What We Know About Monetary Policy Transmission in the Czech Republic

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Collection of Empirical Results

CNB CZECH NATIONAL BANK

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What We Know About Monetary Policy Transmission in the Czech Republic: Collection of Empirical Results

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Abstract

This paper concentrates on describing the available empirical findings on monetary policy transmission in the Czech Republic. Besides the overall impact of monetary policy on inflation and output, it is useful to study its individual channels, in particular the interest rate channel, the exchange rate channel, and the wealth channel. The results confirm that the transmission of monetary impulses to the real economy works in an intuitive direction and to an intuitive extent. Our analyses show, however, that the global financial and economic crisis might have somewhat slowed and weakened the transmission. We found an indication of such a change in the functioning of the interest rate channel, where elevated risk premiums played a major role.

JEL Codes: C11, C32, E44, E52, E58.

Keywords: Bayesian, monetary policy transmission, time-varying parameters, VAR model.

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Nontechnical Summary

This note collects the empirical findings and results of analyses conducted on monetary policy transmission in the Czech Republic. We did not only study the overall impact of interest rate changes on inflation and output, but also explored individual transmission channels – the interest and exchange rate channels, the credit channel, and the asset price/wealth channel. We carry out a number of empirical analyses. Typically, our sample starts in mid-1990s and ends approximately in 2010, depending on particular analysis. The actual estimations were carried out in 2010-2012.

Based on our sample, our estimations show that the reaction of prices to the unexpected interest rate change reaches its maximum after about 5–6 quarters. The reaction of prices to unexpected exchange rate shock is faster and reaches its maximum after one year.

Focusing at interest rate channel, about 60% of the transmission of monetary policy interest rate changes to client interest rates occurs within one month for loans with short fixations and deposits with short maturity. As for the exchange rate pass-through analysis, 1% appreciation leads to a decrease in import prices of 0.7%, consumer prices decrease by 0.11%.

Overall, the results show that the transmission of monetary impulses to the real economy worked in an intuitive and reasonable direction and to an intuitive and reasonable extent. At the same time, the results are in line with previous analyses conducted using the data from the Czech Republic. Our analyses showed, however, that the initial phase of the global financial and economic crisis might have somewhat slowed and weakened the transmission. Elevated risk premiums were the main cause of changes in the functioning of the interest rate channel.

On the other hand, some analyses suggest that the mechanisms investigated are showing a tendency toward their pre-crisis functioning. The more fundamental changes in transmission mechanisms that have probably been taking place in economies with more sophisticated financial markets, will most likely affect the Czech Republic only indirectly.

1. Introduction

The world financial and economic crisis not only was a difficult period of time for financial institutions, enterprises, and households, but also posed a great challenge for authorities responsible for economic policies. Besides various anti-crisis measures taken by many countries' governments, a number of central banks around the globe started a phase of monetary policy loosening in order to at least partly compensate for the collapse of both domestic and foreign demand. In line with this tendency, the Czech National Bank (CNB) in the second half of 2008 entered an unprecedented period of cutting interest rates, which reached a record low (technical zero) level in November 2012 (a 2W repo rate of 0.05%).

Simultaneously, commercial banks' ability to draw liquidity from the Czech central bank contributed to a decline in the spread between the CNB's repo rate and money market interest rates. This spread, after having soared at the end of 2008, decreased and stabilized at levels lower than before, but still rather higher than in the pre-crisis period. This stability meant that the CNB's continuously declining policy rates were basically fully transmitted to inter-bank interest rates.

The CNB's interest rate cuts were aimed at delivering the desired positive monetary impulse mainly by influencing commercial banks' interest rates on loans and deposits. These rates are pivotal for households' and firms' consumption, saving, and investment decisions. In reality, however, the observed client rates did not exhibit such an obvious downward trend as the policy rates did. The reason was simple: the CNB had to push against the naturally pro-cyclical behavior of the financial sector. As a matter of fact, the riskiness of economic agents tends to increase during crises. This is usually reflected in more prudent behavior by commercial banks when deciding on loans and setting interest and other credit conditions. In the case of domestic mortgages there was also a temporary but marked increase in Czech government bond yields following the deterioration in the public finance situation. This prevented mortgage interest rates from declining more robustly and instantly. Interest rates on loans for firms, on the other hand, showed a clear tendency to follow the declining rates on the money market.

Thus, as the economic crisis deepened, it was legitimate to ask whether monetary policy transmission in the Czech Republic was still functioning properly or whether it had weakened compared to what we used to know/think about it based on our pre-crisis experience. Repeated questions from CNB board members on the functioning of the transmission mechanism proved that policy makers were seriously considering this issue in the context of possible further interest rates cuts and potential unconventional monetary policy instruments.

Following this internal debate in the CNB, this note collects the empirical findings and results of analyses conducted on monetary policy transmission in the Czech Republic. Typically, our sample starts in mid-1990s and ends approximately in 2010, depending on particular analysis. The actual estimations were carried out in 2010-2012. As the transmission of monetary policy is a complex process,[1] the description of which is not trivial, we not only studied the overall impact of monetary

[1] To be able to conduct monetary policy effectively, the central bank needs to know the direction and lags with which monetary policy decisions influence the targeted variables. However, because of the many different channels of monetary policy transmission, these transmission parameters are not only uncertain, but also changing. Ascertaining the existence and strength of particular channels is therefore crucial for the conduct of monetary policy.

policy, but also explored individual transmission channels. Thus, we employed an aggregated approach to identify the total effect of monetary policy on inflation and output as well analyzing the individual transmission channels (the interest and exchange rate channels, the credit channel, and the asset price/wealth channel).

The results confirmed that the transmission of monetary impulses to the real economy worked in an intuitive and reasonable direction and to an intuitive and reasonable extent. Moreover, the results are broadly in line with previous analyses presented in the available literature and consistent with the basic characteristics of the Czech economy. The analyses showed, however, that the initial phase of the global financial and economic crisis might have somewhat slowed and weakened the transmission. This change in the functioning of the interest rate channel was caused by elevated risk premiums. Some analyses (a time-varying parameter VAR model and exchange rate pass-through analysis) suggest, however, that the mechanisms investigated are showing a tendency toward their pre-crisis functioning. The more fundamental changes in transmission mechanisms that have probably been taking place in economies with more sophisticated financial markets will most likely affect the Czech Republic only indirectly.

The rest of the paper is organized as follows. Section 2 provides a stylized description of monetary policy transmission and summarizes the empirical literature on monetary policy transmission in the Czech Republic. In section 3, VAR, Bayesian VAR, and time-varying parameter VAR are used to estimate the magnitudes and lags of monetary policy transmission as well as the effects of the global economic and financial crisis on the transmission. Section 4 focuses on three individual channels of transmission – the interest rate channel, the exchange rate channel, and the wealth channel. Finally, section 5 concludes.

2. Setting the Stage: Stylized Description of Monetary Policy Transmission and Empirical Literature Review

2.1 Stylized description of transmission from policy interest rates to inflation

The basic scheme of transmission from policy interest rates to inflation is depicted in Chart 2.1.1. Therefore, we focus on conventional monetary policy in this description and abstract from various unconventional measures. The primary transmission channels include the interest rate channel, the exchange rate channel, the credit channel, and the asset price channel. The importance of particular channels in a particular economy depends on the openness of the economy, its financial system development, as well as the role of the banking sector. Other channels of monetary policy transmission include the expectations channel and the risk-taking channel.

Through the *interest rate* channel, monetary policy influences the real economy by changing key interest rates. As consumption, saving, and investment decisions are typically based on long-term interest rates, the first necessary condition for effective monetary policy is a functioning channel of transmission of monetary policy interest rates to financial market interest rates.[2] The changes in

[2] This transmission typically works due to the no-arbitrage relationship between these two types of interest rates. The elevation of spreads between monetary policy interest rates and financial market interest rates during the global financial crisis was in contradiction with this assumption. This paper abstracts from this issue.

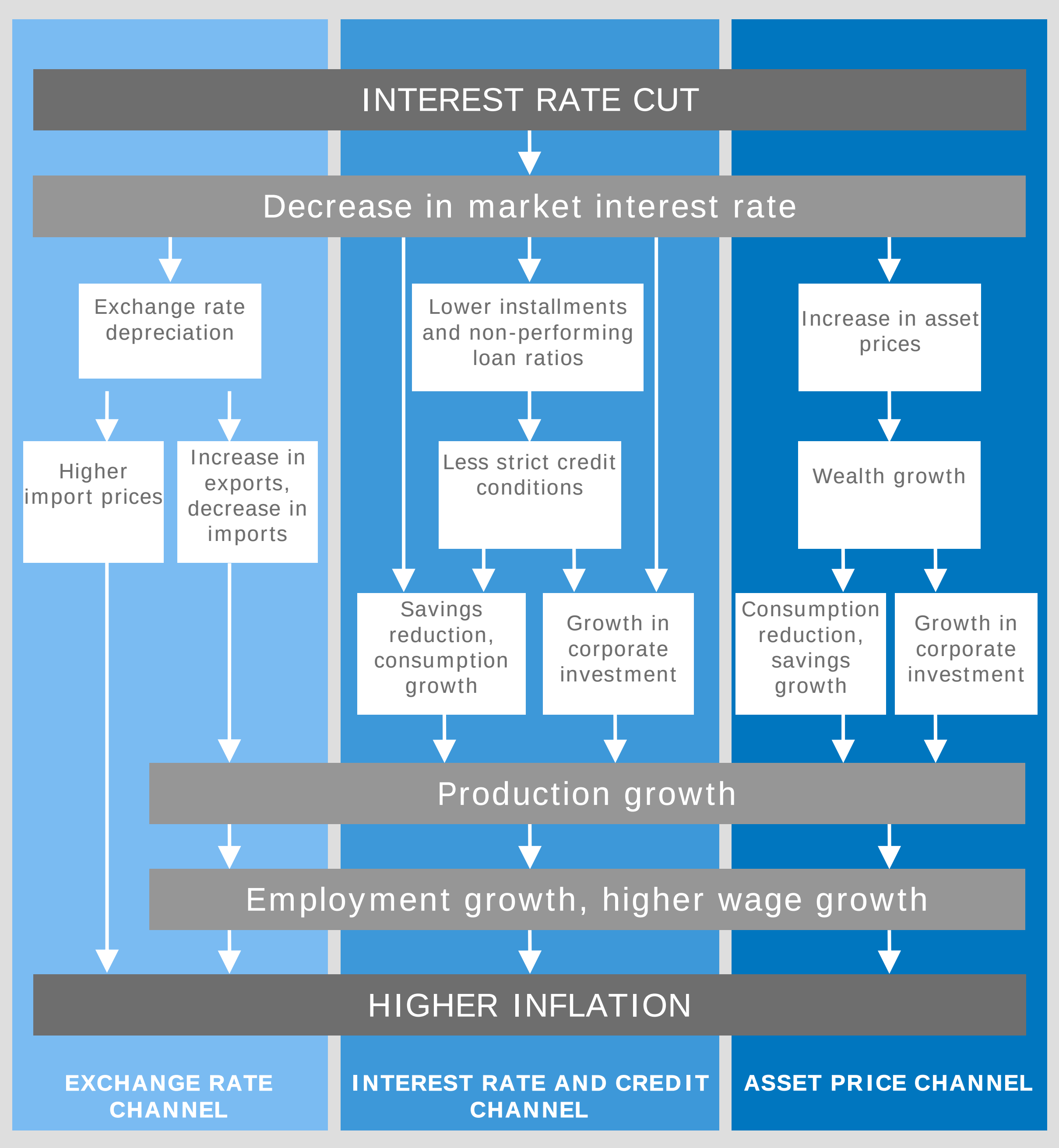
financial market interest rates influence the costs of interbank borrowing, to which banks subsequently react by adjusting their deposit interest rates (the alternative bank financing cost). At the same time, the changes in the cost of bank financing influence the interest rates on loans provided by banks. In the end, client interest rates on deposits and loans enter the optimization process of economic agents in terms of intertemporal substitution or valuation of economic projects.

The existence of imperfect information and substitution between financial assets in bank-based economic systems cause the transmission of interest rate changes to be inhomogeneous across economic agents (*credit channel*).[3] The heterogeneity of effects on firms is caused mainly by the availability and value of collateral, and is reflected in the availability and conditions of loans (*financial accelerator, balance-sheet channel*); firms with worse financial positions are affected by a monetary policy tightening more than firms with good financial positions. In the case of banks, the *credit channel* is linked to *agency costs* and the strength of bank balance sheets, which determine the external premium of bank financing and banks' access to external sources and influence the changes in the credit supply after monetary policy changes.

The effect of the *exchange rate* on inflation is especially substantial for very open economies. An exchange rate shock has a direct effect on consumer inflation through prices of imported consumer goods. Indirect effects include the price effects of substitution between domestic and foreign goods, changes in the domestic prices of raw materials and intermediate goods, and changes in the monetary policy stance.

[3] For the purposes of this publication, the existence of the credit channel in the Czech Republic was analyzed using the methodology of Iacoviello and Minetti (2008) using monthly data for 2004–2009. However, the results do not appear to be robust enough, therefore credit channel analysis is not included in the following text.

Figure 2.1.1: Primary Transmission Channels Between Change in Key Monetary Policy Interest Rates and Inflation



The *asset price channel* can cause asset price adjustments induced by interest rate changes to influence the value of households' and firms' balance sheets, which is reflected in their confidence in the economy. The effectiveness of this channel for households is conditional on their perceptions about whether growth in real estate prices and financial asset prices increases wealth and is a source of consumption spending. In the case of firms, growth in a firm's stock market value makes investment capital relatively cheaper (*Tobin's Q*).

