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Evaluating Changes in the Monetary Transmission Mechanism in the Czech Republic

Michal Franta, Roman Horváth, and Marek Rusnák*

Abstract

We investigate the evolution of the monetary policy transmission mechanism in the Czech Republic over the 1996–2010 period by employing a time-varying parameters Bayesian vector autoregression model with stochastic volatility. We evaluate whether the response of GDP and the price level to exchange rate or interest rate shocks changes over time, with a focus on the period of the recent financial crisis. Furthermore, we augment the estimated system with a lending rate and credit growth to shed light on the relative importance of financial shocks for the macroeconomic environment. Our results suggest that output and prices have become increasingly responsive to monetary policy shocks, probably reflecting financial sector deepening, more persistent monetary policy shocks, and overall economic development associated with disinflation. On the other hand, exchange rate pass-through has weakened somewhat over time, suggesting improved credibility of inflation targeting in the Czech Republic with anchored inflation expectations. We find that credit shocks had a more sizeable impact on output and prices during the period of bank restructuring with difficult access to credit. In general, our results show that financial shocks are less important for the aggregate economy in an environment of a stable financial system.

JEL Codes: E44, E52.
Keywords: Monetary policy transmission, sign restrictions, time-varying parameters.

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Nontechnical Summary

We examine the evolution of the monetary transmission mechanism in the Czech Republic. The Czech economy has witnessed many important changes in the past two decades, and it is likely that the nature of monetary transmission has evolved as well.

Using data for 1996–2010, we estimate the so-called Bayesian time-varying vector autoregression model with stochastic volatility for the Czech economy. The empirical model is flexible and is particularly suited to examining how the interactions of various economic variables change over time. Given the focus of this paper, we specifically examine the evolution of monetary policy effects on prices and GDP. For example, the model can shed light on the question of whether the responses of prices to monetary policy is stronger nowadays, as compared to the period right after the inflation targeting regime was adopted (e.g. in 1998). In addition, the model explicitly accounts for the fact that the volatility of the economic environment can evolve over time, too. This is an important feature that makes the model realistically account, for example, for the higher uncertainty during the current global financial crisis.

Our model consists of both macroeconomic and financial variables. As a consequence, we can examine the degree of interaction between the macroeconomic and financial sectors as well as its evolution over time. First, our results suggest that output and prices have become increasingly responsive to monetary policy shocks, probably reflecting financial sector deepening and the overall economic development of the Czech economy. The responsiveness of output and prices to monetary shocks did not increase further during the crisis, but remained largely constant at the pre-crisis level. Therefore, we do not find evidence that the effect of monetary policy on the aggregate economy declined during the crisis. We find that exchange rate pass-through has weakened somewhat over time. This is probably associated with improved credibility of inflation targeting in the Czech Republic with anchored inflation expectations.

Next, our results show that the effect of credit growth on the aggregate economy is stronger in an environment of a less stable financial system. More specifically, we find that the effect of credit on GDP and prices is more sizeable around the year 2000, when the Czech banking sector was in the process of restructuring. The effect of the credit shock is not stronger during the current financial crisis. In this regard, it is vital to note that the Czech financial system remained largely stable during the crisis, with a well capitalized and liquid banking sector. This supports our finding that a negative financial shock (of identical magnitude) is more harmful to the economy when the financial system is not stable than otherwise.

1. Introduction

The transition to a market-based economy, the deepening of trade spurred by trade liberalization and integration into the European Union, the consolidation of the banking industry at the turn of the century, and finally, the recent financial crisis are all reasons to believe that the structure of the Czech economy has been changing over time. These changes in the structure of the economy are likely to have had an impact on the monetary transmission mechanism, i.e., the effects of monetary policy on the aggregate economy (Cogley and Sargent, 2005). Furthermore, changes in the conduct of monetary policy caused by the introduction of an inflation targeting regime have probably influenced the strength of monetary policy as well. The inflation targeting regime was adopted in the Czech Republic as a disinflation strategy. It might well be the case that the transmission mechanism was different at the beginning of the new regime in 1998 than several years later, after inflation had fallen to levels consistent with price stability, thus anchoring inflation expectations (Holub and Hurnik, 2008).

Against this background, it is somewhat surprising that the evidence about changes in the monetary transmission mechanism in the Czech Republic is rather scarce. Moreover, although it is of utmost importance to policymakers to know the strength of monetary transmission at times of crisis, strikingly, the effects of monetary policy actions on the economy during this period have not been investigated comprehensively so far. The contribution of our paper is to provide the stylized facts about changes in the strength of monetary policy actions over time, especially during the recent crisis. We aim to investigate the qualitative as well as quantitative implications of these changes by estimating a recently developed time-varying Bayesian vector autoregression model – TVP BVAR (Primiceri, 2005). In addition, the recent global financial crisis has reminded us of the critical role the financial sector plays in macroeconomic fluctuations. For this reason, we augment our baseline macroeconomic TVP BVAR model with several financial variables to assess the importance of financial shocks over time.

Our results suggest an increasing responsiveness of output and prices to monetary policy shocks. We attribute this finding to financial sector deepening and more persistent monetary policy shocks as well as to overall economic development and disinflation. Inflation targeting started as a disinflation strategy in 1998 with a nearly double-digit inflation rate. Over the years the target has been reduced to 2%, i.e., a value typically considered as being in line with price stability. In this regard, we find that exchange rate pass-through has weakened over time. This is probably related to the improved credibility of inflation targeting in the Czech Republic and anchored inflation expectations.

We find that credit shocks had a more sizeable impact on output and prices during the period of bank restructuring in about the year 2000. This period was characterized by higher non-performing loans inherited from the transition toward a market-oriented economy in the 1990s and rather difficult access to credit. Therefore, our results imply that financial shocks are less important for the evolution of the aggregate economy in an environment of financial stability.

The paper is organized as follows. The related literature is discussed in Section 2. Section 3 introduces our econometric model. Section 4 gives the results. Concluding remarks are offered in Section 5. Appendix A with additional results follows.

2. Related (Time-Varying) VAR Literature

2.1 VAR Models and the Price Puzzle

Ever since the seminal contribution by Sims (1980) the vector autoregression model has been the major tool for investigating the monetary policy transmission mechanism. The stylized facts about the monetary transmission mechanism for the US economy are summarized in an authoritative survey by Christiano et al. (1999). They conclude that following a contractionary monetary policy shock economic activity declines quickly in a hump-shaped manner, while the negative reaction of the price level is more delayed and persistent. Similarly, Peersman and Smets (2001) provide evidence for the euro area as a whole, while Mojon and Peersman (2001) investigate the effects of monetary policy shocks in the individual countries of the euro area.

Many of the early results, however, were plagued by a counterintuitive finding that the price level increases following a monetary policy tightening. This observation was noted by Sims (1992) and named the price puzzle by Eichenbaum (1992). The solution initially proposed to alleviate the price puzzle was to add commodity prices into the system (Sims, 1992; Christiano et al., 1999). On the other hand, Giordani (2004) stresses the importance of including a measure of potential output in the VAR. A different approach is pursued by Bernanke et al. (2005), who point out that central banks look at practically hundreds of time series and therefore, in order to avoid omitted variables bias and ensure correct identification of monetary policy shocks, an econometrician should use a richer dataset as well. Because the inclusion of other variables in the VAR is limited due to degrees of freedom considerations, they make use of factor analysis and augment the standard VAR with factors approximated by principal components. Other solutions, especially in an open economy framework, make use of alternative identification strategies such as non-recursive identification (Kim and Roubini, 2000; Sims and Zha, 2006) or identification by sign restrictions (Canova and Nicolo, 2002; Uhlig, 2005). Finally, some studies put forward that the price puzzle is limited to studies that do not estimate the transmission mechanism over a single monetary policy regime (Elbourne and de Haan, 2006; Borys et al., 2009; Castelnuovo and Surico, 2010). Note that we address these issues by looking at the time-specific impulse responses, which are identified using sign restrictions (more on this below).

2.2 Time-Varying VARs

It has long been recognized that the structure and functioning of the economy changes over time, and so there is a need to account for that evolution in the estimation procedure as well (Koop et al., 2009). Two main approaches to modeling changes in the transmission mechanism have appeared in applied work. First, the sample can be split and the model estimated over subsamples. Second, we can directly model the change of coefficients within the system (e.g. using the structural break or random walk assumption). For example, the former approach of splitting the sample into subsamples to investigate changes in the monetary policy rule was employed by Clarida et al. (2000). However, the date on which the sample should be split is often not clear. Importantly, it is more likely that the economy is changing gradually as opposed to undergoing sudden abrupt changes (Koop et al., 2009). As a result, a second vein of literature typically makes less restrictive assumptions about the behavior of the economy. It typically uses time-varying coefficients models by employing the Kalman smoother for the full sample, as opposed to time-invariant estimation procedures, which use only the information contained in the relevant subsample.

Furthermore, even within explicit modeling of the evolution of parameters over time, there are several different approaches that can be used. For example, Stock and Watson (1996) estimate a model with a small number of structural breaks. Alternatively, the Markov switching VAR model as employed by Sims and Zha (2006) might be considered. However, time-varying parameters VAR models have gained popularity recently. The reason for this popularity lies in the flexibility of this approach. For example, the system does not have to jump from one regime to another, as is often the case with Markov switching VAR models.

The modeling of time variation using the random walk assumption in multivariate models goes back to Canova (1993). More recently, Cogley and Sargent (2001) estimate time-varying parameters vector autoregressions (TVP VAR) with constant volatility of shocks to contribute to the discussion about the "bad policy" versus "bad luck" literature originated by Clarida et al. (2000). The limitation of the Cogley and Sargent model is the constant volatility assumption, which neglects possible heteroskedasticity of shocks and any nonlinearities in the relations among the variables of the model. Consequently, Cogley and Sargent (2005) allow for time-varying variance, although the simultaneous relations among the variables (covariances) are still modeled as time invariant. As was later pointed out by Primiceri (2005), this limits their analysis to reduced-form models (usable for data description and forecasting), and prevents any structural interpretation. To reconcile this issue, Primiceri (2005) stresses the importance of allowing for time variation in the variance-covariance matrix of innovations and estimates the TVP VAR model with stochastic volatility.

