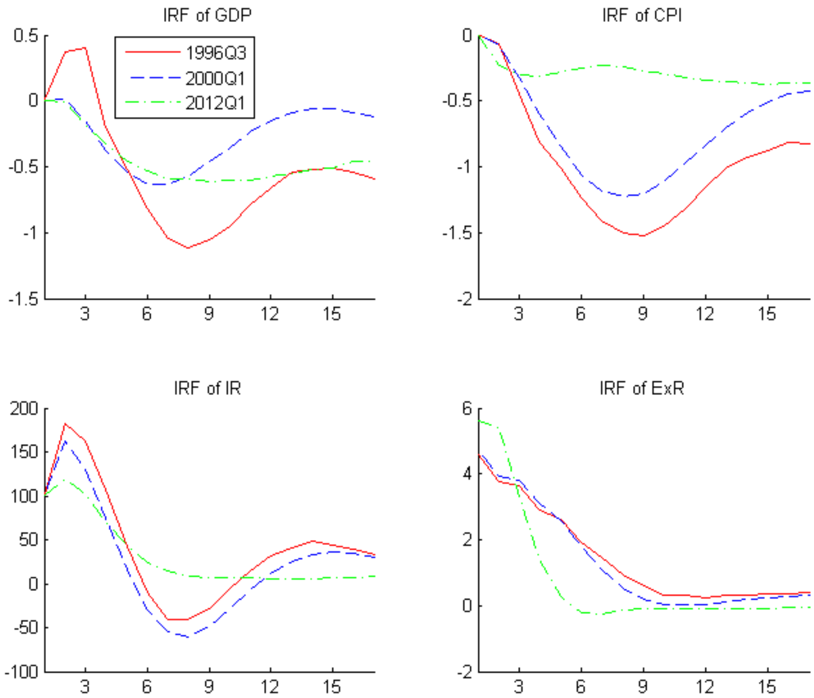
Note: Difference in impulse responses at the time periods 1996Q3, 2000Q1 and 2012Q1 for one, four, eight and 12 quarters ahead.

analysis on the posterior probability for the difference in the impulse responses.

We consider the statistical difference in the impulse responses between different time periods by calculating the ratio of the MCMC draws of the responses between two time periods. More specifically, we estimate the posterior probability that the response at one given time period (first considered response) is smaller than at another given time period (second considered response). We consider again the three time periods referred to above and present the differences in the impulse responses to the monetary policy shock in Table 2.2. Posterior probability values close to 50% indicate a weak difference between the two periods. Values above (below) 50% imply that the first response is smaller (bigger) than the second response. The posterior difference for GDP to a monetary policy shock between the three considered time periods is stronger for one-quarter ahead and becomes weaker for the other quarters ahead. Regarding prices, we estimate a strong difference in responses between 1996Q3 and 2012Q1 as well as between 2000Q1 and 2012Q1 for the 8th quarter and 12th quarter ahead. The evidence for the exchange rate responses is rather strong between 1996Q3 and 2012Q1 as well as 2000Q1 and 2012Q1 for the 4th quarter and 8th quarter ahead while the responses between 1996Q3 and 2000Q1 is weaker.

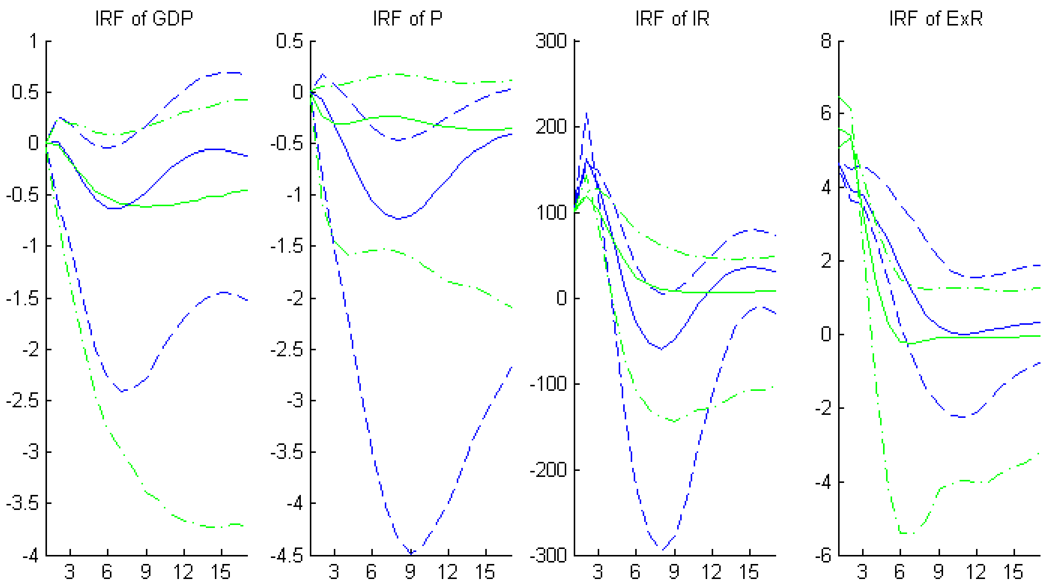
Figure 2.2: Responses at Different Time Periods to a Monetary Policy Shock

(a) Responses without Percentiles



Median impulse responses to a 100 BPs monetary policy shock at 1996Q3, 2000Q1 and 2010Q3.

(b) Responses with Percentiles



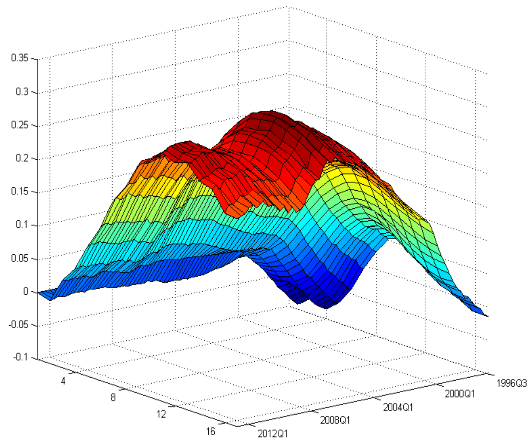
Median impulse responses (solid line) to a 100 BPs monetary policy shock with 16-th and 84-th percentiles (dashed line) of the posterior distribution at 2000Q1 (blue) and 2012Q1 (green).

2.5.2 Impulse Responses to Exchange Rate Shocks

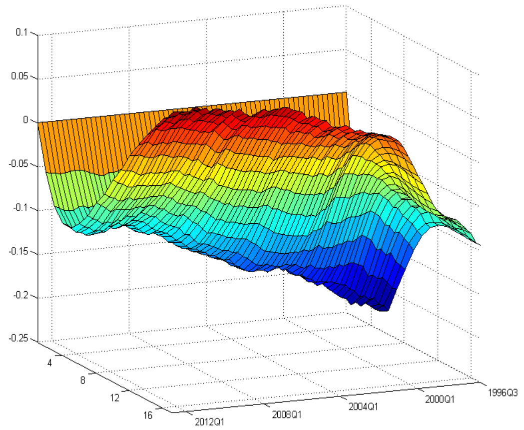
In this section, we analyse the median impulse responses to a 1% appreciation in the nominal effective exchange rate over time (17 quarters, time period: 1996:1-2012:3) (see Figure 2.3). The estimated pass-through of an exchange rate shock is in line with the theory and highlights the importance to account for time variation.

Figure 2.3: Time-Varying Impulse Responses to an Exchange Rate Shock

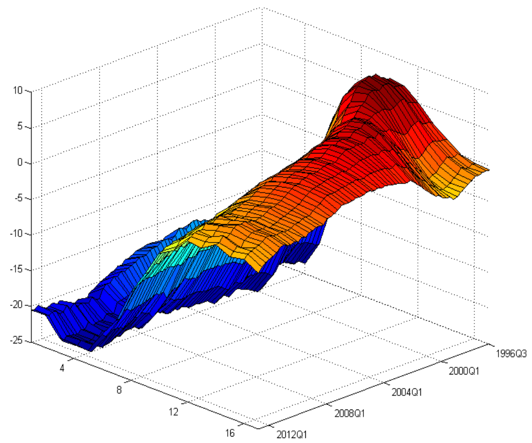
(a) GDP



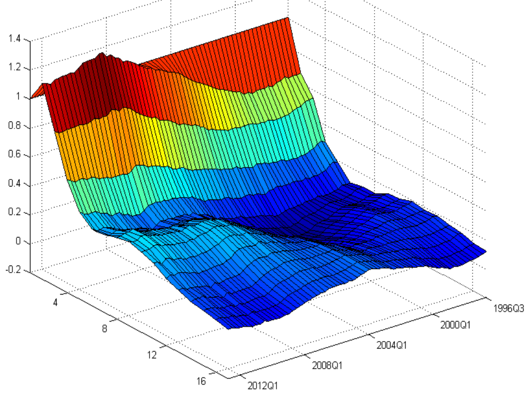
(b) Prices



(c) Interest Rate



(d) Exchange Rate



Median impulse responses to a 1% exchange rate appreciation.

Regarding the effect of the exchange rate shock on *output*, it can be a mixed one depending on whether the expenditure-switching channel (negative effect on output, since exports decline due to appreciation) or the interest rate channel (positive effect on output, since interest rates decline following an appreciation) dominates. Our empirical findings suggest that the expenditure-switching channel prevails from 1996 until 2000 (Figure 2.3a), albeit its effect seems to be insignificant (Figures C.2c, C.2d). Since 2000, however, it appears to become less costly to absorb exchange rate shocks with respect to output. A possible explanation for this time variation is that at the beginning of the sample, Poland did not have a free floating exchange rate. In such a context, the interest rate channel is less important since domestic money market rates follow foreign interest rates (Cevik and Teksoz (2012)). As for the positive impact since 2000, the rise in GDP may not only result from the stimulating impact of decreasing interest rates after an exchange rate appreciation in a flexible exchange rate regime, but it may also indicate economic convergence which is not captured by the model. To ensure that the positive effect on output is not driven by the lag of foreign variables, we follow Franta et al. (2011) and estimate a quarterly VAR with exogenous foreign variables.[17] Also in this specification, GDP increases after an exchange rate shock, confirming the robustness of our results.

Concerning *prices*, our results confirm the general finding in the literature of decreasing inflation following an appreciation of the zloty (Figure 2.3b). However, our findings suggest that prices respond with a slightly decreasing pass-through to an exchange rate shock, with the median impulse response declining to about 0.2% in 1996Q4 and to about 0.1% in 2012Q1 after six quarters.[18] We also investigate the time-varying effect of an exchange rate pass-through on import and producer prices. To our knowledge, this has not been attempted in the economic literature yet. As expected, import prices reveal a stronger decline than consumer or producer prices (Appendix D, Figures D.3, D.4 and D.5). Furthermore, both import and producer prices converge faster to zero than consumer prices. Concerning import prices, it seems that the pass-through is strongest between 1996 and 2000 (Figures D.3b, D.5a and D.5b), whereas for producer prices, the pass-through appears to have increased since 2000 (Figure D.4b). This decline across the pricing chain is well documented in the literature and also estimated by other studies on Poland (Bitans (2004), Ca'Zorzi et al. (2007), McCarthy (2007)).

Regarding the impact on the *interest rate*, exchange rate shocks seem to be accommodated

[17]Specifically, we add the following four variables: EA GDP at market price, chain linked volumes, 2005=100, seasonally adjusted; EA Commodity Price Index; EA Euribor 3-month, average of observations through period; EA HICP, overall monthly index, seasonally adjusted.

[18]A similar result is estimated by Bitans (2004). He uses a recursive VAR - however, with constant coefficients - and accounts for two subsamples (1993-1999 and 2000-2003).

by interest rate decreases (in the range of roughly -20 basis points within the first year, see Figure 2.3c). As illustrated above, this in turn might also stimulate output. The impact converges to zero after about two years. Finally, the impact of the exchange rate shock on the *exchange rate* itself dies out quickly, approximately after one year (Figure 2.3d).

Comparison of impulse responses at different horizons and points in time

A comparison of the responses at the 4th and 8th quarter is given in Appendix C, Figures C.2c and C.2d. Consumer prices seem to respond significantly negative to an exchange rate shock, but the effect on GDP is only significant after two years between the period 2000 and 2008 (Figures C.2c and C.2d).

Table 2.3: Posterior probability for the difference in the impulse responses to an exchange rate shock at different time periods

Horizon	1 Q (%)	4 Q (%)	8 Q (%)	12 Q (%)
GDP				
1996/2000	13.0	77.3	73.5	72.8
1996/2012	12.3	79.4	80.9	80.0
2000/2012	12.4	60.7	68.5	68.7
HICP				
1996/2000	44.1	56.0	52.4	59.8
1996/2012	44.1	69.2	57.9	60.6
2000/2012	45.6	70.1	58.0	51.6
IR				
1996/2000	46.1	47.5	35.6	50.8
1996/2012	45.2	47.6	32.6	42.4
2000/2012	49.4	51.5	43.4	40.8
ExR				
1996/2000	53.9	50.6	44.7	47.7
1996/2012	51.5	61.1	61.1	57.8
2000/2012	48.1	62.1	68.3	61.2

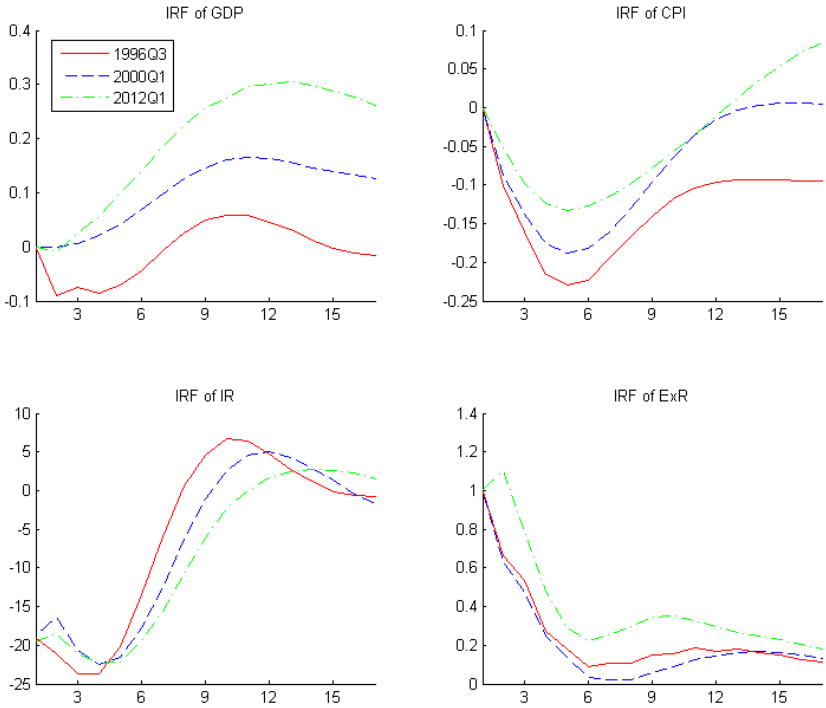
Note: Difference in impulse responses at the time periods 1996Q3, 2000Q1 and 2012Q1 for one, four, eight and 12 quarters ahead.

As before, for a better illustration of the difference in the impulse responses across time, we plot the median impulse responses at 1996Q3, 2000Q1 and 2012Q1 (see Figure 2.4a). Figure 2.4b plots the impulse responses at 1996Q3 and 2012Q1 with their percentiles. Especially for GDP, there seems to be a difference between the impulse responses at 1996Q3 and 2012Q1.

We also evaluate the statistical difference in the impulse responses to the exchange rate shock at different time periods. The estimates of the posterior probability indicate time variation between those periods as well. As shown in Table 2.3, the responses for GDP between 1996Q3 and 2000Q1 as well as between 1996Q3 and 2012Q1 reveal a clear difference. The differences in responses for prices and the interest rate are weaker. Concerning the exchange rate, we estimate a slightly stronger difference in the impulse response between 2000Q1 and 2012Q1.

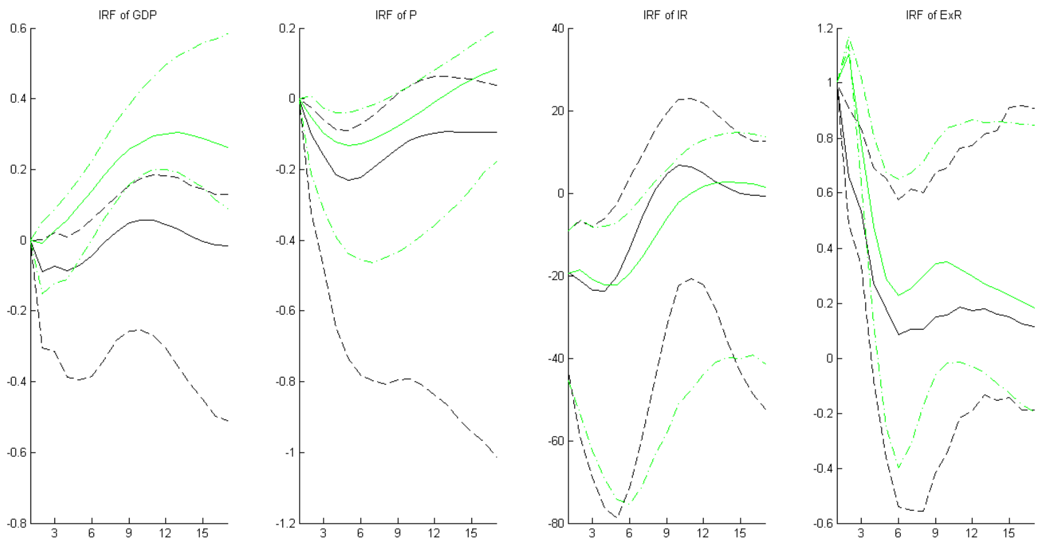
Figure 2.4: Responses at Different Time Periods to an Exchange Rate Shock

(a) Responses without Percentiles



Median impulse responses to a 1% exchange rate appreciation at 1996Q3, 2000Q1 and 2012Q1.

(b) Responses with Percentiles



Median impulse responses (solid line) to a 1% exchange rate appreciation with 16-th and 84-th percentile (dashed line) of the posterior distribution at 1996Q3 (black) and 2012Q1 (green).

2.6 Robustness Checks on Priors

Since data for Poland are only available with a short time horizon, calibrating the priors is a challenge. For a robustness check, we estimate the priors on a subset of the sample (1996Q1 - 2007Q4) and obtained results that support those presented in this paper. We also extend our dataset with data for GDP and prices constructed by Darvas (2009)[19] and estimate the priors on two different training samples. Both are based on data from 1993Q1 until 2007Q4, whereas for the second training sample the initial years (1993Q1-1995Q4) are dropped. The results also confirm the findings presented in this paper.

As a final robustness check, we change the prior for B_0 to a hierarchical prior which combines the Minnesota prior and the TVP-prior. This is because the TVP-prior could suffer from over-parameterization and the risk of over-fitting increases with a short time horizon. Mitigating these issues is possible with the help of the Minnesota prior that provides for a shrinkage. The results obtained confirm those presented in this paper.[20]

2.7 Conclusion

By applying the TVP-VAR developed by Primiceri (2005), this paper represents the first attempt at analysing the impact of monetary policy and exchange rate shocks in a fully time-varying model in Poland. Our findings show that the reaction of macroeconomic variables in the Polish economy to monetary policy and exchange rate shocks has, indeed, varied across time. Next to the exchange rate, prices and output reveal considerable time-varying effects across our sample from 1996 until 2012. Overall, our results suggest that the Polish economy has become more resilient to these shocks over time.

More specifically, a monetary policy shock (tightening) - which does affect negatively and significantly GDP after around two years - seems to have a stronger impact on output at the end of the 90s (maximum decrease of about 1%) than between 2000 and 2008 (decrease of about 0.5%). Since the financial crisis in 2008, output seems to react somewhat stronger again. Following the same monetary policy shock on prices, we estimate a strong decline until 2001 (maximum decline of about 1.4%). From 2004 onwards, the effect on prices

[19] He constructs quarterly GDP data based on mainly annual GDP series. The price index is a core inflation measure as in Darvas (2001). These data are used for the period from 1993Q1 until 1995Q4.

[20] We gladly provide all our robustness checks upon request.

has become weaker. Interestingly, interest rate responses are rather stable across time and the effect on the nominal effective exchange rate converges much faster to zero after 2004.

The exchange rate shock, defined as an appreciation of the nominal effective exchange rate, has a considerable time-varying effect on output. From 1996 to 2000, the expenditure-switching channel prevails. Thereafter, the interest rate channel seems to dominate, leading to a positive effect on output. Following an exchange rate appreciation, consumer prices appear to decline, although it seems that this pass-through is somewhat decreasing across time (in 1996Q4 -0.2%, in 2012Q1 -0.1% after six quarters). Among the three price indices considered, import prices show the strongest reaction to an exchange rate shock.

We would like to stress the different robustness checks conducted for testing the consistency of our results. The various checks, inter alia in the prior specifications and in the data sample, confirm the findings presented in this paper. The use of the TVP-VAR with stochastic volatility is also supported by a marginal likelihood estimation based on the harmonic mean estimator that compares the TVP-VAR with a constant BVAR. Moreover, a sophisticated model selection algorithm is used to ensure the correct specification of the prior beliefs about the amount of time variation.

