Empir Econ (2014) 46:827–842
DOI 10.1007/s00181-013-0699-0

Evaluating changes in the monetary transmission mechanism in the Czech Republic

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Received: 12 October 2011 / Accepted: 6 February 2013 / Published online: 9 May 2013

Abstract We investigate the evolution of the monetary policy transmission mechanism in the Czech Republic over the course of the 1996–2010 time period through the use of 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 has changed over time, focusing on the period of the recent financial crisis. Our results suggest that prices have become increasingly responsive to monetary policy shocks. However, in terms of credible intervals, the stability of the monetary policy transmission mechanism in the Czech Republic cannot be rejected. Furthermore, it is demonstrated that the exchange rate pass-through has largely remained stable over time.

Keywords Monetary policy transmission · Sign restrictions · Time-varying parameters

Electronic supplementary material The online version of this article (doi:10.1007/s00181-013-0699-0) contains supplementary material, which is available to authorized users.

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JEL Codes E52 · E44

1 Introduction

There are many reasons to believe that the structure of the Czech economy has changed over time, including this nation's transition to a market-based economy; the deepening of trade in the Czech Republic that has been spurred by trade liberalisation and the integration of this country into the European Union; the consolidation of the Czech banking industry that occurred at the turn of the century and the recent financial crisis. It is likely that these changes in the structure of the Czech economy have impacted the monetary transmission mechanism in the Czech Republic; in other words, these changes have likely altered the effects of monetary policy on the aggregate economy (Cogley and Sargent 2005). Furthermore, changes in the conduct of monetary policy that have been caused by the introduction of an inflation targeting regime have also most likely influenced the strength of Czech monetary policy. This inflation targeting regime was adopted in the Czech Republic as a disinflation strategy. In fact, 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 expectations had been anchored because inflation had fallen to levels that were consistent with price stability (Holub and Hurnik 2008).

Given this context, it is somewhat surprising that relatively little evidence exists regarding changes in the monetary transmission mechanism in the Czech Republic. Moreover, although it is extraordinarily important for policymakers to know the strength of monetary transmission during times of crisis, strikingly, the effects of monetary policy actions on the Czech economy during the recent financial crisis have not yet been comprehensively investigated. Our paper contributes to the extant literature by providing stylised facts about changes in the strength of monetary policy actions over time, particularly during the recent crisis. We seek to investigate both the qualitative and quantitative implications of these changes by estimating a recently developed Bayesian vector autoregression model with time-varying parameters (the TVP BVAR model) (Primiceri 2005). The transmission of exchange rate shocks is also examined in this study.

The estimations of this study demonstrate high uncertainty with respect to results; this uncertainty arises because of the small data sample that is used in this exercise. From an examination of the central tendencies that are reflected by the medians of the posterior distributions of impulse response functions, our results suggest an increasing responsiveness of prices in the Czech Republic to monetary policy shocks. We attribute this finding to the deepening of the Czech financial sector and to the overall economic development and disinflation that have occurred in this country. The strongest impact of monetary shocks on prices occurs between 1 and 2 years after the shock. We also find that exchange rate pass-through has remained largely stable over time. Finally, we do not observe any weakening of monetary transmission during the recent financial crisis.

The rest of this paper is organised as follows. The related literature is discussed in Sect. 2. Section 3 introduces our econometric model. Section 4 provides our results.

Concluding remarks are offered in Sect. 5. Subsequently, Sect. A provides additional results.

2 Related (time-varying) VAR literature

2.1 VAR models and monetary transmission

Since the publication of the seminal contributions of Sims (1980), the vector autoregression model has become the major tool for investigating monetary policy transmission mechanism. The stylised facts about the monetary transmission mechanism in the US economy are summarised in a survey by Christiano et al. (1999). These researchers conclude that following a contractionary monetary policy shock, economic activity declines quickly in a hump-shaped manner; by contrast, the negative reaction with respect to price level is more delayed and persistent. Similarly, Peersman and Smets (2001) provide evidence regarding the monetary policy transmission mechanism for the euro area as a whole, whereas Mojon and Peersman (2001) investigate the effects of monetary policy shocks in the individual countries of the euro area.