Recently, TVP VAR has been used widely to study changes in the transmission of various phenomena, such as monetary policy (Canova et al., 2007; Benati and Surico, 2008; Baumeister and Benati, 2010), fiscal policy (Kirchner et al., 2010; Pereira and Lopes, 2010), financial shocks (Eickmeier et al., 2011), oil price shocks (Baumeister and Peersman, 2008; Shioji and Uchino, 2010), yield curve dynamics (Mumtaz and Surico, 2009; Bianchi et al., 2009), and exchange rate dynamics (Mumtaz and Sunder-Plassmann, 2010).

2.3 Evidence for the Czech Republic

As far as modeling of monetary transmission in the Czech economy is concerned, Borys et al. (2009) use a battery of VAR models and identification strategies to show that the monetary transmission mechanism works relatively well when estimated on a single monetary policy regime. Havranek et al. (2011) employ a block-restriction VAR model and examine the interactions of macroeconomic conditions and the financial sector. They find that monetary policy has a systematic effect on financial stability and that some financial variables improve the forecasts of inflation and economic activity. Darvas (2009) is the first to estimate Czech monetary transmission in a time-varying framework in 1993–2008. He finds that the nature of monetary transmission mostly does not change over time (although the output response was somewhat stronger in 2008 than in 1996). However, his model does not account for changes in the variance of shocks (such as those in 1997, i.e., a period of exchange rate turbulence in the Czech Republic, or in the recent 2008–2009 financial crisis). To account for this possibly important feature, we estimate the Bayesian time-varying parameter model with stochastic volatility. As discussed above, neglecting heteroskedasticity of shocks might confound changes in the magnitude of shocks with changes in the transmission mechanism, thus yielding inconsistent estimates. In addition, we consider the effects of exchange rate shocks and financial shocks on macroeconomic fluctuations.

3. TVP BVAR

3.1 The Model

Following Primiceri (2005), we set up the model

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

where y_t is an $n \times 1$ vector of endogenous variables which are observable, c_t is an $n \times 1$ vector of time-varying intercepts, $B_{i,t}$, $i = 1, \ldots, p$, are $n \times n$ matrices of time-varying VAR coefficients, and u_t are unobservable shocks with time-varying variance-covariance matrix Ω_t for $t = 1, \ldots, T$.

Since it has been recognized recently that it is of paramount importance to allow not only the coefficients, but also both the error variances and the covariances to vary over time, we will use a triangular reduction of Ω_t, such that

$$A_t \Omega_t A_t' = \Sigma_t \Sigma_t' \tag{2}$$

or

$$\Omega_t = A_t^{-1} \Sigma_t \Sigma_t' (A_t^{-1})', \tag{3}$$

where A_t is the lower triangular matrix

$$A_t = \begin{bmatrix} 1 & 0 & \ldots & 0 \\ \alpha_{21,t} & 1 & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ a_{n1,t} & \ldots & a_{n(n-1),t} & 1 \end{bmatrix} \tag{4}$$

and Σ_t is the diagonal matrix

$$\Sigma_t = \begin{bmatrix} \sigma_{1,t} & 0 & \ldots & 0 \\ 0 & \sigma_{2,t} & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ 0 & \ldots & 0 & \sigma_{n,t} \end{bmatrix}. \tag{5}$$

Thus we have

$$y_t = c_t + B_{1,t}y_{t-1} + \ldots + B_{p,t}y_{t-p} + A_t^{-1}\Sigma_t \varepsilon_t, \tag{6}$$

where ε_t are independent identically distributed errors with $var(\varepsilon_t) = I_n$.

We rewrite (6), by stacking all the right-hand-side coefficients in a vector B_t, to obtain

$$y_t = X'B_t + A_t^{-1}\Sigma_t \varepsilon_t, \tag{7}$$

where $X' = I_n \otimes [1, y'_{t-1}, \ldots, y'_{t-p}]$.[1]

Next, we need to specify the law of motion for the parameters of the model. The VAR coefficients B_t and the elements of A_t are assumed to follow a random walk, while for the variance of shocks Σ_t we will use a stochastic volatility framework and assume that its elements follow a geometric random walk. Formally, the dynamics of the parameters are specified as follows:

$$B_t = B_{t-1} + \nu_t \tag{8}$$

$$\alpha_t = \alpha_{t-1} + \zeta_t \tag{9}$$

$$log\ \sigma_t = log\ \sigma_{t-1} + \eta_t. \tag{10}$$

Note that our model is, in fact, a state space model with equation (7) as the measurement equation and the state equations defined by (8), (9), and (10).

The innovations (ε_t, ν_t, ζ_t, η_t) are assumed to be jointly normal with the variance covariance matrix

$$V = var\left(\begin{bmatrix} \varepsilon_t \\ \nu_t \\ \zeta_t \\ \eta_t \end{bmatrix}\right) = \begin{bmatrix} I_n & 0 & 0 & 0 \\ 0 & Q & 0 & 0 \\ 0 & 0 & S & 0 \\ 0 & 0 & 0 & W \end{bmatrix}, \tag{11}$$

where I_n is an n-dimensional identity matrix and Q, S, and W are positive definite matrices.[2]

3.2 Priors

In this section, we specify the prior distributions for the parameters of the model. The mean and the variance of B_0 are chosen to be the OLS point estimate and four times its variance from the time-invariant VAR:

$$B_0 \sim N(\widehat{B}_{OLS}, 4 \cdot var(\widehat{B}_{OLS})).$$

The prior for A_0 is obtained similarly:

$$A_0 \sim N(\widehat{A}_{OLS}, 4 \cdot var(\widehat{A}_{OLS})).$$

Next, for $log\ \sigma_0$ the mean of the prior distribution is set to be the logarithm of the OLS estimate of the standard errors from the same time-invariant VAR, and the variance-covariance matrix is arbitrarily chosen to be proportional to the identity matrix:

$$log\ \sigma_0 \sim N(log\ \widehat{\sigma}_{OLS}, 4I_n).$$

[1] Symbol $\otimes$ denotes the Kronecker product.

[2] We assume S to be block diagonal, i.e., that the contemporaneous relationships among the variables evolve independently. For example, there are three blocks of S in the VAR consisting of four variables.

Finally, the priors for the hyperparameters are set as follows:

$$Q \sim IW(k_Q^2 \cdot \tau \cdot var(\widehat{B}_{OLS}), \tau),$$
$$W \sim IG(k_W^2 \cdot (1 + dim(W)) \cdot I_n, (1 + dim(W))),$$
$$S_l \sim IW(k_S^2 \cdot (1 + dim(S_l)) \cdot var(\widehat{A}_{l,OLS}), (1 + dim(S_l)),$$

where τ is the size of the training sample, S_l denotes the corresponding blocks of S, while $\widehat{A}_{l,OLS}$ stand for the corresponding blocks of $\widehat{A}_{OLS}$. The parameters k_Q, k_W, and k_S are specified below. The degrees of freedom of the scale matrices for the inverse-Gamma prior distribution of the hyperparameters are set to be one plus the dimension of each matrix. Moreover, following the literature (Cogley and Sargent, 2001) the scale matrices are chosen to be constant fractions of the variances of the corresponding OLS estimates on the training sample multiplied by the degrees of freedom.

3.3 Identification and Structural Interpretation

While in a closed economy the recursive identification scheme seems to be plausible for identifying the effects of monetary policy shocks (Christiano et al., 1999), in an open economy setting such identification might confound monetary policy shocks with exchange rate shocks (Kim and Roubini, 2000). Our identification strategy largely follows Jarocinski (2010) and combines zero restrictions and sign restrictions (Canova and Nicolo, 2002; Uhlig, 2005; Rubio-Ramirez et al., 2010). We assume that output and prices do not respond contemporaneously to monetary policy and exchange rate shocks. We remain agnostic about the sign of the subsequent response to the shock. In addition, we exploit sign restrictions in order to distinguish between monetary policy shocks and exchange rate shocks. We assume that a monetary policy shock is associated with an increase in the interest rate and exchange rate appreciation, while an exchange rate shock manifests itself as a rise in the interest rate and exchange rate depreciation. Such restrictions are theoretically motivated by the uncovered interest parity condition (Vonnák, 2010). Our identifying restrictions are summarized in Table 1.

Table 1: Identifying Sign and Zero Restrictions

	Output		Prices		Interest Rate		Exchange Rate	
Horizon	Impact	Lag 1	Impact	Lag 1	Impact	Lag 1	Impact	Lag 1
Monetary Policy Shock	0	?	0	?	+	+	+	+
Exchange Rate Shock	0	?	0	?	-	-	+	+

Note: The exchange rate is defined such that an increase denotes appreciation. ? stands for no restriction.

The identification restrictions are implemented using Givens rotations as in Fry and Pagan (2011). In general, the sign restrictions are checked for a set of possible transformations of structural residuals into reduced form residuals. For an orthonormal matrix Q, it holds that:

$$A_t^{-1} \Sigma_t \epsilon_t = A_t^{-1} \Sigma_t Q' Q \epsilon_t, \tag{12}$$

where $Q\epsilon_t$ represent another vector of uncorrelated structural residuals of unit variance. To ensure the contemporaneous zero restrictions on output and prices, the following specific form of the Givens rotations is employed:

$$Q = \begin{pmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & \cos(\theta) & -\sin(\theta) \\ 0 & 0 & \sin(\theta) & \cos(\theta) \end{pmatrix} \tag{13}$$

Parameter θ represents a random draw for the uniform distribution on the interval $\langle 0, \pi \rangle$. The form of the Givens rotations ensures that structural shocks to the third and fourth variable do not contemporaneously affect the first two variables.

We extend the baseline VAR to include two variables capturing the credit market – specifically, we include the lending rate and credit. We intend to investigate the effects of credit shocks on the economy. The motivation for this exercise stems from several reasons. First, an adverse credit shock figured prominently as one of the likely triggers of the recent crisis (Borio and Disyatat, 2011). Second, with the growing implementation of macro-prudential policies in central banks designed to counteract possible boom/bust cycles, it seems important for policy makers to gauge the effects of possible regulatory policies, which might cause a reduction of credit, on macroeconomic aggregates (Goodhart et al., 2009).