For future work on Poland, provided data availability allows, it would be interesting to apply the time-varying factor augmented VAR framework (TVP-FAVAR, Koop and Korobilis (2010)). This would allow to compare our results with those found on the basis of a richer dataset. Furthermore, it would be interesting to compare the effects of monetary policy and exchange rate shocks in Poland with those in other CEE countries and the Euro Area.

A Data sources

This paper uses quarterly data on Poland and covers a time horizon between 1996:1 and 2012:3. We estimate the model in levels. Like Sims et al. (1990) state, this accounts for possible discrepancy which may arise in case of incorrectly assumed cointegration restrictions. Also, if there are unit roots in the data, it will not influence the likelihood function, since nonstationarity is of no concern in a Bayesian framework. In the following, the used time series are described:

Gross domestic product (GDP): Log of gross domestic product at market prices, chain linked volumes, reference year 2005, in million zloty, working day and seasonally adjusted by TRAMO/SEATS, quarterly series. Source: European Central Bank.

Consumer price (CPI): Log of HICP, overall index (2005=100), monthly index converted to a quarterly series (averaging over three respective months), neither seasonally nor working day adjusted. Source: Eurostat.

Short-term interest rate (IR): Money market interest rate, deposit liabilities, 3 months (80-100 days) maturity, in percent, denominated in Polish zloty. Source: Eurostat.

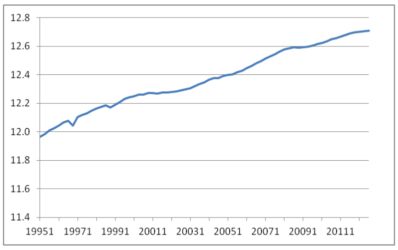
Exchange rate (ExR): Log of ECB nominal effective exchange rate, Euro area-17 countries vis-a-vis the EER-40 group of trading partners (AU, CA, DK, HK, JP, NO, SG, KR, SE, CH, GB, US, BG, CZ, LV, LT, HU, PL, RO, CN, DZ, AR, BR, CL, HR, IS, IN, ID, IL, MY, MX, MA, NZ, PH, RU, ZA, TW, TH, TR and VE) against Polish zloty. Monthly index (reference period: 99Q1=100) converted to a quarterly series (averaging over three respective months). Source: European Central Bank.

Import price (ImpP): Log of import prices of goods and services, overall index, quarterly series (reference year 2000), in national currency, seasonally and working day adjusted. Source: Eurostat.

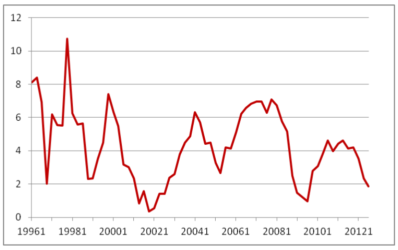
Producer price (ProdP): Log of industry producer prices, overall index, total output prices (industry [except construction, sewage, waste management and remediation activities]), quarterly series (reference year 2005), in national currency, gross data. Source: Eurostat.

Figure A.1: Quarterly Data, Poland

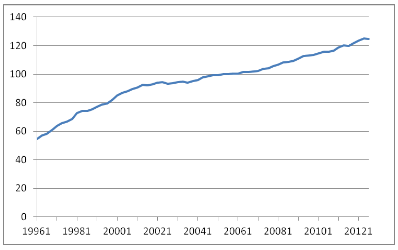
(a) GDP in levels



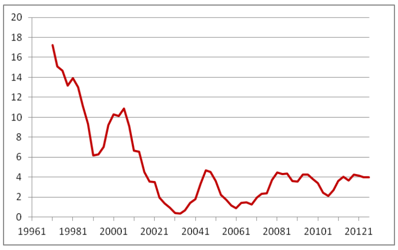
(b) GDP, yearly growth rate



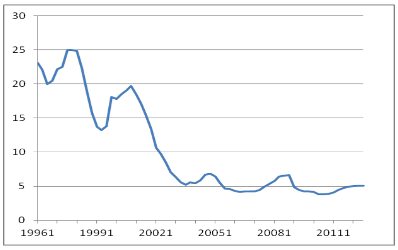
(c) CPI in levels



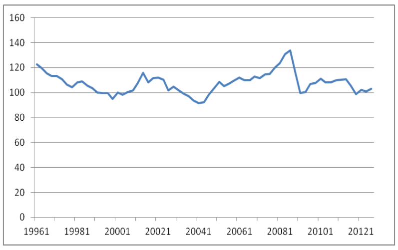
(d) CPI, annual rate of change



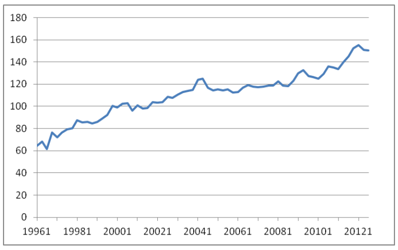
(e) IR in percent



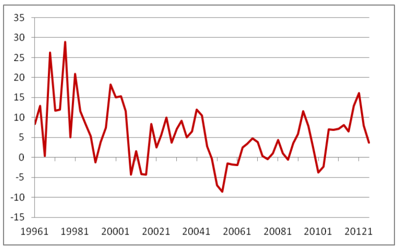
(f) ExR in levels



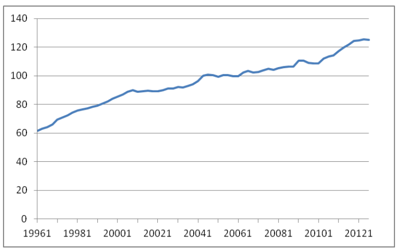
(g) Import Price Index in ln levels



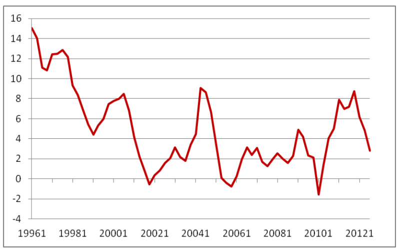
(h) Import Price Index, annual rate of change



(i) Producer Price Index in ln levels



(j) Producer Price Index, annual rate of change



B Posterior Probability Estimates for k_Q, k_W and k_S

Table 2.4: Posterior Probability Estimates for k_Q, k_W and k_S based on the RJMCMC Method

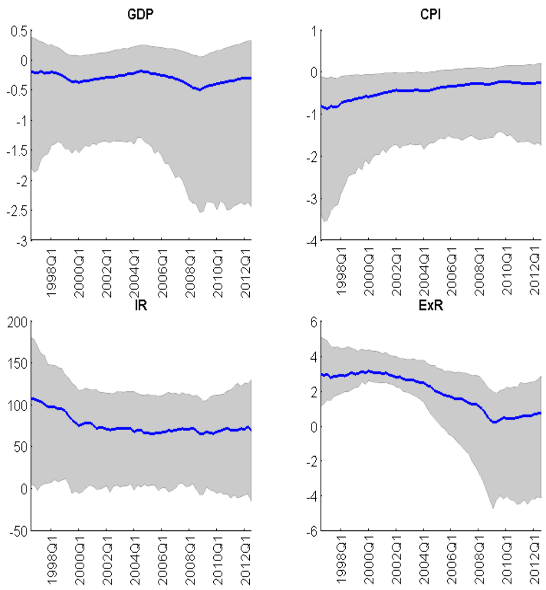
Model	k_Q	k_W	k_S	Posterior probability
1	0.0100	0.0100	0.0010	0
2	0.0500	0.0100	0.0010	0
3	0.1000	0.0100	0.0010	0.001
4	0.0100	0.0250	0.0010	0
5	0.0100	0.0250	0.0100	0.001
6	0.0100	0.1000	0.0010	0
7	0.0100	0.1000	0.0100	0.0859
8	0.0100	0.0100	0.0100	0
9	0.0500	0.0250	0.0010	0
10	0.0500	0.0250	0.0100	0.003
11	0.0500	0.0100	0.0100	0
12	0.0500	0.1000	0.0010	0
13	**0.0500**	**0.1000**	**0.0100**	**0.9091**
14	0.1000	0.0250	0.0010	0
15	0.1000	0.1000	0.0010	0
16	0.1000	0.1000	0.0100	0
17	0.1000	0.0250	0.0100	0
18	0.1000	0.0100	0.0100	0

Note: Posterior probability estimates are based on the reversible jump Markov chain Monte Carlo method for the set of 18 models. These are constructed from all possible combinations of $k_Q = \{0.01; 0.05; 0.1\}$, $k_W = \{0.001; 0.01\}$ and $k_S = \{0.01; 0.025; 0.1\}$.

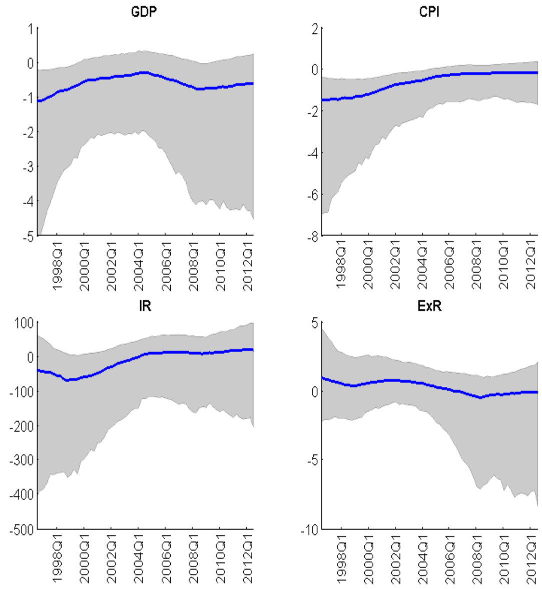
C Impulse Responses to a Monetary Policy and Exchange Rate Shock at the Fourth and Eighth Quarter with Consumer Prices

Figure C.2: Impulse Responses at Different Horizons with Consumer Prices

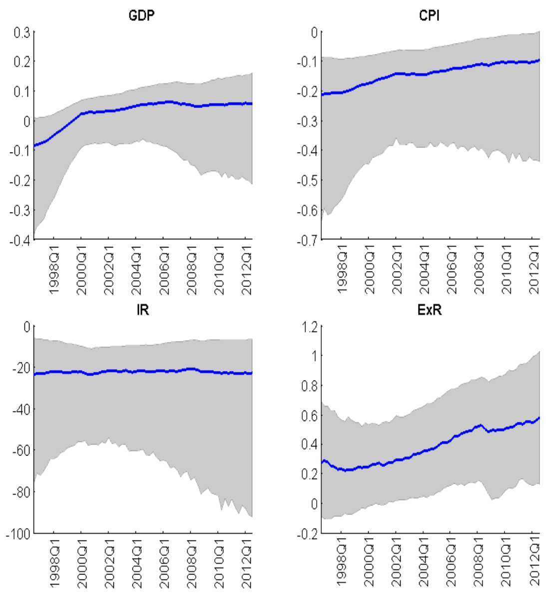
(a) Responses at the 4th quarter to a Monetary Policy Shock



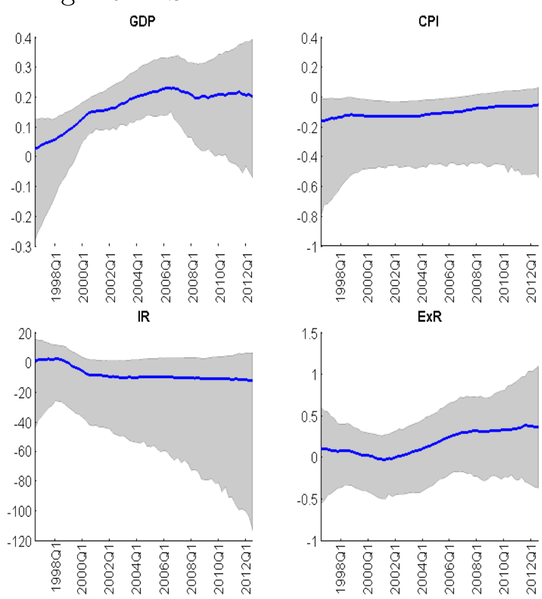
(b) Responses at the 8th quarter to a Monetary Policy Shock



(c) Responses at the 4th quarter to an Exchange Rate Shock



(d) Responses at the 8th quarter to an Exchange Rate Shock



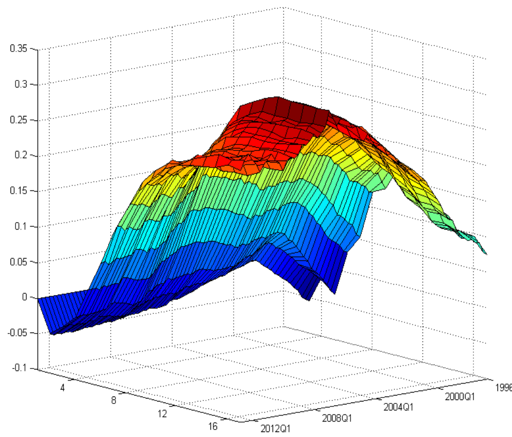
Median impulse responses (blue solid line) to a 100 BPs monetary policy shock or 1% exchange rate shock with 16-th and 84-th percentiles (grey area) of the posterior distribution of the responses at the 4th and 8th quarter, respectively.

D Impact of an Exchange Rate Shock with Import- and Producer Prices

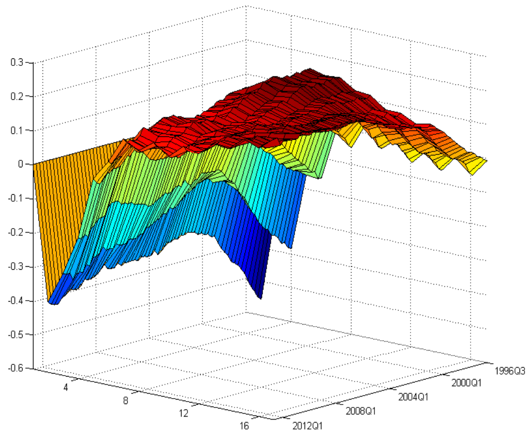
D.1 Estimation of the Exchange Rate Shock with Import Prices

Figure D.3: Time-Varying Impulse Responses to an Exchange Rate Shock with Import Prices

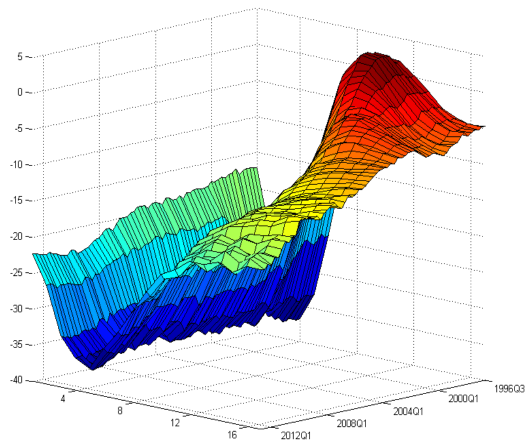
(a) GDP



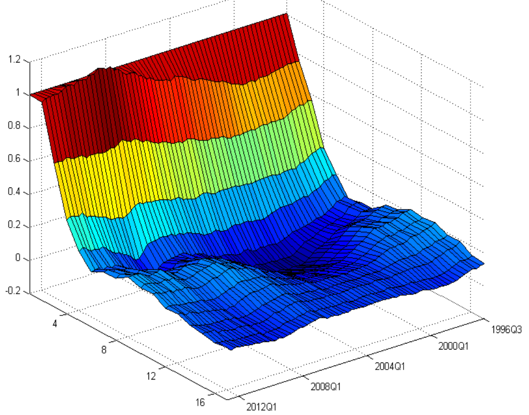
(b) Import Prices



(c) Interest Rate



(d) Exchange Rate

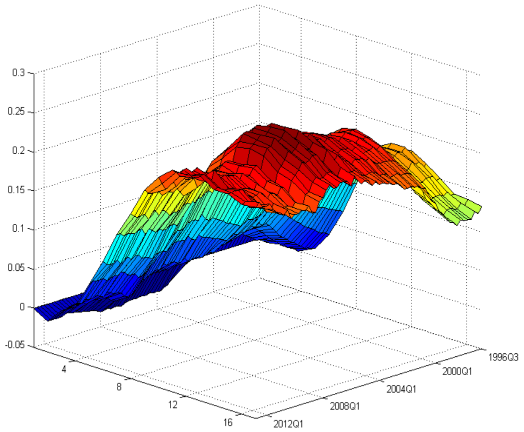


Median impulse responses to a 1% exchange rate appreciation.

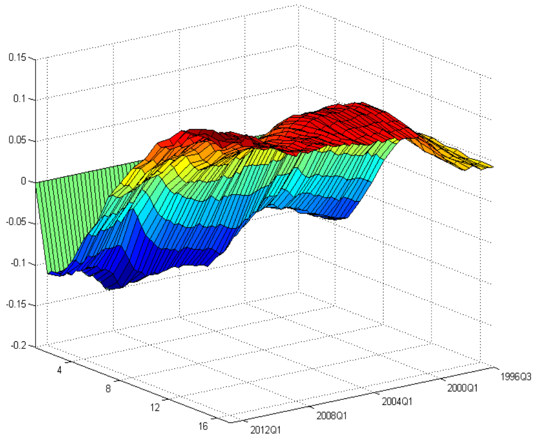
D.2 Estimation of the Exchange Rate Shock with Producer Prices

Figure D.4: Time-Varying Impulse Responses to an Exchange Rate Shock with Producer Prices

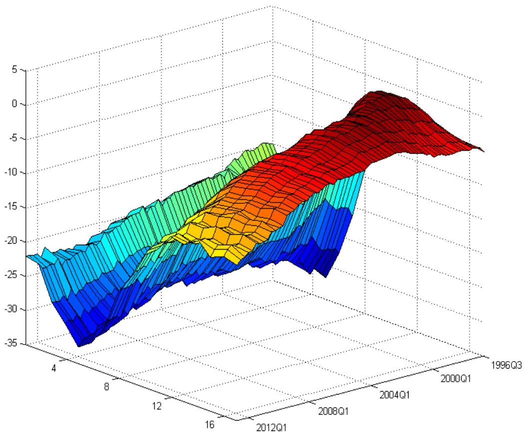
(a) GDP



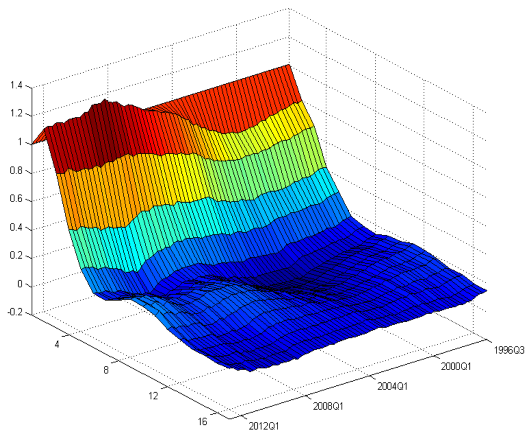
(b) Producer Prices



(c) Interest Rate



(d) Exchange Rate

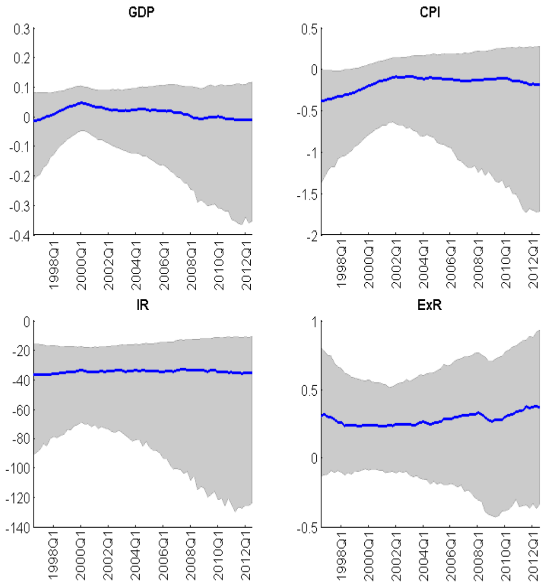


Median impulse responses to a 1% exchange rate appreciation.

