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THE IMPACT OF MONETARY POLICY AND EXCHANGE RATE SHOCKS IN POLAND

EVIDENCE FROM A TIME-VARYING VAR

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In 2014 all ECB publications feature a motif taken from the €20 banknote.

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Abstract. This paper follows the Bayesian time-varying VAR approach with stochastic volatility developed by Primiceri (2005), to analyze whether the reaction of output and prices to interest rate and exchange rate shocks has changed across time (1996-2012) in the Polish economy. The empirical findings show that: (1) output appears more responsive to an interest rate shock at the beginning of our sample. Since 2000, absorbing this shock has become less costly in terms of output, notwithstanding some reversal since the beginning of the global financial crisis. The exchange rate shock also has a time-varying effect on output. From 1996 to 2000, output seems to decline, whereas for periods between 2000 and 2008 it has a positive significant effect. (2) Consumer prices appear more responsive to an interest rate shock during the first half of our sample, when Poland experienced high inflation. The impact of an exchange rate shock on prices seems to slightly decrease across time.

Keywords: Bayesian time-varying parameter VAR, monetary policy transmission, exchange rate pass-through

JEL classifications: C30, E44, E52, F41.

Non-technical Summary

In an increasingly globalised world, small open economies are exposed to exogenous monetary policy shocks triggered, for example, by an event occurring many miles away. They can be also exposed to exchange rate movements reflecting internal or external factors such as political instability or sudden changes in investors' sentiment. Understanding the impact of such monetary policy and exchange rate shocks is of key importance for central bankers. Although a large number of papers have analysed these effects, it is only recently that a flexible estimation framework allowing for time variation has been applied. Accounting for time variation is very important, especially for economies that have experienced significant structural and institutional changes over time.

This paper investigates whether there is significant time variation in the responses of the Polish economy to monetary policy and exchange rate shocks. Specifically, we estimate a time-varying VAR model with stochastic volatility and analyse the impact of exogenous monetary policy and exchange rate shocks which can be triggered by domestic or external events. We study their effects on real GDP, consumer prices, short-term interest rates and exchange rates, using quarterly data for the period between 1996Q1 and 2012Q3.

The evidence presented confirms the importance of allowing for a full time-varying framework. The empirical findings show that the response of macroeconomic variables to an interest and/or exchange rate shock do vary over time. Output seems more responsive to an interest rate shock at the beginning of our sample, where a 100 basis points monetary policy tightening has a cumulative output cost of about 1% after two years. Notwithstanding some reversal since the beginning of the global financial crisis, from 2000 to 2007, a monetary policy shock is less costly to absorb. Prices appear more responsive to an interest rate shock during the first half of the sample, when Poland experienced high inflation. With regard to an exchange rate shock (a 1% appreciation of the zloty), from 1996 to 2000 output seems to react negatively, whereas for the period between 2000 and 2008 such a shock has a positive significant effect. Thereafter, this effect on output mitigates. The response of prices to an exchange rate shock decreases slightly across time. The same shock is also estimated for import and producer prices. We find that the effect on import prices is larger than on consumer or producer prices, confirming a decrease along the pricing chain that has also been found in other studies.

Overall, our 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.

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 analyzed 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 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

[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 5d). 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`

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 organized as follows. A brief literature overview is given in section 2. Section 3 describes the econometric model and estimation strategy. Section 4 briefly summarizes the marginal likelihood results. The results of the TVP-VAR are presented in section 5. Section 6 summarizes our robustness checks and section 7 concludes.

2. General Survey of Related Literature

Since the seminal work by Sims (1972, 1980 and 1986)VAR models are widely used in multivariate time series analysis, especially for investigating both monetary policy and exchange rate shocks.

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 3.2).

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]

[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.

[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)).

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]

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

[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-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)).

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 prefered tool for analysing the effect of such shocks in Poland.

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:

$$y_t = c_t + B_{1,t} y_{t-1} + \ldots + B_{l,t} y_{t-l} + u_t, \tag{3.1}$$

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

$$VAR(u_t) \equiv \Omega_t = A_t^{-1} H_t (A_t^{-1})', \text{with } H_t = \Sigma_t \Sigma_t'. \tag{3.2}$$

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

$$\Sigma_t = \begin{bmatrix} \sigma_{1,t} & 0 & \ldots & 0 \\ 0 & \sigma_{2,t} & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ 0 & \ldots & 0 & \sigma_{n,t} \end{bmatrix} \quad A_t = \begin{bmatrix} 1 & 0 & \ldots & 0 \\ \alpha_{21,t} & 1 & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ \alpha_{n1,t} & \ldots & \alpha_{n(n-1),t} & 1 \end{bmatrix} \tag{3.3}$$

The time-varying VAR can then be summarized as:

$$y_t = X_t' \tilde{B}_t + A_t^{-1} \Sigma_t \varepsilon_t, \tag{3.4}$$

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 3.4 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:

$$B_t = B_{t-1} + \nu_t, \tag{3.5}$$

$$\alpha_t = \alpha_{t-1} + \xi_t, \tag{3.6}$$

$$log\sigma_t = log\sigma_{t-1} + \eta, \tag{3.7}$$

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:

$$\begin{bmatrix} \varepsilon_t \\ \nu_t \\ \xi_t \\ \eta_t \end{bmatrix} \sim N(0, V), \text{ with } V = \begin{bmatrix} I_n & 0 & 0 & 0 \\ 0 & Q & 0 & 0 \\ 0 & 0 & S & 0 \\ 0 & 0 & 0 & W \end{bmatrix}, \tag{3.8}$$

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.

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 3.5, 3.6 and

[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 6 for further details.

3.7). 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 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 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.

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).

[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 $k_S = \{0.01; 0.025; 0.1\}$.

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 1). We use zero and sign restrictions as follows[13]:

- 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 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 $\leq$.

Technically, the sign restrictions are implemented using the QR-decomposition method for finding Q. We have a four variable VAR, implying a 4 x 4 Q matrix.

[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.

Due to zeros in the first two rows of the sign restriction matrix, the decomposition is restricted to the last two columns. Thus, the Q matrix has the following form:

$$Q = \begin{pmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & \left[\begin{array}{cc} QR(\theta[1,1]) & QR(\theta[1,2]) \\ QR(\theta[2,1]) & QR(\theta[2,2]) \end{array} \right] \\ 0 & 0 & & \end{pmatrix}. \tag{3.9}$$

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 Q 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.

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 likehood 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 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.

5. Results of the TVP-VAR

In what follows, sections 5.1 and 5.2 present, respectively, the estimated median impulse responses of the monetary policy and exchange rate shocks

[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 $\tau = 0.99$.

(see also Appendix C) and include an analysis on the posterior probability for the difference in the impulse responses. In section 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 summarizes the estimation on convergence diagnostics.