2.2 Time-varying VARs

It has long been recognised that the structure and functioning of the economy changes over time; therefore, there exists a need to account for this evolution in estimation procedures (Koop et al. 2009). Two main approaches to modelling changes in the transmission mechanism over time have appeared in empirical investigations. First, a sample can be split and a model may be estimated over individual subsamples. Second, we can directly model coefficient changes within the system (e.g., through the use of the structural break or random walk assumptions). For example, Clarida et al. (2000) utilise the first approach of splitting the sample into subsamples to investigate changes in the monetary policy rule. However, the dates on which a sample should be split to create the most informative subsamples are often unclear. Importantly, it is more likely that an economy will change gradually instead of undergoing sudden abrupt shifts (Koop et al. 2009). As a result, a second vein of relevant literature typically includes less restrictive assumptions about the behaviour of an economy. Studies from this body of literature typically create time-varying coefficient models by applying the Kalman smoother over the entire sample; by contrast, time-invariant estimation procedures use only the information that is contained in each relevant subsample.

Furthermore, several different approaches can be used for the explicit modelling of the evolution of parameters over time. For example, Stock and Watson (1996) estimate a model with a small number of structural breaks. Alternatively, the Markov switching VAR model of Sims and Zha (2006) represents another possible modelling technique. However, TVPs VAR models have recently become increasingly popular. This popularity is driven by the flexibility of this approach. For example, in contrast to Markov switching VAR models, VAR models with TVPs do not have to jump between different regimes.

Canova (1993) is the first researcher to model time variation using the random walk assumption in multivariate models. More recently, Cogley and Sargent (2001) contribute to the 'bad policy' versus 'bad luck' literature stream originated by Clarida et al. (2000) by estimating vector autoregressions that involve time-varying parameters (TVP VAR models) and a constant volatility of shocks. The limitation of the Cogley and Sargent model is its constant volatility assumption, which neglects the possible heteroskedasticity of shocks and any nonlinearities that may exist in the simultaneous relations among the variables of the model. Consequently, Cogley and Sargent (2005) allow for time-varying variance, although the simultaneous relations among the variables (the covariances) are still treated as time-invariant features. As noted by Primiceri (2005), this approach limits the analysis of Cogley and Sargent (2005) to reduced-form models that are usable for data description and forecasting and does not allow for any structural interpretations. To address this issue, Primiceri (2005) stresses the importance of allowing for time variation in the variance–covariance matrix of innovations and estimates a TVP VAR model with stochastic volatility.

Recently, TVP VAR approaches have been widely used to study changes in the transmission of various phenomena, such as monetary policy (Canova et al. 2007; Benati and Surico 2008), 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

With respect to the modelling of monetary transmission in the Czech economy, Borys et al. (2009) utilise a battery of VAR models and identification strategies to demonstrate that the monetary transmission mechanism functions relatively well if estimated for a single monetary policy regime. Havranek et al. (2012) employ a block-restriction VAR model and examine the interactions between macroeconomic conditions and the Czech financial sector. These researchers find that monetary policy produces a systematic effect on financial stability and that the use of certain financial variables improves forecasts of inflation and economic activity. Darvas (2013) estimates Czech monetary transmission in a time-varying framework for the 1993–2008 time period. He finds that the nature of monetary transmission remains largely unchanged 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 shifts that may have occurred in 1997, a year of exchange rate turbulence in the Czech Republic, or during the recent 2008–2009 financial crisis). To incorporate this potentially important consideration into a model, we estimate a Bayesian TVP model that includes stochastic volatility. As discussed above, a failure to appropriately consider the heteroskedasticity of shocks might cause changes in the magnitude of shocks to be confounded with changes in the transmission mechanism, yielding inconsistent estimates. In addition, we consider the effects of exchange rate shocks on macroeconomic fluctuations.