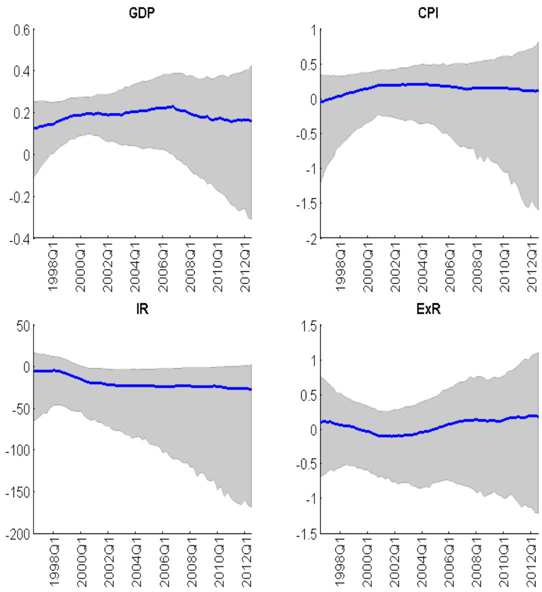
D.3 Estimation at the Fourth and Eighth Quarter to an Exchange Rate Shock with Import and Producer Prices

Figure D.5: Impulse Responses at the 4th and 8th Quarter to an Exchange Rate Shock

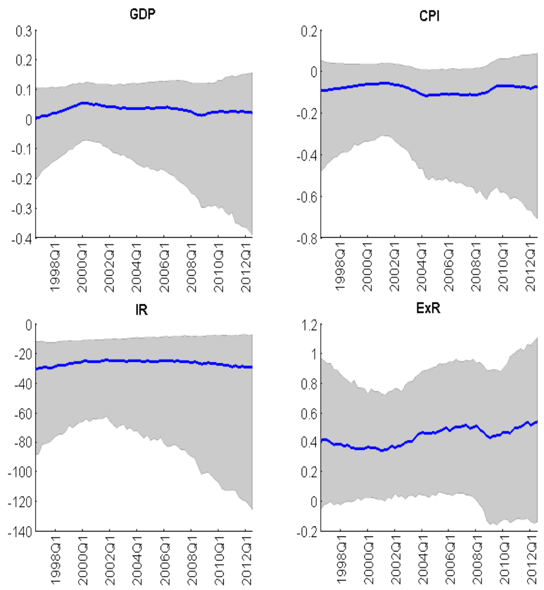
(a) Responses at the 4th quarter with Import Prices



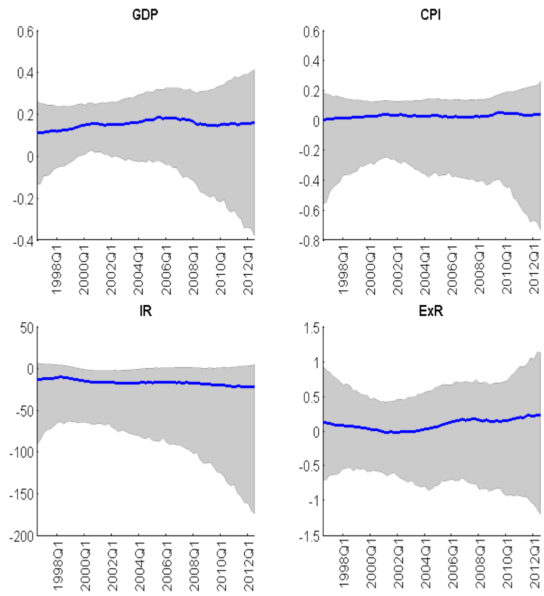
(b) Responses at the 8th quarter with Import Prices



(c) Responses at the 4th quarter with Producer Prices



(d) Responses at the 8th quarter with Producer Prices



Median impulse responses (blue solid line) to a 1% exchange rate appreciation with 16-th and 84-th percentiles (grey area) of the posterior distribution of the responses at the 4th and 8th quarter, respectively.

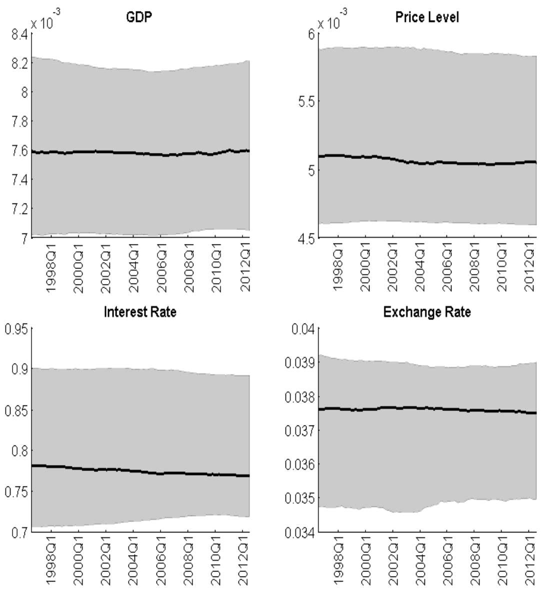
E Time-Varying Posterior Estimates of the Stochastic Covariance

The stochastic covariance matrix of the residuals comprises two matrices. First, the time-varying diagonal matrix Σ_t which denotes the stochastic volatility of the structural shock. The second matrix, the time-varying lower triangular matrix A_t captures the size of the simultaneous impact on the other variables of the variable which is shocked.

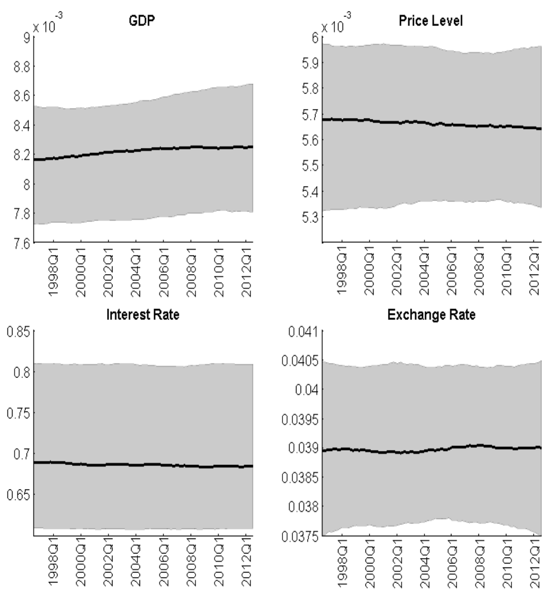
Concerning Σ_t, not much time variation is visible. Figure E.6 below shows the estimated stochastic volatility of the structural shock on GDP, prices, the interest rate and the exchange rate. It plots the posterior mean and the 16th and 84th percentile of the standard deviation of the shock. The second matrix, the time-varying simultaneous relations are plotted in Figure E.7. The simultaneous effect on the interest rate of the price shock is clearly time varying.

Figure E.6: Volatility of Structural Shock

(a) Volatility of the Structural Shock to a Monetary Policy Shock

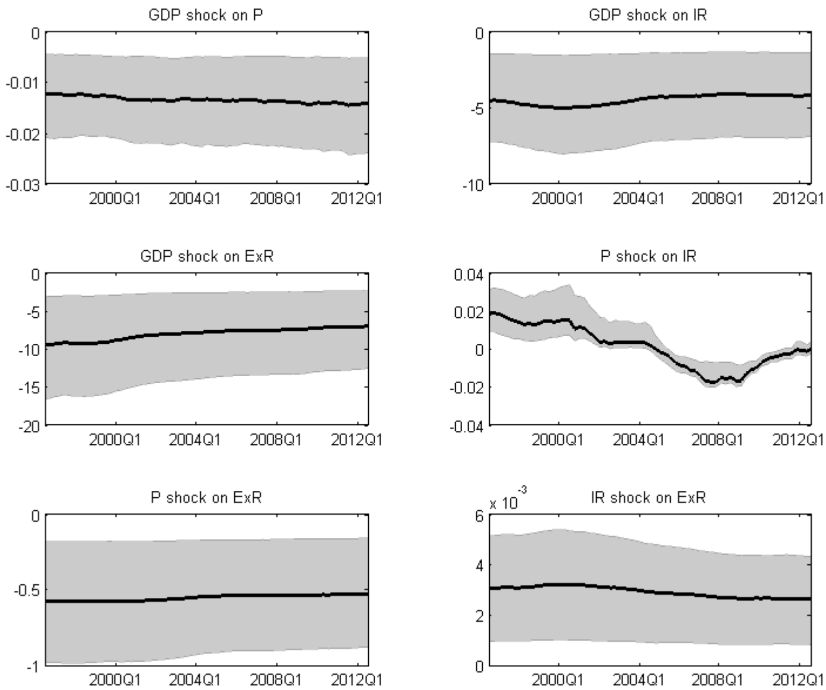


(b) Volatility of the Structural Shock to an Exchange Rate Shock



Posterior mean (solid line), 16-th and 84-th percentiles (in grey) of the standard deviation of residuals of the GDP, price, interest rate and exchange rate equation.

Figure E.7: Posterior Estimates for the Simultaneous Relation $\tilde{\alpha_{it}}$



Posterior estimates for the simultaneous relations. Posterior mean (solid line), 16-th and 84-th percentiles (in grey).

F Convergence Diagnostics

This section gives convergence diagnostics of the Markov chain Monte Carlo algorithm. We follow Primiceri (2005) to calculate the convergence diagnostics. These autocorrelations measures are based on the Econometric Toolbox illustrated by LeSage (1999). For space reasons, the convergence diagnostics are only given for estimates of the point 2012Q1.[21]

We refer to three measures of convergence diagnostics: (i) 10-th-order sample autocorrelation of the draws; (ii) inefficiency factors (IFs) for the posterior estimates of the parameters, it is an estimate of $(1 + 2\sum_{k=1}^{\infty} \rho_k)$, with ρ_k as the k-th-order autocorrelation of the chain, adequate estimates are below or above the value of 20; (iii) and the Raftery and Lewis (1992) diagnostics, calculating the necessary number of runs to obtain a certain precision (the desired precision = 0.025, necessary probability for obtaining this precision = 0.95, calculated for the 0.025 quantile of the marginal posterior distribution).

[21] Compared to other points in time, the respective estimates are very similar.

F.1 Convergence Diagnostics of the Markov Chain Monte Carlo Algorithm

The (a) panel of Figure F.8 refers on the horizontal axis throughout the points 1-36 to B (time varying coefficients), points 37-42 correspond to A (time varying simultaneous relations), and points 43-46 refer to Σ (time varying volatilities). Respectively, the hyperparameter panels (b), (c) and (d) of Figure F.8, relate throughout the points 1-1296 to Q, points 1297-1332 to S and points 1233-1348 to W.

We start with a short summary of the 10-th-order autocorrelation. It is useful to scrutinise the autocorrelation function of the draws, to evaluate how well the randomly selected chain mixes. For an efficient algorithm, the draws need to be independent from each other. This is verified by low values of the autocorrelation function (see Figure F.8a and F.8b). The autocorrelation estimates for Σ exhibit some correlation indicating inefficiency (see below for discussion).

The diagnostics concerning the inefficiency factors (IFs) calculates values very much below 20, thus suggesting efficiency. An overview is also given in Table 2.5 below. Concerning the IFs of A and B, the statistics show very low estimates. However, the IFs referring to Σ indicate some inefficiency. Considering the higher dimensionality of our problem, however, these results seem satisfactory (Kirchner et al. (2010)). Also Franta et al. (2011) illustrate that some inefficiency should be of a minor concern when the total number of runs required by the Raftery and Lewis (1992) statistics is well below the actual number used in this study. As can be seen in Figures F.8a and F.8d, the suggested number of iterations is below the actual number used. Furthermore, the impulse responses are calculated with respect to normalised shocks, hence, the inefficiency problem should not matter (Franta et al. (2011)).

To sum up, the total number of suggested iterations is far below the number used in this paper and, on average, we obtain satisfying IFs as well as autocorrelation estimates. Hence, the convergence diagnostics are sufficient.

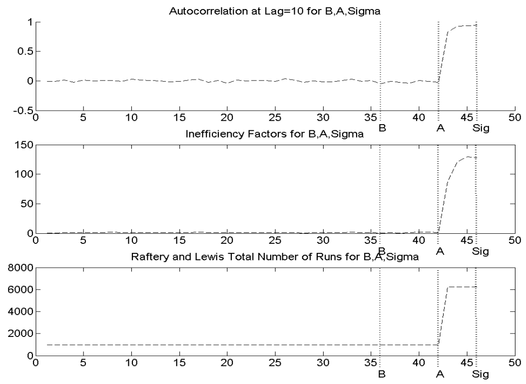
Table 2.5: Distribution of the Inefficiency Factors

	Median	Mean	Min	Max	10-th Percentile	90-th Percentile
A	1.2355	1.4926	0.7585	3.5888	0.7605	3.3658
B	1.3324	1.5356	0.5428	3.7816	0.8969	2.5113
Σ	146.1012	146.5849	145.2834	148.8536	145.2834	148.5536

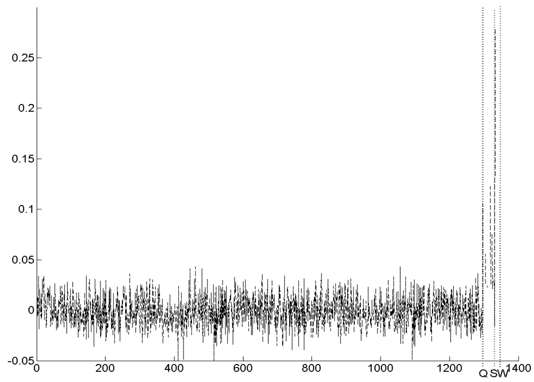
Overview of the inefficiency factors (IFs) for the posterior estimates of different sets of time varying parameters. A: time varying simultaneous relations; B: time varying coefficients; Σ: time varying volatilities.

Figure F.8: Convergence Diagnostics

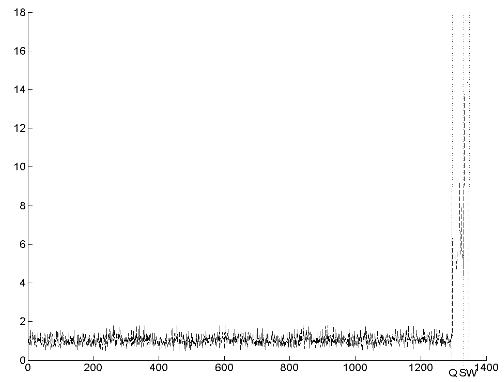
(a) Summary of Convergence Diagnostics for A, B and Σ



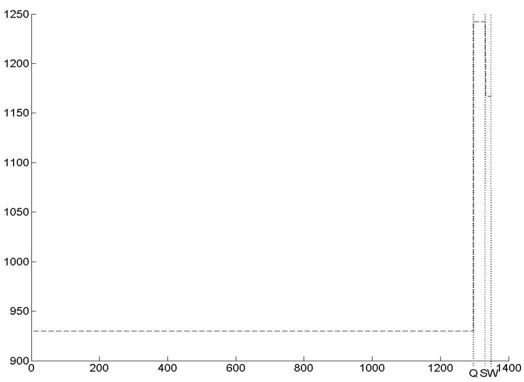
(b) 10-th-order Autocorrelation for Q, S and W



(c) IFs for Q, S and W



(d) Raftery and Lewis total number of runs for Q, S and W



Panel (a) refers on the horizontal axis throughout the points 1-36 to B (time varying coefficients), points 37-42 to A (time varying simultaneous relations), and points 43-46 to Σ (time varying volatilities). The hyperparameters in panels (b), (c) and (d) relate throughout the points 1-1296 to Q, points 1297-1332 to S and points 1233-1348 to W.

Chapter 3

Monetary Policy Transmission before and after the Financial Crisis: A nonlinear VAR model for the Euro Area

3.1 Introduction

The interest rate for main refinancing operations (MRO) is one of the most publicised and forecasted economic indicators in the Euro Area. Interest rate changes influence economic development and are of great importance for financial markets. Especially against the background of the recent Euro crisis and thus the resulting challenges for a successful conduct of monetary policy, central banks must have an accurate assessment of how its policy decisions are transmitted through the economy. This requires a sound knowledge on how monetary policy unfolds.

This paper provides new empirical evidence on the transmission mechanism of monetary policy in the Euro Area. I analyse whether there are possible differences in the transmission mechanism before and after the occurrence of the financial crisis. Furthermore, I provide new empirical evidence on the Euro Area transmission mechanisms by employing a nonlinear VAR model. In contrast to the linear VAR, the nonlinear VAR model allows to distinguish between the effects of expected and unexpected policy changes, which have a very different effect on the economy. It is especially important, for an accurate assessment of monetary policy effectiveness, to differentiate between these expectations.

Previous research mostly focused on standard VAR approaches using synthetic Euro Area data to trace the effect of monetary policy in the Euro Area (see Van Aarle et al. (2003) and Peersman and Smets (2003) among others). Furthermore, based on a VAR framework, there is surprisingly little research which investigates possible differences in the Euro Area transmission mechanism before and after the financial crisis. Since the influence of this crisis is likely to have impaired the central banks reaction function as well as changed the monetary policy transmission mechanism, the data in this analysis is divided into two samples ranging from January 1999 to September 2008 and from October 2008 to December 2014.[1] This allows to investigate whether there are possible differences in the transmission mechanism of monetary policy before and after the financial crisis in 2008.

To address these questions, I follow the empirical framework of Hamilton and Jorda (2002). The estimation is based on two steps. In the *first* step, the short-term policy rate is predicted via the autoregressive conditional hazard model (ACH) and the ordered probit model (OP). This allows to capture both when and how the central bank changes its interest rate. In the *second* step, the nonlinear VAR model, build on the ACH and OP model, studies the impact of unexpected and expected monetary policy changes on key macroeconomic variables such as output and prices. The results obtained deliver new and more thorough insights on the Euro Area monetary policy transmission, emphasising the importance of this approach. More specifically, I investigate the effects of monetary policy on industrial production, prices, money growth and the exchange rate by using a linear VAR as well as the nonlinear VAR model. My estimates show that monetary policy shocks can have different results compared to previous results based on the standard VAR approach for the Euro Area. Especially for the pre financial crisis sample, an unexpected increase in the policy interest rate leads to much larger effects on industrial production, consumer prices, money growth and the exchange rate than a standard VAR or an expected interest rate decrease which is not implemented. For both samples, the expected but not implemented change has only a small effect on e.g. industrial production or prices, as the market simply seems to expect a decrease at the next monetary

[1] Against the background of a rather short time horizon from October 2008 until January 2014, it is necessary for the VAR estimation to extend the data series in the post financial crisis sample until December 2014. This step provides a sufficient sample length for a time series investigation. For further details please refer to section 3.7.1.

policy meeting. Consequently, the nonlinear VAR model allows to distinguish and analyse the influence of these two different effects where a usual linear VAR mixes these two expectations. Compared to the pre financial crisis sample the post financial crisis impulse responses seem to be generally less significant. This could be either due to a shorter time horizon or it indicates a less effective monetary policy during the post crisis sample. One can argue that the policy interest rate is close to the Zero Lower Bound (ZLB) for some parts of that period. Under such circumstances, the central bank's ability to reduce its short-term policy rate meaningfully ended.

I contribute to the literature in two aspects. First, recent empirical evidence on the possible influence of the financial crisis as well as evidence based on aggregated Euro Area data following a monetary policy shock is rather scarce. Therefore, it is important to shed more light on these issues. I address this by investigating whether there are possible changes in the monetary transmission mechanism which may have occurred with the beginning of the global financial crisis. To this end I split the data in September 2008, the time of the outbreak of the financial turmoil. Hence, the pre financial crisis sample starts with the introduction of the Euro in January 1999 and spans until September 2008. The post financial crisis sample lasts from October 2008 until December 2014. This allows me to shed light on whether the financial crisis led to changes in the Euro Area monetary policy transmission mechanism. To do this, I am interested in the effects of a standard VAR as one can argue that this is one of the most common tools to investigate monetary policy in the Euro Area. To my knowledge, this is the first work estimating a standard VAR for the post financial crisis period on aggregated Euro Area data. Second, by following Hamilton and Jorda (2002), I allow for an important differentiation among an unexpected tightening and an expected not implemented decrease in the policy rate. As pointed out, the usual VAR response mixes the effects of these two very different expectations. My results show that taking these into account lead to very different implications of the ECB's monetary policy. Hence, I deliver new and more thorough insights on its effects on the overall economy. Moreover, in contrast to Hamilton and Jorda (2002), I implement exogenous variables into the VAR specification of the nonlinear VAR model. This controls for changes in world demand and inflation.

The structure of my paper is as follows: section 3.2 outlines previous work on interest rate modeling, forecasting and the transmission mechanism of monetary policy effects. Section 3.3 briefly introduces the ACH and OP model. The following two sections give details on the data and empirical specification. Section 3.6 presents the estimation results of the ACH and OP model. Section 3.7 refers to the transmission mechanism of monetary policy shocks. It is divided into five subsections: Subsection 3.7.1 outlines the VAR and nonlinear forecasting method, subsections 3.7.2 to 3.7.4 describe the three different policy shocks for the pre and post financial crisis sample as well as their possible differences. Subsection 3.7.5 briefly summarises the effect of monetary policy at the ZLB within the IS-LM framework and it provides a theoretical link between my results and the effectiveness of monetary policy when the economy is in the liquidity trap. Section 3.8 concludes.

3.2 Survey of Related Literature

This paper is related to the literature on interest rate setting and forecasting as well as on monetary policy transmission. Therefore, the literature overview is divided into two subsections. First, I briefly summarise related literature on econometric models of interest rate setting and forecasting. Second, related literature on monetary policy transmission is outlined.

3.2.1 Interest Rate Setting and Forecasting

Abundant research is carried out to model interest rate changes. Early work was based on the classical regression model for estimating the size of an interest rate change (compare Froyen (1975), Lombra and Torto (1977) and Smirlock and Yawitz (1985) among others).[2] Note that these approaches disregard the discreteness of interest rate changes as they are based on the assumption that the policy interest rate follows a linear and continuous process. Central banks announce whether interest rates change during their regular (often monthly) meetings, defining somewhat an upper limit on possible changes during a year. Adjustments usually occur in a series of small 25 basis point steps instead of fewer rather

[2] A further vital economic approach is the estimation via the 'Taylor rule'. Taylor (1993) illustrates that central bank behavior can be described by means of a policy reaction function. The interest rate is the policy instrument, depending on both inflation and current output gap. However, as Svensson (1999) states, the reaction function of central banks is usually not a Taylor-type reaction function, instead, it is influenced by much more information. In particular, the more open the economy is, the more potential variables have an impact. A further issue: by applying the Taylor rule, one needs information of future inflation or current output gap which is not known at the time of the central bank's decision on monetary policy (Carstensen (2006)). Using information published before the meeting is more credible and additionally serve as indicators for future inflation.

larger ones. This emphasises the importance of a discrete choice model such as the ordered probit model (OP). It captures such characteristics and served useful in other studies analysing discrete variables and asymmetrical changes. For example, Hausman et al. (1992) examine an OP model on transaction stock prices, and Dueker (1999) whether US prime rate changes are asymmetric. By means of an OP model, the Bank of England's monetary policy is studied by Eichengreen et al. (1985) and Davutyan and Parke (1995). The ECB's policy reaction function is estimated by Carstensen (2006) and Gerlach (2007) and Choi (1999) look at the Federal Reserve's reaction function.