5.1. **Impulse Responses to Monetary Policy Shocks.**

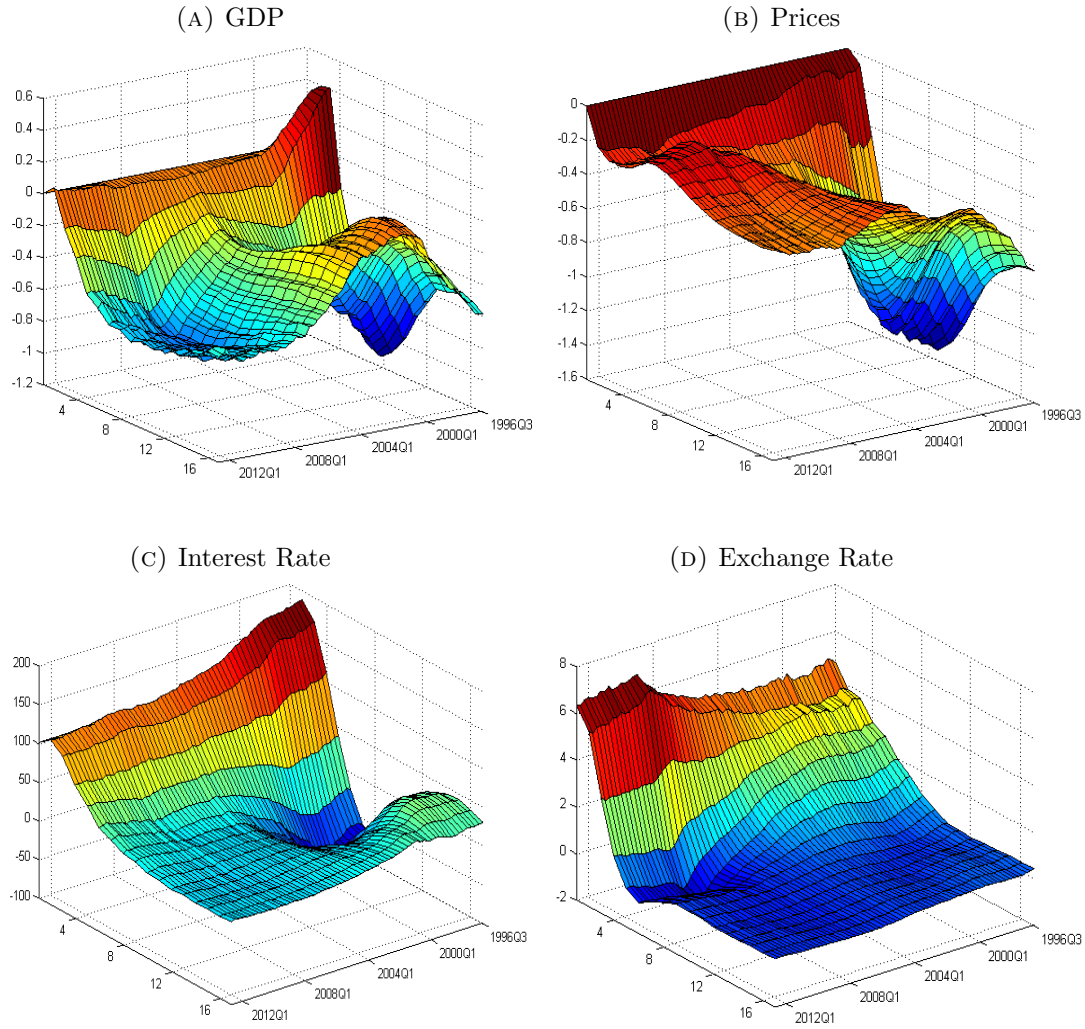
Figure 1 presents the median impulse responses (over 17 quarters and the time period: 1996:1-2012:3) to a 100 basis points (BPs) interest rate increase in the given period across the sample. This monetary policy shock has the expected impact on GDP (↓), prices (↓), interest rate (↑) and exchange rate (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 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 6a, 6b).

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 1b). The impact on prices seems to be strongest between 1996 and 2001, a period during which Poland experienced high inflation (Figure 5d). 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 1c). Interestingly, this effect has not changed since the beginning of the global financial crisis.

FIGURE 1. Time-Varying Impulse Responses to a Monetary Policy Shock



Median impulse responses to a 100 BPs monetary policy shock.

Finally, the impact of a monetary policy shock on the *nominal effective exchange rate* is, as expected, initially positive (Figure 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.

5.1.1. *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 6a and 6b 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 6a and 6b). The influence on the nominal effective exchange rate seems to be different across time (Figures 6a and 6b), converging more quickly towards zero from 2004 onwards.

TABLE 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

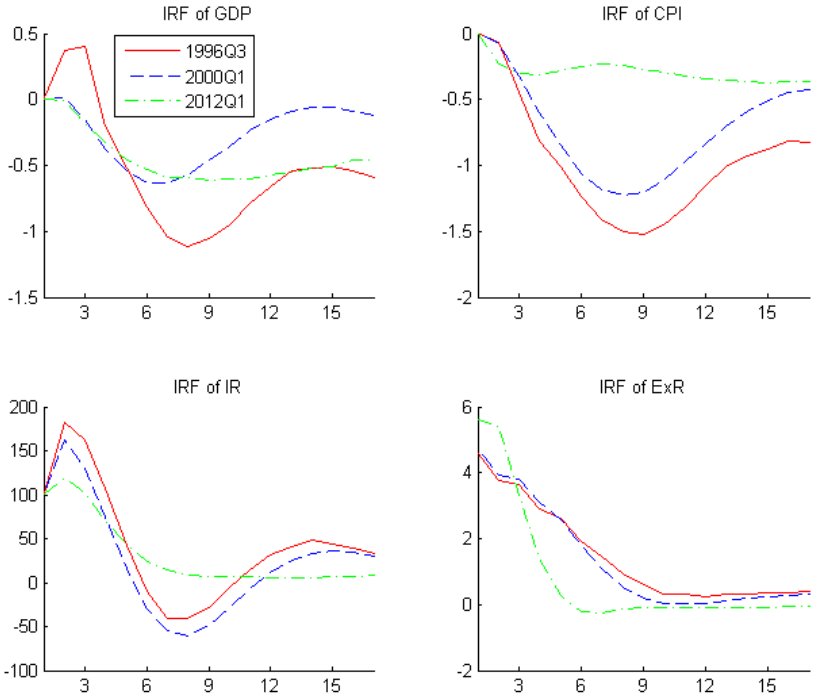
Note: Difference in impulse responses at the time periods 1996Q3, 2000Q1 and 2012Q1 for one, four, eight and 12 quarters ahead.