3 The TVP BVAR approach

3.1 The model

In accordance with Primiceri (2005), we construct the following model:

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

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

Since it has recently been recognised that it is important to allow not only the coefficients but also the error variances and the covariances to vary over time, we will employ a triangular reduction of Ω_t such that

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

or, equivalently,

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

where A_t is the lower triangular matrix

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

and Σ_t is the diagonal matrix

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

Thus, we have

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

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

We rewrite (6) by stacking all of the coefficients from the right side of this equation in a vector B_t and thereby 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]

We must next 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; for the standard deviation of shocks Σ_t we will utilise a stochastic volatility framework and assume that the elements of this framework follow a geometric random walk. Formally, the dynamics of the parameters may be 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 a state space model with (7) as the model's measurement equation and the state equations of (8)–(10).

The innovations $(\varepsilon_t,\ \nu_t,\ \zeta_t,\ \eta_t)$ are assumed to be jointly normal with the following variance–covariance matrix:

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

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 of this OLS point estimate from the time-invariant VAR:

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

The prior for A_0 is obtained in a similar manner:

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

Next, for $\log\sigma_0$, the mean of the prior distribution is established as the logarithm of the OLS estimate of the standard errors from the aforementioned time-invariant

[1] The symbol $\otimes$ denotes the Kronecker product.

[2] We assume that S is block diagonal, i.e. we presume that the contemporaneous relationships among the variables evolve independently. For example, there are three blocks of S in the VAR that consists of four variables. Furthermore, to reduce the dimensionality of the estimation, we impose the condition that the matrix W must be diagonal (Kirchner et al. 2010).

VAR, and the variance–covariance matrix is arbitrarily chosen to be proportional to the identity matrix:

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

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

$$\begin{aligned}
Q &\sim IW\left(k_Q^2 \cdot \tau \cdot \mathrm{var}\left(\widehat{B}_{\mathrm{OLS}}\right),\ \tau\right),\\
W &\sim IG\left(k_W^2 \cdot (1+\dim(W)) \cdot I_n,\ (1+\dim(W))\right), \text{ and}\\
S_l &\sim IW\left(k_S^2 \cdot (1+\dim(S_l)) \cdot \mathrm{var}\left(\widehat{A}_{1,\mathrm{OLS}}\right),\ (1+\dim(S_l))\right),
\end{aligned}$$

where τ is the size of the training sample, S_l denotes the corresponding blocks of S, and $\widehat{A}_{1,\mathrm{OLS}}$ represent the corresponding blocks of $\widehat{A}_{\mathrm{OLS}}$. 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 equal to one plus the dimension of each matrix. Moreover, in accordance with the published literature (Cogley and Sargent 2001), the scale matrices are chosen to be constant fractions of the products obtained by multiplying the variances of the corresponding OLS estimates on the training sample by the degrees of freedom of the appropriate scale matrix.

3.3 Identification and structural interpretation

In a closed economy, the recursive identification scheme appears to be a plausible approach for identifying the effects of monetary policy shocks (Christiano et al. 1999); however, in an open economy, this type of identification scheme might confound monetary policy shocks with exchange rate shocks (Kim and Roubini 2000). We, therefore, use sign restrictions to identify these two types of shocks (Canova and Nicolo 2002; Uhlig 2005; Rubio-Ramirez et al. 2010). Our identification approach is consistent with the method of Farrant and Peersman (2006). Two types of shocks are identified: monetary policy shocks and exchange rate shocks. The imposed sign restrictions are summarised in Table 1. Note that although we estimate the model using differences, the sign restrictions are imposed on levels.

The sign of a response for output and price level is imposed for four quarters because there are assumed to be lags in the transmission of the two types of shocks. For the interest rate, there is assumed to be a lag of two quarters that reflects interest rate

Table 1 Identifying sign restrictions

	GDP	Price level	Interest rate	NEER
Monetary policy shock	< 0	< 0	> 0	> 0
Exchange rate shock	< 0	< 0	< 0	> 0

Note The exchange rate is defined such that an increase denotes appreciation smoothing. Finally, exchange rate reaction is restricted to only the immediate impact of a shock.