Also Hamilton and Jorda (2002) develop a model for the federal funds rate target. Their approach is the first to consider the asymmetries in range as well as the discrete changes. More specifically, this approach accounts for two characteristics: *first*, based on today's information there is insecurity about the next interest rate change, since the intervals between interest rate changes are not equal in length. Their autoregressive conditional hazard model (ACH) captures the length between interest rate changes and thus calculates the probability of an interest rate change during the upcoming week. *Second*, based on an interest rate change, the size of this change is estimated by means of the OP model. To investigate the transmission of a monetary policy shock, a nonlinear forecasting model, based on the ACH and OP results, is developed.[3] Their contribution marks a forecast improvement and significantly enhances VAR predictions. In contrast to a standard linear VAR, it differentiates between an expected interest decrease not implemented and an unexpected increase in the policy rate. They find that these two expectations lead to very different implications of a monetary policy shock. As a result, it is important to take it for an accurate assessment of the monetary policy transmission into account (for further details, see below in section 3.2.2). I use their seminal framework to examine the ECB's monetary policy and its implications for the Euro Area.

[3]A lot of literature is concentrating on discrete event forecasts. It is based on a given information set, establishing nonlinear functions which require numerically intensive methods (see e.g. Cargnoni et al. (1997), Piazzesi (2001), Lunde and Timmermann (2004), Dueker (2005)). However, Hamilton and Jorda (2002) propose an easier calculation based on the model of Engle and Russell (1998). Compare 3.3.1 for a detailed description on developing a data generating process where the nonlinear function becomes simpler, not requiring extensive numerical calculations.

3.2.2 Monetary Policy Transmission

In monetary policy analysis, VAR models have been for some decades time a very popular tool for analysing monetary policy transmission. There is extensive literature concentrating on the US monetary transmission mechanism (see e.g. Bernanke and Blinder (1992), Sims (1992) and Evans and Marshall (1998)). Regarding the Euro Area, most of the research on monetary policy transmission is on cross-country differences.[4] In contrast, the monetary transmission on the aggregate level is not as large. In terms of a structural VAR refer to Monticelli and Tristani (1999), Van Aarle et al. (2003) and Peersman and Smets (2003).[5] Cecioni et al. (2011) use a Bayesian VAR for analysing the European monetary transmission mechanism. However, so far most of the Euro Area analyses rely on synthetic data from 1980 onwards.[6] During this time a common central bank and hence a common monetary policy was not yet established. Studies using actual Euro Area data are Weber et al. (2009), Cecioni et al. (2011) and Fahr et al. (2013). Weber et al. (2009) analyse whether the creation of the Euro Area changed the transmission mechanism by means of a standard VAR with a Cholesky decomposition. Their time horizon ranges from 1999Q1 until 2006Q4. Following a monetary policy tightening, GDP temporarily declines and prices have a delayed response and then remain permanently at a lower level. Cecioni et al. (2011) and Fahr et al. (2013) estimate a Bayesian VAR. More specifically, Fahr's et al. (2013) estimation is based on recursive identification with 12 variables from January 1999 until December 2011 at monthly frequency.[7] Cecioni et al. (2011) use an identification based on recursive identification or sign restrictions for the period from January 1999 to July 2007 or to August 2009. According to previous findings in the literature, Cecioni et al. (2011) estimate a temporary fall in output and a permanent decline in prices following a monetary tightening. The difference between the two sample periods (until 2007 or 2009) is not large. By comparing these two samples, they

[4] A structural VAR is, e.g., used by Mojon and Peersman (2001) and Peersman (2004) and more recently Barigozzi et al. (2013) apply a dynamic factor model. A detailed overview on the country specific effects as well as on aggregated Euro Area level is given by Angeloni et al. (2002).

[5] Generally these studies find a temporarily reduction in output, with the peak effect occurring after about one year. Prices respond with a delayed decrease and then stay permanently at a lower level.

[6] Studies using synthetic European data are for example Gerlach and Schnabel (2000), Peersman and Smets (2003), Smets and Wouters (2003), Andrés et al. (2006), Sousa and Zaghini (2008), Barigozzi et al. (2013) and Castelnuovo (2013). The synthetic data is generated by aggregating the national data of 11 national countries which first adopted the Euro. For further information of the aggregation method, see Fagan et al. (2005)

[7] Following a standard policy shock in Fahr et al. (2013), output is only slightly and temporarily influenced, whereas the effect on prices is insignificant.

investigate whether there are possible differences in the transmission mechanism occurring from the financial crisis in 2008. However, as generally argued in the literature, the financial crisis has led to a rethinking of monetary policy frameworks. Concerning the ECB, Gerlach and Lewis (2010) provide evidence on a policy shift after the bankruptcy of Lehman Brothers in September 2008. This highlights the importance of a sample split in September 2008. Therefore, Cecioni's et al. (2011) approach does not fully capture the influence of the financial crisis on the transmission mechanism.

Against this background, recent empirical evidence on the possible influence of the financial crisis as well as evidence based on aggregated Euro Area data following a monetary policy shock is rather scarce. I address this by analysing whether there are possible differences in the transmission mechanism due to the occurrence of the financial crisis by means of a standard VAR and the nonlinear VAR. My results show that taking these into account lead to very different implications of the ECB's monetary policy. Therefore, this paper gives new and more thorough insights on monetary policy effects in the Euro Area.

3.3 The Model

To model the interest rate setting of the ECB, I follow Hamilton and Jorda (2002). In what follows I briefly summarise in section 3.3.1 the ACH model and in section 3.3.2 the OP model.

3.3.1 The Autoregressive Conditional Hazard Model

Based on today's information, there is insecurity about the next interest rate change, since the intervals between these changes are not equal in length. The autoregressive conditional hazard model (ACH) captures the duration and thus the dynamics of the intervals by calculating the probability of an interest rate change during the next period. For estimating this probability of an interest rate change during the next period, the autoregressive conditional duration model (ACD) of Engle and Russell (1998) is extended to the ACH model.[8] The ACD model is designed to explain the dynamics of events

[8] Engle and Russell (1998) construct a new statistical model, able to deal with data arriving at asymmetrical intervals. Since the model evaluates the expected time between events - the average interval

occurring as well as the average interval of time between events.

The duration between events is denoted by u_n. It summarises the length between the nth and the $(n+1)$th point in time the central bank changed the interest rate. The expectation of u_n conditional on past durations $u_{n-1}, u_{n-2}, \ldots$ is described by ψ_n. The ACD(r,m) is then written as:[9]

$$\psi_n = E(u_n|u_{n-1}, u_{n-2}, \ldots) = \sum_{j=1}^{m} \alpha_j u_{n-j} + \sum_{j=1}^{r} \beta_j \psi_{n-j}. \tag{3.1}$$

With the help of recursion one can write $u_{1-j} = \bar{u}$, where $\bar{u}$ describes the average length between interest rate changes; and $\psi_{1-j} = \bar{\psi}$, with $\bar{\psi}$ denoting the average expected duration between events which can be written as:

$$\bar{\psi} = \frac{\sum_{j=1}^{m} \alpha_j \bar{u}}{1 - \sum_{j=1}^{r} \beta_j}. \tag{3.2}$$

Every event 'interest rate change' leads to a new value in equation (3.1). In case of no event, ψ_n stays the same. To explain this, let $N(t)$ describe the cumulative amount of changes in the interest rate at month t. Hence, $N(t)$ is a counting process since the events 'interest rate change' generate a discrete time series. In case of an interest rate change in the interval $(0, t]$, $N(t)$ counts further. For example, $N(0) = 0$, $N(t) = N(t-1)$ if no change occurs in the interval $(t-1, t]$, and $N(t) = N(t-1) + 1$ if the interest rate is changed at time t.[10] Equation 3.1 can then be rewritten in terms of the counting process and thus in calender time with $\psi_{N(t)}$ for expected duration or respectively $u_{N(t)}$ for the

- it is referred to as the autoregressive conditional duration model (ACD). The duration between events is viewed as a random process and the approach introduces a new class of point processes with dependent arrival rates.

[9]Equation (3.1) is reminiscent of ARMA models, where ψ_n - the expected duration - is a function of past durations of lag order r and expected durations of lag order m. Equation (3.1) can therefore be expressed as an ARMA(max$\{m,\ r\}$, r) illustration. $(\alpha_j + \beta_j)$ depict the jth autoregressive coefficient. To obtain stationarity, $\sum_{j=1}^{m} \alpha_j + \sum_{j=1}^{r} \beta_j < 1$ has to hold.

[10]For illustrative purpose, assume that the first change occurs in month three and the second in month seven. Then $N(t)$ has the following structure:

$$N(t) = \begin{cases} 0 & \text{for } t \ = 1, 2 \\ 1 & \text{for } t \ = 3, 4, 5, 6 \\ 2 & \text{for } t \ = 7, 8, \ldots \end{cases}$$

actual duration between events:

$$\psi_{Nt(t)} = \sum_{j=1}^{m} \alpha_j u_{N(t)-j} + \sum_{j=1}^{r} \beta \psi_{N(t)-j}, \tag{3.3}$$

This also allows to incorporate updated information, which is crucial for forecasting the occurrence of the next interest rate change.

The hazard rate h_t denotes the conditional probability of an interest rate change based on the information set Υ_{t-1} known at time t - 1. The hazard rate can be written as:

$$h_t = \Pr[N(t) \neq N(t-1)|\Upsilon_{t-1}]. \tag{3.4}$$

Note that the hazard rate can also be written in terms of expected duration $\psi_{N(t-1)}$ which is specified in equation (3.3):

$$h_t = \frac{1}{\lambda[\psi_{N(t-1)} + \delta' z_{t-1}]}. \tag{3.5}$$

The term z_{t-1} characterises the vector of explanatory variables. It describes the information set known at time $t-1$, and δ denotes its parameter vector.[11] The function λ assures a proper estimation of the hazard rate. More specifically, it may occur that the denominator takes on a value that is too small and hence h_t may lie outside of $(0, 1)$. In this case, h_t is specified to a constant just below unity. It therefore provides a smooth transfer and ensures $0 < h < 1$.[12]

It is now straight forward to estimate the log likelihood function. If the ECB changes its interest rate during month t, then $x_t = 1$ and zero otherwise. As denoted in equation (3.4), the likelihood function is the probability of observing x_t conditional on the information set Υ_{t-1}:

$$g(x_t|\Upsilon_{t-1}; \theta_1) = (h_t)^{x_t}(1-h_t)^{1-x_t}, \tag{3.6}$$

[11] The ACH model comprises the ACD representation as a special case. See Hamilton and Jorda (2002) for further details.

[12]

$$\lambda(v) = \begin{cases} 1.0001 & v \leq 1 \\ 1.0001 + \frac{2\Delta_0(v-1)^2}{\Delta_0^2+(v-1)^2} & 1 < v \leq 1+\Delta_0 \\ 0.0001 + v & v \geq 1+\Delta_0, \end{cases}$$

with $\Delta_0 = 0.1$.

with $\theta_1 = (\delta', \alpha', \beta')'$. Computing the log of equation (3.6) generates the conditional log likelihood function:

$$L_1(\theta_1) = \sum_{t=1}^{T} [x_t log(h_t) + (1 - x_t) log(1 - h_t)], \tag{3.7}$$

which is maximised with respect to the unknown parameters (θ_1). Additionally, it is necessary for the numerical optimisation algorithms to restrict $\alpha_j \geq 0, \beta_j \geq 0$, and $0 \leq \beta_1 + \ldots + \beta_r \leq 1$.

As mentioned, I assess different combinations of explanatory variables. For evaluating their respective performance, the Akaike and Schwarz information criteria are used (for further details, see section 3.3.2).

It is important to clarify the underlying frequency for result interpretation. The time units affect the magnitude of ψ, the expected duration. For example, if the chosen time interval is very short, the probability of a change during an interval becomes small. The probability rises in case of a longer interval. But, one has to ensure that no more than one event occurs during each interval; this would lead to an invalid estimation of equation (3.5). I take a monthly frequency for two reasons: (1) the Governing Council of the ECB meets every four weeks to decide upon their monetary policy and (2) for the Euro Area most data is published on a monthly frequency. Using weekly data is therefore meaningless since there is not enough updated information in between that may contribute to a change.[13] Regarding the ECB, using a monthly model is hence a proper assumption.[14]

3.3.2 The Ordered Probit Model

So far the focus was whether the ECB intends to alter the interest rate on MROs next month or if it is to stay constant. As a next step, I examine the size of the interest rate change conditional on a change. This is done with the help of the ordered probit model (OP).[15]

[13] In contrast, Hamilton and Jorda (2002) use a weekly frequency for their US-estimation as they have enough data on a weekly frequency available.

[14] Note, this ensure that the hazard rate h is specified between ($0 < h < 1$).

[15] The literature characterises such a time series as a marked point process, whereas the point process describes the probability of an interest rate change at time t. 'Marks' refer to the magnitude by which

Since the policy interest rate changes in discrete fixed amounts, usually in multiples of 25 basis points, there are only k different discrete amounts by which the ECB changes the policy rate (see section 3.5 for the specification of the exact number of states k). The OP model is designed to model discrete changes, therefore it is a suitable approach to estimate the magnitude of a change given the information set Υ_{t-1} and conditional on a change.

I assume that there is an unobserved variable y_t^* which is related to the observed interest rate changes y_t. y_t^* depends on a vector of exogenous variables w_{t-1}, characterising the available information at period $t-1$ such that

$$y_t^* = \mathrm{w}'_{t-1}\pi + \epsilon_t, \tag{3.8}$$

where $\epsilon_t|\mathrm{w}_{t-1}$ i.i.d $N(0,1)$. The k possible interest rate changes by which the ECB implements a change can be denoted by $s_1 < s_2 < \ldots < s_k$. The relationship between the unobserved continuous variable y_t^* and the observed interest changes y_t can then be written as:

$$y_t = \begin{cases} s_1 & \text{if } y_t^* \in (-\infty, c_1] \\ s_2 & \text{if } y_t^* \in (c_1, c_2] \\ \vdots & \\ s_k & \text{if } y_t^* \in (c_{k-1}, \infty), \end{cases} \tag{3.9}$$

where the partition boundaries c have the structure $c_1 < c_2 < \ldots < c_k$ and $c_0 = -\infty, c_k = \infty$. They define the interval boundaries of y_t^*. The conditional probability of an interest rate change by the amount s_j in period t is given by:

$$\Pr(y_t = s_j|\boldsymbol{w}_{t-1}, x_t = 1) = \Pr(c_{j-1} < \boldsymbol{w}'_{t-1}\pi + \epsilon \leq c_j),$$

where $j = 1, 2, \ldots, k$. It is conditioned on a change ($x_t = 1$) as well as on w'_{t-1}.[16] This probability is described by means of the standard normal cumulative distribution function $\Phi(z)$. It is then possible to rewrite these conditional probabilities. Note that $x_t = 1$ in

the ECB decreases or increases the interest rate.

[16] These probabilities are mainly based on the partition boundaries c and the particular distribution of ϵ. As assumed in equation (3.8), the error term is normally distributed.

case of a change and zero otherwise:

$$\Pr(y_t = s_j | \boldsymbol{w}_{t-1}, x_t = 1) =$$

$$\begin{cases} \Phi(c_1 - \mathrm{w}'_{t-1}\pi) & \text{for } j = 1 \\ \Phi(c_j - \mathrm{w}'_{t-1}\pi) - \Phi(c_{j-1} - \mathrm{w}'_{t-1}\pi) & \text{for } j = 2, 3, \ldots, k-1 \\ 1 - \Phi(c_{k-1} - \mathrm{w}'_{t-1}\pi) & \text{for } j = k. \end{cases} \tag{3.10}$$

The exogenous variables, w_{t-1}, influence the size of an interest rate change. More specifically, a 'higher' interval becomes more likely the larger the value of $\mathrm{w}_{t-1}\pi$ (vector of exogenous variables multiplied by the respective coefficient). For example, a rise in the exogenous variable GDP induces the ECB to increase its interest rate, leading to a positive OP-coefficient of this exogenous variable.

The maximum likelihood estimation calculates the parameters of the OP model. The log of equation (3.10) is taken and $l(y_t|\mathrm{w}_{t-1})$ denotes the log of the probability of observing y_t given $x_t = 1$ and w_{t-1}:

$$l(y_t|\mathrm{w}_{t-1}; \theta_2) =$$

$$\begin{cases} \log\left[\Phi(c_1 - \mathrm{w}'_{t-1}\pi)\right] & \text{if } y_t = s_1 \\ \log\left[\Phi(c_j - \mathrm{w}'_{t-1}\pi) - \Phi(c_{j-1} - \mathrm{w}'_{t-1}\pi\right] & \text{if } y_t = s_2, \ldots, s_{k-1} \\ \log\left[1 - \Phi(c_{k-1} - \mathrm{w}'_{t-1}\pi)\right] & \text{if } y_t = s_k, \end{cases} \tag{3.11}$$

where $\theta_2 = (\pi', c_1, c_2, \ldots, c_{k-1})'$. The conditional log likelihood of the OP model is then given by:

$$L_2(\theta_2) = \sum_{t=1}^{T} x_t l(y_t|\mathrm{w}_{t-1}; \theta_2), \tag{3.12}$$

which is maximised with respect to θ_2 and $c_j > c_{j-1}$ for $j = 1, 2, \ldots, k-1$.

The following equation gives the joint log likelihood of the ACH and OP model. More specifically, it describes when and how the interest rate is changed by the ECB. The unknown parameters, θ_1 and θ_2, maximise this expression:[17]

$$L = \sum_{t=1}^{T} \log f(x_t, y_t|\Upsilon_{t-1}; \theta_1, \theta_2) = L_1(\theta_1) + L_2(\theta_2). \tag{3.13}$$

The respective additive terms are given in equation (3.7) and (3.12).[18]

As mentioned before, by means of the Akaike and Schwarz information criteria, I evaluate the performance of different combinations of explanatory variables. These are standard tools in the literature for assessing the performance of different explanatory variable combinations. I employ these criteria to find a suitable model specification which best characterises the setting of the interest rate on MROs. More specifically, I search for the combination of explanatory variables which yields the best model for explaining the interest rate setting. These criteria provide a valuation of finite samples and comprise the trade-off between including more explanatory variables and the overall goodness of fit.[19]

3.4 Data

Next to the editorials of the ECB's *Monthly Bulletin*, also Hamilton and Jorda (2002) and Gerlach (2007) serve as a guideline for variable selection. As discussed in section 3.3.1, I use monthly data. Earlier research papers, such as Gerlach (2007) and Chevapatrakul et al. (2007), which focus on analysing monetary policy strategies, also use monthly data.

According to the ECB's two pillar strategy, I broadly classify the examined variables into two groups. The first group of explanatory variables relates to macroeconomic developments (second pillar of the ECB) while the second group includes variables on financial

[17]The unknown parameters in the ACH model comprise the explanatory variables as well as α. Regarding the OP model, I estimate in addition to the explanatory variables the partition boundaries (see section 3.5 for a detailed description).

[18]Like Engle (2003) illustrates, if θ_1 and θ_2 have no parameters in common it is the same as if equation (3.13) is maximised or if one calculates a separate maximisation of equation (3.7) and (3.12) respectively. Even if θ_1 and θ_2 have common parameters, consistent but inefficient estimates are obtained by calculating the joint likelihood function in equation (3.13) separately.

[19]The general objective is to minimise them and thus find a specification which describes the underlying data the best. The Akaike and Schwarz information criteria are defined as follows:

$$\begin{aligned} AIC &= -2\tfrac{L}{T} + 2\tfrac{k}{T} \\ SBC &= -2\tfrac{L}{T} + log(T)\tfrac{k}{T}, \end{aligned} \tag{3.14}$$

where k denotes the number of estimated parameters, L depicts the maximised log likelihood value and T refers to the number of observations.

and monetary aggregates (first and second pillar of the ECB).

Variables belonging to the first group are inflation-, output- and employment measures. A detailed overview on the considered variables is given in Table 3.1. I use a *core inflation* indicator[20], a 12-month expected *inflation forecast* to account for the forward looking behavior of the ECB, a *GDP deflater* and a *unit labour cost* index to examine to which extent labour market earnings put pressure on prices.

Regarding the output measures, I use *industrial production*, *output-gap* and the *Euro Stoxx* (50).[21] Further, the *consumer* and *economic sentiment* indicators are used since the ECB discusses them in their monthly editorials. It is sensible for them to use these indices as a proxy for the actual state of the economy.[22]

To take the developments on labor markets into account, I refer to a *unemployment rate forecast*.