For a better illustration of the difference in the impulse responses to a monetary policy shock across time, we also present Figures 2a and 2b. These allow for a comparison of impulse responses at specific points in time. Figure 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 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 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. 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.

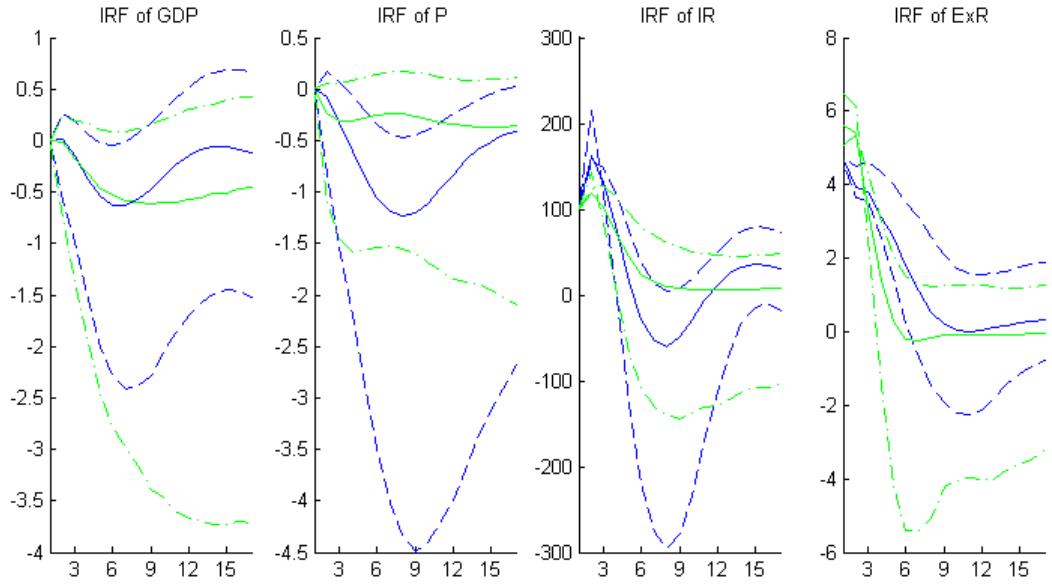
FIGURE 2. Responses at Different Time Periods to a Monetary Policy Shock

(A) Responses without Percentiles



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

(B) Responses with Percentiles



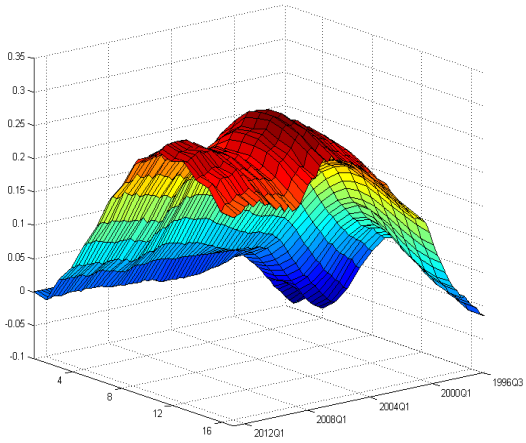
Median impulse responses (solid line) to a 100 BPs monetary policy shock with 16-th and 84-th percentiles (dashed line) of the posterior distribution at 2000Q1 (blue) and 2012Q1 (green).

5.2. **Impulse Responses to Exchange Rate Shocks.**

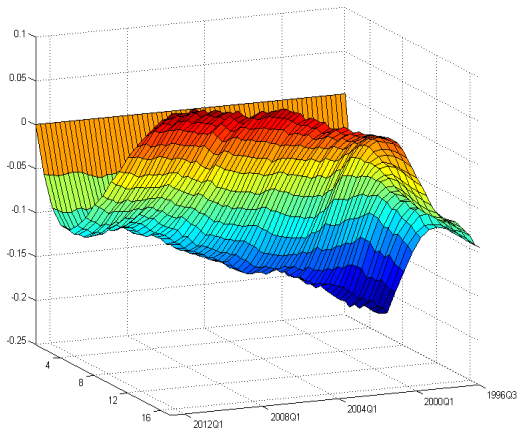
In this section, we analyse the median impulse responses to a 1% appreciation in the nominal effective exchange rate over time (17 quarters, time period: 1996:1-2012:3) (see Figure 3). The estimated pass-through of an exchange rate shock is in line with the theory and highlights the importance to account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

FIGURE 3. Time-Varying Impulse Responses to an Exchange Rate Shock

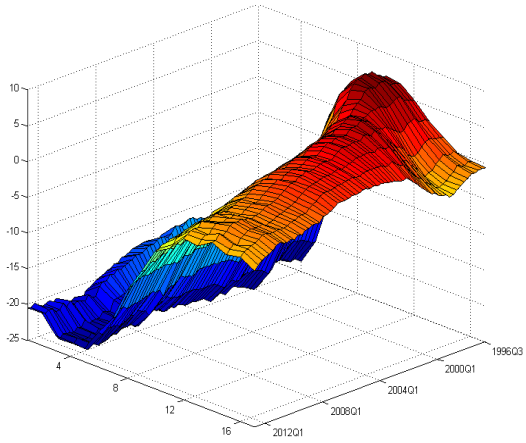
(A) GDP



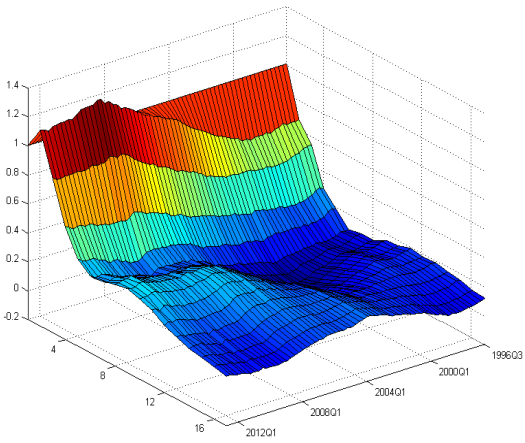
(B) Prices



(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 3a), albeit its effect seems to be insignificant (Figures 6c, 6d). 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 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 7, 8 and 9). 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 7b, 9a and 9b), whereas for producer prices, the pass-through appears to have increased since 2000 (Figure 8b). 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 by interest rate decreases (in the range of roughly -20 basis points within the first year, see Figure 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 3d).

5.2.1. *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

[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).

Appendix C, Figures 6c and 6d. Consumer prices seem to respond significantly negatively to an exchange rate shock, but the effect on GDP is only significant after two years between the period 2000 and 2008 (Figures 6c and 6d).