In accordance with the approach of Fry and Pagan (2011), the identification restrictions are implemented using Givens rotations. In general, the sign restrictions are examined to determine the applicability of a set of possible transformations of structural residuals into reduced-form residuals. For an orthonormal matrix Q, the following relation must hold:

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

where $Q\epsilon_t$ represent another vector of uncorrelated structural residuals with unit variance. Givens rotations are employed as follows:

$$Q = Q_{12}(\theta_1) \times Q_{13}(\theta_2) \times Q_{14}(\theta_3) \times Q_{23}(\theta_4) \times Q_{24}(\theta_5) \times Q_{34}(\theta_6), \tag{13}$$

where

$$Q(\theta_m)_{ij} = \begin{pmatrix} 1 & \cdots & 0 & \cdots & 0 & \cdots & 0 \\ \cdots & \ddots & \cdots & \cdots & \cdots & \cdots & 0 \\ 0 & \cdots & \cos(\theta_m) & \cdots & -\sin(\theta_m) & \cdots & 0 \\ \vdots & \vdots & \vdots & 1 & \vdots & \vdots & \vdots \\ 0 & \cdots & \sin(\theta_m) & \cdots & \cos(\theta_m) & \cdots & 0 \\ \cdots & \cdots & \cdots & \cdots & \cdots & \ddots & \cdots \\ 0 & \cdots & 0 & \cdots & 0 & \cdots & 1 \end{pmatrix}. \tag{14}$$

In the above expression, i and j denote the row and column of the matrix, respectively. The parameters θ_m, $m = 1, \ldots, 6$ are drawn from a uniform distribution on the interval $\langle 0, \pi \rangle$. Additional details are provided by Fry and Pagan (2011).

3.4 Data and estimation strategy

We use data that are obtained at a quarterly frequency, and our sample spans the time period from the first quarter of 1996 to the fourth quarter of 2010. We use the seasonally adjusted real 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 regarded as the minimum set of factors that must be considered in the analysis of a small and open economy.

All of the variables except for the interest rate are transformed to log differences. The model is thus estimated in a stationary form. Two lags are used for this estimation. This number of lags represents a compromise between the enormous expansion of the parameters that are required to consider an additional lag and the minimum number of lags that are needed to capture the dynamics of a system. Since we are working with a small sample, we do not select a training sample; instead, we use the entire

1996–2010 time period to elicit the priors of the model. This strategy is suggested by Canova (2007) to address situations in which a training sample is not available.[3]

With respect to the prior for the time variation of the coefficients, we opt for a prior value of $k_Q = 0.05$, which effectively indicates that we are attributing 5 % of the uncertainty surrounding the OLS estimates to time variation (Kirchner et al. 2010). Furthermore, we set the conditions of $k_S = 0.025$ and $k_W = 0.025$.

In total, 400,000 iterations of the Gibbs sampler are produced to determine the estimation results. However, the first 200,000 iterations of the Gibbs sampler are discarded to allow for convergence. Moreover, to address possible autocorrelations among the draws, we only retain every 10th iteration. Consequently, the results that we present are based on the 20,000 remaining iterations. We discuss the details of the convergence diagnostics in Appendix A.4.

In the following sections of this paper, we present the median responses to shocks that have been normalised to monetary policy shocks of one percentage point and exchange rate shocks that equate to a 1 % appreciation; this normalisation allows these responses to be comparable across periods. Impulse responses from the reduced-form VAR in differences are transformed to levels by considering cumulative impulse responses. The transformation to levels produces an impulse response with respect to the mean growth value of a variable.

4 Results

The responses to monetary policy shocks and exchange rate shocks are presented in Figs. 1 and 2, respectively, in Sect. 4.1 and in Figs. 3 and 4, respectively, in Sect. 4.2. Section A contains certain additional results, such as 68 % credible intervals for responses to shocks up on impact and over a four-quarter time horizon. These credible intervals demonstrate high uncertainty related to results. The following discussion of results is based on the medians of the posterior impulse response distribution at a particular horizon. With respect to the 68 % credible intervals, the changes in the effects of monetary policy and exchange rate shocks over time are not significant.

