Choques nas bonificações de risco, política monetária e transpasso do tipo de câmbio na República Checa, Hungria e Polônia

Balázs Vonnák[*]

Este artigo explora o papel da política monetária em uma economia pequena e aberta na que os choques no tipo de câmbio têm um efeito importante. Usou-se o modelo VAR estimado para a República Checa, Hungria e Polônia. Implementaram-se, além do mais, restrições contemporâneas e de signos para identificar o efeito da política monetária e dos choques nas bonificações de risco. Usaram-se estimados do mesmo modelo para o Canadá, Suécia e o Reino Unido como ponto de referência para economias desenvolvidas com baixa inflação. Os resultados indicam que, em geral, o impacto dos choques nas bonificações de risco faz com que seja quase impossível para a política de tipos de juros modificar o tipo de câmbio com o fim de minimizar as consequências inflacionárias. Por outro lado, a baixa inflação pode diminuir o transpasso do tipo de câmbio, contribuindo a que na política monetária seja possível obviar os choques no tipo de câmbio.

Classificação JEL: E31, E52, F31.

Palavras chave: política monetária, choques nas bonificações de risco, transpasso do tipo de câmbio, VAR estrutural, restrição de signos.

Choques en las primas de riesgo, política monetaria y traspaso del tipo de cambio en la República Checa, Hungría y Polonia

Balázs Vonnák[*]

Este artículo explora el papel de la política monetaria en una economía pequeña y abierta en la que los choques en el tipo de cambio tienen un efecto importante. Se usó el modelo VAR estimado para la República Checa, Hungría y Polonia. Se implementaron además restricciones contemporáneas y de signos para identificar el efecto de la política monetaria y de los choques en las primas de riesgo. Se usaron estimados del mismo modelo para Canadá, Suecia y el Reino Unido como punto de referencia para economías desarrolladas con baja inflación. Los resultados indican que, por lo general, el impacto de los choques en las primas de riesgo hace que sea casi imposible para la política de tipos de interés modificar el tipo de cambio con el fin de minimizar las consecuencias inflacionarias. Por otra parte, la baja inflación puede disminuir el traspaso del tipo de cambio, contribuyendo a que en la política monetaria sea posible obviar los choques en el tipo de cambio.

Clasificación JEL: E31, E52, F31.

Palabras clave: política monetaria, choques en las primas de riesgo, traspaso del tipo de cambio, VAR estructural, restricción de signos.

Risk premium shocks, monetary policy and exchange rate pass-through in the Czech Republic, Hungary and Poland

Balázs Vonnák*

This paper investigates the role of monetary policy in a small open economy, where exchange rate shocks are important. VAR models are estimated for the Czech Republic, Hungary and Poland. Contemporaneous and sign restrictions are imposed in order to identify the effect of monetary policy and risk premium shocks. Estimates from the same model for Canada, Sweden and the UK are used as a benchmark for developed economies with low inflation. The results suggest that the typical size of a risk premium shock renders it almost impossible for the interest rate policy to smooth the exchange rate with the aim of minimizing inflationary consequences. On the other hand, low inflation may decrease the exchange rate pass-through, which helps the monetary policy ignore exchange rate shocks.

JEL classification: E31, E52, F31.

Keywords: monetary policy, risk premium shocks, exchange rate pass-through, structural VAR, sign restriction.

I.   INTRODUCTION

After the collapse of the communist regimes in Central and Eastern Europe (CEE henceforth), several ex-communist countries were looking for a proper monetary regime. In the beginning, they faced high inflation as a consequence of price liberalization, as well as slow growth, due to the process of restructuring the economy. Since these economies were small and open ones, stabilization of the exchange rate seemed to be the most straightforward way to contain inflationary pressures.

After the first, more or less successful years of tight exchange rate management, the limits, or, in some cases, the pitfalls of these regimes became obvious. The Czech Central Bank widened the intervention band of the koruna in 1996 from ±0.5 percent to ±7.5 percent. After the attack against and the depreciation of the currency in 1997, the intervention band was abandoned. Since the beginning of 1998, the Czech Central Bank operates an inflation-targeting regime. In 1995, the Hungarian government and the Central Bank decided to introduce a crawling band regime, with a bandwidth of ±2.25 percent. In 2001, the band was widened to ±15 percent and the Central Bank announced targets for the inflation. The intervention band was fully eliminated only in 2008. Poland replaced its crawling peg by a crawling band regime in 1995. The band was widened in several steps from ±2 to ±15 percent. They switched to free float and inflation targeting in 2000.