Next to these inflation and output measure, the ECB also focuses on money and credit growth indicators. I use *M3* as money growth is an important measure for the Council to decide upon possible monetary policy changes. It emphasises inflationary pressure arising from monetary sources. For the Council this variable is the single most important indicator of monetary developments. I also consider the *previous change* of the interest rate on MROs.[23] The 6-month spread between the Bubill (non-interest-bearing treasury bill) and the interbank Eonia rate (overnight interbank interest rate) is also taken into account. Expectations in short-term market rates of a change in the official interest rate can hence be captured in changes of the *Bubill spread* (compare Goodhart (1996), Sack (2000) and Rudebusch (2002)). I also investigate the Council's reaction towards changes in the *real effective exchange rate*.[24]

One can divide the suggested explanatory variables into two information sets regarding

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[20]It seems suitable to apply the common Harmonised Index of Consumer Prices (HICP) as an inflation measure for the Euro Area. However, world inflation rates were to some extent influenced by international price level shocks. These arose in particular in the energy sector during the last decades. But, these shocks may be viewed as temporary effects on prices and can thus can be ignored by central banks. Therefore, the *core inflation* measure provides an appropriate indicator.

[21]The output gap variable is considered e.g. by Chevapatrakul et al. (2007) and Svensson (1999).

[22]Note, these time series are non-reversible, that is they are not due to revisions. Compare Gerlach (2007) for further details.

[23]An overview of the dates and magnitude of the changes in the interest rate on MROs is given in Appendix A Table 3.6.

[24]This indicator is defined such that a negative growth rate implies a depreciation of the Euro vis-à-vis the rest of the world.

their ability to predict changes in the policy interest rate. The narrower information set incorporates measures such as future expected values of inflation, changes in the Bubill spread and M3 growth, whereas the wider information set contains variables ahead of this. Nevertheless, it is crucial to analyse the influence of the latter category as in practice pressure to raise interest rates arise from various sources.[25] Basically one can expect that variables belonging to the wider information set reveal a lower significance.

In general, final revised figures are used because forecast-series of past decades are not available anymore. This is due to the fact that they are regularly updated and revised. Most of the considered explanatory variables are lagged by one month, unless otherwise stated, accounting for the respective availability of information. For example, inflation for the month of January is not released before the middle of February.[26]

[25]Such sources encompass money, financial and labor markets, demand and output, prices as well as monetary policy (compare Chevapatrakul et al. (2007)).

[26]For simplification, the value of January is still used even if the ECB meets at the beginning of February.

Table 3.1: Candidate Explanatory Variables in the Specification of the ACH and OP Model

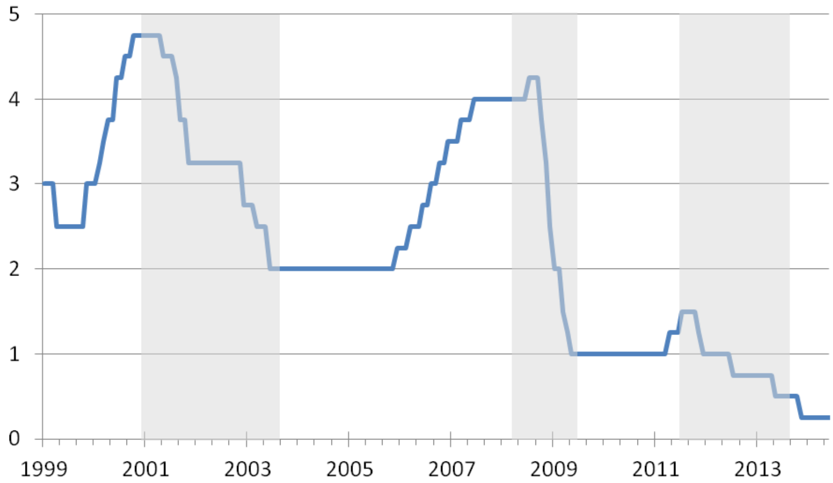
Variable	Transformation	Source	Sign ACH*	Sign OP**
Inflation Measures				
GDP Deflater, EA17	Yearly average of yoy change (%), 3 lags	Eurostat	[-]	[+]
HCPI Index, less food and energy, EA	Yearly average of yoy change (%)	Eurostat	[-]	[+]
Unit Labour Cost Index, Whole Economy, EA17	Quarterly change (%), 3 lags	ECB	[-]	[+]
12-month-ahead CPI Forecast, EA	Yearly average of annual log change (%)	ECB	[-]	[+]
Output measures				
Industrial Production Index (total), EA17	Yearly average of yoy change (%)	ECB	[-]	[+]
Output-gap (in % of potential GDP), EA		IMF	[-]	[+]
Economic Sentiment, EMU	3-month average of yoy change (%)	EuCom	[-]	[+]
Consumer Sentiment, EMU	3-month average of yoy change (%)	EuCom	[-]	[+]
Euro Stoxx (50), EA	3-month average of monthly change (%)	Stoxx	[-]	[+]
Employment Measure				
Unemployment Rate Forecast, EA	6-month-ahead average of yoy change (%)	EuCom	[-]	[-]
Monetary Variables				
Bubill Spread	Spread between Bubill and Eonia	Bundesbank	[-]	[+]
Real Effective Exchange Rate (Basket of 42 Currencies)	12-month average, yoy change (%)	ECB	[-]	[-]
M3, EA	3-month average, yoy change(%)	ECB	[-]	[+]
MRO, EA	Change	ECB		[+]

*, ** These two columns show the expected signs.
* For ACH estimation: a [-] sign implies that this variable increases the probability of an interest rate change and vice versa for [+].
** For OP estimation: a [-] sign implies that this variable increases the probability of an interest rate decrease and vice versa for [+].

3.5 The ECB's MRO's Interest Rate

One focus of my paper is to characterise the setting of the ECB's main policy interest rate, the main refinancing operations interest rate (MROs). This interest rate provides the bulk of liquidity to the banking system. It is therefore useful to look at its development across time. The considered time period of this study ranges from January 1999, when the ECB started to undertake the common monetary policy for the Euro Area, up to December 2014.[27] From January 1999 until June 2000 the ECB conducted its main refinancing

Figure 3.1: Development of the Policy Interest Rate



Interest Rate on MROs: Fixed Rate/ Minimum Bid Rate in percent from January 1999 through December 2014 (monthly data). The forecast for the policy rate from January 2014 onwards is obtained from Feri Euro Rating GmbH. Grey areas mark official recession periods in the Euro Area (source: OECD recession indicator for the Euro Area from the peak through the period preceding the trough).

operations on the basis of a fixed-rate tender procedure. During the period from June 2000 until October 2008 a variable rate tender system, also called minimum bid rate, was used. The minimum bid rate refers to the minimum interest rate at which counterparties may place their bids. In October 2008, the ECB reinstated to the fixed-rate tender procedure.

In order to test whether there are significant differences in the monetary policy transmission before and after the financial crisis, the analysis of this paper is divided into two samples ranging from January 1999 until September 2008 and from October 2008 until December 2014. During the pre financial crisis sample the interest rate changed 38 times

[27] To have a sufficient length for estimating the second sample, an extension of the data set from February 2014 until December 2014 becomes necessary. See section 3.7.1 for further details.

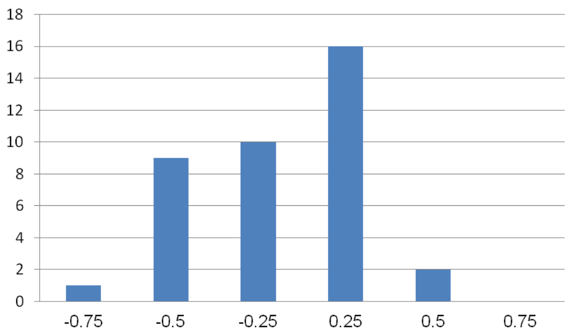
and during the post financial crisis sample 14 changes were implemented (Figure 3.2).

From the development of the interest rate on MROs, one can see that its pattern seems to be related to the development of economic activity (Figure 3.1). Interest rates are adjusted upwards during periods of high economic growth and vice versa during a time of recession.[28] This pattern is in particular apparent for the latest financial crisis which started in mid 2008 and induced the central bank to quickly lower their key interest rates.

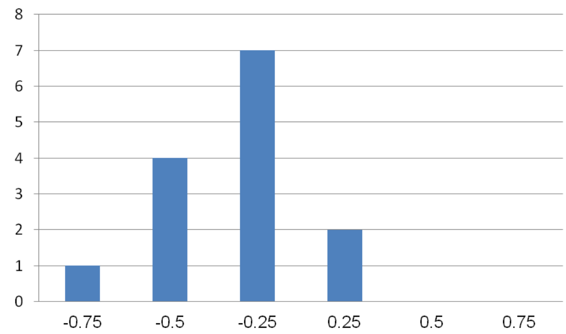
The step-wise pattern of the interest rate changes arises because it is set in multiples of 25 basis points. The ECB seems to follow a policy based on small interest rate changes. If a change occurred, it was mostly a 25 basis point change (Figure 3.2). As explained

Figure 3.2: Magnitude of Interest Rate Changes on MROs (in %)

Pre Financial Crisis Sample: Jan 1999 - Sep 2008



Post Financial Crisis Sample: Oct 2008 - Dec 2014



above, I need to specify the number of states k which are used in the estimation of the OP model. k denotes the different discrete amounts by which the ECB may change the interest rate. Figure 3.2 shows that they did not alter the interest rate by more than 75 basis points in absolute value. Based on this information, I define five states for the pre crisis sample and four states for the post crisis sample, as it is important to guarantee that each state consists of at least one observation.[29] Hence, the possible interest changes s_k are defined as follows:

Pre crisis sample:	$s_1 = -0.5$ and less; $s_2 = -0.25$; $s_3 = 0$; $s_4 = 0.25$; $s_5 = 0.5$ and more
Post crisis sample:	$s_1 = -0.5$ and less; $s_2 = -0.25$; $s_3 = 0$; $s_4 = 0.25$ and more

[28]For example, inflationary pressure might be high during boom periods, whereas a slump is in general characterised by lower inflation rates. Since the crucial aim of the ECB is to achieve and maintain price stability, they accordingly adjust their main monetary policy steering tool, the interest rate on MROs.

[29]Therefore, four or three interval boundaries are respectively estimated for the pre financial crisis sample or the post crisis sample.

3.6 Empirical Results

In the subsequent two sections the estimates of the ACH and OP model are outlined.

3.6.1 Estimates of the ACH Model

The objective of the ACH model is to calculate the probability of the ECB changing the interest rate on MROs during any given month. Since the decisions are presumably influenced by the development and outlook of economic conditions, I incorporate a set of exogenous variables z_{t-1} in the estimation of the hazard rate (compare equation 3.5).

Besides the exogenous variables, the specifications additionally comprise a constant c and the variable $\tau_{N(t-1)-1}$ denotes the duration in months between the last interest rate changes as of month $t-1$.[30]

I expect negative coefficients of the explanatory variables since they are specified to be in the denominator of the hazard rate (compare equation (3.5)). This implicates that negative and significant coefficients contribute to an increase in the probability of an interest rate change in period t.

For the specific task of forecasting whether the ECB changes its interest rate on MROs during the pre financial crisis sample, the absolute value of the spread between the six-month Bubill rate and the Eonia rate and the inflation forecast prove to be very useful (Table 3.2). According to the ECB statements, the inflation forecast delivers a much better performance than other macro indicators. For forecasting an interest rate change during the post crisis sample, a specification without the inflation forecast but with the Bubill spread yields the best performance. This finding may indicate that the aggregated inflation development seems to be of minor importance for the interest rate setting during the post crisis sample. Generally, one can argue that the Euro Area is close to a liquidity trap during some parts of this period.[31] Under such circumstances interest rates are close to zero since inflation rates are very low (refer to section 3.7.5 for a brief discussion of monetary policy at the ZLB as well as for a possible linkage of my results to monetary

[30] I exclude the expected lagged duration ($\psi_{N(t-1)-1}$) as of month $t-1$ from equation 3.5 because the underlying data did not provide the necessary dynamic for receiving feasible estimates. Possible reasons: the available time horizon and the monthly frequency do not yet provide the sufficient dynamic.

[31] In the middle of 2009 the interest rate has been decreased to 1% and since August 2013, the interest rate on MROs is at 0.5% or below.

Table 3.2: ACH Parameter Estimates for the Pre Financial Crisis Sample (Jan 1999 - Sep 2008)

Parameter	Variable	Estimate	Standard Error
α	$u_{N_{(t-1)-1}}$	0.2548*	0.1306
δ_1	Constant	38.1140***	13.4075
δ_2	$\lvert InflF12_{t-1}\rvert$	-16.5658***	5.8118
δ_3	$\lvert SP_{t-1}\rvert$	-13.8801***	4.7037

Note: Dependent variable: probability of an interest rate change; number of observations: 117. Log likelihood is -50.3768. Variable definitions: $|InflF12_{t-1}|$ is the absolute value of the 12-month ahead average of the inflation forecast at time $t-1$. $|SP_{t-1}|$ is the absolute value of the spread between the six-month Bubill rate and the Eonia rate (overnight interbank interest rate).

policy in a liquidity trap). Table 3.3 summarises the estimates on the post financial crisis sample.

The hazard rate of the pre financial crisis sample reveals about a one in seven chance of an interest rate change during the next month ($h = 0.1306$).[32] The hazard rate of the post financial crisis sample is 0.123.

Regarding explanatory variables on inflation, output, money growth and the exchange rate, I receive, as expected, negative and also significant coefficients for the pre financial crisis sample. For the post financial crisis sample, I still obtain negative coefficients but the significance is somewhat less pronounced. This finding may also mirror the fact that the Euro Area is close to or at the ZLB during some parts of the post crisis sample (see below in section 3.7.5 for a more detailed discussion).

As a robustness check, the post crisis sample is also estimated on data from October 2008 until January 2014. This excludes the forecast horizon from February 2014 until December 2014. The estimated coefficients and their significance confirm those presented.

[32]The hazard rate during the pre financial crisis sample is calculated as follows:

$$\frac{1}{(0.2548)(4.1538) + 38.114 + (-16.5658)(1.7447) + (-13.8801)(0.1883)} = \frac{1}{7.6564} = 0.1306. \qquad (3.15)$$

with the two explanatory variables on the absolute value of the inflation forecast and the spread. The average absolute spread is 0.1883, the average inflation forecast equals 1.7447 and the mean duration $\bar{u}$ is 4.1538. The coefficient estimates are given in Table 3.2.

Table 3.3: ACH Parameter Estimates for the Post Financial Crisis Sample (Oct 2008 - Dec 2014)

Parameter	Variable	Estimate	Standard Error
α	$u_{N_{(t-1)-1}}$	0.9671**	0.4487
δ_1	Constant	5.4495**	2.1641
δ_2	$\lvert SP_{t-1}\rvert$	-8.4437***	3.1430

Note: Dependent variable: probability of an interest rate change; number of observations: 75. Log likelihood is -28.408. Variable definitions: $|SP_{t-1}|$ is the absolute value of the spread between the six-month Bubill rate and the Eonia rate (overnight interbank interest rate).

3.6.2 Empirical Estimates of the OP Model

Next, I summarise the results of the OP model, which estimates the size of a change given an interest rate change occurs.

The crucial point of the OP model is to determine the possible categories by which the Governing Council may change the interest rate on MROs. Four partition boundaries for the pre financial crisis sample and three for the post crisis sample are estimated (see section 3.5).

For both samples, the OP results reveal the expected signs and significance on the used explanatory variables. More specifically, measures on money growth, economic outcome (industrial production, GDP, economic sentiment, Euro Stoxx) have the expected positive sign and are significant. This implicates that a stronger performance of such variables increases the probability of an increase in the interest rate on MROs. Furthermore, estimated coefficients on the unemployment rate, its forecast as well as the real effective exchange rate have negative coefficients and are less significant. This is in line with economic theory. More specifically, these results suggest a higher probability of a decrease in the ECB's policy interest rate following an appreciation of the Euro and an increase in the unemployment rate. However, the explanatory power of the employment indicators and the exchange is not very strong. This mirrors that the development of these variables generally have a less important impact on the ECB's decision of an interest rate change.

For the specific task of estimating the size of an interest rate change during the pre and post financial crisis sample, the previous change in the interest rate, denoted by $y_{t_{N(t-1)}}$, serves very useful. The positive value reveals that a previous rise in the policy rate increases the probability of an increase within the current month. An according strong

Table 3.4: OP Parameter Estimates for the Pre Financial Crisis Sample (Jan 1999 - Sep 2008)

Parameter	Variable	Estimate	Standard Error
π_1	$y_{t_{N(t-1)}}$	0.6560*	0.4078
π_2	SP_{t-1}	3.4959***	0.6256
c_1		-2.6232***	0.3286
c_2		-2.2845***	0.2932
c_3		1.4248***	0.2011
c_4		2.6716***	0.3237

Note: Dependent variable: size of interest rate change; number of observations: 117. Log likelihood is -62.0857. Variable definitions: $y_{t_{N(t-1)}}$ is the magnitude of the last target change as of date $t-1$. SP_{t-1} is the value of the the six-month Bubill rate minus the Eonia rate.

Table 3.5: Parameter Estimates for the Post Financial Crisis Sample (Oct 2008 - Dec 2014) of the OP-Model

Parameter	Variable	Estimate	Standard Error
π_1	$y_{t_{N(t-1)}}$	3.7908***	0.3361
π_2	SP_{t-1}	3.3146***	0.4472
c_1		-3.6422***	0.2049
c_2		-2.3923***	0.1835
c_3		1.9194***	0.1453
c_4		7.4683	0.2751

Note: Dependent variable: size of interest rate change; number of observations: 75. Log likelihood is -28.5913. Variable definitions: $y_{t_{N(t-1)}}$ is the magnitude of the last target change as of date $t-1$. SP_{t-1} is the value of the the six-month Bubill rate minus the Eonia rate.

explanatory power is also estimated for the spread between the six-month Bubill rate and the Eonia rate, SP_{t-1} for both samples. This spread captures the expectation in short-term market rates of a change in the policy interest rate. If the six-month Bubill rate is above the Eonia rate, the ECB is more likely to raise the interest rate. Tables 3.4 and 3.5 summarise the estimates for the pre and post financial crisis samples respectively.

To check whether my OP results are sensitive to an extension until December 2014, I reestimate the post financial crisis sample also on data from October 2008 until January 2014. The results confirm those presented.

3.7 Monetary Policy Shocks

The method used by this paper allows to compare impulse responses between the linear VAR and the nonlinear VAR model. Based on linear VARs, a large amount of work is carried out on calculating the influence of monetary policy. But, the linear VAR mixes the effects of two expectations: impulse responses based on (a) an expected interest rate decrease which is not implemented by the central bank and (b) an unexpected increase in the policy interest rate. In contrast, the nonlinear VAR model differentiates between the influence of these two very different effects. Section 3.7.1 briefly summarises the used VAR model as well as the non-linear VAR model. For more details on the estimation of the non-linear model refer to section D in the Appendix. Sections 3.7.2 to 3.7.4 describe the results obtained for the pre and post financial crisis samples as well as their possible differences. Section 3.7.5 briefly discusses the effect of monetary policy at the ZLB and provides a theoretical link between my results and the influence of monetary policy in a liquidity trap.

3.7.1 Specification of the VAR set up

I extend Hamilton and Jorda's (2002) model and additionally incorporate exogenous variables. Thus, I estimate the VAR model based on two groups of variables. The first group of variables, x_t, denotes a vector of exogenous foreign variables for month t. It contains monthly data on the log of an oil price index (oil_t), the log of US IP (y_t^{US}) and a US short-term nominal interest rate (s_t^{US}).[33] By including these variables, I capture changes in world demand and inflation and avoid a potential price puzzle.[34]

$$x_t = (\mathrm{oil}_t, \mathrm{y}_t^{US}, s_t^{US})' \tag{3.16}$$

The exogenous variables are determined outside the VAR system. More specifically, these variables influence the second group of variables, the endogenous variables, contemporaneously. But the endogenous variables have no feedback on the exogenous variables.

[33] Crude Oil-Brent, FOB, in USD, per barrel; US IP, total index volumes, 2007=100 prices, seasonally adjusted, in USD; US T-Bill, secondary market, 3 month, middle rate. See Appendix B for more details.

[34] The price puzzle is a widespread empirical finding in the VAR literature arising from the issue that monetary policy makers consider information on inflation which is not included in the set of VAR variables. This may lead to price increases following an interest rate tightening. For further details see Sims (1992).

The endogenous variables are specified by y_t which denotes a vector of observable macro variables.

$$y_t = (\mathrm{IP}_t, \mathrm{P}_t, MRO_t, \mathrm{M3}_t, \mathrm{ExR}_t)' \tag{3.17}$$

The VAR-model is estimated in levels. More precisely, I use monthly data on the log of industrial production (IP), log consumer prices (P), the interest rate on main refinancing operations (MRO) in levels, the log money aggregate M3 (M3) and the log of the real effective exchange rate (ExR).[35] An estimation based on levels allows for implicit co-integrating relationships in the data, and still have consistent estimates of the parameters (Sims et al. (1990)). Standard information criteria specify the lag length of the VAR. Based on the Likelihood Ratio (LR) and the Akaike criterion (AIC), 6 lags are used for the pre and post crisis samples. The pre crisis sample covers the time period between January 1999 and September 2008, whereas the post crisis sample ranges from October 2008 until December 2014. From February 2014 onwards forecasted data on the endogenous and exogenous variables is used.[36]

The ordering of the variables is the same as in equation (3.17). In general, this choice is motivated by Peersman and Smets (2003), Kim and Roubini (2000), Sims (1992) and Christiano et al. (1999). This ordering implies that a Euro Area monetary policy shock has no contemporaneous influence on IP and prices. This feedback rule is based on the assumption that firms do not immediately change output and prices following an unexpected change in the policy interest rate.[37] Furthermore, it is more realistic to assume that output and prices do not simultaneously respond to policy shocks within a month than within a quarter. Since I use a monthly frequency, this assumption should be not as restrictive. IP and consumer prices are summarised in a vector of variables which come before MRO_t and is given by $y_{1t} = (\mathrm{IP}_t, \mathrm{P}_t)$. Instead, the monetary policy shock may have an impact on M3 and the exchange rate. Hence, the ordering suggests

[35] Euro Area IP, excluding construction, 2010=100, seasonally adjusted; Euro Area 17 harmonised index of consumer prices, all items, 2005 = 100, not seasonally adjusted; Euro area money supply M3, current account, amount outstanding, seasonally adjusted; Euro Area real effective exchange rate, CPI deflated, broad group (40 partners).