4.1 Responses to monetary policy shocks

The estimated median responses to monetary policy shocks are presented in Fig. 1 and reflect the sign restrictions that have been imposed: in response to an unexpected increase of one percentage point in the interest rate, output and prices fall, whereas the nominal exchange rate initially appreciates. With respect to time variation, although the overall transmission of a monetary policy shock to interest rates appears to be

[3] This issue arises because there are no pre-1996 data available for the Czech Republic. We attempted to use data for Slovakia and the UK to construct priors; however, a simple VAR analysis revealed that the transmission mechanisms of the Czech Republic, Slovakia and the UK are not similar. For example, lagged GDP (modelled in terms of differences) produces opposite effects on GDP in the Czech Republic than in Slovakia or the UK. This discrepancy is a consequence of the restrictions imposed by the use of only two lags, a condition that is largely dictated by the methodology that is used for this study (a TVP VAR approach).

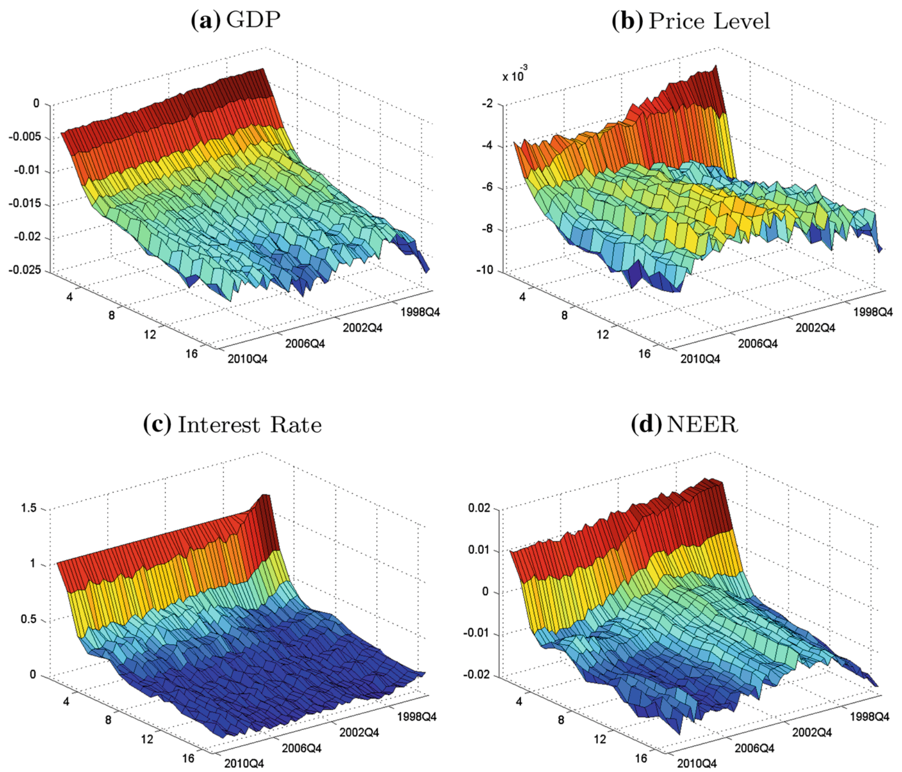

Fig. 1 Time-varying impulse responses to a monetary policy shock of 100 basis points