TABLE 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

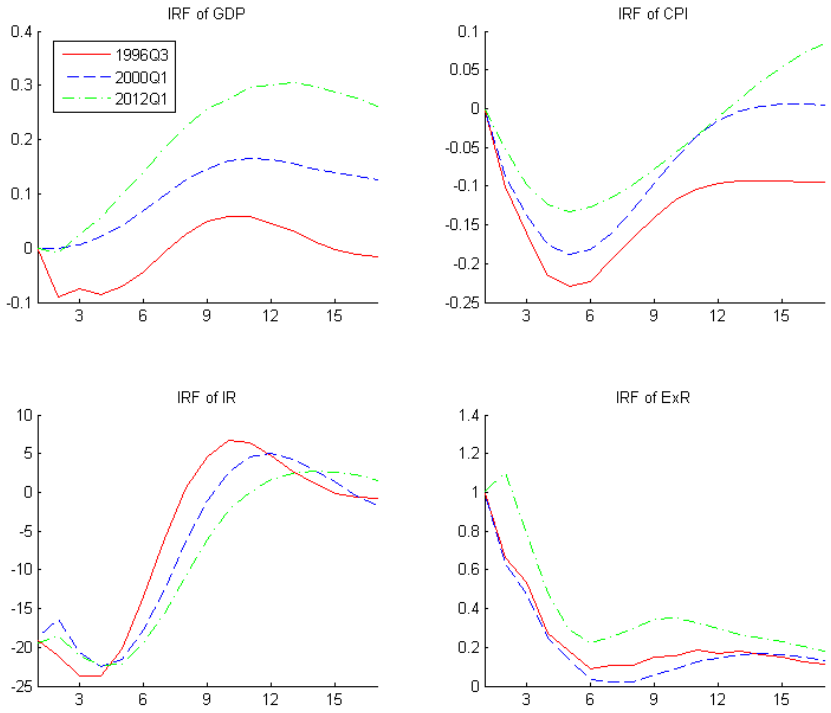
Note: Difference in impulse responses at the time periods 1996Q3, 2000Q1 and 2012Q1 for one, four, eight and 12 quarters ahead.

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 4a). Figure 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 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.

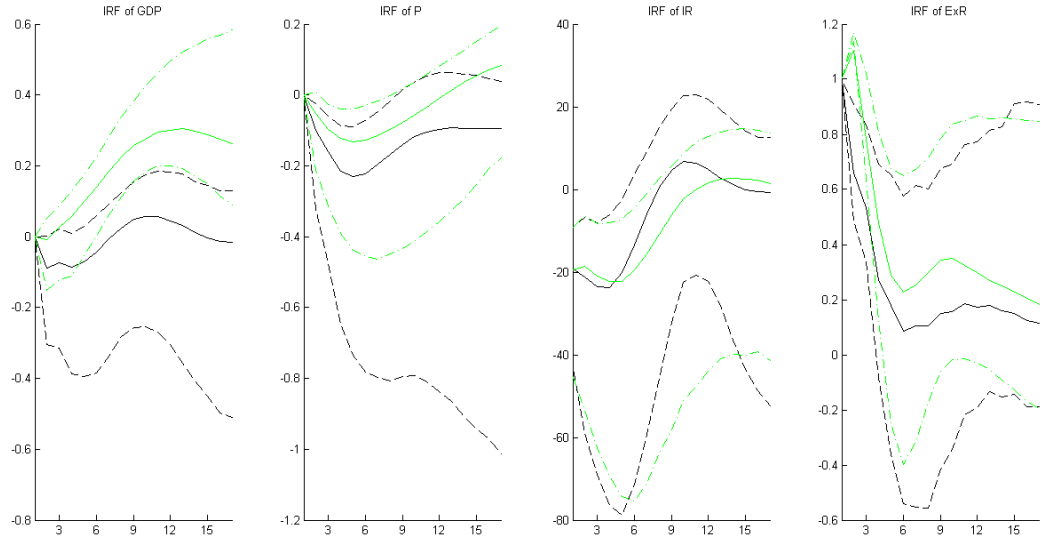
FIGURE 4. Responses at Different Time Periods to an Exchange Rate Shock

(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).

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]

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 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

[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.

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.

Appendix 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.

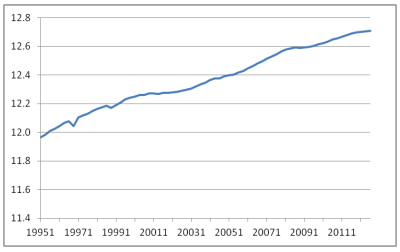
Exchange rate (ExR): Log of ECB nominal effective exchange rate, Euro area-17 countries vis-a-vis the EER-40 group of trading partners (AU, CA, DK, HK, JP, NO, SG, KR, SE, CH, GB, US, BG, CZ, LV, LT, HU, PL, RO, CN, DZ, AR, BR, CL, HR, IS, IN, ID, IL, MY, MX, MA, NZ, PH, RU, ZA, TW, TH, TR and VE) against 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 good and services, overall index, quarterly series (reference year 2000), in national currency, seasonally and working day adjusted. Source: Eurostat.

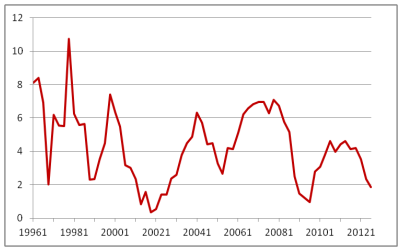
Producer price (ProdP): Log of industry producer prices, overall index, total output prices (industry [except construction, sewage, waste management and remediation activities]), quarterly series (reference year 2005), in national currency, gross data. Source: Eurostat.