[36] For detailed information on the data, see Appendix SectionB. Feri Euro Rating provided the monthly forecasts on IP, HICP, MRO, M3, ExR, oil price index, short term US interest rate and the forecast for US IP. As pointed out, the extension of the second sample is necessary since a sample until January 2014 does not provide a sufficient sample length for estimation of the monthly nonlinear VAR model.

[37] Adjustment costs and planning delays may hinder firms to act within the same month (Kim and Roubini (2000)).

that the interest rate on MROs is not simultaneously influenced by ExR and money stock changes. The respective vector of variables is $y_{2t} = (\mathrm{M3}_t, \mathrm{ExR}_t)$.[38] For robustness, I also change the ordering of the endogenous variables. This does not significantly affect my results presented in sections 3.7.2 to 3.7.4. To identify a monetary policy shock, a standard Cholesky-decomposition is used. The impulse response functions are given by:

$$\frac{\partial \mathrm{E}(y_{t+s}|x_t, MRO_t, y_{1t}, y_{t-1}, y_{t-2}, \ldots)}{\partial MRO_t}.$$

Similarly, it is possible to write the effects for y_{t+s} of new information about MRO_t which is by definition an orthogonalised shock. The corresponding definition of an orthogonalised shock is given by

$$u_t^{MRO} = MRO_t - \mathrm{E}(MRO_t|x_t, y_{1t}, y_{t-1}, y_{t-2}, \ldots),$$

which can be rewritten as

$$u_t^{MRO} = MRO_t - MRO_{t-1} - [\mathrm{E}(MRO_t|x_t, y_{1t}, y_{t-1}, \ldots) - MRO_{t-1}]. \tag{3.18}$$

Two situations lead to a positive u_t^{MRO} in equation 3.18. First, the ECB increased the interest rate on MROs ($MRO_t - MRO_{t-1} > 0$), while the market did not anticipate a change ($\mathrm{E}(MRO_t|x_t, y_{1t}, y_{t-1}, \ldots) - MRO_{t-1} = 0$). Second, the ECB kept its interest rate constant ($MRO_t - MRO_{t-1} = 0$), whereas the market anticipated a decrease in the interest rate ($\mathrm{E}(MRO_t|x_t, y_{1t}, y_{t-1}, \ldots) - MRO_{t-1} < 0$). For further details on the nonlinear forecasting model and its estimation, refer to Appendix D.

3.7.2 Pre Financial Crisis Results on Monetary Policy Transmission

Figure 3.3 plots the influence of the three monetary policy shocks for the pre crisis sample and in Figure E.2, E.3 and E.4 in Appendix E the respective impulse responses with

[38]The assumption that the ExR does not simultaneous influence the policy rate is suitable for large and relatively closed economies such as the Euro Area as a whole. This is, in addition to Peersman (2004), also stated by Eichenbaum and Evans (1995) regarding the US economy. Additionally, the ExR is an arbitrage equation. It depicts the financial market equilibrium and it is a forward-looking asset price which comprises all information available today (Kim and Roubini (2000)).

confidence bands are given. Note that the responses of each shock are normalised to a 100 basis point (BP) shock in the interest rate. The length of the impulse responses is set to 18 months. The *linear VAR* impulse response functions are displayed via the solid lines and estimate the influence of a 100 BP increase in MRO_t. They show that a temporary tightening of the interest rate on MROs seems to be followed by a decrease in money stock (M3). However, this effect is insignificant (see Figure E.2 in Appendix E). The ExR seems to significantly depreciate on impact and during the first month following the shock.[39] Both IP and consumer prices follow with a delayed decrease. The effect on output is slightly significant between 12 and 13 months (see Appendix E Figure E.2). This may reflect that a more credible central bank is generally assumed to obtain its monetary policy goal at lower output costs (Clements et al. (2001)). For prices, I estimate an insignificant decrease. In contrast to an earlier study on European data by Peersman and Smets (2003), who use a similar approach but a different data horizon with synthetic Euro Area data, my VAR result reveal an insignificant and non-permanent decline in prices.

Impulse response functions following the *unexpected increase* in the interest rate are on average comparable to the VAR impulses but quantitatively much larger. They are summarised via the dashed line in Figure 3.3. Figure E.3 in Appendix E reports the responses with confidence bands. Numerically my results are comparable to the US-results of Hamilton and Jorda (2002). Compared to the linear VAR responses, the unexpected shock seems to imply much larger and more long lasting effects on IP, prices, M3 and the ExR. More specifically, following the unexpected shock, both IP and prices significantly decrease. The significant decline in IP occurs after 12 months and prices fall significantly between three and six months and again between 11 and 13 months. Furthermore, the fall in prices, following the unexpected shock, leads to a larger decline in prices than based on the linear VAR shock. Based on the unexpected shock, prices decline by about -0.7% after four months. The respective decline in the linear VAR model seems to stand at about -0.4%. In contrast to the linear VAR, the unexpected shock seems to lead to a significant and permanent decline in M3. Also for the ExR, the unexpected shock confirms a constantly larger depreciation compared to the linear VAR. The effect is significant except for the period between three and 13 months after the unexpected tightening.

[39]Note that an increase of the ExR responses specifies an appreciation of the Euro vis-*à*-vis the other currencys.

The *expected decrease* in the interest rate which is not implemented by the ECB is given by the dotted line in Figure 3.3. Figure E.4 in Appendix E gives the responses with confidence bands. As expected, these seem to have few lasting consequences for the interest rate forecasting as well as for the prediction of the other variables. The market then simply expects a decrease at the next meeting of the monetary authority.

The different results of these two effects, (a) the unexpected increase and (b) the expected decrease which is not implemented, highlights the necessity to treat them separately. My results therefore deliver an important contribution for a better understanding of the different monetary policy shocks and their respective transmission mechanisms in the Euro Area.

Figure 3.3: Pre Financial Crisis Responses to Different Shocks in the Policy Interest Rate

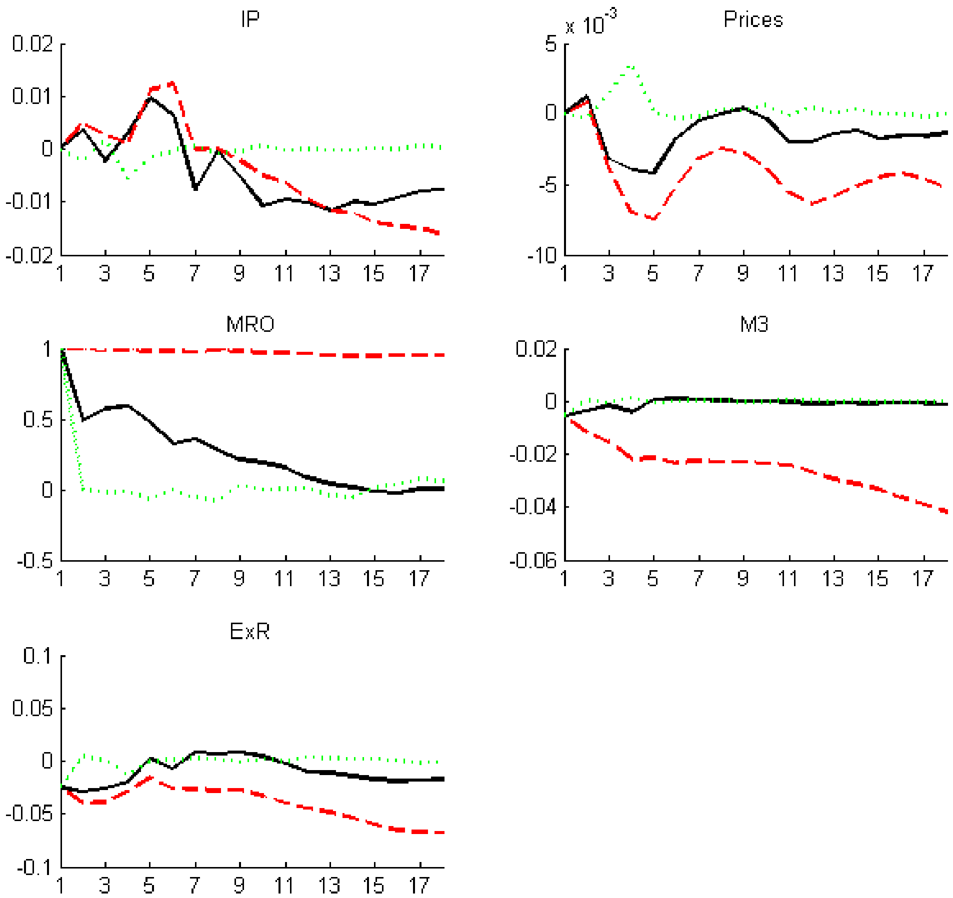


Figure 3.3 plots the influence of the three monetary policy shocks on y_{t+j} for months $j = 0, 1, \ldots, 17$ for the respective variables of y. The shocks are normalised to 100-basis-points. *Solid black line*: VAR impulses. *Dashed red line*: impulses based on an unexpected interest rate increase by the ECB. *Dotted green line*: impulses based on an expected interest rate decrease which is not implemented by the ECB.

3.7.3 Post Financial Crisis Results on Monetary Policy Transmission

Figure 3.4 shows the impulse responses for the post crisis sample of the three monetary policy shocks. The respective impulse responses with confidence bands are given in Figure F.5, F.6 and F.7 in Appendix F.

The impulse responses in Figure 3.4 indicate again that an unexpected shock leads to a larger effect on all endogenous variables than a linear VAR shock. Following a monetary policy tightening, the *linear VAR* responses seem to be insignificant for all endogenous variables except for a slight significant depreciation of the ExR after five months (Figure F.5). In contrast, impulse responses based on the *unexpected shock* reveal a significant decline in output between 12 and 15 months. The effect on prices and M3 is insignificant. The ExR seems to depreciate significantly between nine and 14 months (Figure F.6). As for the pre crisis sample, the effect of an *expected decrease* in the policy rate which is not implemented seems to have only a minor and insignificant impact during the post crisis period.

As a robustness check, I reestimate the post crisis sample on quarterly data based on two lags.[40] This allows for a shorter estimation horizon which is not based on forecasted data. The time horizon lasts from 2008Q4 until 2013Q4. Further, I substitute IP with GDP. In contrast to IP, GDP captures the developments of the whole economy and may therefore give a more thorough overview on output. The results following a linear VAR shock and the unexpected shock are given in Figure F.8 and F.9 respectively in Appendix F.1. They confirm the insignificant effect following the linear VAR shock during the post crisis sample. The effect of the unexpected shock also yields insignificant effects on IP, prices, M3 and the ExR.

3.7.4 Difference in Monetary Policy Transmission between the First and Second Sample

Figure 3.5 compares the *linear VAR* responses for the pre and post financial crisis sample. The black responses refer to the pre crisis sample and the light blue responses to the post

[40]Standard information criteria suggest a lag length of two quarters.

Figure 3.4: Post Financial Crisis Responses to Different Shocks in the Policy Interest Rate

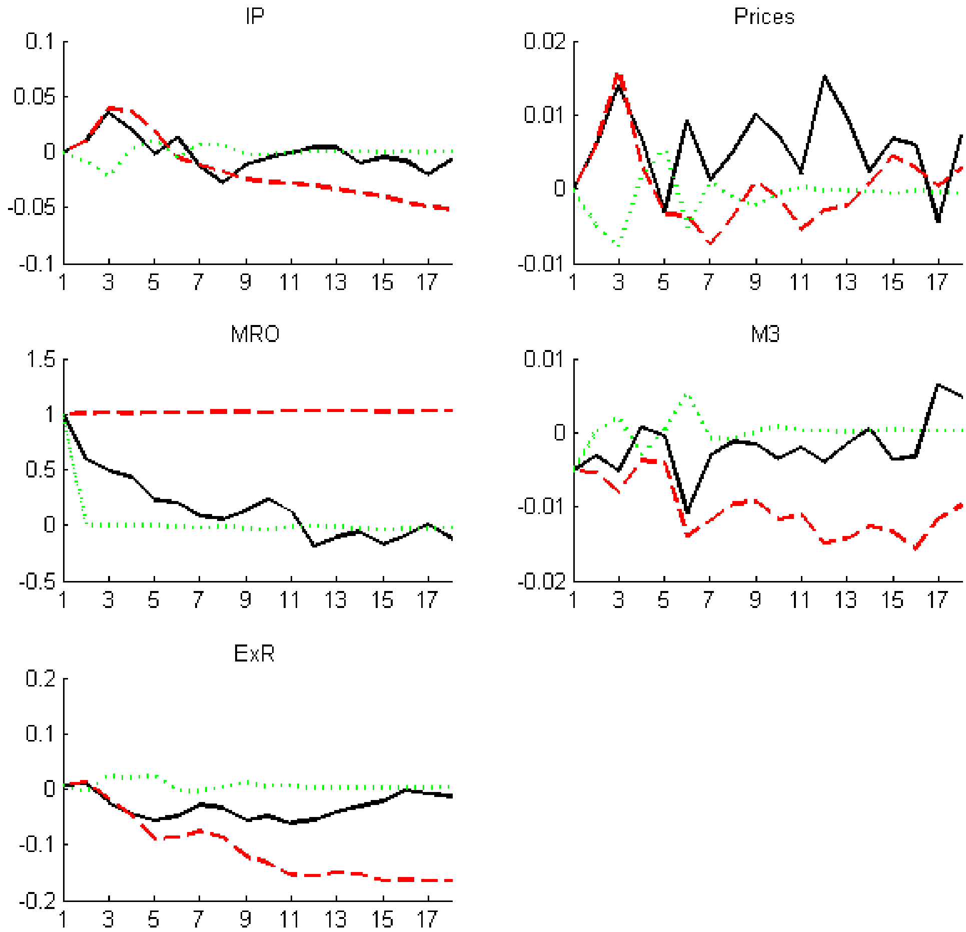


Figure 3.4 plots the influence of the three monetary policy shocks on y_{t+j} for months $j = 0, 1, \ldots, 17$ for the respective variables of y. The shocks are normalised to 100-basis-points. *Solid black line*: VAR impulses. *Dashed red line*: impulses based on an unexpected interest rate increase by the ECB. *Dotted green line*: impulses based on an expected interest rate decrease which is not implemented by the ECB.

crisis period. In contrast to the impulses for the pre crisis, the impulse responses for the post crisis sample seem to be broadly insignificant.

Compare Figure 3.6 for the responses based on the *unexpected shock*. The red lines depict the pre crisis responses and the light blue lines refer to the post crisis sample. Generally, the impulse responses during the pre crisis sample appear to have a more persistent effect than those estimated for the post crisis sample. More specifically, IP seems to significantly contract after 12 months onwards during the pre crisis period and for the post crisis, a significant decline is estimated between 12 and 15 months. Regarding prices, I estimate a significant decline for the pre crisis sample. This is in contrast to the post crisis sample where the response is insignificant. M3 seems to decline permanently significant during

Figure 3.5: VAR Shock: Difference between the Pre and Post Financial Crisis Sample

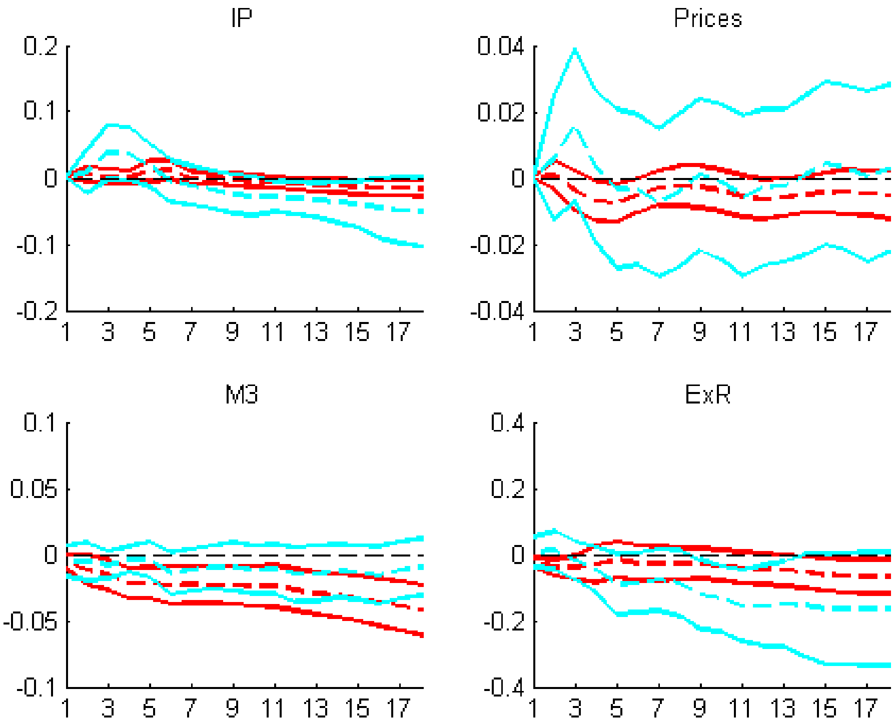


Figure 3.5 plots linear VAR responses for the pre and post crisis sample. The *red* lines refer to the pre crisis sample and the *light blue* lines to the post crisis sample. *Solid lines*: VAR impulses. *Dashed lines*: respective two-standard error bounds.

the pre crisis sample, whereas the post crisis response also reveals an insignificant decline. The ExR significantly depreciates on impact and from 13 months onwards during the pre crisis. In contrast, during the post financial crisis sample the response only depreciates between nine and 14 months.

Overall, compared to the pre crisis, the post crisis responses seem to be less significant. This implies that monetary policy may be less effective during the post financial crisis sample (see section 3.7.5 for a detailed discussion). However, note that a smaller sample may also induce the estimation of wider confidence bands. It will be interesting for future research to see whether the size of the confidence bands shrinks with a longer time horizon available. Thus, my results should be seen as a tentative and preliminary evaluation on the Euro Area monetary transmission mechanism since the occurrence of the financial turmoil in 2008.

Figure 3.6: Unexpected Shock: Difference between the Pre and Post Financial Crisis Sample

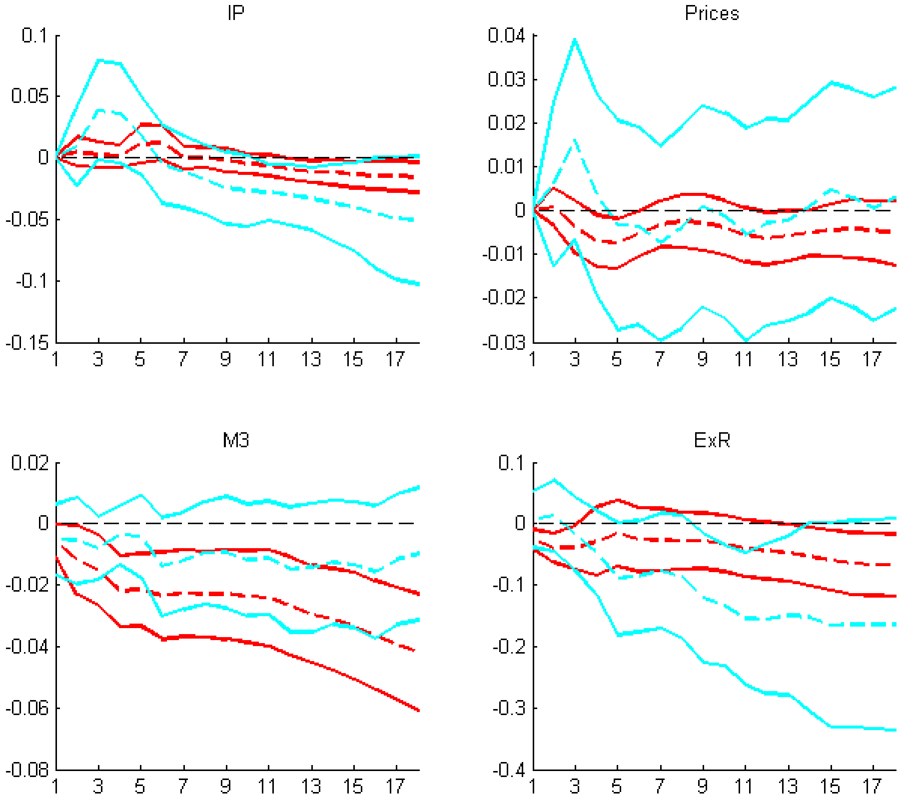


Figure 3.6 plots unexpected monetary policy shock responses for the pre and post crisis sample. The red lines refer to the pre crisis sample and the light blue lines to the post crisis sample. Solid lines: impulses based on the unexpected shock. Dashed lines: respective two-standard error bounds.