2.2 Empirical Studies for the Czech Republic

Empirical studies of monetary policy transmission in the Czech Republic employ a wide range of approaches. Arnoštová and Hurník (2005) use the basic VAR framework to estimate two VAR models – one employing the whole data set available at the time and the other one looking at the period after the monetary policy regime change only (1998–2004). A VAR model was also used for the identification of monetary policy transmission in Holub (2008). This study identifies the transmission of a shock to the exchange rate, whereas in the case of a shock to interest rates the transmission occurred in the direction predicted by economic theory in only one of two alternative estimations and was unexpectedly fast, achieving its maximum after two quarters only.

Bayesian VAR with time-varying parameters was used to evaluate changes in monetary policy transmission in the Czech Republic in Franta et al. (2013). The authors found the exchange rate pass-through to be relatively stable, while the responsiveness of prices to monetary policy shocks increased. A comparison of the results obtained using four estimation techniques (VAR, structural VAR, Bayesian VAR, and factor-augmented VAR) is provided in Borys et al. (2009). The authors find both prices and output to decline after a monetary policy tightening, with the maximum response occurring after about one year. In addition, the authors compare their results with other empirical studies on monetary policy transmission in the Czech Republic. Darvas (2012) uses VAR with time-varying coefficients to compare the transmission of a monetary policy shock for the Czech Republic, Hungary, and Poland on the one hand and the euro area on the other. Havránek and Rusnák (2013) review 67 studies on transmission and use meta-analysis to examine transmission patterns. According to their results, the average transmission lag is 29 months and prices decrease by 0.9% after a one percentage point interest rate hike. Greater financial development is associated with slower transmission, while greater openness is associated with faster transmission. Finally, Rusnák et al. (2013) use meta-analysis to study the occurrence of the price puzzle in the literature on monetary policy transmission. Based on a review of 70 papers, they find evidence of publication bias. Once misspecifications are filtered out, the average impulse response shows a decrease in prices after an interest rate cut.

Other related studies include Pruteanu (2007), who studies the impact of monetary policy on bank lending in the Czech Republic. Babecká-Kucharčuková (2009) focuses on the exchange rate channel of monetary policy transmission in the Czech Republic. She finds that exchange rate shock transmission to prices is relatively fast and the exchange rate pass-through does not exceed 25–30%.

3. Empirical Models of Monetary Policy Transmission

In this section the monetary policy transmission mechanism is formalized using two approaches. Typically, our sample starts in mid-1990s and ends approximately in 2010, depending on particular analysis. The actual estimations were carried out in 2010-2012. The first approach is a vector autoregressive (VAR) model developed for a small open economy. The second approach proposes a Bayesian VAR model with sign restrictions and time-varying parameters (TVP-VAR). The estimated models account for interactions among the key macroeconomic variables and therefore provide a basic knowledge of the intensity and delays of monetary policy impulse transmission. Both the VAR and BVAR models confirm effective transmission from the interest rate to inflation. Monetary policy shock transmission is strongest at a horizon of about 12 months in the Bayesian framework, while it takes 1.5 times longer when the simple VAR model is applied instead. In both models a monetary policy shock (interest rate hike) invokes an immediate appreciation of the exchange rate, which vanishes within a year. Transmission of an exchange rate shock to inflation is somewhat faster compared to interest rate pass-through and has the expected size given the share of imported goods in the consumption basket. The transmission weakened during the global economic and financial crisis, but gained in strength again subsequently.

The empirical literature has converged to a consensus on the aggregate effects of monetary policy transmission in a VAR framework.[4] The interest rate channel and the exchange rate channel are the two most widely examined channels of the transmission mechanism affecting the real economy. The present section first estimates a simple VAR model where the coefficients are determined by the data behavior, i.e., no priors are applied. Relatively short time series and volatility in the data may distort the signals and responses and, as a consequence, lead to wide confidence intervals or insignificant outcomes. For these reasons, the Bayesian framework is then applied. Bayesian models partially overcome this problem thanks to a set of priors and restrictions applied to the data. However, the presence of too many restrictions may lead to a situation where the model's outcome will be a product of theoretical assumptions neglecting the true behavior of the data. From this perspective, Bayesian and classical models tend to complement each other.

3.1 VAR on Quarterly Data

The aggregate modeling of monetary policy transmission follows the previous literature applying VAR to a small open economy. The model was first developed by Mojon and Peersman (2001). Arnoštová and Hurník (2005) applied a slightly modified version to the Czech data. The model has a standard VAR structure (3.1.1):

$$z_t = c_0 + \Psi(L)z_{t-1} + \Phi * x_{t-1} + \mu , \qquad (3.1.1)$$

where $z_t : [y_t, cpi_t, rs_t, neer_t, m_t]$ is a vector of endogenous variables, and $x_t : [wcpi_t, y_t^{EA}, rs_t^{EA}]$ is a vector of exogenous variables. The following notation is used in the vectors: *y* – real GDP, *cpi* – consumer price index, *rs* – three-month money market interest rate (3M PRIBOR or 3M EONIA), *neer* – nominal effective exchange rate against 21 trade partners, *m* – broad money (M2 aggregate), *wcpi* – world commodity price index; superscript EA denotes the euro area. The model is estimated on quarterly data from 1998q1 to 2010q4 collected from the IFS, ARAD, and the CNB's internal database. All variables except the interest rate are converted to indices (2005 average = 100) and taken in logarithms. Seasonal adjustment is applied if appropriate.

The transmission of a shock is based on an impulse response function and requires the application of restrictions on the residuals. Arnoštová and Hurník (2005), Mojon and Peersman (2001), and Elbourne and de Haan (2009) identify short-run structural shocks using Cholesky decomposition: $Ae_t = Bu_t$, where A is a lower triangular matrix with ones along the diagonal, B is diagonal matrix, and e_t and u_t are observed and unobserved residuals, respectively. By construction, the impulse response based on Cholesky decomposition is sensitive to the ordering of the variables in the VAR. To overcome this inconvenience, generalized impulse responses are used instead. Generalized

[4] See Christiano et al. (1999) for an overview.

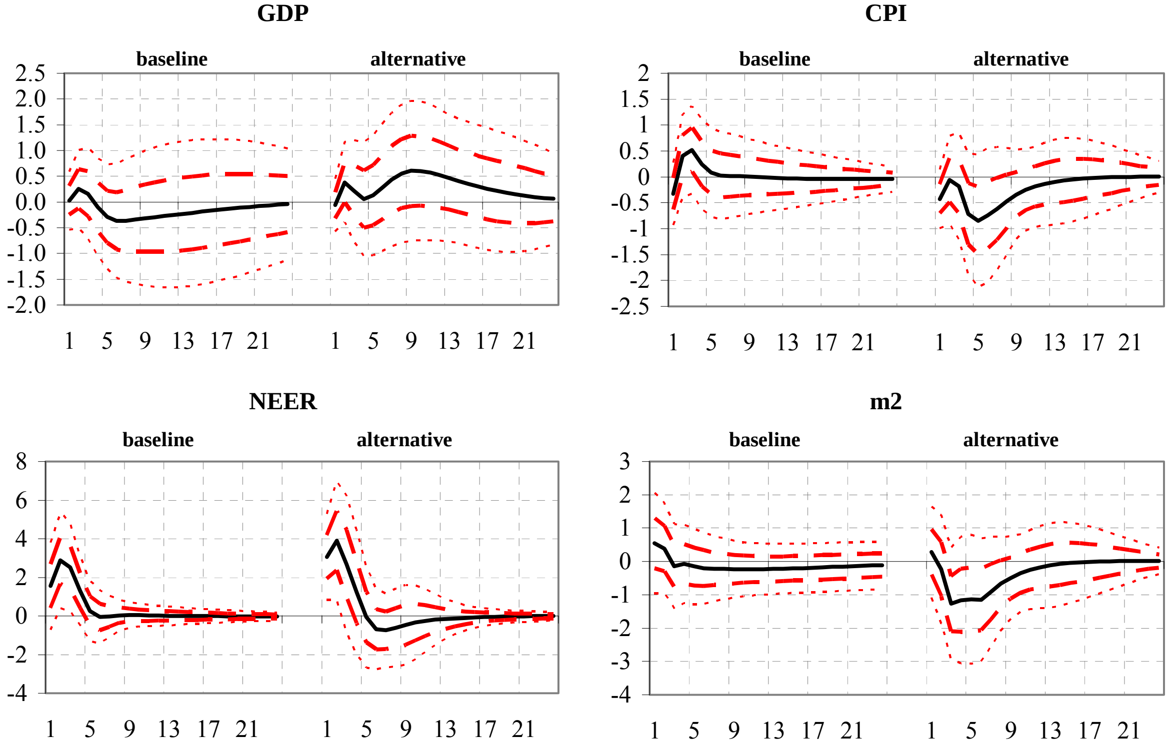
impulse responses have a variable-specific Cholesky factor and do not depend on the ordering of the variables.[5]

Standard unit root tests indicate that all the variables in the model follow an I(1) process; at the same time the model passes the stability test. While the Johansen cointegration test detected a cointegration relation among the variables, the inclusion of an error correction equation in the model makes the impulse responses unstable. For this reason, and following Sims et al. (1990), the selected specification is estimated in levels.

An unexpected rise in the interest rate of 1 pp leads to a quick and statistically significant exchange rate appreciation (of about 3% after two quarters), which vanishes in five quarters after the shock (Figure 3.1.1, baseline). The economic slowdown attains its maximum of about 0.4% in the sixth quarter, but is statistically insignificant. The reaction of the money base is close to zero and statistically insignificant as well. Consumer prices follow the price puzzle: after an appreciation shock prices rise instead of declining. The shock fades out after five quarters. To resolve the price puzzle, an alternative version of model (3.1.1) is estimated in which exogenous variables are taken with two lags. The reasons for the inclusion of lagged values are inertia and expectations about future shocks. For instance, if oil prices achieved high levels some time ago, producers and suppliers of goods to final consumers are likely to increase prices, being constrained by fixed-term contracts and menu costs. Under the alternative assumption about the exogenous variables (Figure 3.1.1, alternative) the price puzzle disappears. The monetary policy tightening is now effective and forces prices to decrease by almost 0.9%. The money base decreases as well. The exchange rate appreciation is higher than in the baseline specification. However, the impact on real GDP is now counterintuitive in terms of its direction, although it does not differ significantly from zero.

[5] The two approaches, however, produce comparable results.

Figure 3.1.1: VAR on Quarterly Data: Impulse Response to Monetary Policy Shock, %



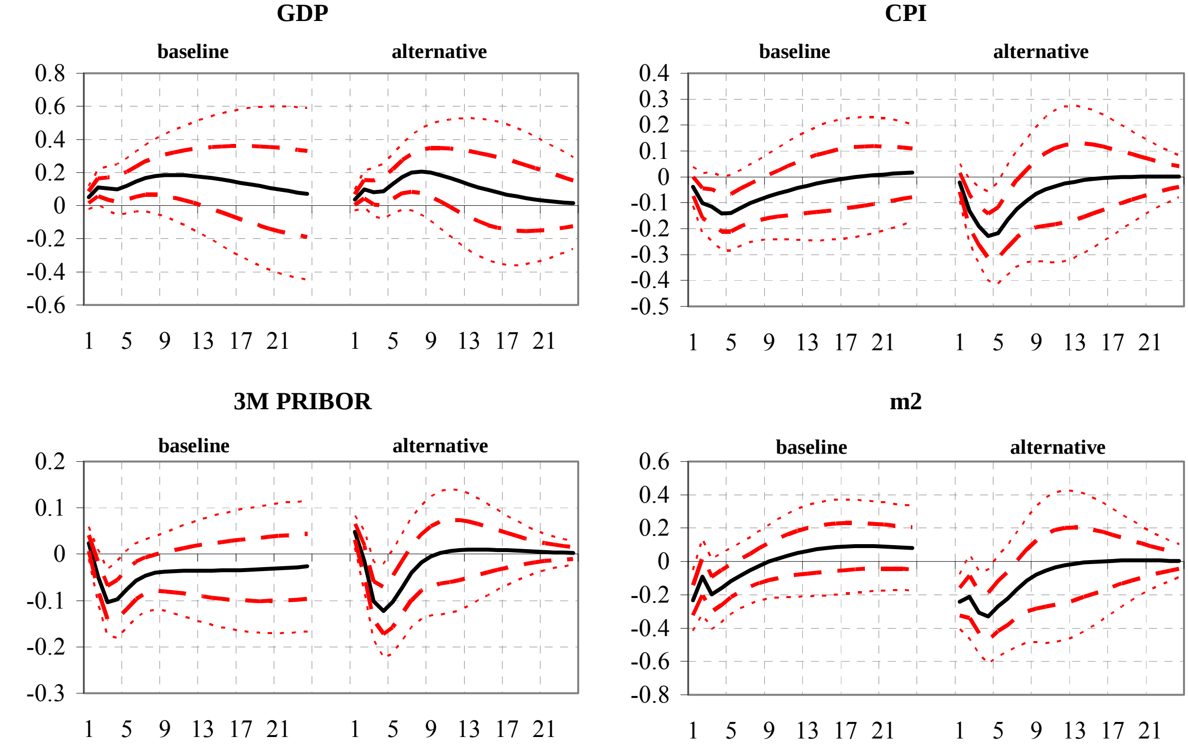
Note: The impulse response to an unexpected interest rate increase of 1 pp is plotted on the vertical axis. The horizontal axis shows periods in quarters. The dashed lines denote confidence bands of ±1 and ±2 standard deviations. An increase in the exchange rate means appreciation.