We identify the effect of credit shocks as follows: output and the price level react only with a lag. Output does not increase one quarter after the shock. The lending rate increases, but the short rate does not, enlarging the spread between the two, and at the same time credit decreases. This identification strategy can be justified by several theoretical models. Overall, the models by Curdia and Woodford (2010), Gertler and Karadi (2011), and Gerali et al. (2010) agree on the effects of an adverse credit shock on real GDP, but disagree on the effects on inflation. Therefore, we leave the reaction of the price level unrestricted. We summarize the identification restrictions in Table 2. Similar restrictions to identify credit shocks were recently applied by Alessi (2011), Busch et al. (2010), Hristov et al. (2011), and Tamási and Világi (2011).

Note that relative to the identification of monetary policy and exchange rate shocks in the previous section (Table 1) we impose some additional restrictions on the loan rate and credit so that we can differentiate credit shocks from monetary policy and exchange rate shocks. The responses to monetary policy and exchange rate shocks are very similar to those from the baseline model and we report them in Appendix A.

Table 2: Identifying Sign and Zero Restrictions

	Output		Prices		Interest Rate		Exchange Rate		Lending Rate		Credit	
Horizon	Impact	Lag 1	Impact	Lag 1	Impact	Lag 1	Impact	Lag 1	Impact	Lag 1	Impact	Lag 1
Monetary Policy Shock	0	?	0	?	+	+	+	+	+	+	-	-
Exchange Rate Shock	0	?	0	?	-	-	+	+	-	-	+	+
Credit Supply Shock	0	-	0	?	-	-	?	?	+	+	-	-

Note: The exchange rate is defined such that an increase denotes appreciation. ? stands for no restriction.

Note that in the case of six variables we use the following Givens rotations:

$$Q = Q_{34}(\theta_1) \times Q_{35}(\theta_2) \times Q_{36}(\theta_3) \times Q_{45}(\theta_4) \times Q_{46}(\theta_5) \times Q_{56}(\theta_6), \tag{14}$$

where

$$Q(\theta_m)_{ij} = \begin{pmatrix} 1 & \cdots & 0 & \cdots & 0 & \cdots & 0 \\ \cdots & \ddots & \cdots & \cdots & \cdots & \cdots & 0 \\ 0 & \cdots & \cos(\theta_m) & \cdots & -\sin(\theta_m) & \cdots & 0 \\ \vdots & \vdots & \vdots & 1 & \vdots & \vdots & \vdots \\ 0 & \cdots & \sin(\theta_m) & \cdots & \cos(\theta_m) & \cdots & 0 \\ \cdots & \cdots & \cdots & \cdots & \cdots & \ddots & \cdots \\ 0 & \cdots & 0 & \cdots & 0 & \cdots & 1 \end{pmatrix}, \tag{15}$$

where i and j denote the row and column, respectively. Parameters $\theta_m, m = 1, \ldots, 6$ are drawn from a uniform distribution $U(0, \pi)$. More details can be found in Fry and Pagan (2011). Our rotation ensures that the reaction of the first two variables to structural shocks to other variables is zero on impact while allowing the imposition of restrictions on the reactions of other variables.

3.4 Data and Estimation Strategy

We use a time-varying parameters Bayesian vector autoregression (TVP BVAR) model with stochastic volatility to estimate the evolution of the monetary policy transmission mechanism in the Czech Republic. We use data at quarterly frequency and our sample spans from 1996:1 to 2010:4. In our benchmark model we use seasonally adjusted GDP as a measure of economic activity, the CPI as a measure of the price level, the 3-month PRIBOR as a measure of short-term interest rates, and the nominal effective exchange rate. These variables are typically considered the minimum set allowing analysis of a small open economy.

Following Sims et al. (1990), we estimated the model in levels. This approach avoids the inconsistency that might occur if we incorrectly impose cointegration restrictions. Furthermore, in a Bayesian framework nonstationarity is not an issue, since the presence of unit roots in the data does not affect the likelihood function (Sims et al., 1990).

To conserve degrees of freedom one lag is used for the estimation. Because we are working with a short sample we do not select a training sample but use the whole 1996:1–2010:4 period to elicit the priors. This strategy is advised by Canova (2007) for cases in which a training sample is not available.

To reduce the dimensionality of the estimation we impose the matrix W to be diagonal. As for the prior about the time variation of the coefficients we opt for a prior value of $k_Q = 0.05$, which effectively means that we are attributing 5% of the uncertainty surrounding the OLS estimates to time variation (Kirchner et al., 2010). Furthermore, we set $k_S = 0.025$ and $k_W = 0.01$.

The estimation results are obtained from 25,000 iterations of the Gibbs sampler after discarding the first 25,000 iterations for convergence. Moreover, to address possible autocorrelation of the draws we keep only every 10th iteration. Consequently, the results we present are based on the 2,500 remaining iterations. We discuss the details of the convergence diagnostics in Appendix A.9.

In what follows, we present the median responses to the normalized responses such that the monetary policy shock is equal to one percentage point and the exchange rate shock is equal to a 1% appreciation, so that the responses are comparable across periods.

4. Results

This section gives the results. The responses to monetary policy shocks and exchange rate shocks are presented in subsection 4.1 and subsection 4.2, respectively. The results on the effects of credit shocks follow. Appendix A contains some additional results such as the estimated volatility of reduced form residuals.

4.1 Responses to Monetary Policy Shocks

The estimated median responses of the monetary policy shocks are presented in 1 and are in line with a wide range of theoretical models: in response to a 1 percentage point unexpected interest rate increase, output and prices fall, while the nominal exchange rate initially appreciates. Notably, the results are free of commonly encountered puzzles such as the price puzzle and delayed overshooting of the exchange rate. As for the time variation, while the transmission of the monetary policy shock to the interest rate as well as the exchange rate seems to be stable, the results suggest changes in the responses of output and the price level. More specifically, the responses of output and prices get stronger over time until the outbreak of the crisis in 2008 and remain relatively stable afterwards. The maximum impact of the monetary policy shock on output and prices is about 8 and 10 quarters, respectively. In contrast to the previous literature, this suggests more persistent effects of monetary policy on the aggregate economy. Borys et al. (2009) survey the previous studies on monetary policy transmission in the Czech Republic and find that output and prices bottom out typically after 4 quarters. In terms of the economic significance of monetary policy shocks, our results are largely comparable to the time-invariant VAR estimates presented in Borys et al. (2009).

Figure 2 gives a closer look at the evolution of the response of output, prices, the exchange rate, and the interest rate to a monetary policy shock at specific horizons. While we do not want to overemphasize the precision of the estimates, the results suggest that output and prices respond to monetary shocks – especially at the horizons of 8 and 12 quarters – about 25% more strongly in the period before the outbreak of the financial crisis than at the beginning of inflation targeting in 1998. This pattern of increasing responsiveness of output and prices might be explained by financial sector deepening and overall economic development coupled with successful disinflation. In addition, the monetary policy shocks become more persistent over time. More persistent policy shocks may induce a stronger reaction of financial markets, including long-term interest rates. As a consequence, this may generate stronger responsiveness of the aggregate economy to monetary policy shocks.

In addition, we present the responses to monetary policy shocks at the 8-quarter horizon over time to assess the statistical significance of our results in Appendix A.4. The confidence bands for TVP VARs are typically not reported, since they are often too large for this type of model. Nevertheless, our results suggest that the intervals are often not as large as commonly thought. See also Appendix A.5 for exchange rate shocks and Appendix A.6 for credit shocks.

Figure 1: Time-Varying Impulse Responses to a 100 Basis Point Monetary Policy Shock

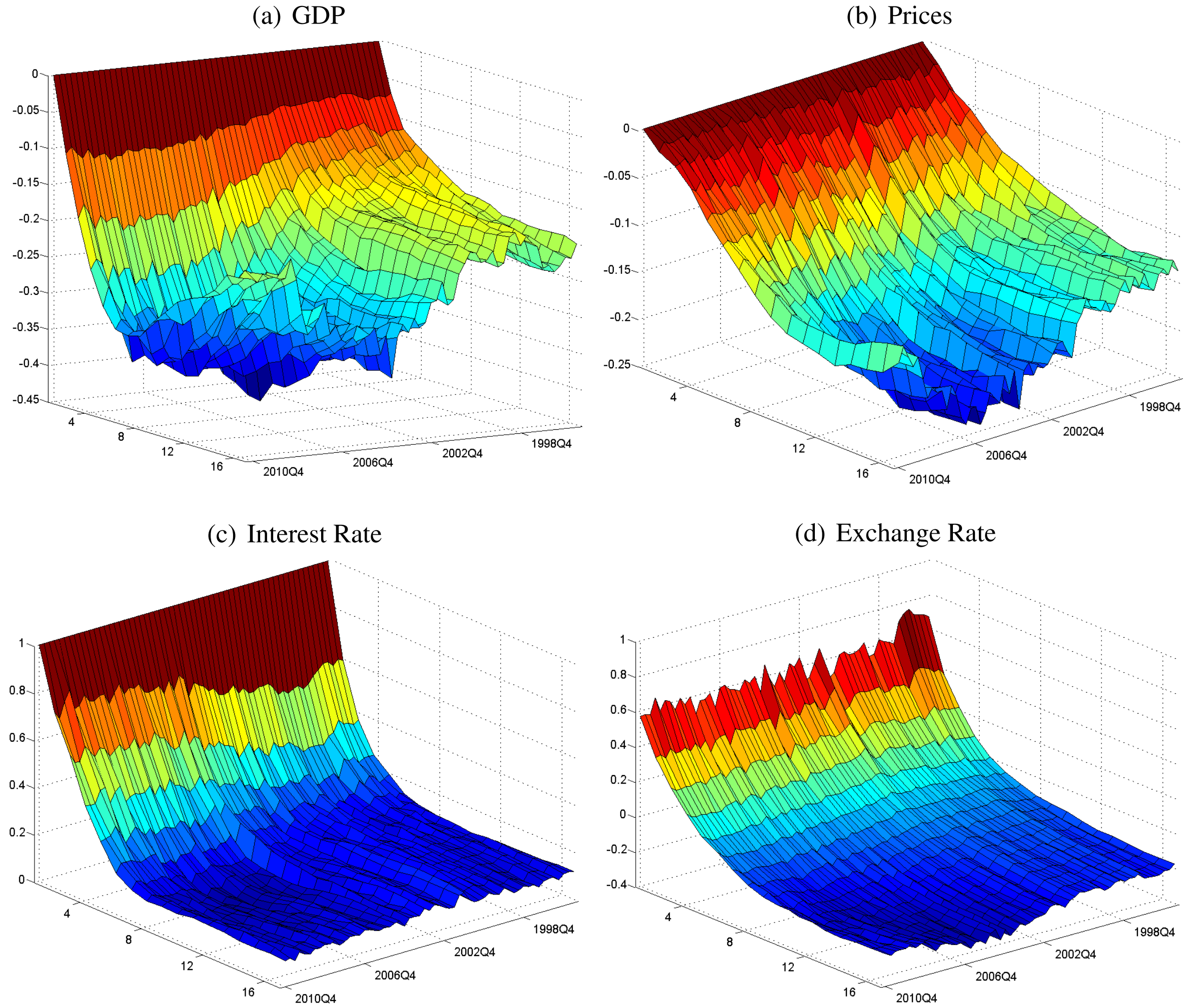
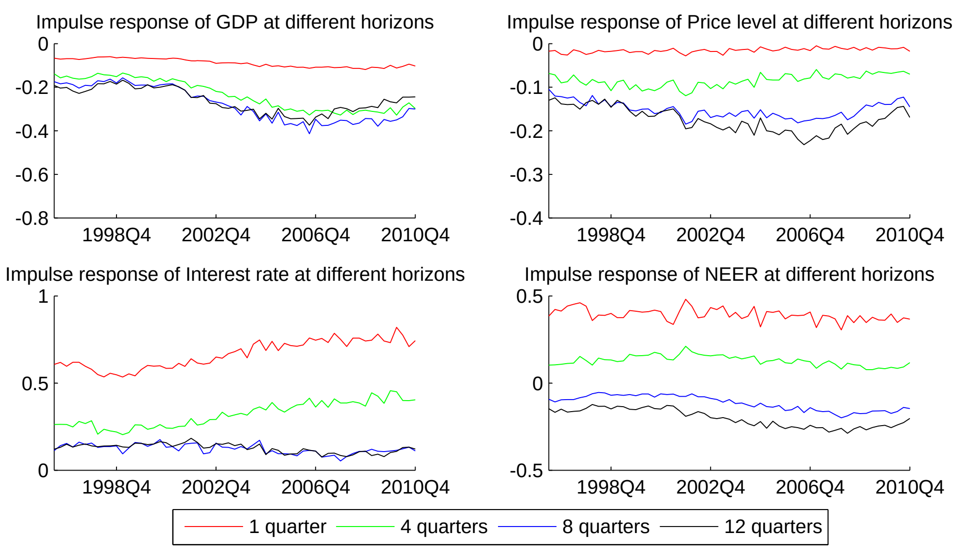