FIGURE 5. Quarterly Data, Poland

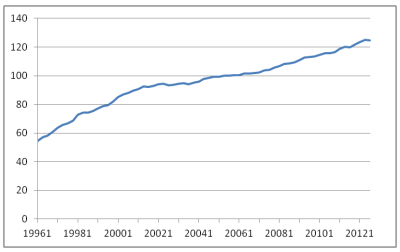
(A) GDP in levels



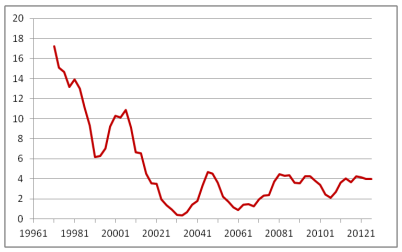
(B) GDP, yearly growth rate



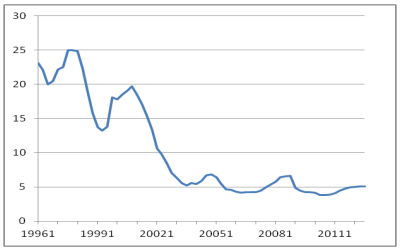
(C) CPI in levels



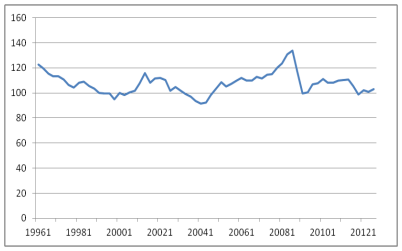
(D) CPI, annual rate of change



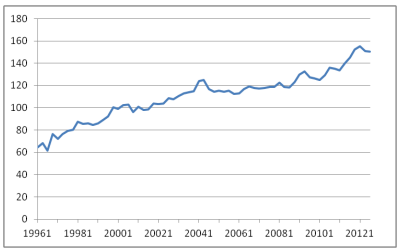
(E) IR in percent



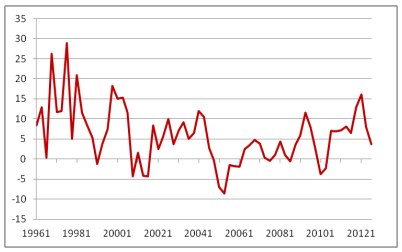
(F) ExR in levels



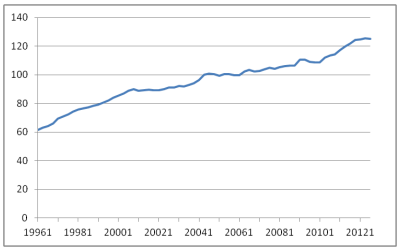
(G) Import Price Index in ln levels



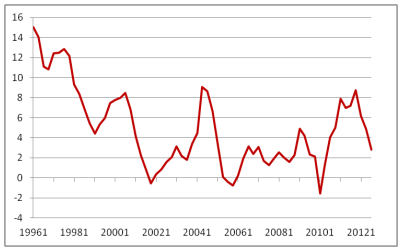
(H) Import Price Index, annual rate of change



(I) Producer Price Index in ln levels



(J) Producer Price Index, annual rate of change



Appendix B. Posterior Probability Estimates for k_Q, k_W and k_S

Table 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

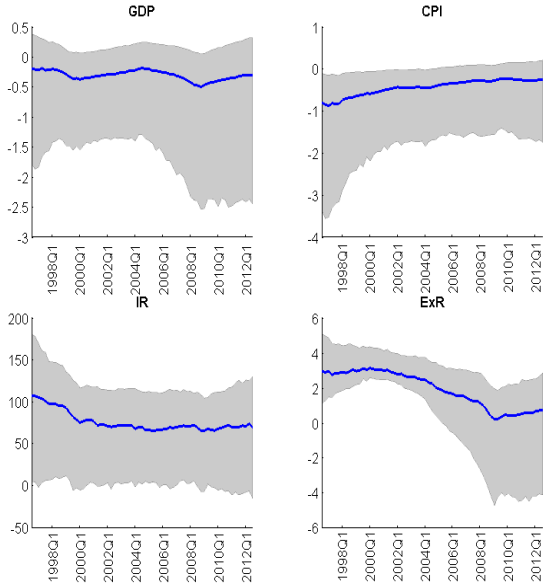
Note: Posterior probability estimates are based on the reversible jump Markov chain Monte Carlo method for the set of 18 models. These are constructed from all possible combinations of $k_Q = \{0.01; 0.05; 0.1\}$, $k_W = \{0.001; 0.01\}$ and $k_S = \{0.01; 0.025; 0.1\}$.

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

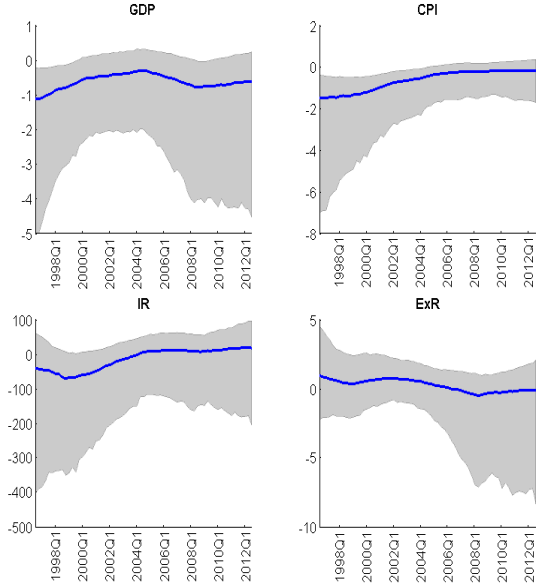
C.1. **Estimation at 4th and 8th Quarter with Consumer Prices.**

Figure 6. Impulse Responses at Different Horizons with Consumer Prices

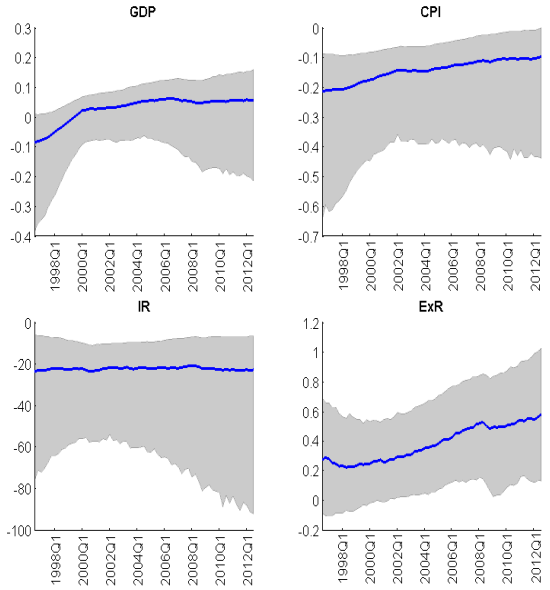
(A) Responses at the 4th quarter to a Monetary Policy Shock



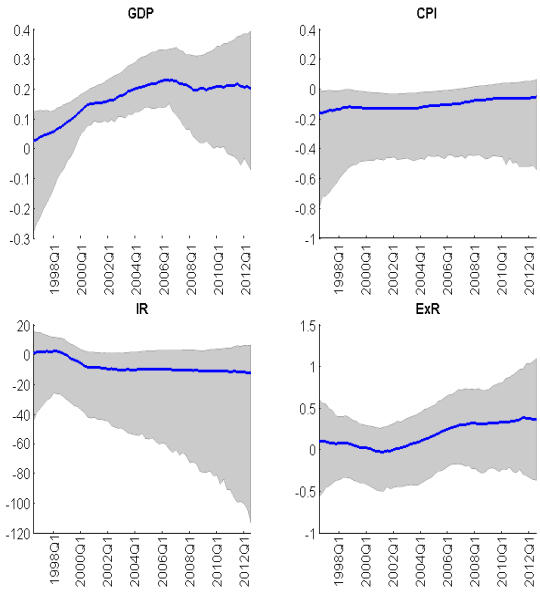
(B) Responses at the 8th quarter to a Monetary Policy Shock