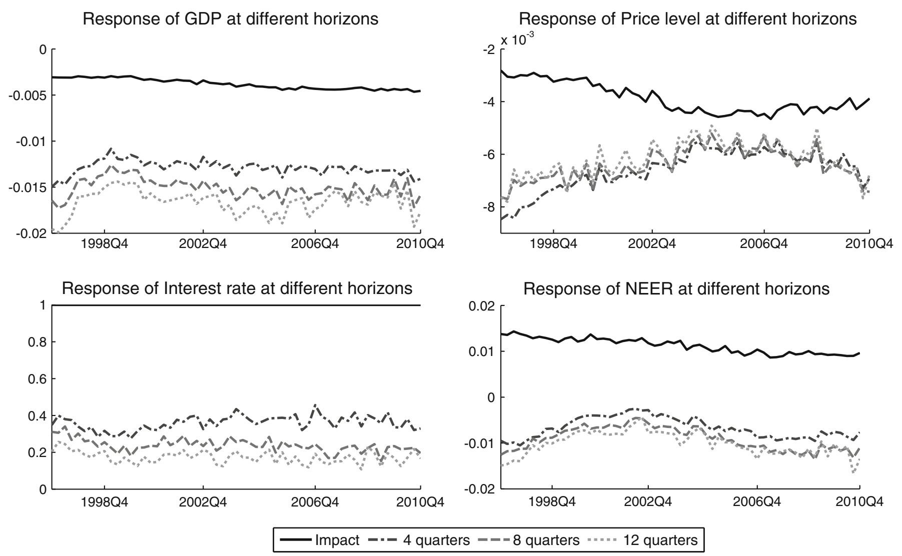

Fig. 2 Responses to monetary policy shocks at different horizons over different time periods