Figure 2: Responses at Different Horizons over Different Periods



4.2 Responses to Exchange Rate Shocks

Exchange rate shocks affect GDP through at least two channels: the expenditure-switching channel (a decrease in net exports following an appreciation) and the interest rate channel (exchange rate shocks are typically accommodated by decreases in interest rates, which can in turn stimulate economic activity). The empirical evidence on the effect of the exchange rate on output is somewhat mixed for the emerging markets and previous studies report both positive and negative effects (Ahmed, 2003; Sanchez, 2007; Vonnák, 2010). The median responses are presented in 3. The results suggest that output increases following an unexpected exchange rate appreciation. The positive reaction of output might be a consequence of the fact that there are no foreign variables in our model. We investigated this issue by running a time-invariant VAR augmented by a set of foreign variables that included commodity prices, the EURIBOR, and German industrial production and price level (Czech National Bank, 2010). However, even in the augmented specification output responds positively. Finally, a plausible explanation is that the rise in output might indicate economic convergence of the Czech Republic that is not captured by our model. As for the reaction of prices, the results suggest that exchange rate shocks pass through to prices relatively quickly, which is consistent with previous microeconomic evidence on the exchange rate pass-through in the Czech Republic (Babecka-Kucharcukova, 2009). Moreover, the results suggest that the pass-through declines over time, which is in line with the international evidence (Mumtaz et al., 2011).

Figure 3: Time-Varying Impulse Responses to a 1% Exchange Rate Shock

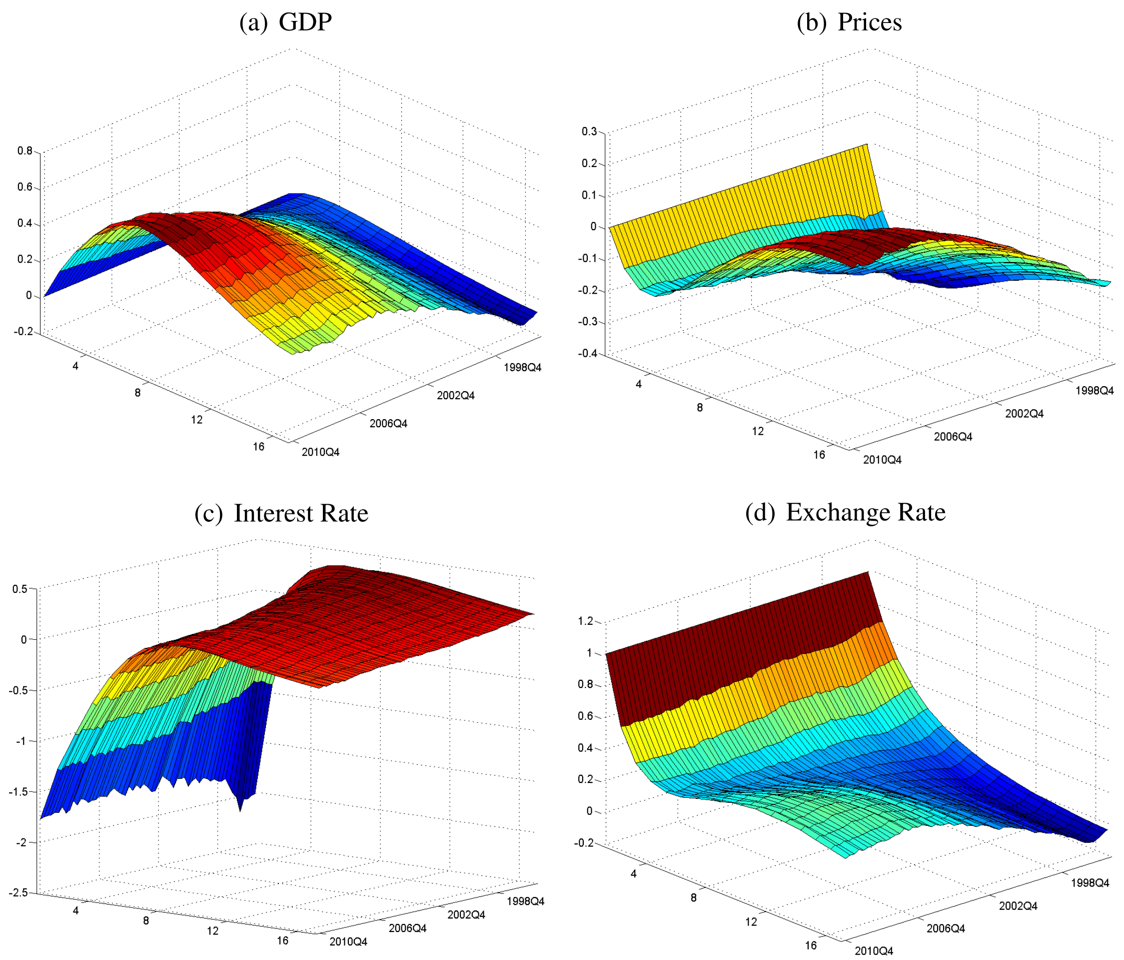
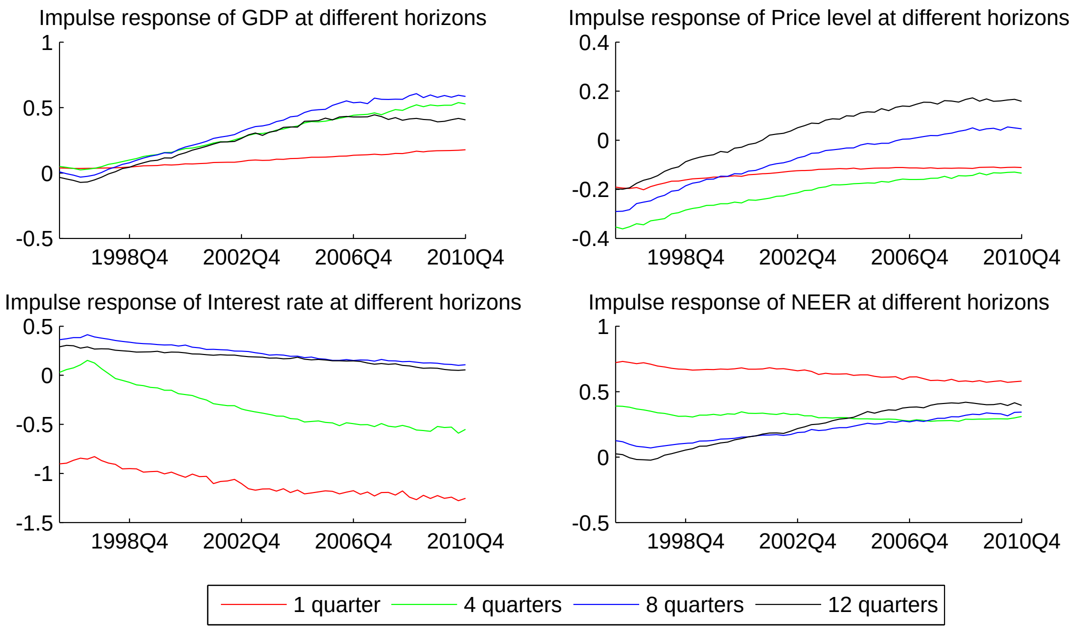


Figure 4 suggests that the reaction of GDP and prices following an exchange rate shock became stronger over time. In addition, the results suggest that the exchange rate shock is more persistent in recent periods than at the beginning of the sample.