3.7.5 Why do the pre and post financial crisis results differ?

The following subsections briefly describe the theoretical effect of monetary policy when the economy is at the Zero Lower Bound (ZLB). Subsection 3.7.5 discusses the effect based on the IS-LM model and subsection 3.7.5 draws a possible link between the obtained results in this paper and the effectiveness of monetary policy at the ZLB.

A brief description of the liquidity trap: the IS-LM view

Monetary policy works mainly through influencing prices and yields of financial assets, which in turn influence economic decisions and thus economic development. When the short-term policy rate is at or near the ZLB it becomes infeasible to use it for providing monetary ease as it can not go below zero. Under such circumstances, investors are indifferent between holding bonds and money,[41] and as a consequence, monetary policy is ineffective at boosting demand (see Krugman (2000) for a detailed discussion). This situation is known as a liquidity trap. In 'normal' times, monetary policy is neutral in the long run. But since prices are to some extent sticky, in the short run a monetary easing may stimulate output and lead to an equiproportional increase in prices. However, when nominal interest rates are close to zero or at the ZLB, so that the economy is in a liquidity trap, monetary policy is ineffective to rise output and prices.

For illustrative purpose, I consider the IS-LM paradigm. In a liquidity trap, money demand becomes more or less infinitely elastic. This implies that the left most part of the LM curve is horizontal at an interest rate near zero. Assume that the IS curve intersects the LM curve in that flat area, as it is shown in Figure 3.7. Then, an expansive monetary policy, which moves the LM curve from LM to LM', has no effect on output and interest rates. Note that the intersection of IS with LM' is still at the same level of output y as the intersection of IS with LM (Figure 3.7). Hence, conventional monetary policy is powerless as the economy is in a liquidity trap. For a more detailed discussion refer to Krugman (2000).

Figure 3.7: IS-LM Model and the Liquidity Trap

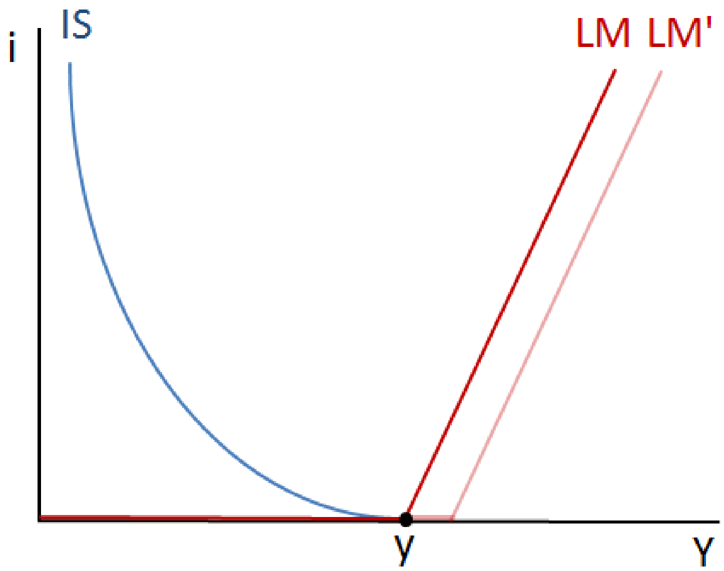


Figure 3.7 describes the liquidity trap in the IS-LM model. The vertical axis denotes the nominal interest rate and the horizontal axis depicts the output level.

[41] Bonds and money become equivalent assets, since both yield the same nominal interest which is essentially zero at the ZLB.

Hypothetical Link between my Results and Monetary Policy in a Liquidity Trap

My results may mirror these circumstances of a liquidity trap in the Euro Area.[42] More specifically, in the wake of the global financial crisis, Euro Area GDP and inflation decreased substantially (compare Figures F.10 and F.11 in Appendix F.2). At the same time, the short-term interest rate on MROs has been lowered from 4.25% in August 2008 to 0.25% in November 2013 (compare Figure 3.1). According to the theory, conventional monetary policy is ineffective when the economy is at the ZLB and thus in a liquidity trap. Since the ECB's policy rate is very low, at most, for the last five years of my post financial crisis sample, my findings may be explained by the liquidity trap. More specifically, the short-term interest rate reached 1% in the middle of 2009. Following a short-lived recovery in 2011, the policy rate temporarily increased to 1.5%. Already at the end of 2011 the policy rate was again gradually reduced. Since November 2013, it remains at an historically low level of 0.25%. With the nominal interest rate being close to the ZLB during some parts of the post crisis sample, the Euro Area economy seems to be for this timeframe in a liquidity trap or close to it. Under such circumstances, the central bank's ability to reduce its short-term policy rate meaningfully ended and as described above, monetary policy is ineffective to increase output and prices. The estimated insignificant effect on output and prices following the monetary policy shock during the post crisis sample might reflect the ineffectiveness of monetary policy at the ZLB. However, it should be stressed that it may still be too early to conclusively judge on the monetary policy transmission in the Euro Area after the financial crisis. When a sample split is implemented with the occurrence of the financial turmoil in 2008, it allows for an overall timeframe of just about five years. For a VAR estimation this is still a rather short time horizon.

3.8 Conclusion

This paper provides new evidence on how monetary policy is transmitted in the Euro Area and whether it has changed with the occurrence of the financial crisis in 2008.

[42] Note, although section 3.7 analyses the transmission of an increase in the policy rate, the same mechanisms holds for a decrease in the respective interest rate. The impulse responses just have the opposite sign.

I am in particular interested in the VAR framework and its explanation on how a monetary policy shock influences key macroeconomic variables. More specifically, I use the approach by Hamilton and Jorda (2002). They developed a non-linear forecasting tool which significantly enhances VAR predictions. It allows to differentiate between an unexpected policy interest rate increase and an expected but not implemented decrease. The usual linear VAR fails to account for this difference which leads to statistically different impulse response functions. Moreover, by means of a VAR setting there is very little evidence on possible differences in the transmission mechanism before and after the financial crisis in 2008. Since it is reasonable to suspect that the financial turmoil impaired the ECB's reaction function, I split the data into two samples. I use monthly data ranging from January 1999 through September 2008 and from October 2008 until December 2014.

Four results of my analysis stand out: (1) it is important to differentiate between an unexpected increase and an expected decrease not implemented as these responses have very different implications for the transmission mechanism whereas the usual linear VAR mixes these effects. (2) Generally, compared to the linear VAR, an unexpected tightening reveals much larger and more long lasting effects on industrial production, consumer prices, money growth and the exchange rate. This result is especially pronounced for the pre crisis sample. 3) Regarding the expected decrease which is not implemented, I observe for both samples rather minor and insignificant impacts. The market might then simply expect a decrease of the policy interest rate at the next central bank's meeting. (4) Tracing a monetary policy shock in the post financial crisis sample generally verifies a less significant effect. The insignificance could indicate that monetary policy is less effective during the second sample.

Given my results, this work contributes to a better understanding of the ECB's monetary policy. It delivers new insights into the possibility of modeling the ECB's interest rate and it contributes to a better understanding of the influence of monetary policy on the overall economy. My results thus underline the importance of this methodology, in particular for the evaluation of monetary policy effects.

A Changes in the Interest Rate on MROs

Table 3.6: Dates of Interest Rate Changes on MROs

Meeting Dates	Date of Change	Interest Rate Change	Interest Rate Level Main Refinancing Operations Fixed Rate Tender	Variable Rate Tender [Minimum Bid Rate]
01.07.1999			3	
04.08.1999	04.09.1999	-0.5	2.5	
11.04.1999	11.05.1999	0.5	3	
02.03.2000	02.04.2000	0.25	3.25	
03.16.2000	03.17.2000	0.25	3.5	
04.27.2000	04.28.2000	0.25	3.75	
06.08.2000	06.09.2000	0.5	4.25	
06.21.2000	06.28.2000			4.25
08.31.2000	09.01.2000	0.25		4.5
10.18.2000	10.06.2000	0.25		4.75
05.10.2001	05.11.2001	-0.25		4.5
08.30.2001	08.31.2001	-0.25		4.25
09.17.2001	09.18.2001	-0.5		3.75
11.08.2001	11.09.2001	-0.5		3.25
12.05.2002	12.06.2002	-0.5		2.75
03.06.2003	03.07.2003	-0.25		2.5
06.05.2003	06.06.2003	-0.5		2
12.01.2005	12.06.2005	0.25		2.25
03.02.2006	03.08.2006	0.25		2.5
06.08.2006	06.15.2006	0.25		2.75
08.03.2006	08.09.2006	0.25		3
10.10.2006	10.11.2006	0.25		3.25
12.07.2006	12.13.2006	0.25		3.5
03.08.2007	03.14.2007	0.25		3.75
06.06.2007	06.13.2007	0.25		4
07.03.2008	07.09.2008	0.25		4.25
10.08.2008	10.15.2008	-0.5	3.75	
11.06.2008	11.12.2008	-0.5	3.25	
12.04.2008	12.10.2008	-0.75	2.5	
01.15.2009	01.21.2009	-0.5	2.0	
03.05.2009	03.11.2009	-0.5	1.5	
04.02.2009	04.08.2009	-0.25	1.25	
05.07.2009	05.13.2009	-0.25	1.0	
04.07.2011	04.13.2011	0.25	1.25	
07.07.2011	07.13.2011	0.25	1.5	
11.09.2011	11.09.2011	-0.25	1.25	
12.14.2011	12.14.2011	-0.25	1.0	
07.11.2012	07.11.2012	-0.25	0.75	
05.08.2013	05.08.2013	-0.25	0.5	
11.13.2013	11.13.2013	-0.25	0.25	

Source: ECB

B Data Sources

This paper uses monthly data on the Euro Area for the time horizon between 1999:01 and 2014:01. The model ist estimated in levels. Like Sims et al. (1990) state, this accounts for possible discrepancy that may arise in case of incorrectly assumed cointegration restrictions. The following time series are used in the VAR estimation:

Industrial Production (IP): Log of industrial production (excluding construction), volumes, 2010 = 100, Euro area-17, monthly series. Source: Eurostat.

Consumer price (CPI): Log of HICP, all items, 2005=100, Euro area-17, neither seasonally nor working day adjusted, monthly series. Source: Eurostat.

Main refinancing rate (MRO): Main refinancing interest rate, middle rate, Euro Area, in percent. Source: European Central Bank.

M3 (M3): Log of money supply M3, outstanding amounts, sa Euro Area. Source: European Central Bank.

Exchange rate (ExR): Log of ECB real effective exchange rate, Euro area-17 countries vis-a-vis the EER-40 group of trading partners (AU, CA, DK, HK, JP, NO, SG, KR, SE, CH, GB, US, BG, CZ, LV, LT, HU, PL, RO, CN, DZ, AR, BR, CL, HR, IS, IN, ID, IL, MY, MX, MA, NZ, PH, RU, ZA, TW, TH, TR and VE) against Polish zloty. Monthly index (reference period: 99Q1=100) monthly series Source: European Central Bank.

Gross domestic product (GDP): Log of gross domestic product at market prices, chain linked volumes, reference year 2005, working day and seasonally adjusted by TRAMO/SEATS, quarterly series. Source: Eurostat.

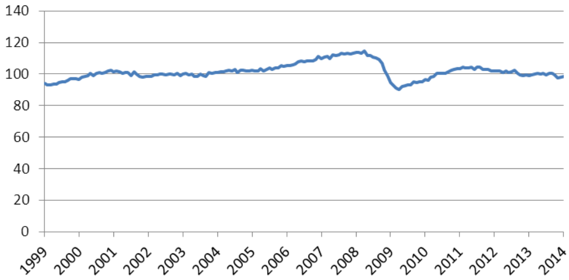
Crude oil-WTI (Oil): Log of crude oil-WTI Spot Cushing, USD/BBL. Source: Datastream.

US industrial production (US_IP): Log of US industrial production, total index, sa, 2007=100. Source: Datastream.

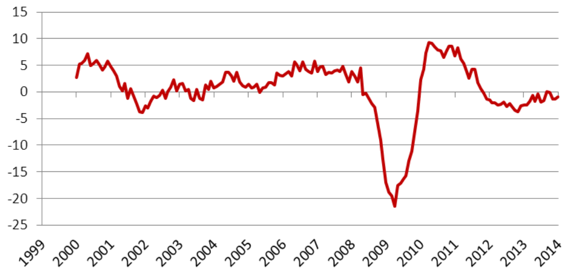
US T-Bill (US T-Bill): Log of US T-Bill rate, secondary market, 3 months, middle rate. Source: Datastream.

Figure B.1: Monthly Data, Euro Area

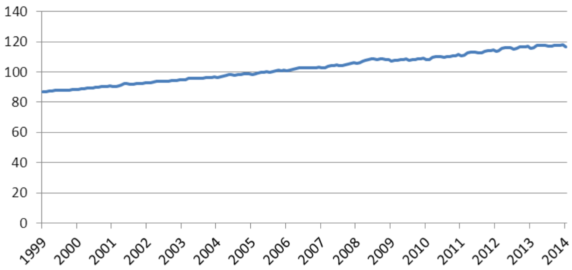
(a) IP in levels



(b) IP, yearly growth rate

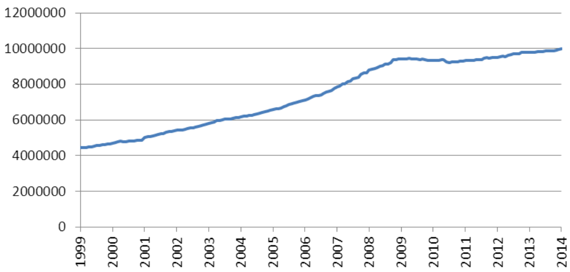


(c) CPI in levels



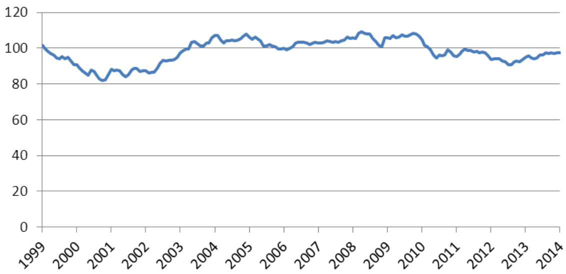
(d) CPI, annual rate of change

(e) M3 in levels

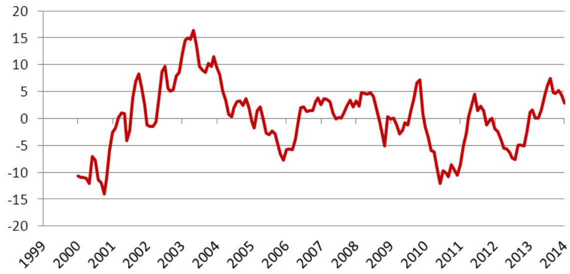


(f) M3, annual rate of change

(g) ExR in levels



(h) ExR, annual rate of change



C Forecasting with the ACH Framework

As Hamilton and Jorda (2002) illustrate, the ACH framework allows calculating the one-period-ahead forecast of the policy interest rate i_{t+1}. This is a closed-form and based on information Υ_t available at time t. The respective forecasting equation is given by:

$$\begin{aligned} \mathrm{E}(i_{t+1}|\Upsilon_t) = (1-h_{t+1})i_t + h_{t+1}\sum_{j=1}^{5}(i_t+s_j) \\ \times[\Phi(c_j - \mathrm{w}_t'\pi) - \Phi(c_{j-1} - \mathrm{w}_t'\pi)], \end{aligned} \tag{3.19}$$

where $\mathrm{w}_t = (y_{t_{N(t)}}, z_t)'$ with $z_t = \mathrm{SP}_t$; the hazard rate h_{t+1} is estimated from equation 3.5; $s_j = (0.25)(j-3)$, c_j are the respective OP parameter estimates and summarised in Table 3.4 and 3.5.[43]

To generate further forecast values of the policy interest rate, four steps are necessary: *First*, with the help of a VAR, containing information of the interest rate i_t, its lag, the lag of the spread or the lag inflation forecast, the forecasts of the SP_t or $\mathrm{InflF12}_t$ are calculated.[44] *Second*, since the forecasting equation 3.19 is due to the information set Υ_{t+j} nonlinear, simulation is required. Equation 3.19 is based on a discrete probability distribution for $i_{t+1}|\Upsilon_t$, from this distribution it is possible to calculate a value for the one-period-ahead forecast $i^{(1)}_{t+1}$. Given this value and assuming a Gaussian error in equation 3.20, it is possible to calculate a one-period-ahead forecast of the explanatory variable $z^{(1)}_{t+1}$ from equation 3.20. This value is a draw from the respective distribution $z_{t+1}|\Upsilon_t$. The discrete probability distribution for $i_{t+1}|\Upsilon_t$ and the just generated forecast value $z^{(1)}_{t+1}$ are used to calculate the two-step-ahead forecast of the policy interest rate $i^{(1)}_{t+2}$, which is a draw from the discrete probability distribution $i_{t+2}|\Upsilon_t$. These steps are repeated to generate a sequence of these single forecast values up to $i^{(1)}_{t+j}$. *Third*, since this is a sequence of single values, it is necessary to repeat these steps from the beginning for generating m simulations of these values $i^{(m)}_{t+j}$ from the distribution $f(i_{t+j}|\Upsilon_t)$. *Fourth*, the final forecast $\mathrm{E}(i_{t+j}|\Upsilon_t)$ is calculated as an average of these M simulations, $M^{-1}\sum_{m=1}^{M} i^{(m)}_{t+j}$.

---

[43]For example, for the pre crisis sample $\mathrm{w}_t'\pi = 0.656 y_{t_{N(t)}} + 3.4959\mathrm{SP}_t$

[44]The respective equations for the pre financial crisis sample with standard errors in parentheses are:

$$\begin{aligned} \mathrm{InflF12}_t &= \underset{(0.0089)}{0.0346} + \underset{(0.0115)}{0.0378} i_t - \underset{(0.0089)}{0.0307} i_{t-1} + \underset{(0.0069)}{0.9758}\mathrm{InflF12}_{t-1}. \\ \mathrm{SP}_t &= \underset{(0.1258)}{0.6656} + \underset{(0.0644)}{0.0406} i_t - \underset{(0.1263)}{0.6851} i_{t-1} + \underset{(0.0763)}{0.3719}\mathrm{SP}_{t-1}. \end{aligned} \tag{3.20}$$

D Further Details on the Estimation of the Nonlinear Forecasting Model

For example, consider the following linear VAR

$$y_t = c + \theta x_t + \Phi_1 y_{t-1} + \Phi_2 y_{t-2} + \ldots + \Phi_6 y_{t-6} + \epsilon_t,$$

with ϵ_t i.i.d. $N(0, \Omega)$. OLS regressions provide the maximum likelihood estimates of the parameters $(c, \theta x_t, \Phi_1 y_{t-1}, \Phi_2 y_{t-2}, \ldots, \Phi y_{t-6})$ of a VAR. In addition, it is necessary to forecast the macro vector of variables y_{2t} that comes after MRO_t given y_{1t} and i_t which is the hypothesised value of the interest rate. This system of equations is estimated via OLS and can be collected and written in vector form as

$$y_{2t} = d + d_1 i_t + D_0 y_{1t} + \theta x_t + B_1 y_{t-1} + B_2 y_{t-2} + \ldots + B_1 y_{t-6} + u_{2t}. \tag{3.21}$$

[45] The next step is to estimate the forecast of $\tilde{y}_{2t|t}(i_t)$ based on equation 3.21 with the hypothesised value i_t and the values for $y_{1t}, y_{t-1}, y_{t-2}, \ldots$. The following vector comprises these forecasts as well as the hypothesized value i_t and y_{1t}

$$\tilde{y}_{t|t}(i_t) = (y'_{1t}, i_t, x_t, \tilde{y}'_{2t|t}(i_t))'. \tag{3.22}$$

For the following system of equations it is then possible to calculate the one-step-ahead VAR forecast based on i_t

$$\begin{gathered}\hat{E}(y_{t+1}|i_t, x_t, y_{1t}, y_{t-1}, y_{t-2}, \ldots) = c + \theta x_t + \\ \Phi_1 \tilde{y}_{t|t}(i_t) + \Phi_2 y_{t-1} + \ldots + \Phi_6 y_{t-5}.\end{gathered} \tag{3.23}$$

Before in Appendix section C, the forecast of the interest rate on MROs (i_{t+1}), based on the ACH model is calculated. This forecast is used to replace the third row of the vector of conditional forecasts in equation 3.23, relating to the VAR forecast value MRO_{t+1}. $\tilde{y}_{t+1|t}(i_t)$ is then the respective vector comprising these forecasts. For calculating the

[45]Since y_{2t} consists of two macro variables, M3 and ExR, d_1 is a 2×1 vector, D_0 a 2×3 matrix, and B_j are 2×5 matrices.

two-step-ahead forecasts conditioned on i_t, the parameters of the VAR are used

$$\begin{aligned}\hat{E}(y_{t+2}|i_t, x_t, y_{1t}, y_{t-1}, y_{t-2}, \ldots) = c + \theta x_t + \\ \Phi_1 \tilde{y}_{t+1|t}(i_t) + \Phi_2 y_{t|t}(i_t) + \ldots + \Phi_6 y_{t-4}.\end{aligned} \tag{3.24}$$

As above in equation 3.23, the third row in equation 3.24 is replaced with the respective forecast i_{t+2}. The vector collecting these forecasts is $\tilde{y}_{t+2|t}(i_t)$. I repeat these steps to estimate the dynamic influences on the macro variables of the forecasts for $i_t, i_{t+1}, \ldots$. The resulting vector, collecting the dynamic influences, is called $\tilde{y}_{t+j|t}(i_t)$.