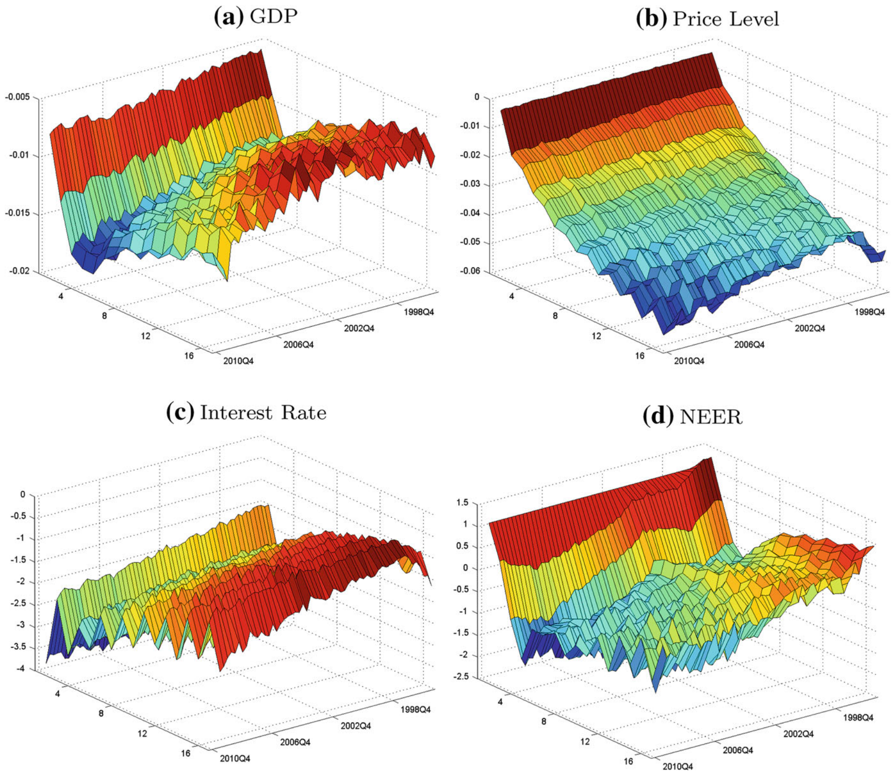

Fig. 3 Time-varying impulse responses to a 1 % exchange rate shock

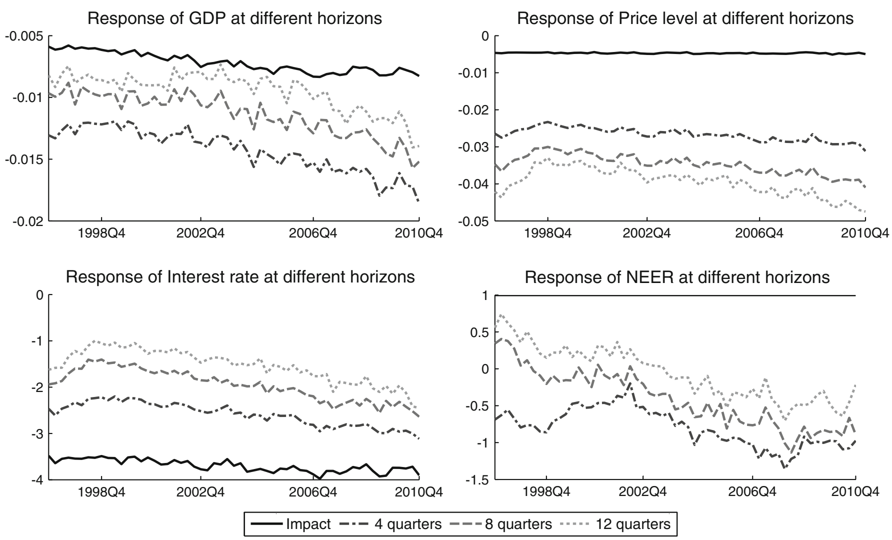

Fig. 4 Responses to exchange rate shocks at different horizons over different time periods

stable, the study results suggest that changes over time occur in the responses of price levels, exchange rates and (to a lesser extent) output to these types of shocks. More specifically, prices react more strongly to the initial impact of a monetary policy shock over time. However, over time, the maximum impact of these shocks becomes weaker, and these shocks produce less persistent effects on prices. This change in the nature of the responses to a monetary policy shock may suggest that economic agents learn about a new monetary policy regime and thereby begin to cope with monetary shocks in an increasingly effective manner over the course of time. The maximum impact of a monetary policy shock on prices occurs at approximately four to eight quarters after the shock has occurred. This result is consistent with the findings of Borys et al. (2009), who survey the previous studies that address monetary policy transmission in the Czech Republic and find that prices typically reach their nadir at four quarters after a shock.

Furthermore, certain interesting changes occur over time with respect to the transmission of monetary policy shocks to the exchange rate. The median impact of these shocks on the exchange rate is becoming stronger over time. At longer time horizons, the exchange rate appears to depreciate less during the 2001–2002 time period than during the other examined periods of this study. This observation most likely relates to expectations during 2001–2002 that the Czech koruna would appreciate as a result of the anticipated conversion of large privatisation revenues from euro to domestic currencies (Geršl 2006).

There are indications that the effects of an unexpected monetary policy shock on the real economy have changed over time. The impact of this type of shock on output appears to have begun to strengthen somewhat following the introduction of an inflation targeting regime in 1998. A weakening of the transmission of this type of shock can be observed at longer time horizons at the end of the examined sample; this phenomenon may reflect weaker monetary policy transmission during the 2008–2009 financial crisis. Interestingly, a similar weakening of medium-term monetary policy shock transmission is not observed with respect to price levels.

Figure 2 provides another examination at the evolution of the responses of output levels, price levels, the exchange rate and the interest rate to a monetary policy shock at specific time horizons. The only aspect of the transmission of this shock that appears to change during the recent financial crisis is that monetary policy shocks appear to have a somewhat greater effect on prices during the crisis than at the other times that are examined.

In addition, we present 68 % credible intervals for the responses to monetary policy shocks with respect to initial impact and over the course of a four-quarter time horizon. To assess the statistical significance of our results, we describe the responses to monetary shock over time in Appendix A.2 (see also Appendix A.3 for exchange rate shocks). The findings from these analyses suggest that changes in the transmission over time should be interpreted with caution because of the rather large uncertainty that exists in the estimates of this study. This issue is typical for studies that utilise time-varying VARs (Primiceri 2005); moreover, the sign restrictions identification approach contributes additional uncertainty to the findings of this study because there are typically many structural models that satisfy a particular set of sign restrictions (Fry and Pagan 2011). By contrast, the recursive approach that is based on the Cholesky decomposition of the variance–covariance matrix is an example of a method that produces only one structural model.