Figure 4: Responses at Different Horizons over Different Periods



4.3 Responses of Credit Shocks

We present the median impulse responses to a negative credit supply shock in 5. Overall, we find that the effects of adverse credit supply shocks on the economy are sizeable. In response to a 1% decrease in credit supply, GDP falls by 0.2% after 2 quarters, and prices decrease by about 0.7%. The lending rate increases, while the central bank reacts by lowering the interest rate, and as a result the credit spread increases. Although the exchange rate appreciates on impact, possibly due to the increase in the lending rate, after two quarters it starts to depreciate, reflecting the deterioration in economic activity.

As for the time variation, the effects of credit supply shocks seem to be less persistent over time. This is likely to be a consequence of improved financial stability of Czech banks associated with improved access to credit. The beginning of our sample is characterized by prudent behavior of Czech banks associated with the restructuring of the Czech banking industry. The Czech banking sector was consolidated in the early 2000s and maintained solid liquidity even during the current global financial crisis. This is also reflected in the development of credit growth – see Appendix A.1 for the figure.

Figure 5: Time-Varying Impulse Responses to a 1% Credit Shock

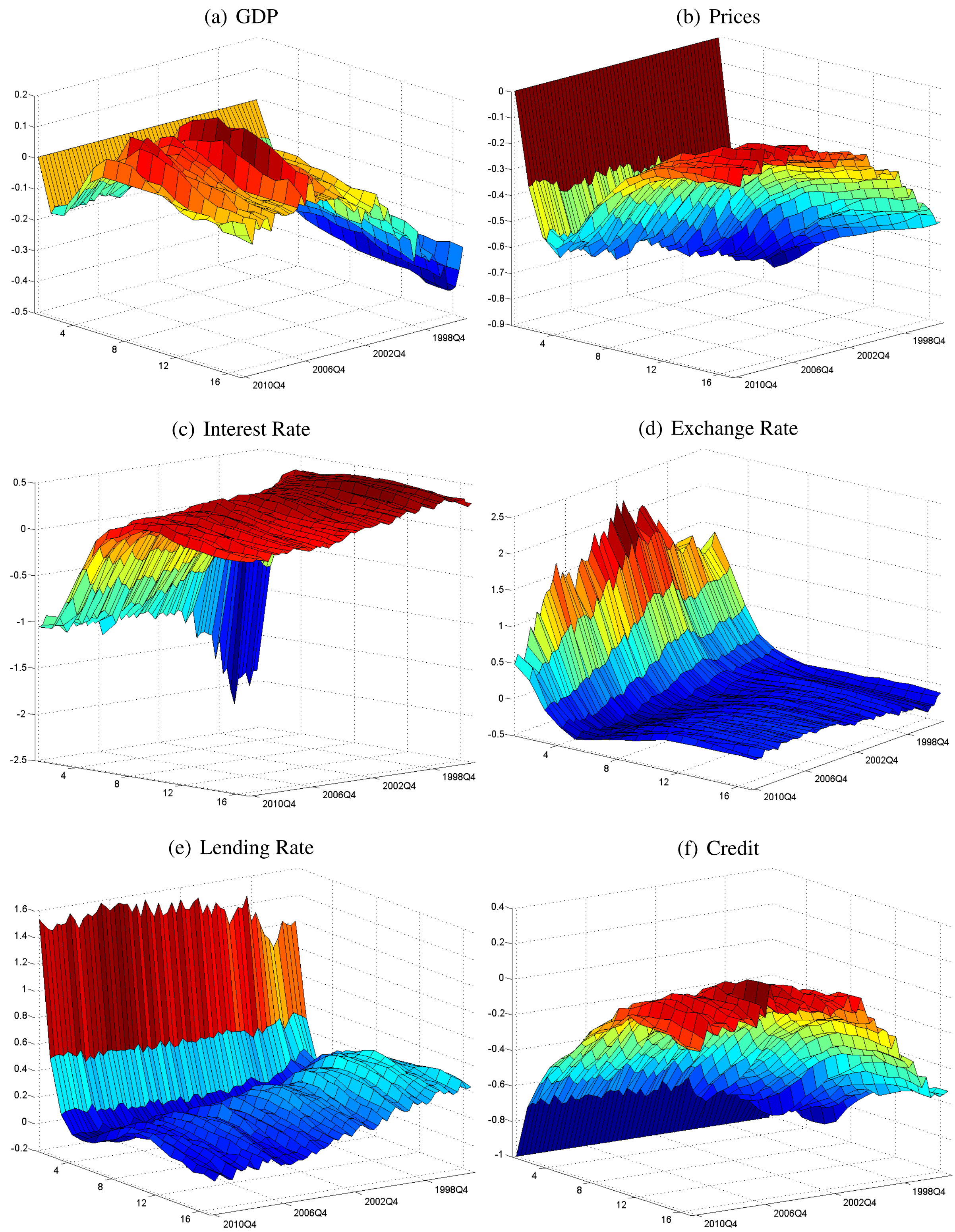
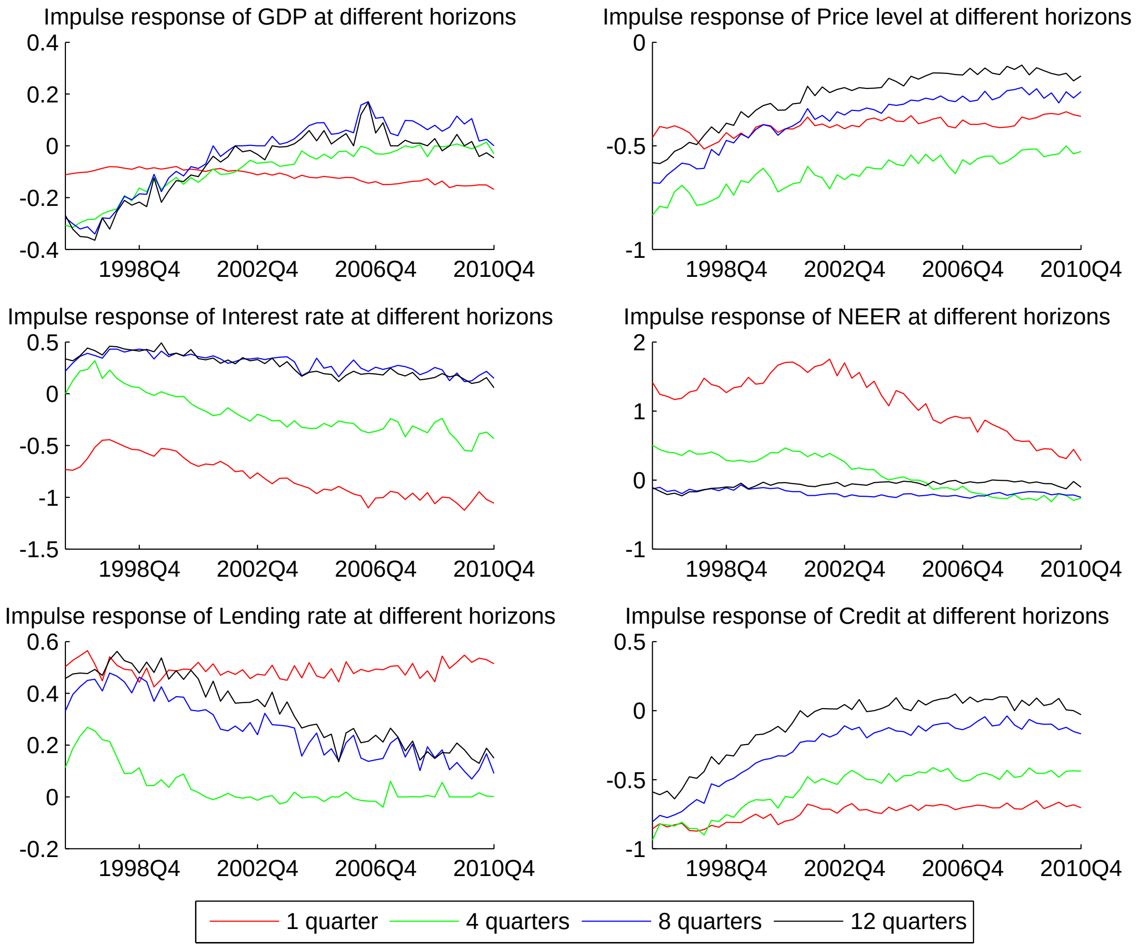


Figure 6: Responses at Different Horizons over Different Periods



5. Concluding Remarks

In this paper, we analyze the evolution of the monetary policy transmission mechanism in the Czech Republic. The Czech economy has witnessed many important economic, institutional, and political changes during the past two decades and has transformed from an inefficient command-driven economy into a market-oriented economy. As concerns monetary policy regime changes, the Czech Republic maintained a fixed exchange rate until May 1997 and adopted inflation targeting in January 1998. Inflation targeting was adopted as a disinflation strategy at a time of nearly double-digit inflation. Inflation has fallen to around 2% in recent years, following gradual reductions in the inflation target. Therefore, it seems reasonable to model the monetary transmission mechanism as time-varying. For this reason, we employ the recently developed Bayesian time-varying vector autoregression with stochastic volatility (Primiceri, 2005). This flexible approach also allows us to model the size of a shock hitting the economy as time-varying to account for periods of more volatile economic developments such as during the current global financial crisis.

The recent financial crisis has reminded us of the important role the financial markets play in macroeconomic fluctuations. To account explicitly for the links between the financial and macroeconomic sector, we include credit growth and the lending rate along with standard macroeconomic variables (output, prices, the interest rate, and the exchange rate) in the VAR system. By doing so, we shed light on the relative importance of financial shocks for the aggregate economy. Importantly, our time-varying framework allows us to assess changes in the relative importance of financial shocks.

Our results suggest an increasing responsiveness of output and prices to monetary policy shocks until the financial crisis. The responsiveness of output and prices to monetary shocks did not increase further during the crisis, but remained largely constant at the pre-crisis level. The increasing responsiveness of the aggregate economy to monetary policy shocks is likely to be consequence of financial market deepening and overall economic development associated with disinflation. In addition, we find that monetary policy shocks become more persistent over time. This may be an additional reason for the greater responsiveness of output and prices to monetary policy shocks, as more persistent shocks are likely to affect the yield curve and, as a consequence, the lending decisions of economic agents, for whom long-term interest rates are typically particularly important. In a similar vein, the credibility of inflation targeting in the Czech Republic and well-anchored inflation expectations might play a role as well (see Holub and Hurnik, 2008). This is supported by the weakening of the exchange rate pass-through over time.