Similarly to the monetary policy shock, an exchange rate appreciation of about 0.3 CZK/EUR does not significantly affect GDP in either the baseline or the alternative specification (Figure 3.1.2).[6] In contrast, the decrease in consumer prices is statistically significant and reaches 0.14% in the baseline specification and 0.23% in the alternative specification, in both cases four quarters after the shock. The interest rate decreases by 0.10–0.12 pp and the money base falls by 0.20–0.33% three or four quarters after the shock.

The results based on an empirical VAR model for a small open economy show the effectiveness of transmission of the CNB's monetary policy. One should nevertheless keep in mind the limits of the present analysis. Most of the empirical results are explainable by economic theory. However, like other studies focused on new EU member states, the model is estimated on relatively short time series and produces wide confidence bands for the impulse responses.

[6] This result contradicts some earlier studies. For instance, the VAR model applied in Holub (2008) finds that the impact of an exchange rate shock on the output gap is statistically significant and occurs in the expected direction, attaining its maximum in the 5th or 6th quarter after the shock.

Figure 3.1.2: VAR on Quarterly Data: Impulse Response to Exchange Rate shock, %



Note: The impulse response to an unexpected exchange rate appreciation of 0.30 CZK/EUR (~ 1%) is plotted on the vertical axis. The horizontal axis shows periods in quarters. The dashed lines denote confidence bands of ±1 and ±2 standard deviations. An increase in the exchange rate means appreciation.

3.2 Bayesian VAR on Monthly Data

The use of VAR models for the investigation of the transmission mechanism in the Czech Republic involves two problems. The first problem is related to the short time series available for the analysis. To avoid the crucial structural change, the data set should not start before 1998, when the Czech National Bank adopted inflation targeting. This constraint implies only a few available observations. Standard VARs contain a large set of parameters to estimate, and this, together with a low number of observations, results in unreliable estimates.

The second issue is general and concerns the identification of structural shocks. The usual recursive identification that prescribes which endogenous variables are allowed to react contemporaneously to a structural shock yields various puzzles. The most famous is the price puzzle (Eichenbaum, 1992), which describes the situation where the price level rises immediately after an unexpected monetary policy tightening. Next, recursive identification is not necessarily appropriate in every case. For example, the restriction on the contemporaneous effect between the exchange rate and interest rates is unrealistic for an open economy targeting inflation.

To address the issues mentioned above, the VAR model for the Czech Republic is estimated using the Bayesian approach and structural shocks are identified using sign restrictions. The Bayesian approach allows us to impose a priori information not directly related to the observed data and thus to mitigate the over-parameterization problem. The sign restrictions identification scheme represents a more realistic identification of structural shocks. Regarding the estimation and identification of shocks we closely follow Vonnák (2010), who estimated a similar model for several CEE countries. Bayesian VAR estimation and identification of shocks based on sign restrictions for the Czech Republic can also be found in Borys et al. (2010). The model employed in this paper differs from Vonnák (2010) and Borys et al. (2009) in the type of prior assumed and in how the sign restrictions are implemented.

The following reduced-form VAR(p,q) is estimated:

$$y_t = a_0 + \sum_{j=1}^{p} A_j y_{t-j} + \sum_{j=1}^{q} B_j x_{t-j} + \varepsilon_t \qquad (3.2.1)$$

where y_t (x_t) for $t = 1,...,T$ is an $M \times 1$ vector of endogenous (exogenous) variables, a_0 is an $M \times 1$ vector of constants, A_j for $j = 1,..., p$ and B_j for $j = 1,..., q$ are coefficient matrices, and the vector of disturbances is i.i.d, $\varepsilon_t \sim N(0, \Sigma)$.

The model is estimated on seasonally adjusted monthly data for industrial production, the CPI, the three-month money market rate (3M PRIBOR), and the exchange rate (CZK/EUR). The variables are in levels and cover the period 1998M1–2011M2. The set of exogenous variables includes the commodity price index and 3M EURIBOR.[7]

The number of lags of the reduced-form model (3.2.1) is based on the autocorrelation of residuals test (Ljung-Box Q autocorrelation test). The model is estimated with five lags of endogenous variables (p=5) and one lag of exogenous variables (q=1).

The VAR model is estimated using the Minnesota prior (Doan, Litterman, and Sims, 1984, and Litterman, 1986), i.e., the prior for the vector of the intercepts and the elements of the coefficient matrices is assumed to be distributed normally with mean α^{PR} and variance V^{PR}:

$$\alpha \equiv \mathrm{vec}\left(\left[a_0 \; A_1 ... A_p \; B_1 ... B_q\right]'\right) \sim N\left(\alpha^{PR}, V^{PR}\right).$$

The Minnesota prior embodies the belief that the time series involved in the analysis behave like a random walk. So, we assume that the prior mean equals one for the coefficients at the first lag of the variable that is on the left-hand side of the equation. For the other coefficients the prior mean equals zero. The prior variance expresses how much the process is assumed to describe the observed time series tightened to the random walk. It is a diagonal matrix parameterized by a set of three parameters (so-called hyperparameters): overall tightness λ, weight θ, and decay ϕ.

[7] Data sources: IMF IFS (index of industrial production, CPI, commodity price index, 3M EURIBOR) and the ARAD database of the Czech National Bank (3M PRIBOR and CZK/EUR). An increase in the exchange rate represents depreciation.

The standard deviation of the prior distribution equals $\sigma_{ii} = \lambda l^{-\phi}$ for the coefficients at the lags of the variable on the left-hand side (i.e., the *i*-th equation and the *i*-th variable) and $\sigma_{ij} = \theta\lambda l^{-\phi} \frac{\hat{\sigma}_i}{\hat{\sigma}_j}$ for the coefficients at all lags of the other variables (i.e., the *i*-th equation and the *j*-th variable, $i \neq j$). The lag of the variable is denoted by l and $\frac{\hat{\sigma}_i}{\hat{\sigma}_j}$ is the ratio of the standard deviations of the AR(1) processes of the respective variable on its own lag. The ratio of standard deviations accounts for the different units of the variables. The values of the hyperparameters are chosen as follows: $\lambda = 0.2^{0.5}$, $\theta = 0.5^{0.5}$, and $\phi = 1$, which is the standard in the literature. The prior for the intercepts and exogenous variables is uninformative – the standard deviation of the prior distribution is equal to $10^{5/2}$.

The posterior distribution of the coefficients in the case of the Minnesota prior is normal:

$$\alpha \mid y \sim N\left(\alpha^{PO}, V^{PO}\right), \tag{3.2.2}$$

where

$$V^{PO} = \left[\left(V^{PR}\right)^{-1} + \left(\hat{\Sigma}^{-1} \otimes (X'X)\right)\right]^{-1}$$
$$\alpha^{PO} = V^{PO}\left[\left(V^{PR}\right)^{-1}\alpha^{PR} + \left(\hat{\Sigma}^{-1} \otimes X\right)' y\right].$$

X denotes a matrix of endogenous and exogenous variables, i.e.,

$$X = \begin{bmatrix} \left(1, y'_0, .., y'_{1-p}, x'_0, ..., x'_{1-p}\right) \\ \left(1, y'_1, .., y'_{2-p}, x'_1, ..., x'_{2-p}\right) \\ \vdots \\ \left(1, y'_{T-1}, .., y'_{T-p}, x'_{T-1}, ..., x'_{T-p}\right) \end{bmatrix} \text{ and } y = \begin{bmatrix} y'_1 \\ \vdots \\ y'_{T-1} \end{bmatrix}.$$

$\hat{\Sigma}$ denotes an OLS estimate of Σ.

The identification of the monetary policy and exchange rate shocks is based on a combination of contemporaneous and sign restrictions. Similarly to the standard recursive approach, we assume that industrial production and consumer prices do not react contemporaneously to monetary policy and exchange rate shocks. Next, a rise in the interest rate and appreciation of the exchange rate for at least six months is assumed after an unexpected monetary policy tightening. On the other hand, and unexpected exchange rate appreciation is assumed to bring about a fall in the interest rate for at least six months. The opposite sign of the reaction of the interest rate for the two shocks enables the identification.[8]

The problem of structural shock identification is equivalent to finding a matrix S such that $SS^{-1} = \Sigma$, i.e., a matrix which inverse transforms uncorrelated structural shocks to the reduced-form residuals. The problem is that such a matrix is not unique and some restrictions need to be imposed. In other words, there are not enough estimated reduced-form parameters to compute all the structural ones (e.g., Greene, 2003). In the system of four endogenous variables, six restrictions need to be imposed on the matrix elements. In our case, the combination of sign restrictions and contemporaneous restrictions does not identify the system exactly. The inertia of industrial production and consumer prices for the monetary and exchange rate shock implies zeros in the intersection of the 3rd and 4th columns and the 3rd and 4th rows. Sign restrictions are implemented following Fry and Pagan (2010) using Givens rotations G, i.e., orthonormal matrices such that $SGG^T S^{-1} = SS^{-1}$. So, in general, to preserve the zero contemporaneous effect on the first two endogenous variables, the following Givens rotation is used:

$$G(\theta) = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & \cos(\theta) & -\sin(\theta) \\ 0 & 0 & \sin(\theta) & \cos(\theta) \end{bmatrix},$$

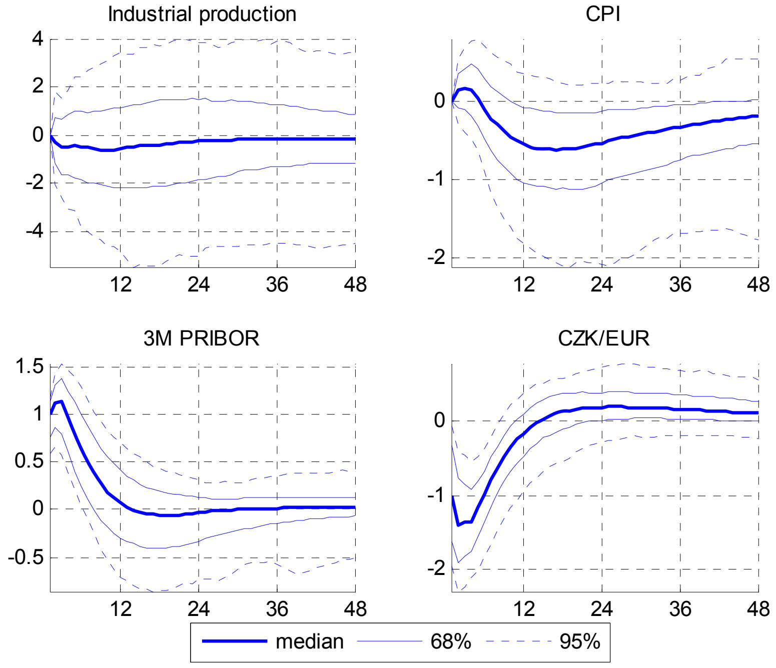
where $\theta \in [0, \pi]$.

The numerical implementation of the estimation and identification proceeds in two steps. In the first step, we take a set of 2,000 draws from the posterior distribution of the model parameters (3.2.2). As a second step, we draw θ from a uniform distribution $U(0, \pi)$ and check the sign restrictions. We repeat the procedure until we get a sample from the posterior distribution satisfying the sign restrictions of size 500. The posterior distribution of the responses is then captured by the median and centered 68% and 95% of the sampled posterior distribution.

Figure 3.2.1 reports the impulse responses of the endogenous variables over 48 months. The monetary policy shock is normalized to unit size. An unexpected monetary policy tightening leads to a fall of the price level, with the bottom occurring after 16 months or so. The median response of the CPI exhibits the price puzzle. The result, however, is insignificant in the sense that the centered 68% of the posterior distribution of the response contains zero. The exchange rate appreciates after a positive monetary policy shock and returns to its initial value after a year.

[8] For a detailed exposition of the concept of sign restrictions identification, see, for example, Canova and De Nicoló (2002) and Uhlig (2005).

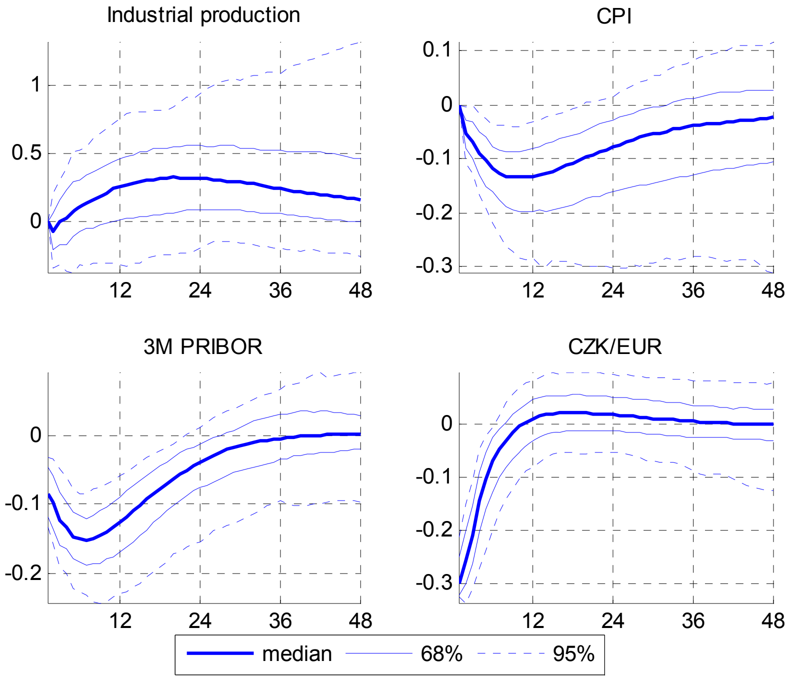
Figure 3.2.1: Bayesian VAR on Monthly Data : Impulse Response to Monetary Policy Shock, %



Note: The horizontal axis denotes the time after the shock in months.