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



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



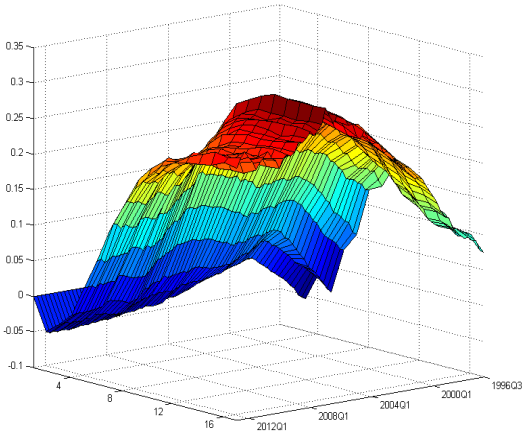
Median impulse responses (blue solid line) to a 100 BPs monetary policy shock or 1% exchange rate shock with 16-th and 84-th percentiles (grey area) of the posterior distribution of the responses at the 4th and 8th quarter, respectively.

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

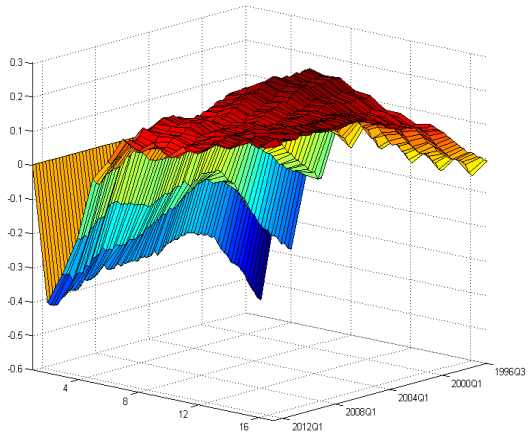
D.1. **Estimation of the Exchange Rate Shock with Import Prices.**

Figure 7. Time-Varying Impulse Responses to an Exchange Rate Shock with Import Prices

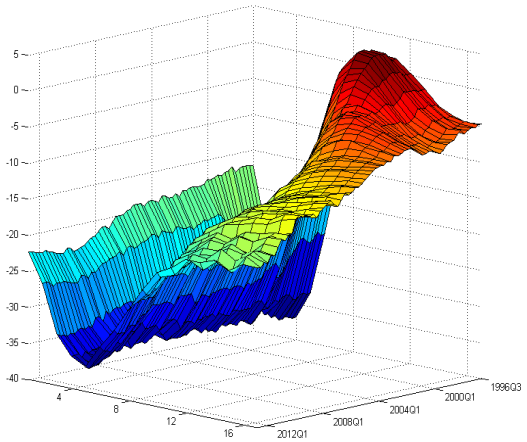
(a) GDP



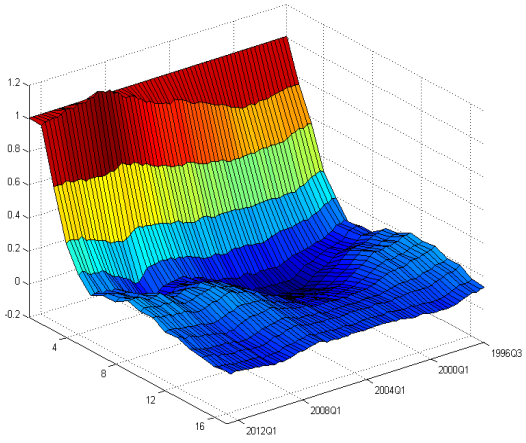
(b) Import Prices



(c) Interest Rate



(d) Exchange Rate

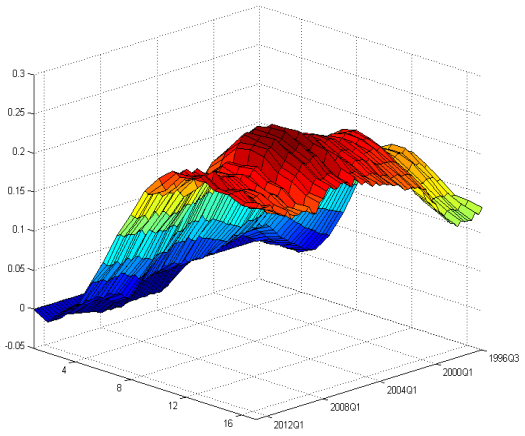


Median impulse responses to a 1% exchange rate appreciation.

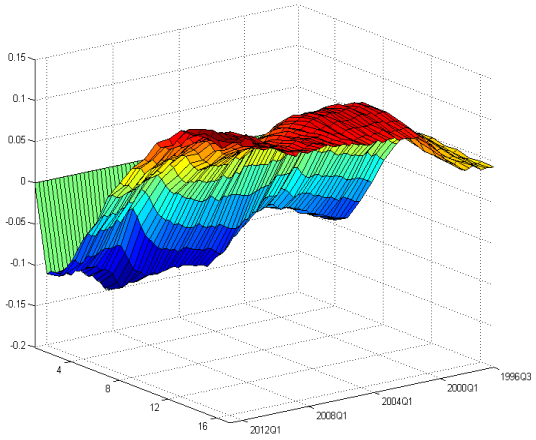
D.2. **Estimation of the Exchange Rate Shock with Producer Prices.**

FIGURE 8. Time-Varying Impulse Responses to an Exchange Rate Shock with Producer Prices