In terms of the interaction of financial and macroeconomic variables, we find that credit shocks had a more sizeable impact on output and prices during the period of bank restructuring in about the year 2000. As concerns the current financial crisis, the effect of credit shocks is not more sizeable. Although this may sound somewhat surprising, it is important to realize that the Czech financial sector has remained largely stable during the crisis. Banking sector capitalization and liquidity are in much better shape than they used to be. The ratio of non-performing loans to total loans has increased somewhat during the current financial crisis, but has remained at a much lower level than in the year 2000. Therefore, the results clearly imply that financial shocks are more important for the evolution of the aggregate economy in an environment of financial instability.

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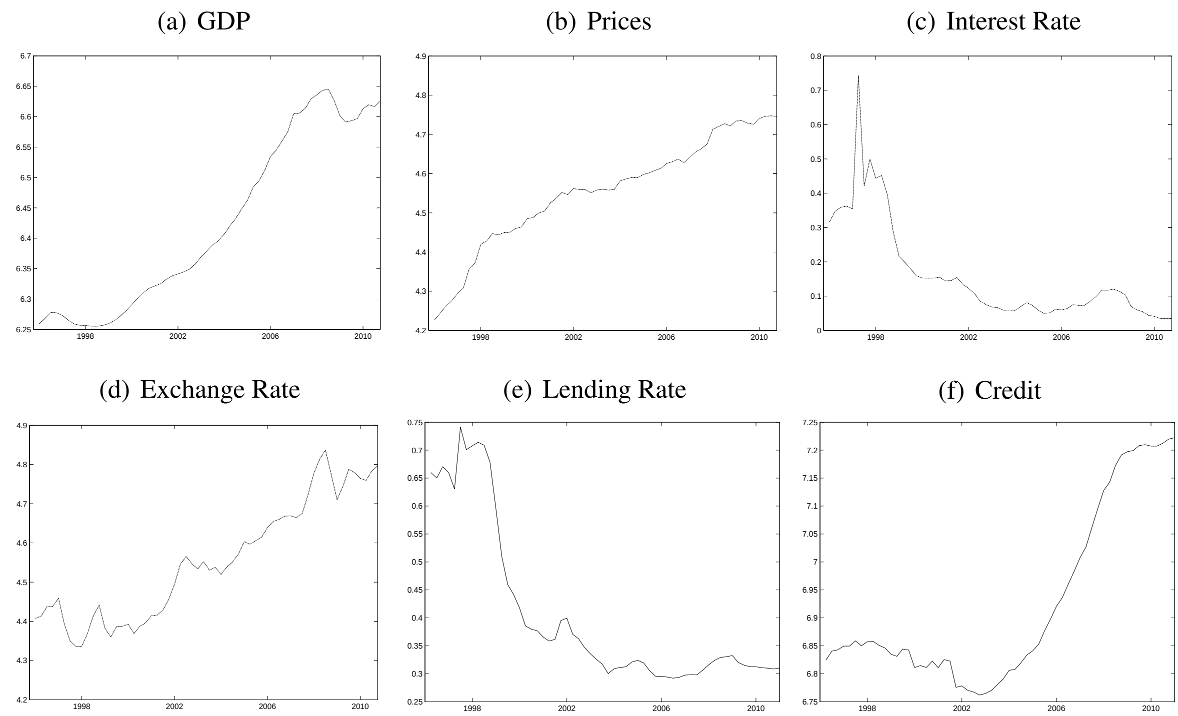
A. Supplementary Figures

A.1 Data

Table 3: Data Used in Estimation

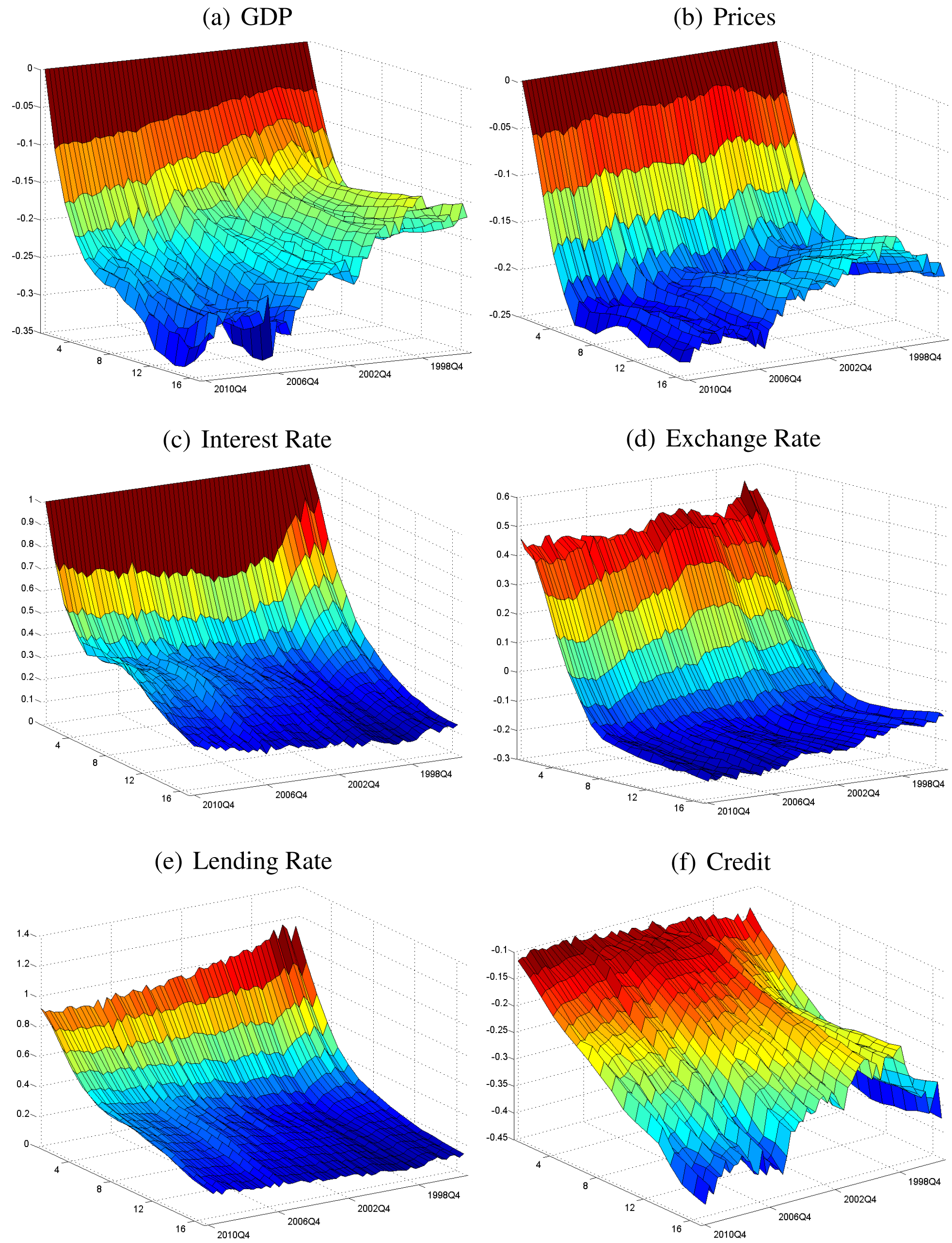
Variable	Time span	Source
GDP s.a.	1996:1–2010:4	IFS 93566..CZF...
Consumer price index	1996:1–2010:4	IFS 93564...ZF...
Money market rate (3-month PRIBOR)	1996:1–2010:4	IFS 93560B..ZF...
Nominal effective exchange rate	1996:1–2010:4	IFS 935..NECZF...
Lending rate	1996:1–2010:4	IFS 93560P..ZF...
Credit	1996:1–2010:4	ARAD

Figure 7: Data Used in Estimation



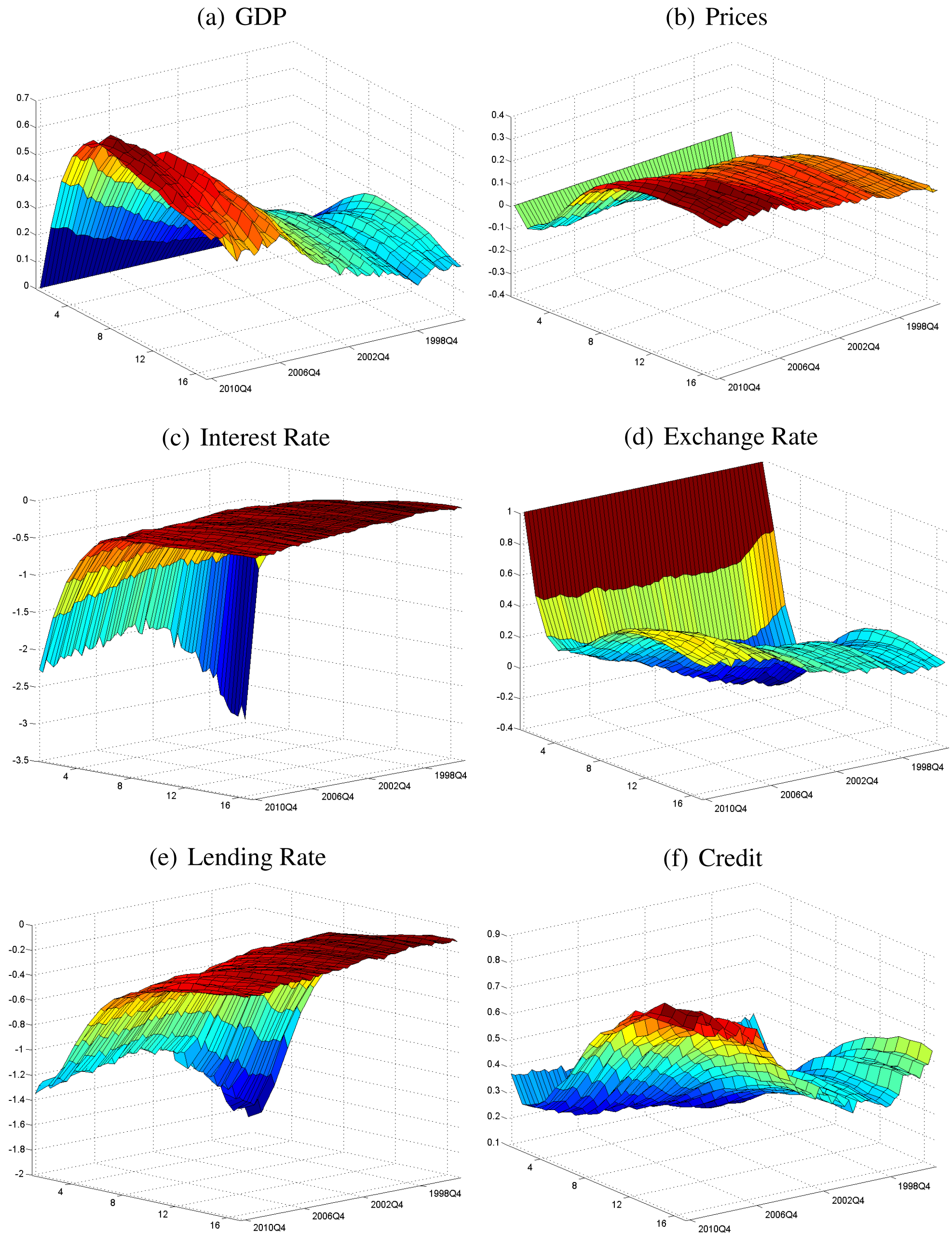
A.2 Responses to Monetary Policy Shock in Augmented Model