4.2 Responses to exchange rate shocks

It is clear that the effects of exchange rate fluctuations on economic activity are far from trivial and are dependent on balance sheet effects. However, exchange rate developments typically affect GDP through one of the following two channels: the expenditure-switching channel (a decrease in net exports following the appreciation of a currency) and the interest rate channel (which refers to the fact that exchange rate shocks are typically accompanied by interest rate decreases that can stimulate economic activity). In Fig. 3, we present the evolution of the impulse responses to a 1 % appreciation shock in the nominal effective exchange rate. Our results suggest that exchange rate appreciation weakens economic activity and that the effects of exchange rate shifts on GDP are strongest at approximately 1 year after a shock. We find that exchange rate shocks produce long-lasting effects on prices and that there is little evidence for any changes in the nature of these price responses over time. Although previously published literature has suggested that exchange rate pass-through has declined over time (Taylor 2000), we do not observe this trend in our data. We presume that our results for exchange rate pass-through reflect the low-inflation environment that began to prevail shortly after the adoption of an inflation targeting regime in the Czech Republic. We also find that exchange rate appreciation shocks are associated with lower interest rates. It is evident that exchange rate appreciation decreases inflation below the target and that a relatively unrestrictive monetary policy is, therefore, appropriate for maintaining an inflation rate that is close to the inflation target.

Figure 4 presents the reaction of variables to an exchange rate shock at different time horizons. These results suggest that the transmission of exchange rate shocks has become stronger over time, including during the period of the recent financial crisis. This effect is particularly evident for output. The pass-through of exchange rate shocks to prices appears to be stable over time. The reaction of interest rates to exchange rate fluctuations appears to be becoming more persistent during recent time periods than this reaction has been in the past.

5 Concluding remarks

In this paper, we analyse the evolution of the monetary policy transmission mechanism in the Czech Republic. The Czech economy has undergone 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. With respect to 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 that featured nearly double-digit inflation rates. Following gradual reductions in the inflation target, inflation in the Czech Republic has decreased to approximately 2 % in recent years. Therefore, it appears reasonable to model the monetary transmission mechanism as a time-varying phenomenon. Given this reasoning, we employ a recently developed Bayesian time-varying vector autoregression approach that includes stochastic volatility (Primiceri 2005). This flexible approach allows us to model the size of a shock hitting the economy in a time-varying manner to account for periods involving volatile economic developments, such as the current global financial crisis.

Our results suggest that prices decline following monetary tightening. The maximum impact of a monetary policy shock on prices occurs at approximately 4 to 8 quarters after the shock; this time span is consistent with the monetary policy horizon of the Czech National Bank. With respect to median responses, we find that the immediate responsiveness of prices to monetary policy shocks is increasing over time; by contrast, the maximum impact of these shocks is becoming weaker over time. Monetary tightening has a dampening effect on GDP, and monetary policy shocks have become somewhat more important over time with respect to producing fluctuations in economic activity. We also find that a monetary policy shock initially results in exchange rate appreciation. Interestingly, our results indicate that at longer time horizons, the exchange rate appears to depreciate less during the 2001–2002 time period than at other examined time periods of this study. During the 2001–2002 time period, markets expected domestic currency to appreciate significantly because the Czech government was expected to convert large privatisation revenues from euro to the Czech koruna. Clearly, this effect would be impossible to detect in the absence of a time-varying VAR model. However, we must also emphasise that the estimates of this study involve a large degree of uncertainty; this issue commonly arises for time-varying VAR models (Primiceri 2005). With respect to exchange rate shocks, our results indicate that exchange rate appreciation has a negative effect on GDP. We find that GDP reaches its bottom at approximately 1 year after an exchange rate appreciation. The effects of exchange rates on prices are long-lasting, and the results of this investigation suggest that exchange rate pass-through has remained largely stable over time.

Acknowledgements We are grateful to anonymous referee, Oxana Babecká, Tomáš Holub, Marek Jarociñski, Balázs Vonnák and seminar participants at the Czech National Bank and at the Economic Challenges in Enlarged Europe conference (Tallinn, Estonia) for helpful comments. We acknowledge financial support from the Grant Agency of Charles University (Grant Number 365111) and Grant Agency of Czech Republic (Grant P402/09/0965). The work was supported by Czech National Bank Project No. A2/2011. The views expressed here are those of the authors and not necessarily those of the Czech National Bank.

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