Figure 3.2.2 demonstrates that after an unexpected exchange rate appreciation of CZK 0.3, the interest rate falls by more than 10 basis points. Moreover, the price level decreases, with the bottom occurring after slightly less than a year. The rise in industrial production is similar to that found in section 2.1. The rise can be explained by real convergence and structural changes related to the inflow of capital, neither of which is captured by the model. Such long-run effects are expected to influence the correlations between the endogenous variables in our dataset in a direction that is counterintuitive from the business cycle perspective.

Figure 3.2.2: Bayesian VAR on Monthly Data: Impulse Response to Exchange Rate Shock, %



Note: The horizontal axis denotes the time after the shock in months.

Several robustness exercises were carried out. First, we tested the set of exogenous variables as in Vonnák (2010): German industrial production, German consumer prices, the German short-term interest rate, and the German nominal effective exchange rate. Almost identical results were obtained for the responses of the interest rate and the exchange rate. Some minor differences can be observed for industrial production and consumer prices (a larger fall in industrial production and a smaller fall in consumer prices after the monetary policy tightening). Next, the results are robust to the choice of time for which the sign restrictions are imposed – 4, 6, 8, and 10-month horizons were tested.

3.3 TVP-VAR

In this section, changes in the transmission mechanism during the recent economic crisis are examined. To that end, we employ a vector autoregression model with time-varying parameters (TVP-VAR). It is important to emphasize that TVP-VARs are heavily parameterized models and thus the uncertainty of the estimates is large because of the short time span of available data – the problem is even more severe than in the case of standard VARs. Therefore, we do not discuss the model and results in detail and provide a few notes only. Moreover, we focus solely on the monetary policy shock.

The model specification closely follows Primiceri (2005) and we refer the reader to that study for details. Darvas (2012) estimates a TVP-VAR model for the Czech Republic. Our approach differs from the two papers in how the monetary policy shock is identified.

The original VAR model (2.2.1) is extended to include the assumption of time-varying parameters (no exogenous variables are further considered):

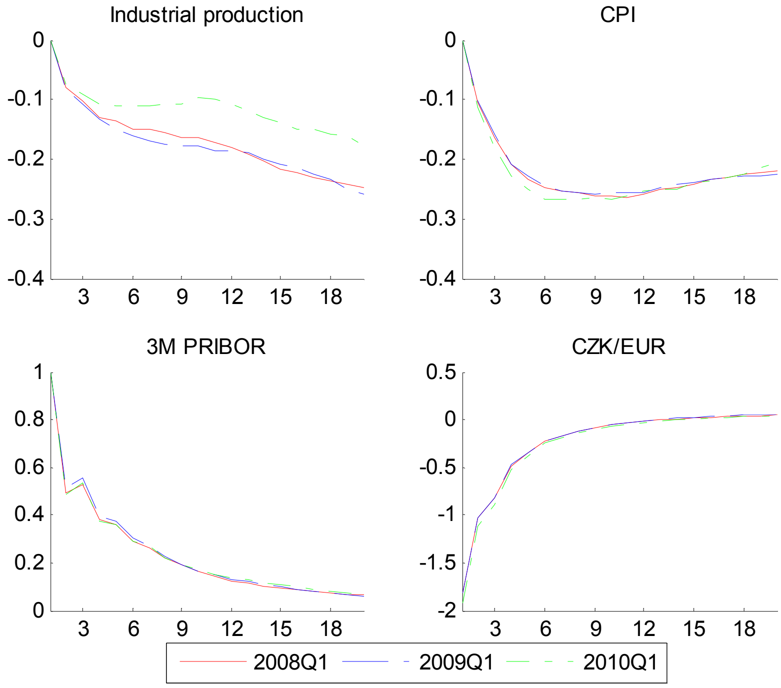
$$y_t = a_t + \sum_{j=1}^{p} A_{j,t} y_{t-1} + \varepsilon_t \qquad \varepsilon_t \sim N(0, \Sigma_t), \tag{3.3.1}$$

where the intercepts and elements of matrices A_j follow a random walk.
The variance of disturbances, Σ_t, is also assumed to change over time and consists of processes following a random walk and a geometric random walk. The monetary policy shock is identified as in the case of the BVAR model, i.e., we combine contemporaneous and sign restrictions. A detailed exposition of the methodology combining TVP-VARs and the sign restriction identification scheme can be found in Franta (2011).

The model is estimated using quarterly data covering the period 1994Q1–2010Q4. The set of endogenous variables contains real GDP, the CPI, the short-term interest rate (3M PRIBOR), and the CZK/EUR exchange rate. The reason for employing quarterly data instead of monthly data is to capture as much of the dynamics as possible with a low number of lags, because each additional lag represents a large increase in the number of parameters to estimate. The model is estimated with two lags $(p = 2)$. As in Darvas (2009) the coefficients on the variables different from the variable on the left-hand side lagged by two periods are assumed to be zero. The responses are based on a sample of 10,000 iterations of the Gibbs sampler. The first 1,000 iterations are discarded to diminish the influence of the initial conditions.

The next figure shows the impulse responses of the normalized monetary policy shocks. It suggests that the transmission mechanism in the Czech Republic has not been affected much by the recent financial and economic crisis. Note that the result is robust to the assumed prior variation in the coefficients, i.e., our prior belief about how much the coefficient can change over time.

Figure 3.3.1: Impulse Responses to Monetary Policy Shock, 2008Q1, 2009Q1, 2010Q1



Note: The horizontal axis denotes the time after the shock in quarters.

3.4 Alternative to TVP-VARs

Instead of using TVP-VARs to infer changes in the transmission mechanism, one can look at the results of BVARs estimated for the period 1998M1–2011M2 and the sub-period 1998M1–2008M9. On the one hand, the necessity of having a huge number of parameters to estimate in TVP-VARs is avoided. On the other hand, the results can detect the change but cannot provide detailed information on its timing and profile.

The following two figures present impulse responses based on a BVAR estimated on the sub-period 1998M1–2008M9 (the Lehman Brothers bankruptcy). Again, the effect of the monetary policy shock and the exchange rate shock is examined.

Figure 3.4.1: BVAR, 1998M1–2008M9: Impulse Responses to Monetary Policy Shock

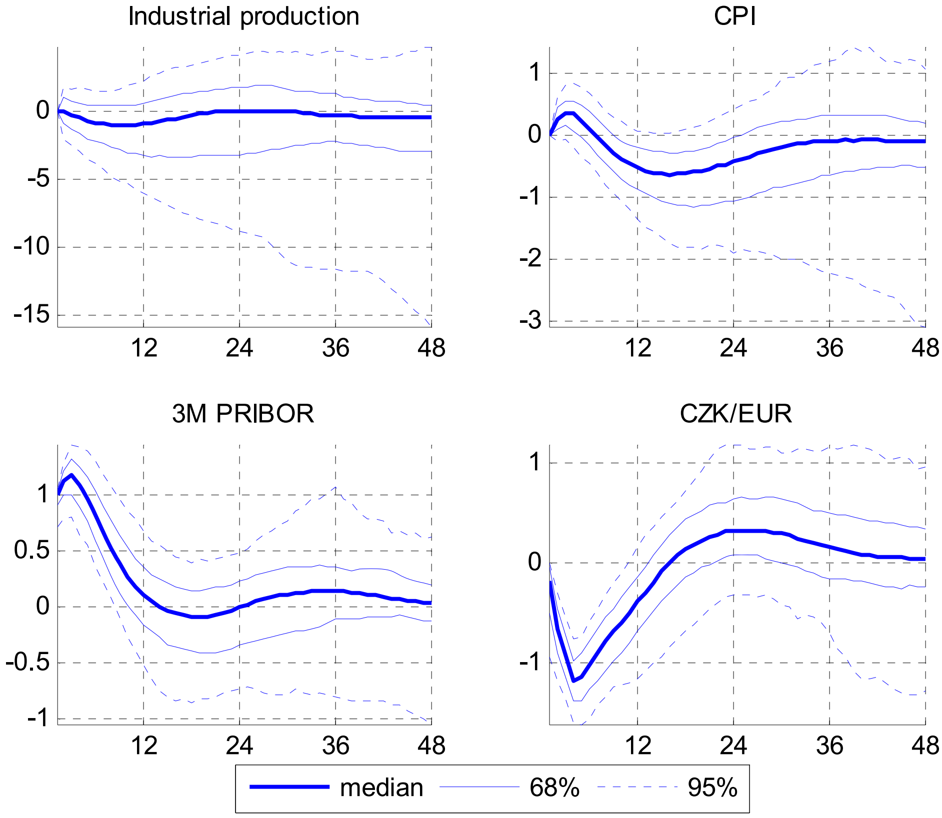
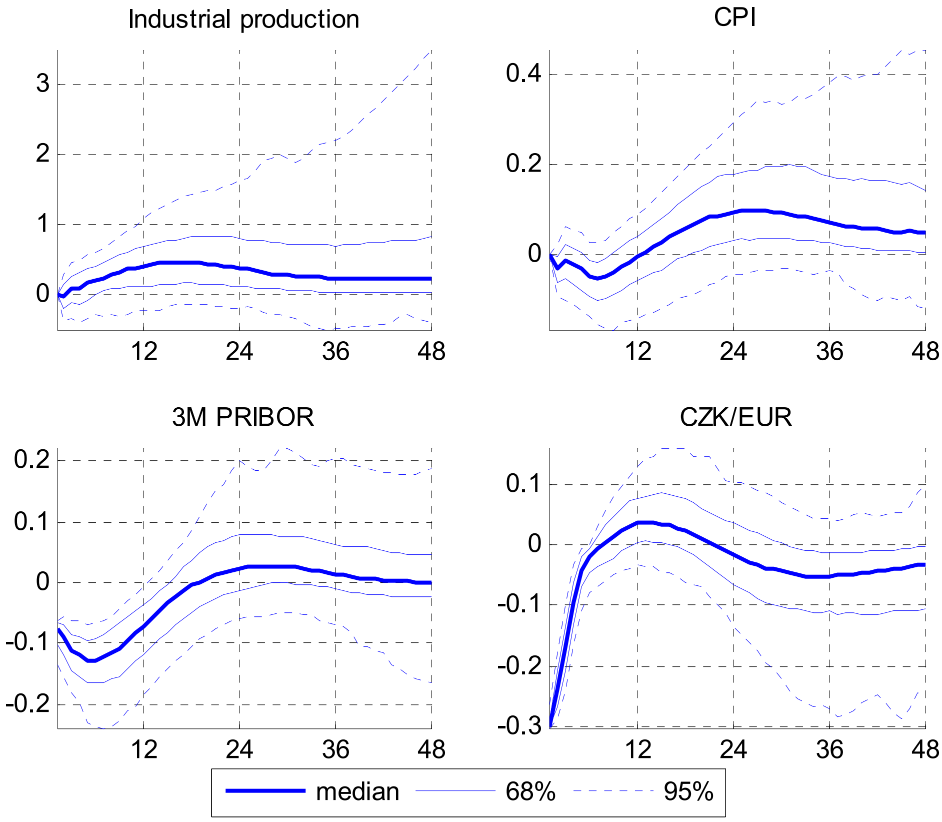


Figure 3.4.2: BVAR, 1998M1–2008M9: Impulse Responses to Exchange Rate Shock



The responses are very similar to those presented in Figure 3.2.1. Thus, no significant change is detected in the transmission of monetary policy and exchange rate shocks in the Czech economy during the recent crisis. The only exception is the effect of an unexpected exchange rate appreciation on the level of consumer prices. Focusing on the pre-crisis period, prices do not decrease as much as when the whole period is taken into account. The reason could be that part of the identified exchange rate shock is accounted for by the trend appreciation of the Czech koruna. This link may have been weakened during the recent recession.

4. Analyses of Individual Channels of Transmission

This section explores three individual channels of transmission – the interest rate channel, the exchange rate channel, and the wealth channel.

4.1 Interest Rate Channel

The key relation for the interest rate channel is the connection between the costs of bank financing on the interbank market, which the central bank influences with its monetary policy, and client interest rates on loans and deposits. The transmission between market rates and client interest rates is affected mainly by the elasticity of loan demand, the degree of market competition, and the existence of asymmetric information. Studies by Crespo-Cuaresma, Egert, and Reininger (2007) and Tieman

(2004) contain results for the aggregated interest rate channel in the Czech Republic. These studies find incomplete transmission of changes in monetary policy rates. Horváth and Podpiera (2009) study interest rate transmission using monthly bank-level data in the Czech Republic. They find quite a fast reaction of client rates to changes in money market interest rates, although this reaction is not complete for certain types of rates. We apply the same approach in the following analysis, which presents updated results.

The link between market rates and bank interest rates (the interest rate pass-through) is typically evaluated using an error correction framework as described by the following equation:[9]

$$\Delta br_{i,t} = \sum_{l=0}^{q} \alpha_{0i,l} \Delta mr_{t-l} + \sum_{k=1}^{p} \alpha_{1i,k} \Delta br_{i,t-k} + \beta_{0i}(br_{i,t-1} - \beta_1 mr_{t-1} - \mu_i) + \varepsilon_{i,t} \qquad (4.1.1)$$

where $br_{i,t}$ denotes the *i*-th bank interest rate at time *t*, mr_t represents the money market rate, and μ is a constant that quantifies the spread of bank interest rates vis-à-vis money market rates. The relation captures both the long-term and short-term dynamics of the money market pass-through to bank interest rates. The long-term pass-through is described by coefficient β_1. If $\beta_1 = 1$, the pass-through is regarded as complete. Coefficient α_0 reflects the short-term dynamics, while coefficient β_0 stands for the speed of adjustment. According to Hendry (1995) (β_1 - α_0)/ β_0 indicates the mean adjustment lag at which the market rate is fully passed through to the bank rate.