Now, I address the question regarding the ECB's influence of an interest rate increase by 25 basis points during month t ($i_t = i_{t-1} + 0.25$) in contrast to the case of no change ($i_t = i_{t-1}$). The respective answer is then normalized in units of a derivative

$$(0.25)^{-1}[\tilde{y}_{t+j|t}(i_t)|_{i_t=i_{t-1}+0.25} - \tilde{y}_{t+j|t}(i_t)|_{i_t=i_{t-1}+0.25}]. \tag{3.25}$$

By replacing the corresponding third row of 3.23 and 3.24 at each repetition in the ACH forecast, I obtain values in 3.25 which are based on t and i_{t-1}. Consequently, these values differ numerically to the standard VAR impulse response function. For computational purpose, equation 3.25 is then averaged over the time span $t = 1, \ldots, T$ and $y_1, \ldots, y_T$.

The next question addresses the predicted interest rate change which is not implemented by the ECB. This refers to the second term in 3.23. The respective answer is modeled as

$$\omega_t[\tilde{y}_{t+j|t}(i_t)|_{i_t=i_{t-1}} - \tilde{y}_{t+j|t}(i_t)|_{i_t=i_{t-1}}], \tag{3.26}$$

where

$$\omega_t = \begin{cases} (i_{t-1} - \hat{i}_{t|t-1})^{-1} & \text{if} |i_{t-1} - \hat{i}_{t|t-1}| > 0.05 \\ 0 & \text{otherwise.} \end{cases}$$

The interest rate forecast on MROs $\hat{i}_{t|t-1}$ for month t is estimated based on the information set $t-1$. ω_t denotes the weight, which respectively drops the no-change-events, since no change was expected. Additionally, positive or negative forecast errors are also rescaled by ω_t into unites similar to equation 3.25. As above, the substitution of the estimated ACH forecasts i_{t+j} for the VAR forecasts MRO_{t+j} induces equation 3.26 to depend on t

and not to be numerically identical to the VAR impulse-response function.

E Pre Financial Crisis Monetary Policy Shocks

Figure E.2: VAR Responses to a Shock in the Policy Interest Rate

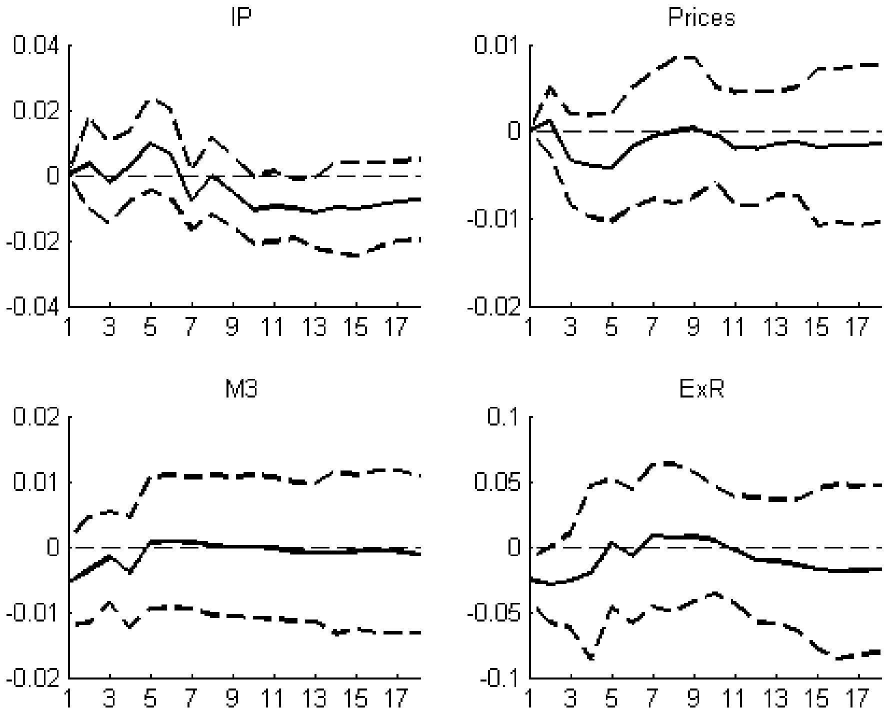


Figure E.2 plots the influence of the monetary policy shock based on the linear VAR on y_{t+j} for months $j = 0, 1, \ldots, 17$ for the respective variables of y. *Solid lines*: VAR impulses. *Dashed lines*: respective two-standard error bounds.

Figure E.3: Unexpected Shock in the Policy Interest Rate

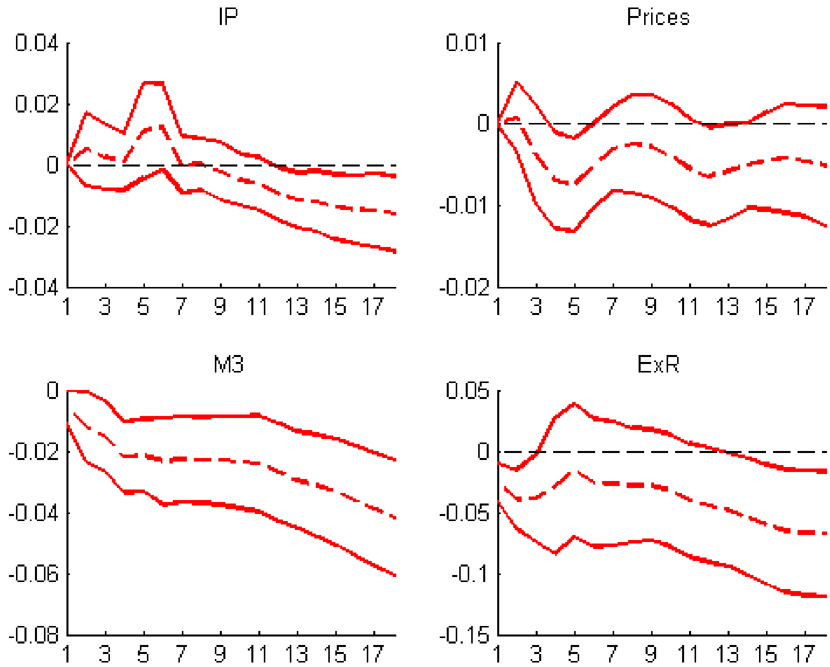


Figure E.3 plots the influence of the unexpected monetary policy shock on y_{t+j} for months $j = 0, 1, \ldots, 17$ for the respective variables of y. *Dashed lines*: impulses of the unexpected interest rate increase. *Solid lines*: respective two-standard error bounds.

Figure E.4: Expected, Not Implemented Shock in the Policy Interest Rate

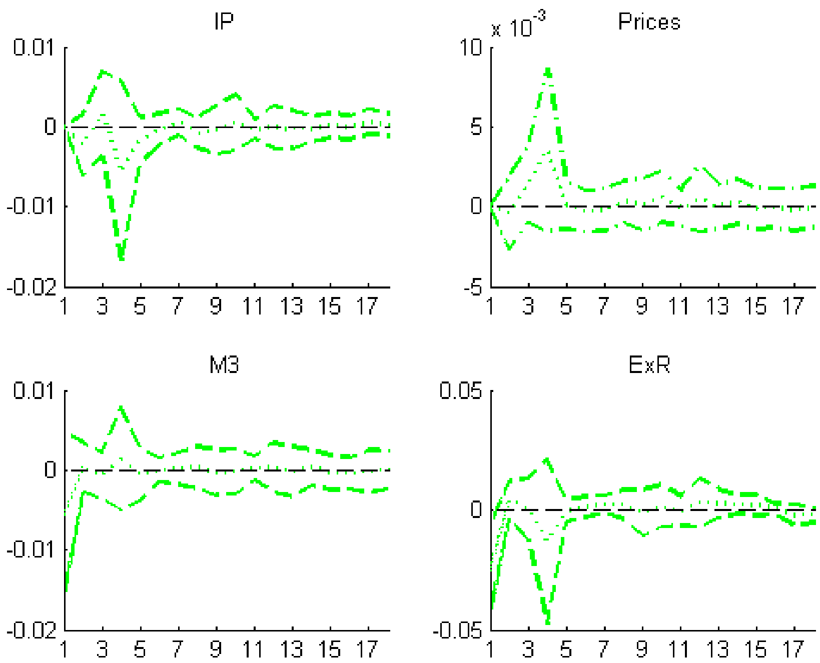


Figure E.4 plots the influence of the expected, not implemented monetary policy shock on y_{t+j} for months $j = 0, 1, \ldots, 17$ for the respective variables of y. *Dotted lines*: impulses of an expected interest rate decrease, not implemented. *Dashed line*: respective two-standard error bounds.

F Post Financial Crisis Monetary Policy Shocks

Figure F.5: VAR Responses to a Shock in the Policy Interest Rate

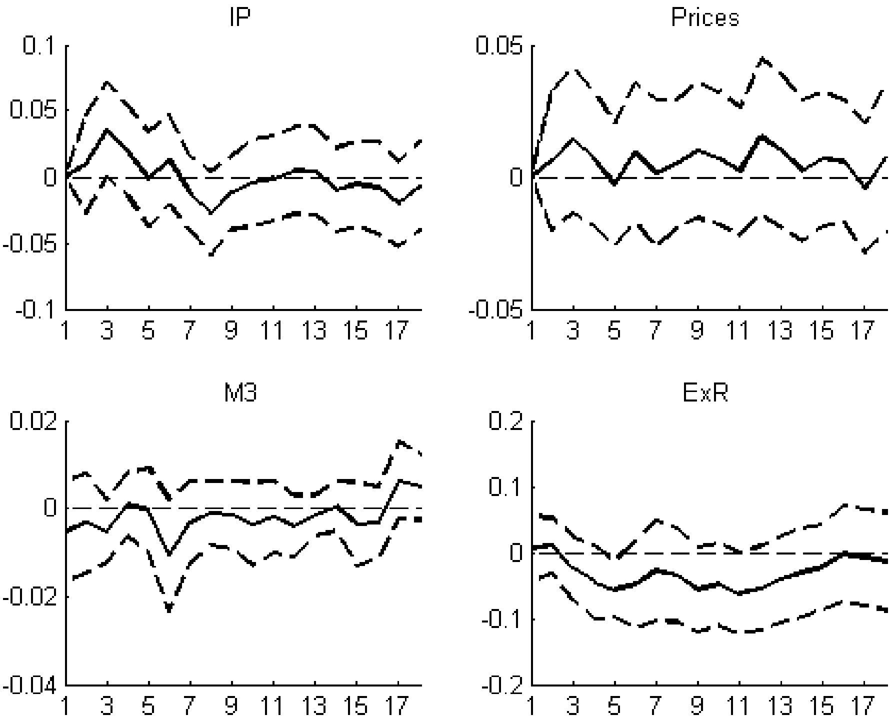


Figure F.5 plots the influence of the monetary policy shock based on the linear VAR on y_{t+j} for months $j = 0, 1, \ldots, 17$ for the respective variables of y. *Solid lines*: VAR impulses. *Dashed lines*: respective two-standard error bounds.

Figure F.6: Unexpected Shock in the Policy Interest Rate

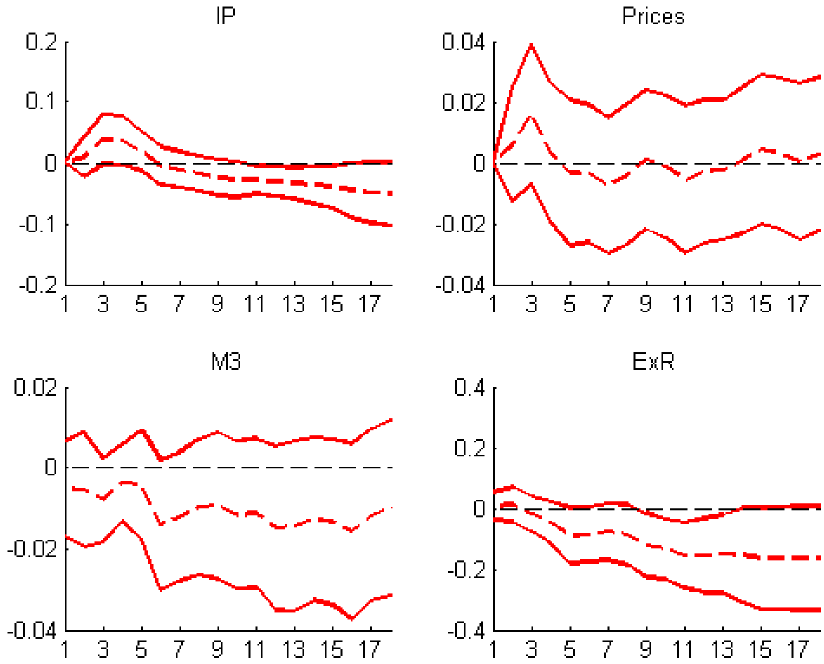


Figure F.6 plots the influence of the unexpected monetary policy shock on y_{t+j} for months $j = 0, 1, \ldots, 17$ for the respective variables of y. *Dashed lines*: impulses of the unexpected interest rate increase. *Solid lines*: respective two-standard error bounds.

Figure F.7: Expected, Not Implemented Shock in the Policy Interest Rate

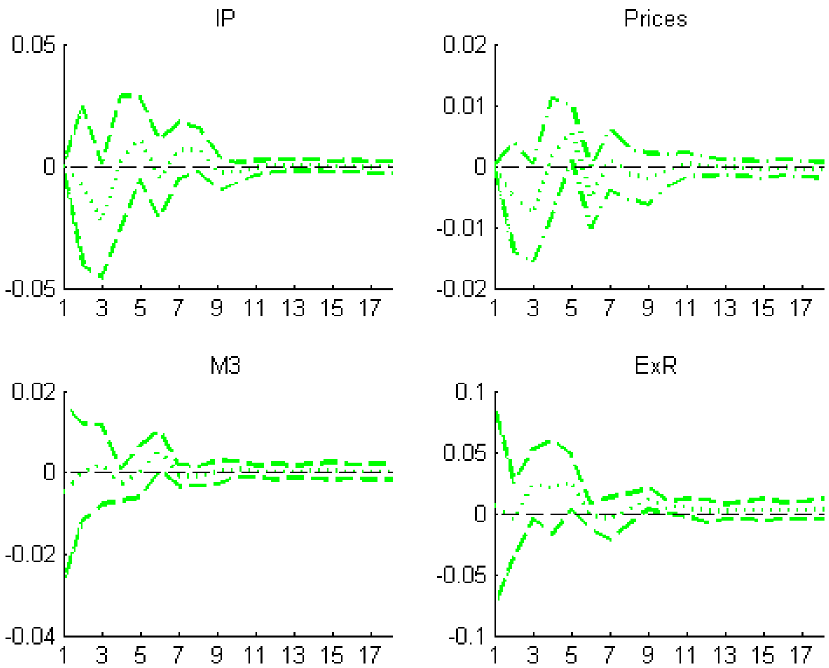


Figure F.7 plots the influence of the expected, not implemented monetary policy shock on y_{t+j} for months $j = 0, 1, \ldots, 17$ for the respective variables of y. *Dotted lines*: impulses of an expected interest rate decrease, not implemented. *Dashed lines*: respective two-standard error bounds.

F.1 Robustness Check with Quarterly Data - Post Financial Crisis Sample

Figure F.8: VAR Responses to a Shock in the Policy Interest Rate

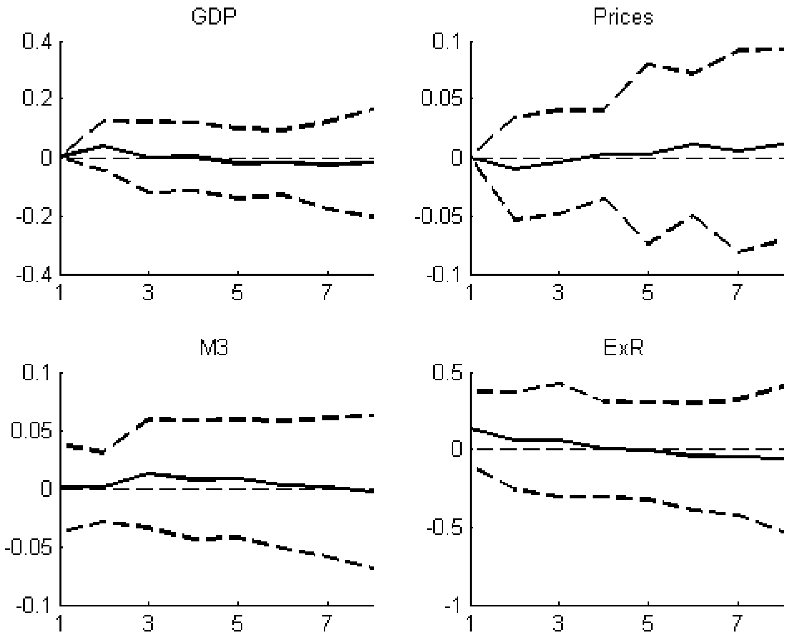


Figure F.8 plots the influence of the monetary policy shock based on the linear VAR on y_{t+j} for quarters $j = 0, 1, \ldots, 7$ for the respective variables of y based on quarterly data. *Solid lines*: VAR impulses. *Dashed lines*: respective two-standard error bounds.

Figure F.9: Unexpected Shock in the Policy Interest Rate

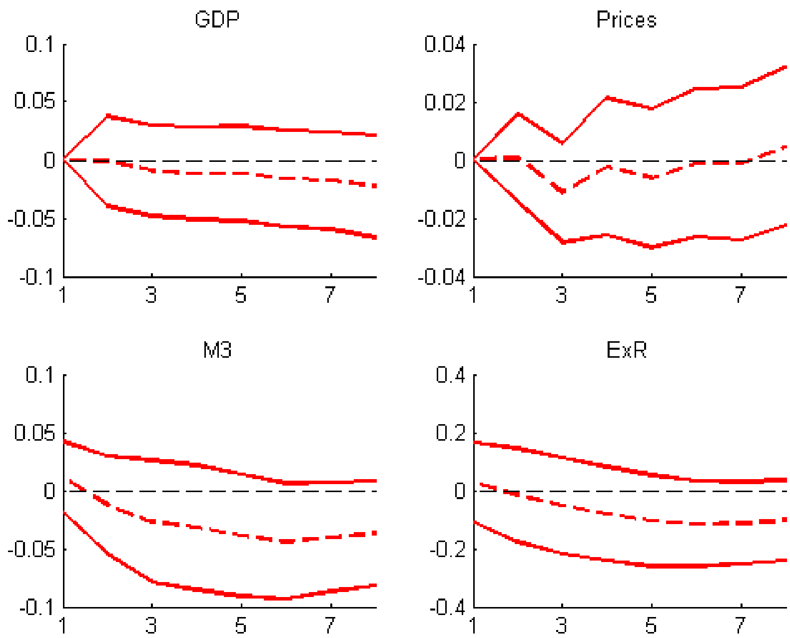
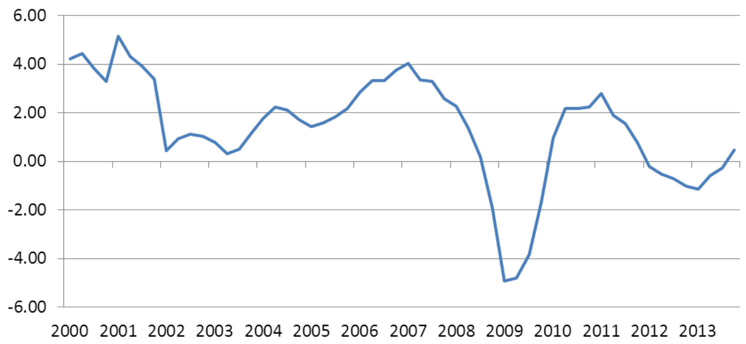


Figure F.9 plots the influence of the unexpected monetary policy shock on y_{t+j} for quarters $j = 0, 1, \ldots, 7$ for the respective variables of y based on quarterly data. *Dashed lines*: impulses of the unexpected interest rate increase. *Solid lines*: respective two-standard error bounds.

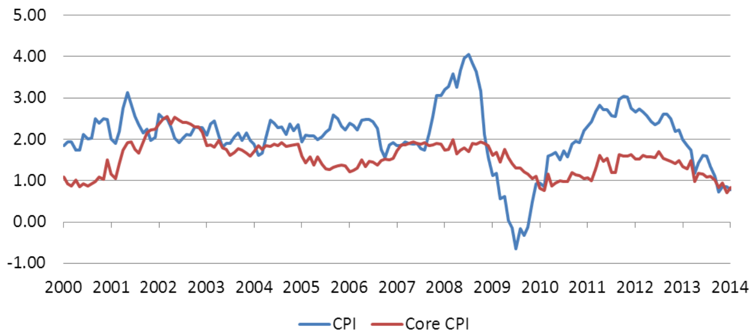
F.2 Development of Euro Area GDP and CPI

Figure F.10: Euro Area GDP



GDP and main components, at 2005 constant Prices, sa, source: Eurostat

Figure F.11: Euro Area Consumer Prices



CPI, all items, harmonised, not sa, source: Eurostat; CPI, excluding energy, food, alcohol and tobacco, not sa, source: Eurostat

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