Figure 8: Time-Varying Impulse Responses to a 100 Basis Point Monetary Policy Shock



A.3 Responses to Exchange Rate Shock in Augmented Model

Figure 9: Time-Varying Impulse Responses to a 1% Exchange Rate Shock



(a) GDP

(b) Prices

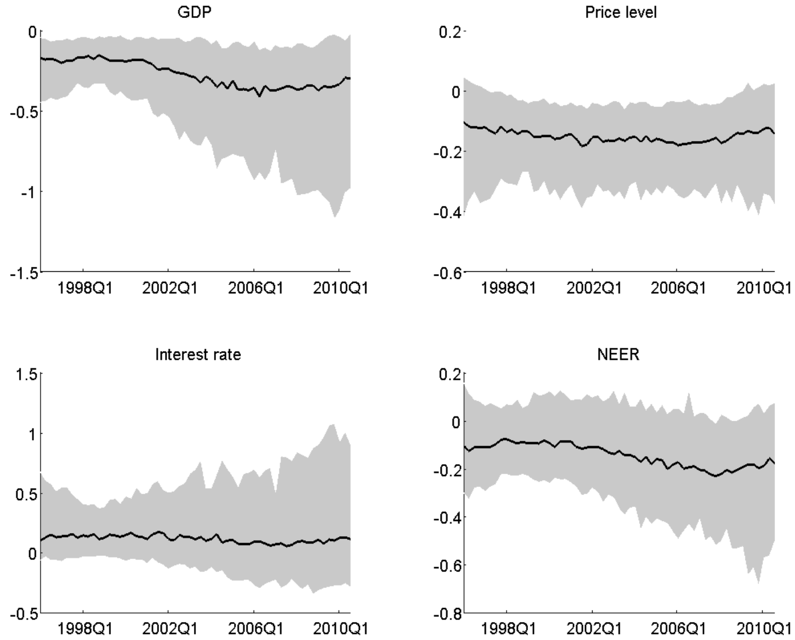
(c) Interest Rate

(d) Exchange Rate

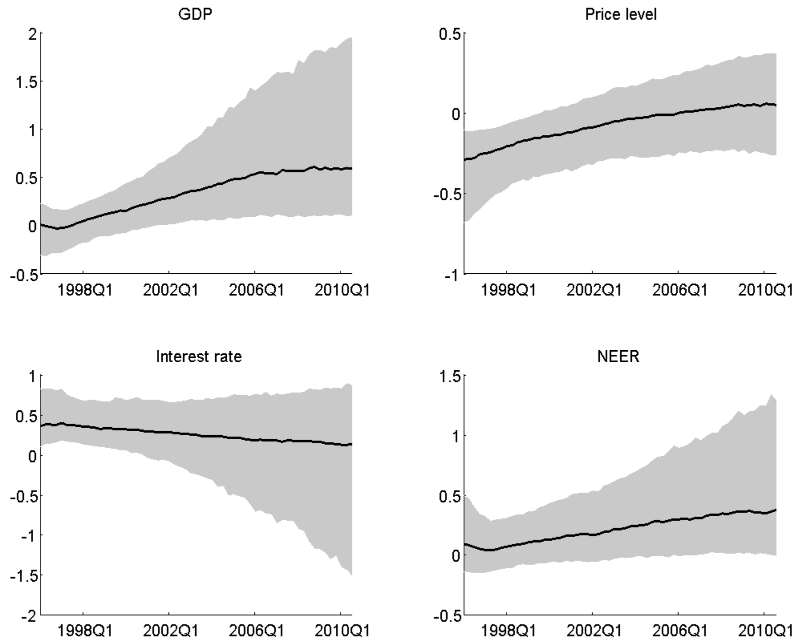
(e) Lending Rate

(f) Credit

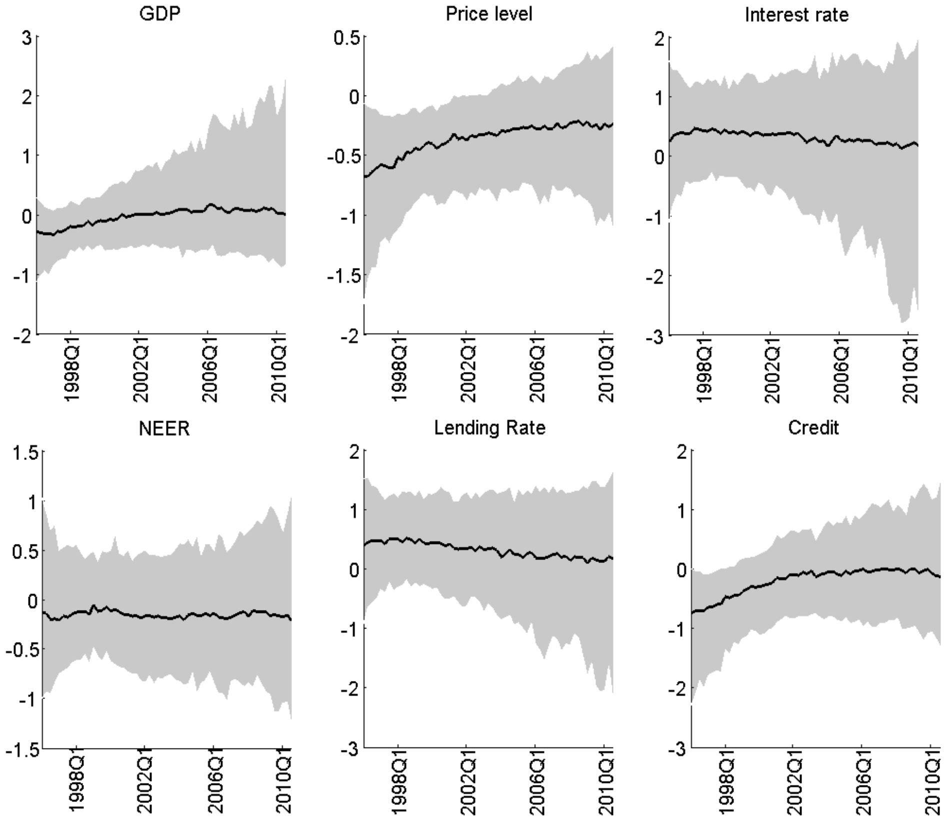
A.4 Responses to Monetary Policy Shock at 8-Quarter Horizon



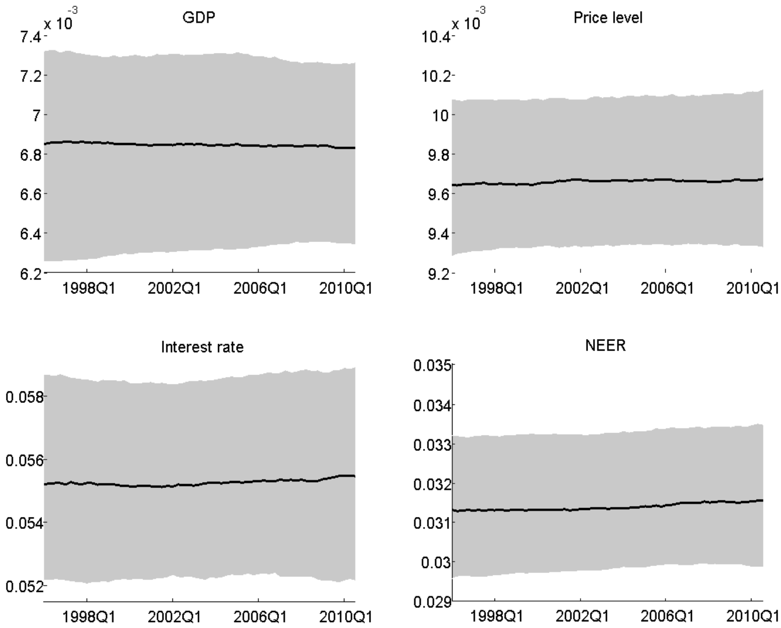
A.5 Responses to Exchange Rate Shock at 8-Quarter Horizon