The analysis accounts for monthly data for the period January 2004 through December 2009. Table 4.1.1 lists the interest rates that we analyzed: six loan rates (four rates on loans to the non-financial sector and two rates on loans to the household sector) and two deposit rates.

[9] Following De Bondt (2005) and De Graeve et al. (2007).

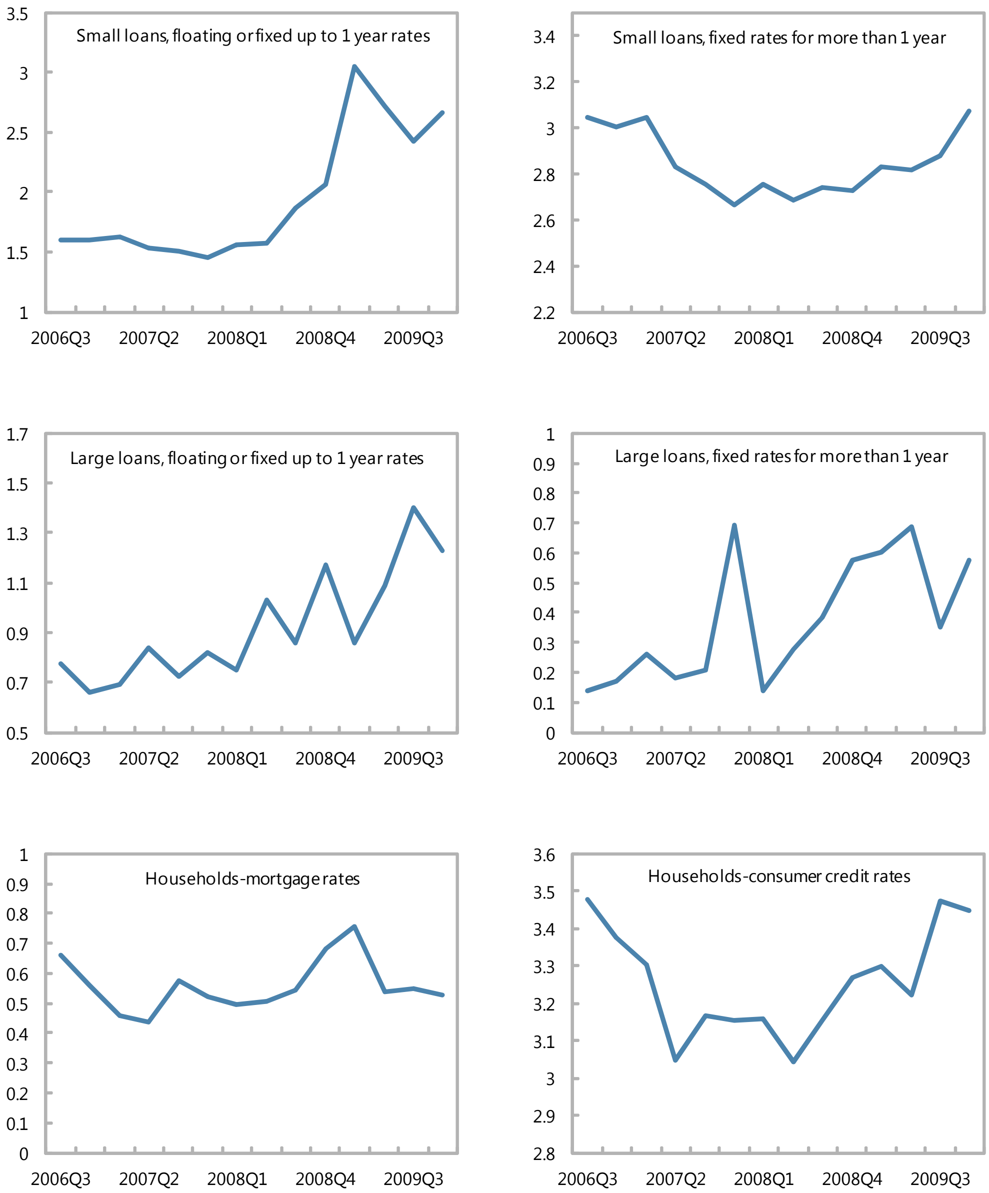
Table 4.1.1: Data Description

Loan rates		**Number of banks in the sample**
For loans to the non-financial sector		
Small loan, floating rate or rate fixed for up to 1 year	Loan amount up to CZK 30 million, rate floating or fixed for up to 1 year	17
Small loan, rate fixed for more than 1 year	Loan amount up to CZK 30 million, rate fixed for more than 1 year	14
Large loan, floating rate or rate fixed for up to 1 year	Loan amount more than CZK 30 million, rate floating or fixed for up to 1 year	17
Large loan, rate fixed for more than 1 year	Loan amount more than CZK 30 million, rate fixed for more than 1 year	9
For loans to the household sector		
Mortgage rate	Loan for house or apartment purchase	10
Consumer rate	Loan generally for durable goods	14
Deposit rates		
Short-term deposit rate	Deposits with maturity above one day and less than two years	20
Long-term deposit rate	Deposits with maturity above two years	8

We used the available data for all commercial banks. Some banks do not have data for some types of loans/deposits (as they did not provide that type of loan or attract that type of deposit). For this reason, the number of banks in the analysis varies across types of loans/deposits.

For each bank and each type of rate described above we used monthly contract-level data (interest rates and the corresponding volume of loans), from which we derived the bank-month median, mean, standard deviation, and skewness. We preferred using medians for our regressions because a normality test performed on the monthly distributions – the skewness/kurtosis test (conceptually similar to the Jarque-Bera test) – systematically rejects the null hypothesis of normal distribution. The computation of the standard deviation turned out to be very illustrative for the period of financial turmoil. Figure 4.1.1 presents the evolution of the asset-weighted quarterly average of the standard deviations for each type of interest rates for the period of interest.

Figure 4.1.1: Banks' Standard Deviations for Interest Rates on Loans (weighted by banks' total assets)



Source: Own calculations.

There are differences apparent in the evolution of the standard deviations across different lending rates. The standard deviations of bank rates indicate differentiation by banks among clients according to their perceived riskiness, so they can proxy for banks' credit standards. For the non-financial sector, the standard deviations remained elevated throughout 2009 without signs of decreasing, while for mortgages the peak was attained in the first quarter of 2009, which coincides with the peak of the financial crisis (characterized by uncertainty in the Czech financial market). This standard deviations for mortgages reverted to their pre-crisis values in mid-2009. This suggests that mortgages are perceived as a safer investment by banks. This is probably due to the existence of large underlying collateral for mortgage loans. In contrast, the perceived risks for loans to non-financial corporations, especially those with floating rates and rates fixed for up to one year, remained high at the end of 2009 as a consequence of uncertainty about the economic outlook.

In order to capture the appropriate cost of funds for each loan rate we compute the average correlation between loan rates on the one hand and money market rates (1M PRIBOR, 3M PRIBOR, 6M PRIBOR, 1Y PRIBOR), each bank's deposit rates, and selected long-term financial market rates (FRA3_6, FRA3_6, IRS5, IRS10, 5Y government bond yield (YTM5), 10Y government bond yield (YTM10) – as a proxy for the opportunity costs of funds – on the other hand.

Table 4.1.2 displays the average correlations between the bank rates for newly granted loans and the term structure of money and financial market rates and bank-specific deposit rates for the period January 2004–December 2009.

Table 4.1.2: Average Correlations Between Bank Rates and Money and Financial Market Rates and Deposit Rates

Loan rates	1M PRIBOR	3M PRIBOR	6M PRIBOR	1Y PRIBOR	FRA3_6	FRA3_9	IRS5	IRS10	YTM5	YTM10	Deposit rate up to 2Y	Deposit rate above 2Y
Non-financial sector												
Up to 30mil CZK floating and fixed to 1y	0.69	0.71	**0.71**	0.70	0.63	0.60	0.36	0.27	0.42	0.36	0.63	0.50
fixed more than 1y	0.35	0.37	0.38	0.40	0.33	0.33	0.33	0.29	0.39	**0.49**	0.39	**0.53**
More than 30mil CZK												
floating and fixed to 1y	0.63	0.65	**0.65**	0.64	0.60	0.59	0.37	0.26	0.38	0.29	0.63	0.42
fixed more than 1y	0.42	0.45	**0.45**	**0.45**	0.42	0.42	0.37	0.33	0.38	**0.45**	0.38	0.39
Household sector												
Mortgage loans	0.32	0.36	0.38	0.40	0.25	0.19	0.19	0.27	0.41	**0.60**	0.30	**0.50**
Consumer loans	0.12	0.13	0.15	0.12	-0.01	-0.02	0.06	0.17	0.23	0.18	0.02	0.09
Deposit rates												
with maturity up to 2Y	**0.87**	0.86	0.85	0.83	0.85	0.83	0.59	0.43	0.46	0.31		
with maturity longer than 2Y	0.43	0.47	0.45	0.47	0.41	0.44	0.46	0.48	0.46	**0.55**		

Note: 1) individual bank correlations are very heterogeneous, ranging from 0.3 to 0.9; 2) average based only on 5 banks´ individual correlations, for which both data on deposit and loan rates were available; 3) based only on 4 banks´ individual correlations. 4) The bold numbers are the highest correlations for each type of interest rate.

Floating rates and rates fixed for up to one year are mostly correlated with the 6M PRIBOR. Rates on small loans to the non-financial sector fixed for more than one year and on mortgage loans are mostly correlated with long-maturity deposit rates and YTM10, respectively. This suggests that the pricing of these loans shifted toward longer maturity benchmark rates. However, given that only eight banks have usable data series for long-maturity deposits, the average correlations between loan rates fixed for more than one year and long-maturity deposit rates took into account only these banks. At the same time, long-maturity deposits rates are, on average, the most correlated with YTM10 rates. This probably results from YTM10 constituting an alternative investment opportunity to long-term deposits. For this reason, and because long-term deposits have usable data series for only a limited number of banks, we use YTM10 as a benchmark for the cost of funds for loans with rates fixed for more than one year and mortgage loans. In the case of large loans with rates fixed for longer than one year, the correlations with 1YPRIBOR and YTM10 are almost equal (see Table 4.1.2). This is due to the heterogeneity in the correlations, with some bank loan rates leaning toward the 1Y PRIBOR and others' toward YTM10. We choose YTM10 since large banks were more correlated with this rate than with the 1Y PRIBOR.

The correlations between rates on consumer loans and PRIBOR rates, financial market rates, and deposit rates are very small and in numerous cases negative, so they are not considered for the cointegration analysis. It appears that with very few exceptions, banks' pricing of consumer loans does not change much with changes in money market and financial rates.

Based on the correlation results presented in Table 4.1.3, we pursued the cointegration analysis between loan rates and the corresponding benchmark rates as theoretically described by equation (4.1.1) for the period January 2004–December 2009. The long-term relationship is estimated using Dynamic Ordinary Least Squares (DOLS)[10] and the short-term dynamics using Swamy's (1970) random coefficient model, which captures the dynamic heterogeneity. Table 4.1.3 presents the results.

Table 4.1.3: Transmission of Changes in Money Market Rates to Client Rates During January 2004–December 2009

	Loan/deposit rates	Proxy of cost of funds	Immediate pass-through α_0	Final pass-through β_1	Speed of adjustment β_0	Adjustment speed in month $(\beta_1-\alpha_0)/\beta_0$
Loans	**corporations - small**					
	Floating or rate fixed to 1Y rate	6M PRIBOR	0.64^{***}(0.15)	0.75^{***}(0.04)↓	-0.43^{***}(0.09)↓	1m
	Fixed for more than 1Y rate	YTM10	0.33(0.27)↓	1.19^{***} (0.15)	-0.4^{***}(0.11)	3m
	corporations - large					
	Floating or rate fixed to 1Y rate	6M PRIBOR	0.62^{***}(0.19)	0.82^{***}(0.03)↓	-0.6^{***}(0.09)	1m
	Fixed for more than 1Y rate	YTM10	0.67 (0.99)	0.83^{***}(0.11)	-0.64^{***}(0.13)↓	2m
	household - mortgage loans	YTM10	-0.09(0.07)	0.91^{***}(0.04)↑	-0.28^{***}(0.03)	3m
Deposits	With maturity longer than 2Y	1M PRIBOR	0.63^{***}(0.09)	0.93^{***}(0.02)↓	-0.35^{***}(0.07)	1m
	With maturity longer than 2Y	YTM10	-0.04(0.26)	0.73^{***}(0.07)↓	-0.47^{***}(0.09)↓	2m

Note: Symbols ***,**, and * denote statistical significance of the parameters at the 1%, 5%, and 10% significance level. Standard errors are in parentheses. The symbol ↓ denotes parameters that are statistically significantly lower (in absolute value at the 10% significance level) than the parameters estimated on the 2004–2008 sample. The symbol ↑ denotes the parameter that is statistically significantly higher (at the 10% significance level) than the parameter estimated on the 2004–2008 sample. Other parameters are not statistically significantly different. The speed of adjustment was rounded to whole months.

The significance and sign of the error correction coefficient – the so-called "speed of adjustment" – confirms the presence of a mechanism to bring bank rates back to their long-term equilibrium for all except interest rates on consumer credit. Cointegration analysis rejects the existence of a cointegration relationship between interest rates on consumer credit and money market rates. This is

[10] Kao and Chiang (2000) conclude that DOLS may be more promising than OLS and FMOLS for the estimation of panel cointegration. Practically, besides a bank dummy to account for the fixed heterogeneity the regression contains the leads and lags of the first differences of the explanatory variable. We considered a maximum of four lags and leads and then eliminated the insignificant variables.

probably due to elevated asymmetric information costs, which seem to be the most pertinent in this market.