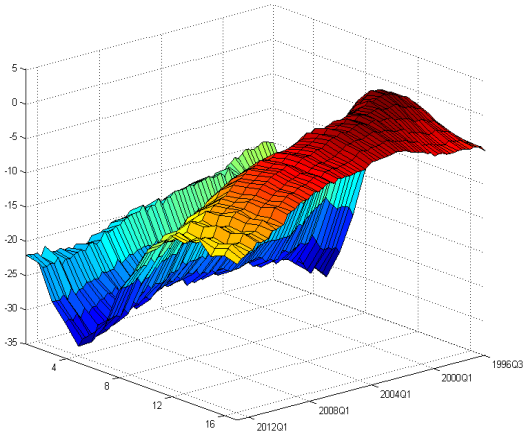
(A) GDP



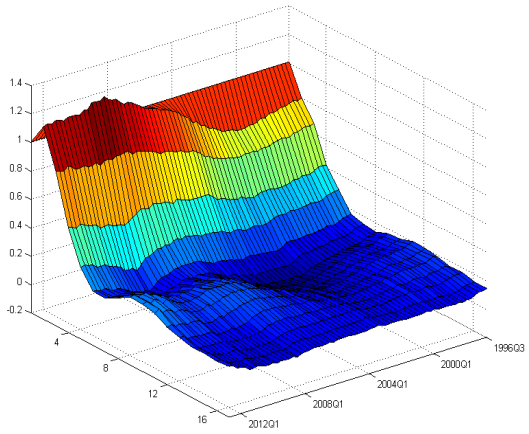
(B) Producer Prices



(C) Interest Rate



(D) Exchange Rate

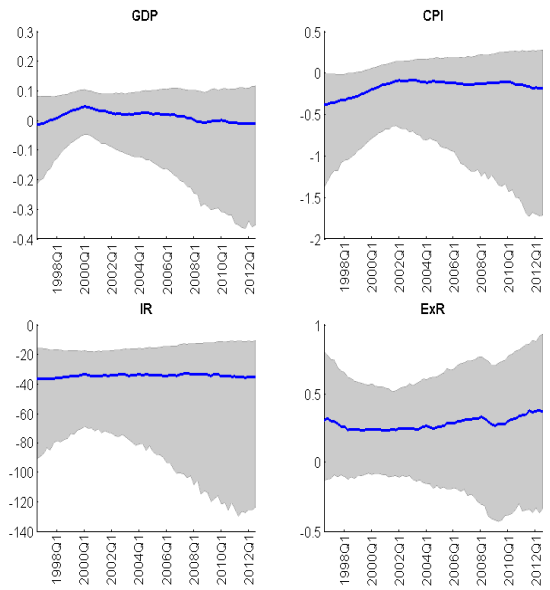


Median impulse responses to a 1% exchange rate appreciation.

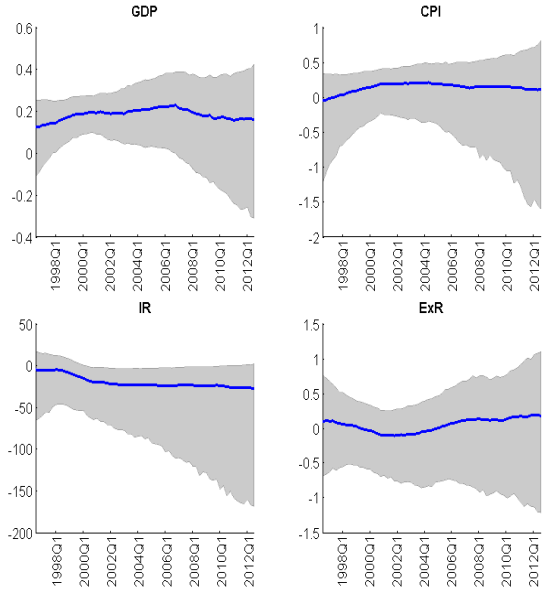
D.3. **Estimation at the Fourth and Eighth Quarter to an Exchange Rate Shock with Import and Producer Prices.**

FIGURE 9. Impulse Responses at the 4th and 8th Quarter to an Exchange Rate Shock

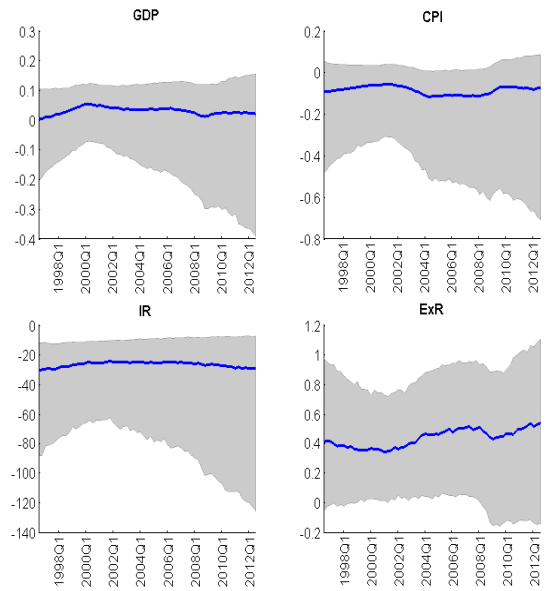
(A) Responses at the 4th quarter with Import Prices



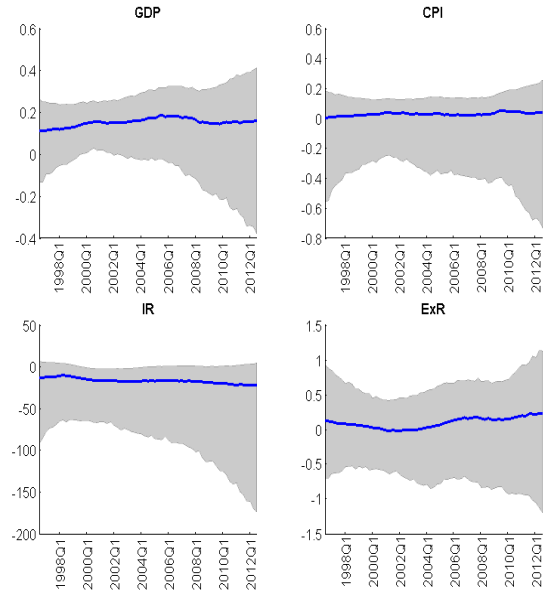
(B) Responses at the 8th quarter with Import Prices



(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.

Appendix E. Time-Varying Posterior Estimates of the Stochastic Covariance

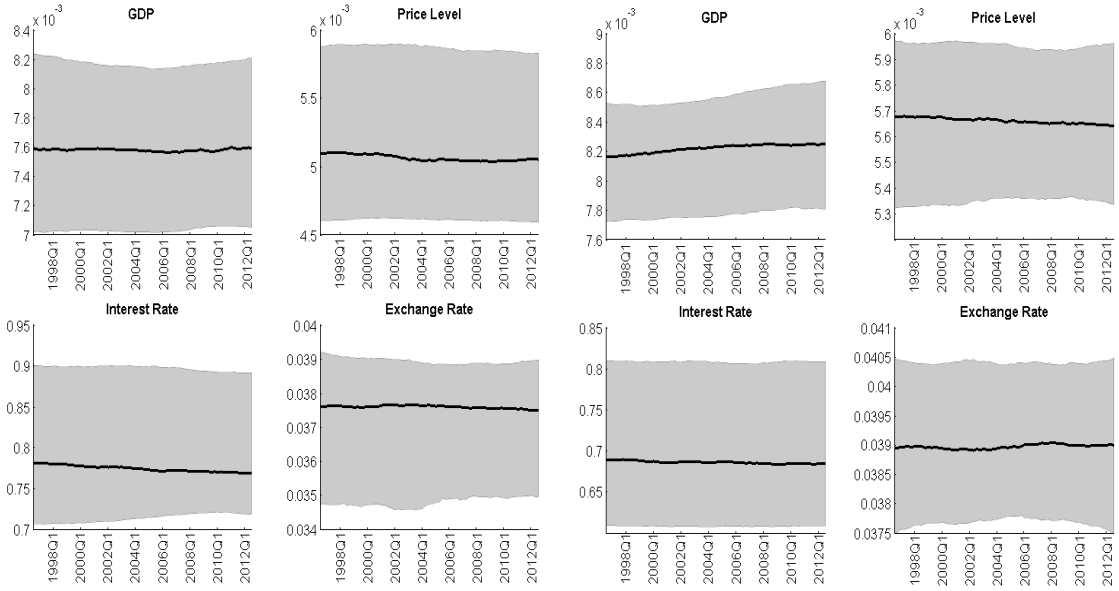
The stochastic covariance matrix of the residuals comprises two matrices. First, the time-varying diagonal matrix Σ_t which denotes the stochastic volatility of the structural shock. The second matrix, the time-varying lower triangular matrix A_t captures the size of the simultaneous impact on the other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