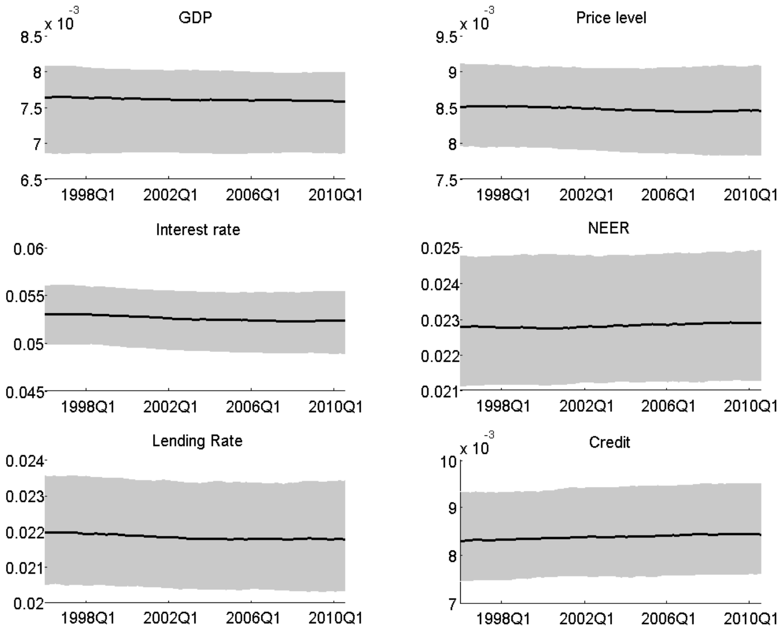
A.6 Responses to Credit Shock at 8-Quarter Horizon



A.7 Volatility of Reduced Form Residuals (Baseline Model)



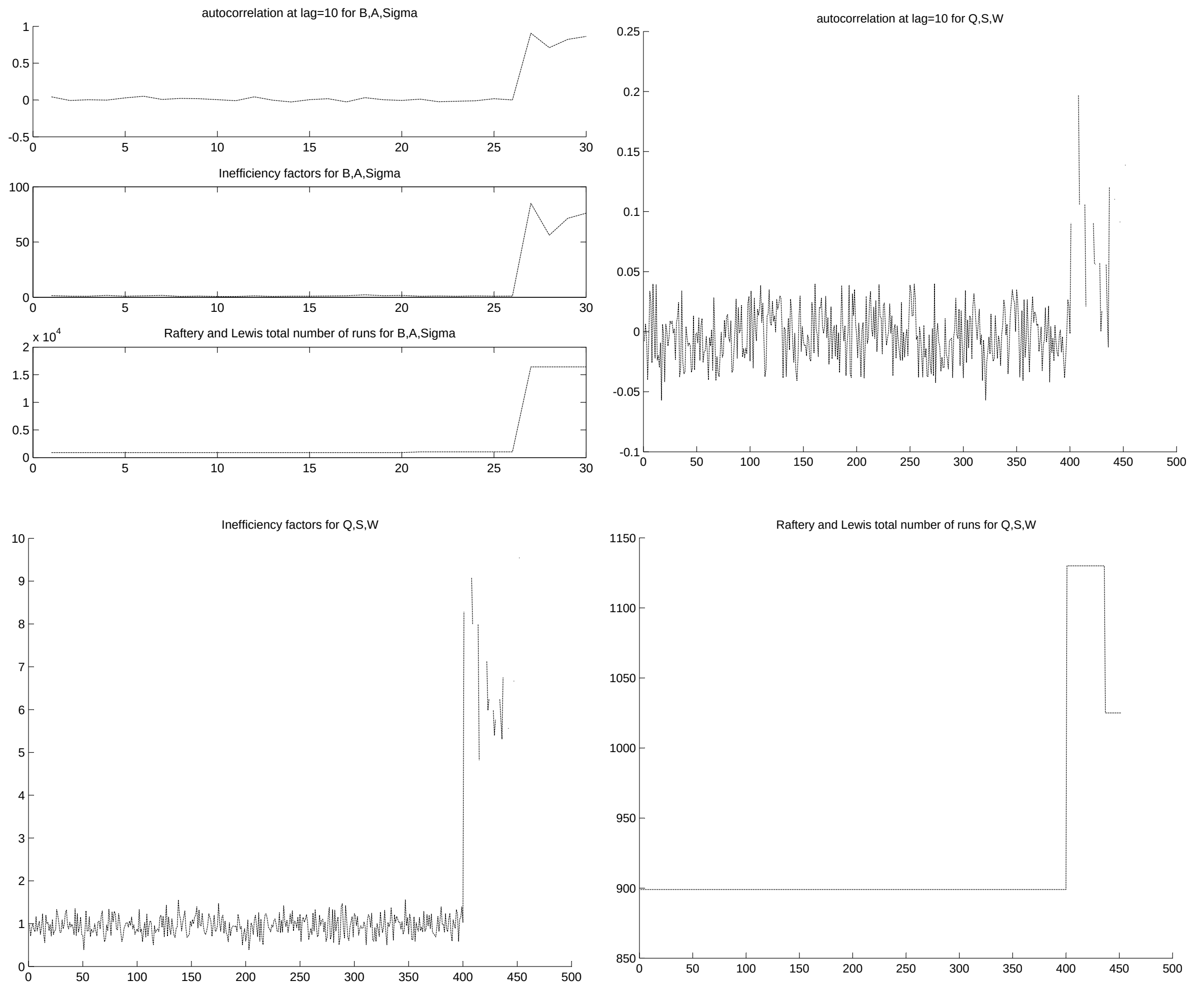
A.8 Volatility of Reduced Form Residuals (Augmented Model)



A.9 Convergence Diagnostics

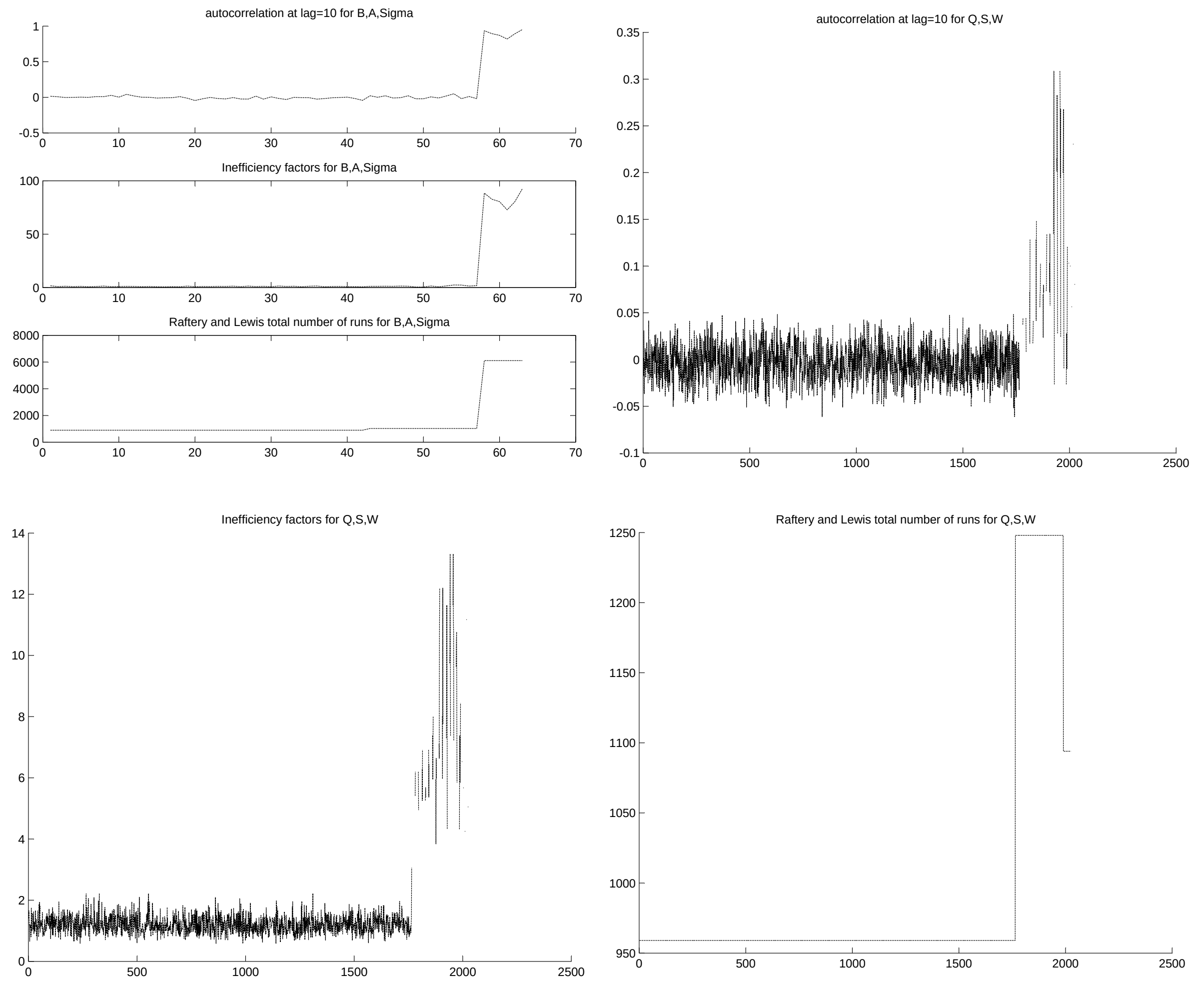
Following Primiceri (2005), the convergence of the Markov chain Monte Carlo algorithm is assessed by various autocorrelation measures and by Raftery and Lewis (1992) diagnostics.[3] To save space, only the convergence diagnostics of the coefficients for 2008Q4 and all hyperparameters are presented. The most straightforward way is to look at the autocorrelation of the Markov chain. Low autocorrelation suggests independence of draws and thus efficiency of the sampling algorithm. The autocorrelation of the chain at a lag equal to 10 is presented. The second measure is the so-called inefficiency factor, which is defined as $1 + 2\sum_{k=1}^{\infty} \rho_k$, where ρ_k represents the k-th autocorrelation of the chain. According to Primiceri (2005), values below 20 are viewed as satisfactory. Finally, the estimate of Raftery and Lewis (1992) provides the number of runs of the sampling algorithm needed to achieve a certain precision (for the 0.025 and 0.975 quantiles of the marginal posterior distributions, the desired accuracy of 0.025 is required to be achieved with a probability of 0.95). The statistics suggest difficulties with the efficiency of the Σ estimates. However, the total number of runs required by the Raftery and Lewis diagnostics is well below the number of iterations we use. Moreover, this is not an issue as we present the responses to the normalized shocks.

Figure 10: Convergence Diagnostics (Baseline Model)



[3] The diagnostics are based on the Econometric Toolbox described in LeSage (1999).

Figure 11: Convergence Diagnostics (Augmented Model)



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