The results also show a significant long-run pass-through parameter (β_1) for all the interest rates. Complete long-run pass-through is observed only for rates on small loans fixed for more than one year, for mortgage loan rates, and for deposits with maturity of up to two years. The long-run pass-through of interest rates on large loans is incomplete, possibly due to the dominant share of relationship-based banking. In the case of small loans with floating rates or rates fixed for up to one year and long-term deposits, incomplete transmission is apparent.

Regarding the immediate pass-through (α_0), within one month we find positive and significant coefficients for floating rates or rates fixed for up to one year for both small and large loans, and for rates on deposits with short maturity. This confirms that floating rates follow money market rates, although the pass-through is not complete. Approximately 60 percent of the transmission for these loans takes place within a month. On the other hand, bank loan rates which are set for more than one year and rates on deposits with long maturity do not respond at all within one month. When accounting only for the pre-crisis period, the results show that small loan rates fixed for more than one year also react within a month. This suggests that during the period of financial distress, rates fixed for more than one year were the most affected by the uncertainty in the financial markets and macroeconomy.

The impact of the financial crisis is analyzed by comparing the updated results with the estimates from the 2004–2008 sample using the same model. Statistically significantly different coefficients are labeled (Table 4.1.3). During the financial crisis, the overall transmission decreased especially for floating and short-term fixed interest rates for corporate clients and deposit interest rates. The decrease in market rates transmitted less to these interest rates then it would have done prior to the crisis. The transmission remained unchanged but slowed (the speed of adjustment declined) for corporate loans with interest rates fixed for longer periods. At the same time, the results suggest that the relation between mortgage interest rates and the yield on long-term government bonds strengthened. This might have been due to the crowding-out effect of the issuance of government bonds, which caused mortgage interest rates to remain high for a longer period.

In connection with the financial and economic crisis, banks perceived higher riskiness and tightened their credit standards. In the second half of 2008, there is an apparent increase in variability of client interest rates across all types of loans (see Figure 4.1.1). The differentiation is systematically higher for small loans to corporate clients and for consumer credit to households than for large loans to corporate clients and mortgages to households. In the last period, a decline in the dispersion and a return to pre-crisis levels was recorded only for mortgages and large loans to corporate clients with maturity exceeding one year.

4.2 Exchange Rate Channel

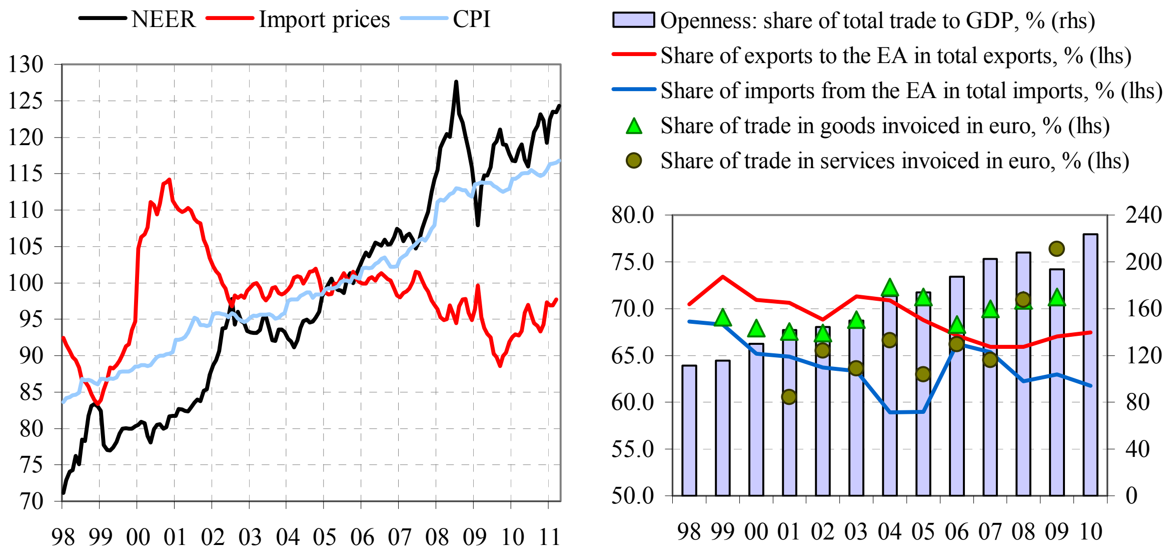
A shock to the exchange rate is transmitted to consumer price inflation directly through a change in prices of imported goods in the consumption basket, as well as indirectly via imported inputs and imported intermediate goods.[11] Therefore, the elasticity of the price of an imported good expressed in the domestic currency to an exchange rate shock depends on the willingness of the seller of the good on the local market to adjust the markup. In the absence of price rigidity (fully adjustable markups), the exchange rate pass-through to the price of an imported good is unity.[12] However, it should be much less than unity for the producer price of a domestically fabricated good, since only the price of imported components incorporates the exchange rate. The reference basket used for the CPI index is composed of tradable and nontradable goods and services, which consist partly of items with regulated prices. For this reason, consumer prices react much less to an exchange rate shock than do producer prices[13] (Figure 4.2.1). At the same time, the Czech economy has a high degree of openness, mostly toward the euro area. Despite a slight decline in the euro area's weight in total trade, exports to and imports from the euro area remain important. Together with the significant share of contracts invoiced in euro (which has indeed increased further recently), this implies high exchange rate pass-through to import prices. The observable correlation between the nominal effective exchange rate (NEER) and import prices (see the figure below) supports this hypothesis. However, the conclusion about the relation between the NEER and the CPI is not so obvious.

[11] An additional indirect impact on inflation is generated via the so-called expenditure-switching channel, where a change in the price of an imported good expressed in the local currency has an impact on aggregate demand and prices of domestically produced goods.

[12] The pass-through is zero when producer currency pricing is applied. Pure producer currency pricing or pure local currency pricing is rarely observable, and the macro data usually incorporate a combination of producer and local currency pricing, so the exchange rate pass-through to import prices is positive but less than unity.

[13] Assuming that the markup is the same as for an imported good.

Figure 4.2.1: Stylized Facts

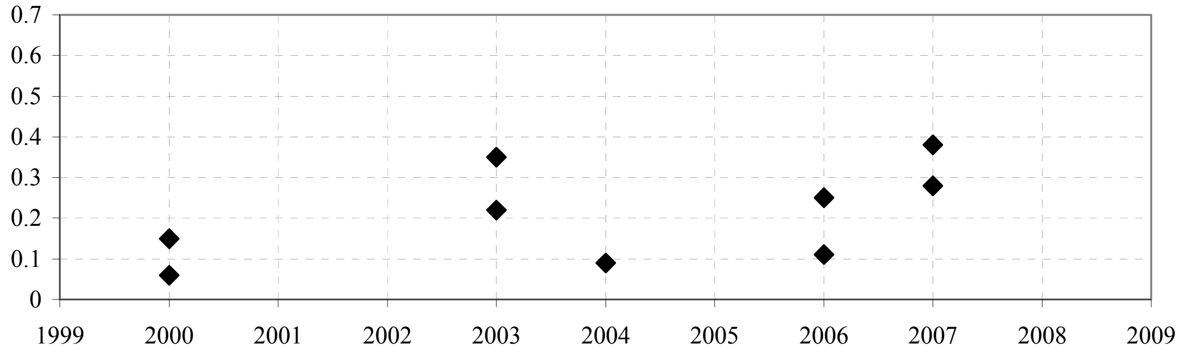


Note: The nominal effective exchange rate (NEER) with respect to 21 trade partners. The last observation refers to April 2011 for the exchange rate and the CPI, and to March 2011 for import prices. An increase in the NEER means appreciation. There is a structural break in the currency invoicing series due to a methodological change in 2004.

Source: ARAD, Eurostat, Analyses of the Czech Republic's Current Economic Alignment with the Euro Area (2010), International Role of the Euro (various issues), CNB staff calculation.

Previous studies measuring the exchange rate pass-through to Czech inflation have found that the reaction to a 1% shock varies from close to zero to almost 0.4% (Figure 4.2.2). Therefore, the transmission of changes in the exchange rate to inflation is incomplete, i.e., less than one hundred percent of the shock. According to Devereux and Yetman (2010), the pass-through is not full owing to price rigidity. Microeconomics explains incomplete transmission by mark-up and pricing-to-market strategies (Krugman, 1987). Other impediments to consumer prices reacting to a shock include menu costs and the costs of distribution and transportation. From the macroeconomic perspective, the pass-through is low in countries that have adopted inflation targeting (Taylor, 2000). While previous studies give some insight into the degree of shock transmission, their results are hardly comparable due to substantial heterogeneity in data, time periods, and empirical methodologies.

Figure 4.2.2: Exchange Rate Pass-Through to Czech Inflation Based on Literature Review



Note: The response of inflation to a 1% exchange rate shock (in %) is on the vertical axis. The horizontal axis plots the last year of the data sample used for the pass-through estimation. Square points denote the following studies: Babecká Kucharčuková (2009), Beneš et al. (2003), Bitāns (2004), Darvas (2001), Holub (2008), Korhonen and Wachtel (2005), Maria-Dolores (2008), Mihaljek and Klau (2001).

The present analysis is based on the McCarthy (2007) model with pricing chain, supply, and demand variables. The pricing chain contains import prices (*pm*), producer prices (*ppi*), and consumer prices (*cpi*). The demand shock enters the model via the index of industrial production (*y*); the supply shock is approximated by the seasonally adjusted price of Brent crude oil in USD/b (*oil*) and the index of industrial production for the euro area (y^{EA}). The exchange rate shock is modeled through a shock to the nominal effective exchange rate against 21 trade partners (*neer*).[14] Due to non-stationarity in the time series and the presence of a cointegrated relation, the model is estimated in vector error correction (VEC) form as described in (4.2.1) and (4.2.2).[15]

$$z_t = A * ECT_t + B(L)z_{t-1} + \Xi * x_t + c_o + \varepsilon_t \,, \qquad (4.2.1)$$

where A is a vector of coefficients on the residuals of the cointegrated equation (ECT_t). B is a matrix of coefficients for the vectors of endogenous variables z_t; $z_t = [\Delta neer_t, \Delta pm_t, \Delta ppi_t, \Delta cpi_t, \Delta y_t]$, Δ is a first-difference operator. Ξ is a matrix of coefficients for the vectors of exogenous variables x_t; $x_t = [oil_t, y_t^{EA}, dummy]$; c_0 is a constant term and ε_t is a residual. *dummy* denotes a binary variable equal to unity in the period of economic overheating and subsequent collapse (August 2007–February 2010), when the

[14] The pricing chain model is estimated mainly for advanced economies, where longer time series are available. For this reason the model may include additional endogenous and exogenous variables, but given the relatively short time span, and wishing to preserve more degrees of freedom, we do not consider other variables in the present analysis.

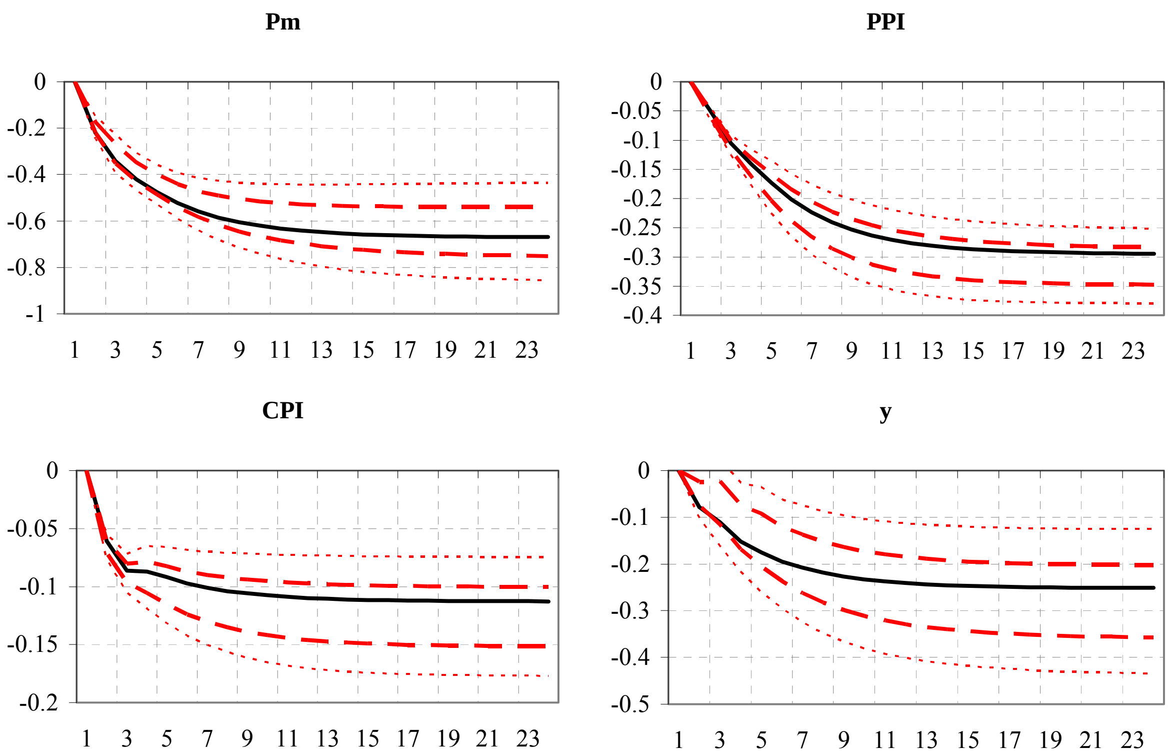
[15] Following the Akaike information criterion the model is estimated with two lags. This is also the optimal lag length according to the likelihood ratio test. The Schwarz and Hannan-Quinn information criteria suggest only one lag, which is not optimal for monthly data. The final prediction error test finds five or more lags. The Johansen cointegration test indicates two cointegrated relations, but when the model is estimated with two error correction terms at least half of the long-run coefficients are insignificant and need to be constrained to zero. For this reason, the model is estimated with only one cointegrated equation.

transmission mechanism was most likely to have been different from the average transmission observed over the estimated period. Superscript *EA* marks the euro area variables. The cointegrated equation has the following form:

$$cpi_t = \omega_0 + \omega_1 ppi_t + \omega_2 pm_t + \omega_3 neer_t + \omega_4 y_t + \eta_t . \tag{4.2.2}$$

All variables enter the model in logarithms and have monthly frequency. The estimation period is January 1998–March 2011. The variables are converted into indices: average 2005=100. The data sources are Eurostat, ARAD, and the internal CNB database. Figure 4.2.3 shows the reaction of the macroeconomic variables to the exchange rate shock.