Concerning Σ_t, not much time variation is visible. Figure 10 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 11. The simultaneous effect on the interest rate of the price shock is clearly time varying.

Figure 10. Volatility of Structural Shock

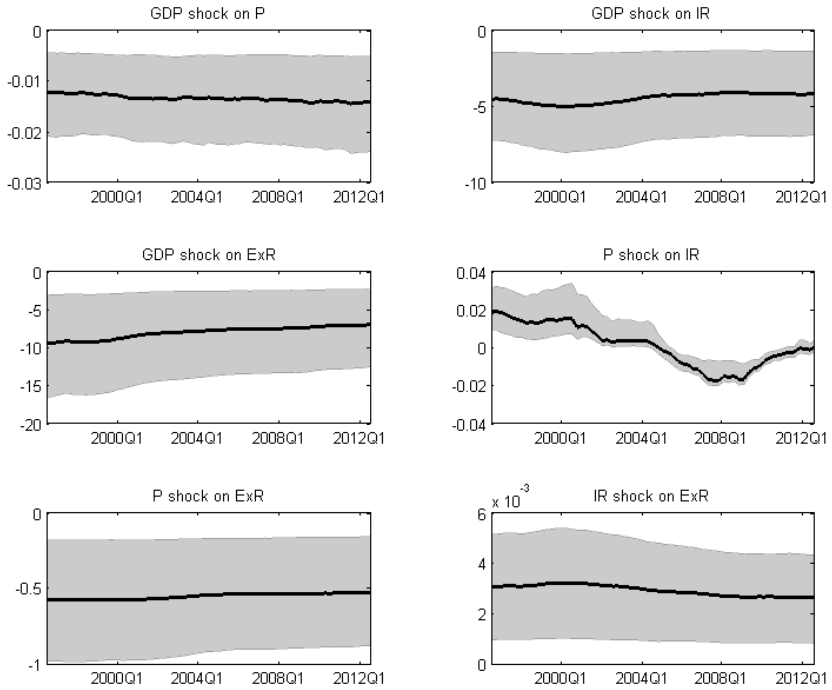
(a) Volatility of the Structural Shock to a Monetary Policy Shock

(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 11. 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).

Appendix 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).

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

The (a) panel of Figure 12 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 12, 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 scrutinize 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 12a and 12b). 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 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 12a and 12d, the suggested number of iterations is below the actual number used. Furthermore, the impulse responses are calculated with respect to normalized shocks, hence, the inefficiency problem should not matter (Franta et al. (2011)).

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

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.

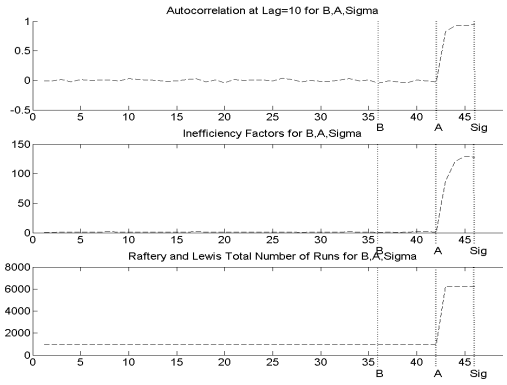
TABLE 5. Distribution of the Inefficiency Factors

	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

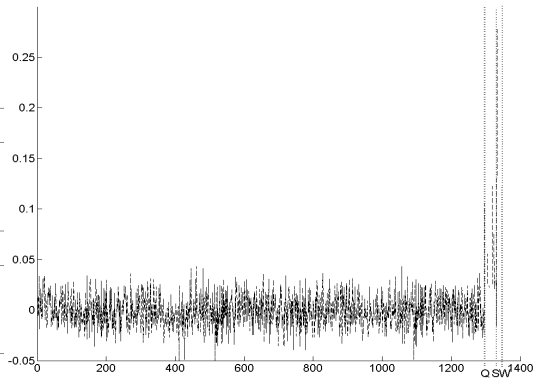
Overview of the inefficiency factors (IFs) for the posterior estimates of different sets of time varying parameters. A: time varying simultaneous relations; B: time varying coefficients; Σ: time varying volatilities.

FIGURE 12. Convergence Diagnostics

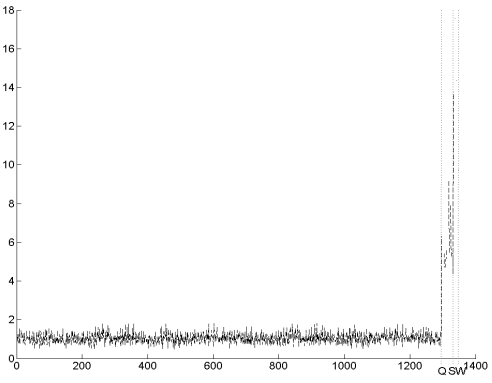
(A) Summary of Convergence Diagnostics for A, B and Σ



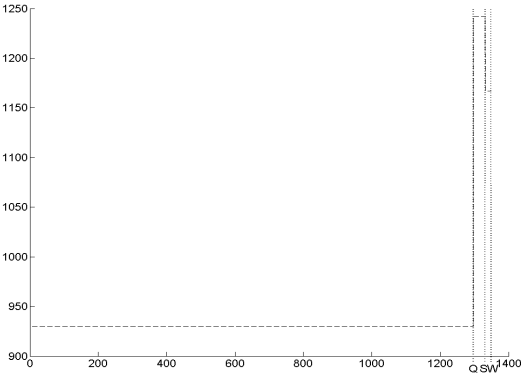
(B) 10-th-order Autocorrelation for Q, S and W



(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.

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