Figure 4.2.3: Accumulated Impulse Response to 1% Shock to Exchange Rate

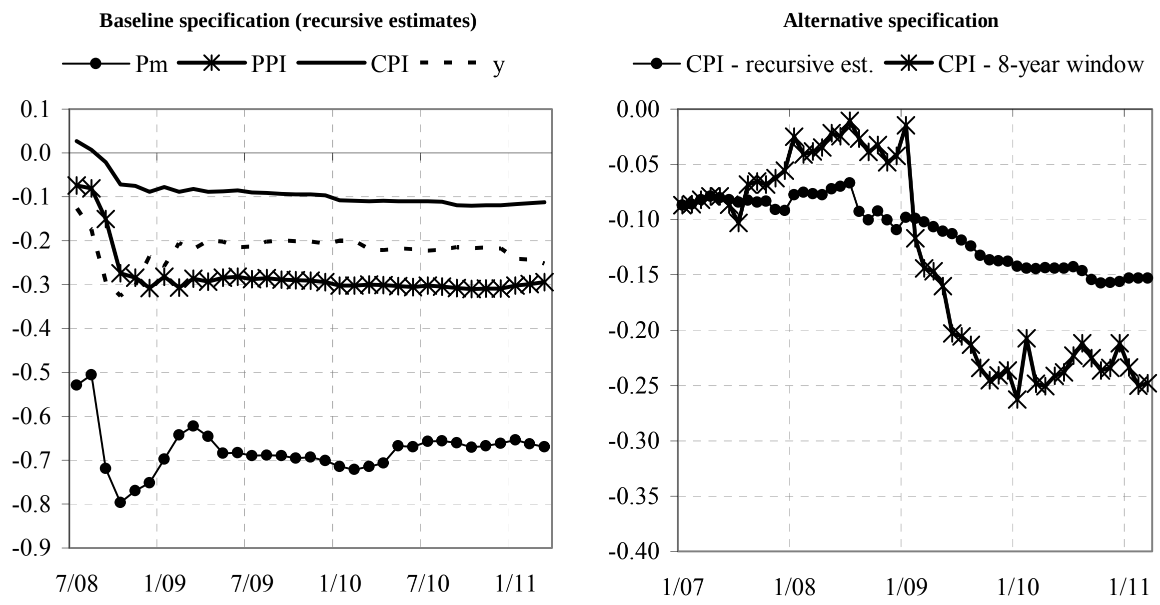


Note: The impulse response to an unexpected exchange rate appreciation of 1% (~0.3 CZK/EUR) is plotted on the vertical axis. Periods (in months) are shown on the horizontal axis. The dashed lines denote confidence bands of ±1 and ±2 standard deviations computed from 14 alternative specifications differing from each other in the number of endogenous and exogenous variables and their lags. An increase in the exchange rate means appreciation.

The exchange rate pass-through fades along the pricing chain. As expected, import prices are the main channel of shock transmission to domestic inflation. The pass-through to import prices is close to 0.70%, consistent with the high degree of openness of the Czech Republic. The pass-through to producer prices (and to industrial production) is roughly half as large as the transmission of the same shock to import prices. The lowest response to the shock is found for Czech inflation. An exchange rate appreciation of 1% may lower consumer prices by 0.11%. In line with previous studies, complete transmission of the shock is fast, taking between 4 and 6 months.[16] A sensitivity analysis based on specifications with a smaller number of exogenous and endogenous variables or a different number of lags shows that the pass-through to inflation varies between 0.07% and 0.18%. The upper bound result is close to the results in section 3.

The stability of the impulse responses over time is tested by applying recursive estimates (Figure 4.2.4, left window). Due to the large number of parameters to be estimated[17] the selected specification produces unstable results when the period is shortened. Until almost the end of 2008, the impulse responses are volatile and often change sign. Starting from 2009, the reaction to the exchange rate shock stabilizes from roughly 0.10% for consumer prices to about 0.70% for import prices. Given the high volatility of the results before 2009 it is impossible to make any conclusion about the pass-through during the crisis. The impact of the crisis on the transmission mechanism is tested using an alternative specification with a smaller number of estimated parameters, i.e., without Czech and euro area industrial production. While this specification may suffer from omitted variable bias, it produces more stable results and is suitable for estimations where the time period is relatively short. Impulse response stability is tested starting from the period ending in January 2007. In addition to recursive estimates, the alternative specification is estimated with a moving window (Figure 4.2.4, right window).

Figure 4.2.4: Pass-Through Stability Over Time



Note: The accumulated impulse response to an unexpected exchange rate appreciation of 1% (~0.3 CZK/EUR) is plotted on the vertical axis. The horizontal axis shows the end of the estimation period. An increase in the exchange rate means appreciation.

[16] For earlier exchange rate pass-through estimates for the Czech Republic based on the pricing chain approach, see Babecká-Kucharčuková (2009).

[17] For a model estimated with two lags, one cointegrated vector, five endogenous variables, and three exogenous variables, 79 parameters have to be estimated.

The response of consumer prices based on the alternative specification indicates a possible disruption to the transmission mechanism during the crisis (a similar result is found in section 3). The estimates based on short time series additionally show a recent increase in the pass-through to 0.25%. This result should be interpreted with caution due to the volatility of the results. As Figure 4.2.4 (the moving window estimates) shows, the inclusion and suppression of one time period changes the size of the impulse response by up to 0.05%.

In summary, about 0.11% of the initial shock to the exchange rate is fully transmitted to consumer prices after 1–1.5 years. The estimation of alternative specifications leads to an increase in the pass-through to 0.18%. The speed and size of the pass-through are in line with the results based on macroeconomic models described in sections 3.1 and 3.2. During the economic crisis the transmission was lower.

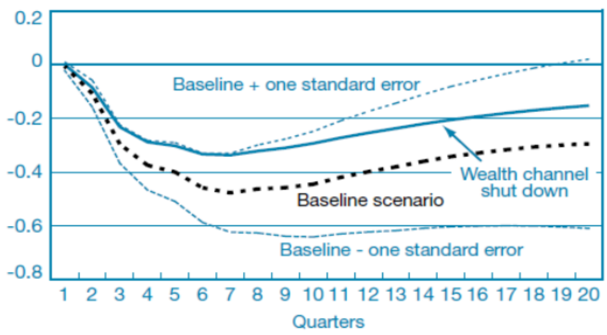
4.3 Wealth Effect

The wealth channel captures the indirect effects of monetary policy on inflation through the impact of changes in interest rates on asset prices, which affect households' perceptions of their wealth and sources for financing consumption.[18] This channel has not been analyzed for the Czech economy so far. However, part of the wealth effect is described in the study of Dvořáková and Seidler (2013), who find a positive effect of both housing wealth and stock market wealth on consumption. A positive effect of housing wealth on consumption is also documented in Šonje et al. (2012), who focus on four European post-transition countries including the Czech Republic. Finally, Seč and Zemčík (2007) use Czech microdata to find that higher real estate prices result in a consumption increase for apartment owners and that higher rents imply a consumption reduction for renters

The wealth effect, as estimated by a counterfactual experiment, was analyzed for the USA by Ludvigson, Steindel, and Lettau (2002) and for the euro area by Siokis (2005). Figures 4.3.1 and 4.3.2 show the empirical results obtained from those studies. The data for neither the USA nor the euro area confirm a statistically significant wealth effect.

[18] Changes in interest rates influence asset prices via both asset-pricing models due to the discount rate and the available liquidity of economic agents and their demand for investment instruments. Financial resources can be obtained for consumption either by selling assets or as collateral against the value of a loan.

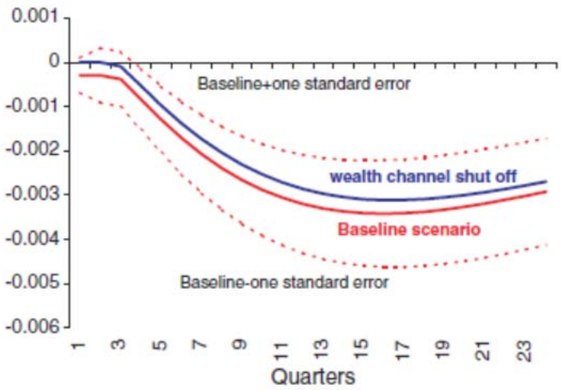
Figure 4.3.1: Results of Wealth Effect Analysis for USA by Ludvigson, Steindel, and Lettau (2002)



Notes: The figure shows the twenty-quarter response of the variables to a one-standard-deviation (81-basis-point) innovation in the federal funds rate. The vertical axis represents the percent deviations of the variables (the basis-point deviations of the federal funds rate). The sample period is 1966:1–2000:3.

Source: Ludvigson, Steindel, and Lettau (2002)

Figure 4.3.2: Results of Wealth Effect Analysis for Euro Area by Siokis (2005)



Notes: The figure shows impact of an interest rate shock on consumption.

Source: Siokis (2005)

Methodology

The analysis examines how innovations in interest rates influence household wealth, and how those changes in wealth influence consumer spending. The methodology is based on Ludvigson, Steindel, and Lettau (2002), who estimate the wealth effect in a two-step procedure. Firstly, a baseline model that includes a wealth proxy is estimated. Secondly, a counterfactual experiment is applied to the model – see Bernanke, Gertler, and Watson (1997). The baseline model is a structural vector autoregression model (SVAR, see equation 4.3.1) with five variables: household consumption c_t, gross disposable income hdd_t, wealth w_t, the interest rate IR_t, and inflation π_t.[19] For the analysis we use quarterly data for the period 1998Q1–2009Q4.

$$B_0 z_t = k + B_1 z_{t-1} + B_2 z_{t-2} + ... + B_p z_{t-p} + u_t \tag{4.3.1}$$

where $z_t = (\pi_t, hdd_t, c_t, w_t, IR_t)'$. The vector of disturbances u_t represents the structural innovations; these disturbances are assumed to be serially uncorrelated and uncorrelated with each other. The matrix B_0 governs the contemporaneous relations among the variables in the system (4.3.2).

[19] Lower-case letters denote log variables.

$$B_0 = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 \\ \beta_{21} & 1 & 0 & 0 & 0 \\ \beta_{31} & \beta_{32} & 1 & \beta_{34} & 0 \\ \beta_{41} & \beta_{42} & 0 & 1 & \beta_{45} \\ \beta_{51} & \beta_{52} & \beta_{53} & 0 & 1 \end{bmatrix} \tag{4.3.2}$$

To identify the structural innovations we consider specific restrictions similarly to Bernanke and Blinder (1992) and Ludvigson, Steindel, and Lettau (2002). The interest rate responds contemporaneously to macroeconomic developments (consumption), but changes in interest rates (given production lags) can only affect these variables with a one-period lag, thus $\beta_{35} = 0$. The wealth w_t measured at the beginning of the period is not contemporaneously influenced by consumption c_t, which is given by the flow within the period, thus $\beta_{43} = 0$. Ludvigson, Steindel, and Lettau (2002) assume that the log of aggregate consumption is close to a random walk, consistent with a permanent-income type of behavior. Consequently, following the approach of Ludvigson, Steindel, and Lettau (2002) we allow interaction between wealth w_t and the interest rate within the same period ($\beta_{45} \neq 0$). At the same time we restrict the influence of asset prices on monetary policy ($\beta_{54} = 0$).[20]

The subsequent *counterfactual experiment* involves estimating the model with the same set of variables, but without the consumption-wealth channel, as in Ludvigson, Steindel, and Lettau (2002). It is newly assumed that the contemporaneous response of consumption to wealth is set to zero ($\beta_{34} = 0$) and also that the lagged response of consumption to wealth variables is set to zero in the equation of consumption included in equation (4.3.1).[21] The assumed restrictions are described by matrix B_0 (4.3.3). The model according to the restrictions in (4.3.2) is the baseline model and the model according to the restrictions in (4.3.3) was estimated without the consumption-wealth channel (model without CW). The difference between the total effect of changes in interest rates on consumption (baseline model) and the estimated effect of the use of the counterfactual experiment (model without CW) is interpreted as a contribution to the wealth effect of the monetary policy transmission mechanism.

The lag structure was performed according to standard criteria, resulting in a delay of two quarters. The time series of real consumption, real gross disposable income, and real estate prices are integrated of order one, based on the Johansen cointegration test. For this reason, it is possible to use the VAR model despite the presence of non-stationarity.

[20] This is relevant mainly in the case of financial assets (shares and obligations). For real estate assets this assumption does not need to be so strict, but accepting this assumption allows us to compare the results of Ludvigson, Steindel, and Lettau (2002). We thus stick to the restriction set according to Ludvigson, Steindel, and Lettau (2002).

[21] The order of the variables is $\pi_t, hdd_t, c_t, w_t, IR_t$, so consumption (c) is the third row for equation (4.3.1).

$$B_0 = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 \\ \beta_{21} & 1 & 0 & 0 & 0 \\ \beta_{31} & \beta_{32} & 1 & 0 & 0 \\ \beta_{41} & \beta_{42} & 0 & 1 & \beta_{45} \\ \beta_{51} & \beta_{52} & \beta_{53} & 0 & 1 \end{bmatrix} \quad (4.3.3)$$

Wealth is approximated by prices of dwellings obtained from tax statements.[22] The choice is based on the observed volatility and information content of the variable and the fact that in the Czech Republic there are no reliable data for the wealth variable consisting of both prices and quantity. The wealth variable should be defined more broadly in the optimal case. There is a possible approximation obtained from the Financial Accounts Statistics compiled by the CNB. These statistics provide financial assets divided by sectors (households, non-financial firms, etc.) into financial instruments. The drawback of this series is its length, starting only in 2004. In alternative specifications household wealth is approximated by the equity index, i.e., another type of asset, and the exchange rate, changes in which may influence the perceived purchasing power of households.

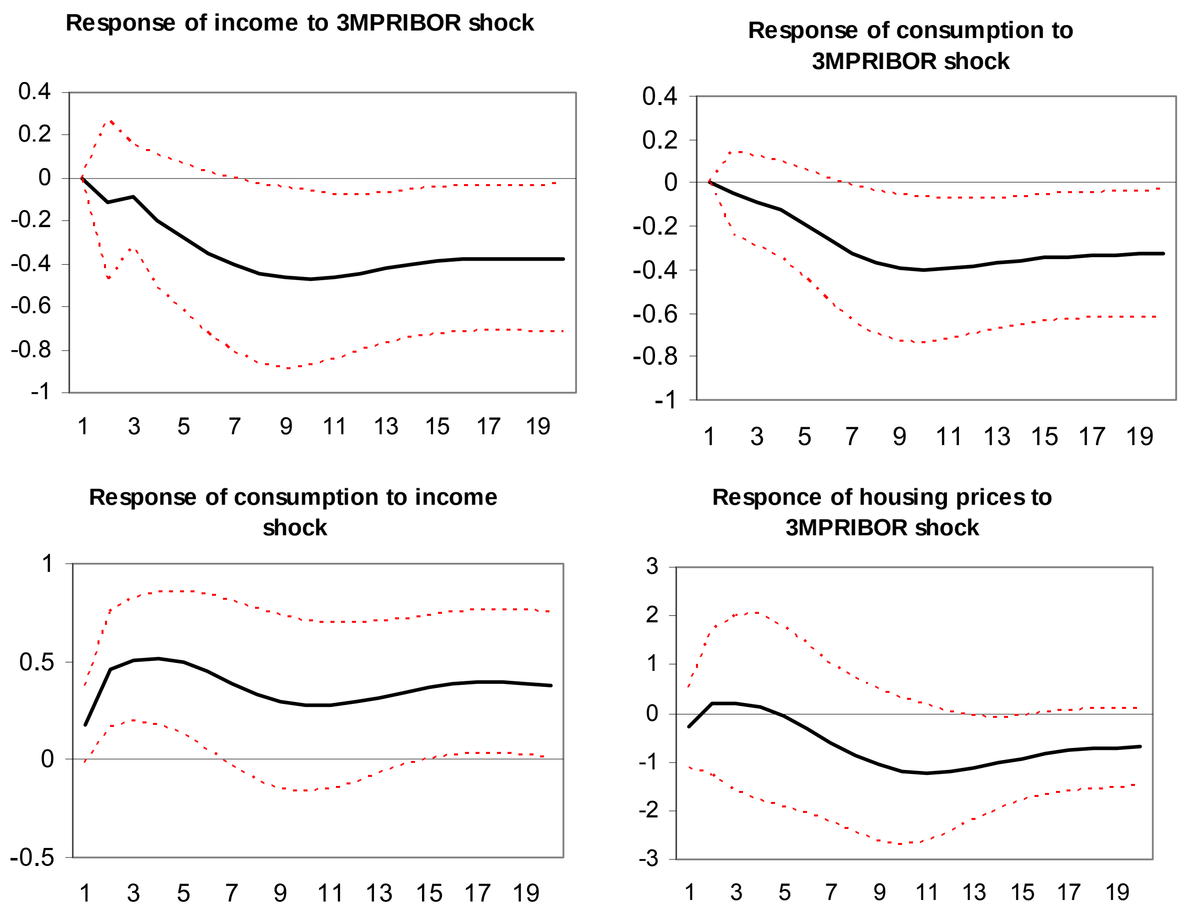
Results

Selected impulse responses are presented in Figure 4.3.3. The responses to a temporary positive shock to interest rates of one percentage point are generally very slow, showing a slight decline in real gross disposable income and real household consumption. However, the model shows a significant and rapid response of household consumption to changes in gross disposable income (captured by the response to the one-standard-deviation shock), with the response peaking two to three quarters after the shock. The interest rate shock is reflected in housing prices only very slowly and slightly (peaking in approximately ten quarters). This may reflect the relatively lengthy and complex process of buying property.

The impulse responses of the structural VAR with five variables and restrictions under (4.3.2) are presented in Figure A.1 in the appendix. The positive shock to interest rates is a one-time shock and disappears after 10 quarters. The reactions to it are generally very slow, in the form of a slight decline in real gross disposable income and real household consumption, or insignificant in terms of a decline in wealth. A positive shock to wealth (housing prices) does not induce an intuitive response in consumption, i.e., prevent its growth (see Figure A.1 in the appendix). Unlike Ludvigson, Steindel, and Lettau (2002), in our case the positive wealth shock leads to a decline in interest rate. This confirms the hypothesis that the central bank does not target asset prices directly, and the result is consistent with the restrictions set in matrix B_0 (4.3.3) as the parameter $\beta_{54} = 0$ applies.

[22] These are transfer prices. The alternative specification of the model (exchange rate and equity index) leads to results which are not significantly different.

Figure 4.3.3: Selected Impulse Responses (%)

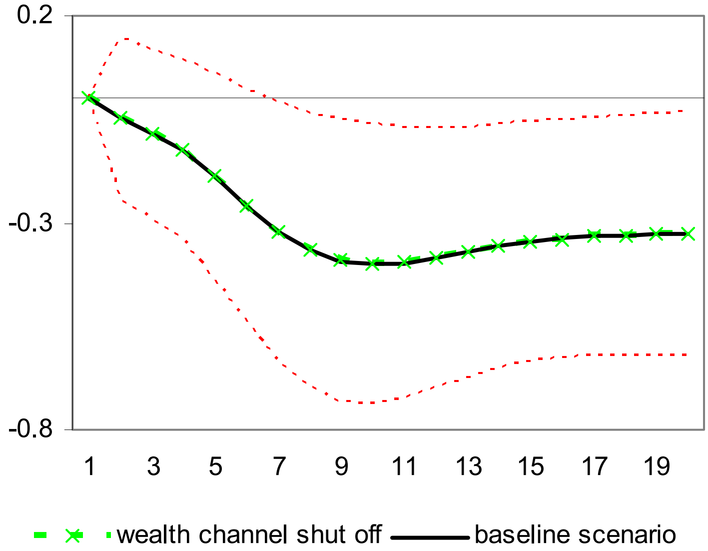


Note: Responses in % for a shock to interest rates of 1%. Time period in quarters on the horizontal axis. The vertical axis shows the size of the impulse response. The dashed lines indicate confidence intervals of size ± 2 standard deviation errors.

Figure 4.3.4 shows that the empirically observed impact of monetary policy shocks on household consumption is almost identical regardless of whether or not the effect of endogenous changes in wealth is taken into account by the model. The presence of the wealth channel has not been confirmed in this analysis. This is a result similar to the finding of Siokis (2005) for the euro area, where for the period 1977–2002 the difference between the impulse responses of household consumption with and without the wealth channel was not found to be statistically significant.

Selected impulse responses obtained by the baseline and the counterfactual experiment are shown in Figure 4.3.4. Statistically significant differences of the impulse responses from the baseline scenario and the counterfactual experiment would be confirmation of the existence of the wealth channel. In our case, the responses are almost identical, i.e., we have not confirmed the presence of the wealth effect.

Figure 4.3.4: Comparison of Impulse Responses of Household Consumption to Interest Rate Shock in Model With and Without Wealth Channel



Note: Responses of consumption (in %) to a 3M PRIBOR shock of 1%. Time period in quarters on the horizontal axis. The dashed red lines indicate confidence intervals of size ± 2 standard deviation error estimates.

5. Concluding Remarks

The presented collection of empirical results provides a look at the transmission mechanism in the Czech Republic using different empirical strategies. Typically, our sample starts in mid-1990s and ends approximately in 2010, depending on particular analysis. The actual estimations were carried out in 2010-2012 and therefore, the empirical results reflect the sample we used. Irrespective of the method employed, the results are broadly in line with previous research.

Estimations using VAR, Bayesian VAR, and time-varying parameter VAR show that the fall in prices after an unexpected interest rate hike reaches its maximum after about 5–6 quarters. The response of prices to an exchange rate shock reaches its maximum after one year according to both the VAR and Bayesian VAR estimates. The impact of the crisis is estimated in two ways – using the time-varying parameter VAR and estimating the Bayesian VAR on the full sample as well as the pre-crisis subsample. Both approaches suggest that the transmission of monetary policy shocks was not substantially affected by the crisis in the Czech Republic.

Looking closer at the interest rate channel, approximately 60% of the transmission of changes in monetary policy rates to client rates occurs within one month for loans with floating rates, loans with rates fixed for up to one year, and for deposits with short maturity. As for the impact of the global financial and economic crisis, the evidence suggests that the transmission weakened and slowed in the first few years after the crisis hit.

A detailed analysis of the exchange rate pass-through shows that a 1% appreciation leads to a decrease in import prices of 0.7%. The pass-through to producer prices is about half as large. Most importantly, consumer prices decrease by 0.11%. The transmission of an exchange rate shock to prices weakened during the crisis.

Finally, regarding the significance of the wealth channel, the presented analysis did not confirm the presence of a wealth channel in the transmission mechanism in the Czech Republic.

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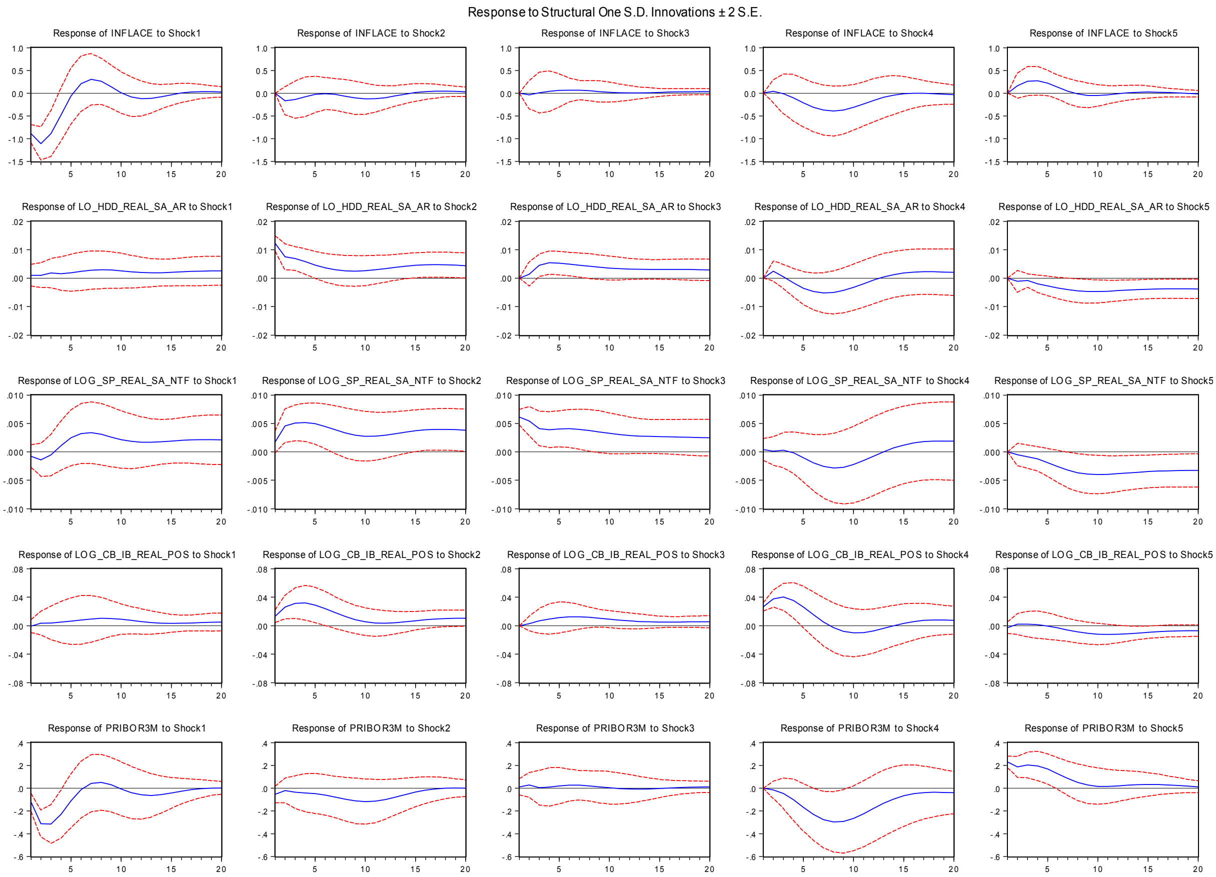
Appendix

Figure A. 1: Impulse Response of Structural VAR With Five Variables – Baseline



Note: Shock 1 (shock to inflation), shock 2 (shock to gross disposable income), shock 3 (shock to household consumption), shock 4 (shock to house prices), shock 5 (shock to interest rates).

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