Even though all these CEE countries under investigation have a floating exchange rate and an inflation target, the question remains open: how flexible should be the monetary policy towards exchange rate movements? Opponents of free-floating regimes argue that in these countries the exchange rate channel dominates the interest

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rate channel, thus the interest rate policy should aim at managing the exchange rate. They recommend a marked policy rate response to exchange rate shocks. On the other hand, tight exchange rate management, even if conducted by an interest rate policy instead of intervention in the foreign exchange market, may bring about speculation and instability of the exchange rate. Moreover, a higher volatility of the exchange rate may also reduce the pass-through of exchange rate movements. According to this view, letting the exchange rate absorb the shocks, even those originating from financial markets, has no harmful consequence in the long run, as economic agents learn to hedge themselves. This attitude has the major benefit of relieving the interest rate policy of pursuing short-term exchange rate targets.

Although the inflationary effect is only one aspect of exchange rate volatility, in this paper I ignore all the trade-enhancing and other long-run welfare consequences of exchange rate dynamics. Instead, I try to focus purely on the role monetary policy plays in the pass-through of exogenous exchange rate shocks to consumer prices. In particular, I address the following questions: what are the effects of risk premium shocks? How should monetary policy react to these shocks? How successful were these countries in preventing exchange rate disturbances from spreading to overall consumer prices?

In order to estimate the dynamic interaction among the key variables, I adopt a VAR approach. I identify monetary policy and exchange rate risk premium shocks by imposing both contemporaneous and sign restrictions on impulse responses. Then, I investigate the reaction of the interest rate policy to the latter shock and evaluate its performance in terms of price stability. In order to avoid methodological biases and to capture specific features of exchange rate pass-through in transition economies, I estimate the same model for three developed countries as well: Canada, Sweden and the UK. I chose these countries because they are small open economies with a floating exchange rate, therefore comparable with the CEE countries under investigation.

The structure of the paper is the following: First, the econometric framework is introduced. Section III presents the estimated effect of monetary policy and risk premium shocks on the key variables. In section IV, the issue of exchange rate pass-through is discussed. Section V concludes.

II.   THE EMPIRICAL MODEL

This section describes the econometric modeling framework.

A.   DATA AND SPECIFICATION OF THE VAR MODEL

As endogenous variables of the VAR, one should include some measure of real activity, a price index, and because of the focus of the paper, the exchange rate and the policy instrument. This can be regarded as the minimum set. On the other hand, the short time span does not allow including too many endogenous variables without increasing the risk of overfitting.

I used the industrial production as a measure of real activity. I used it instead of the GDP to avoid the possible measurement error of the latter in these countries. Another advantage of using industrial production is that it is available at monthly frequency, which facilitates identification. I used the overall consumer price index to capture price dynamics.

To measure the policy stance, I used 3-month T-bill rates in the case of Hungary and Poland, and 3-month money market rates for the Czech Republic. Finally, I used nominal effective exchange rate series for all countries. A more detailed description of the data sources can be found in the appendix.

An intercept, as well as foreign exogenous variables, were included in the VARs. I used contemporaneous and one-period lagged German industrial production, consumer prices, short interest rate, and nominal effective exchange rate as control for external effects. In the case of Canada, the foreign country was the U.S.

The industrial production and consumer price series are seasonally adjusted. I used variables in levels. When making statistical inference, I adopt a Bayesian approach; hence, the issue of unit roots needs not to be addressed[1].

Since the transition process and particularly the change in monetary policy regimes may have caused structural breaks in the data generating process, the selection of the sample may be crucial for the results we obtain. Here we face a trade-off: shorter

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1   See Sims and Uhlig (1991).

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periods are more likely to be free of structural breaks but contain fewer observations. For the Czech Republic I started the sample in 1998:M1, when the Central Bank switched to inflation targeting. The Polish sample starts in 1997:M1, earlier than the introduction of the free-floating and inflation-targeting regime, but at a time when the intervention band was already wide enough to let the exchange rate absorb shocks. Since in Hungary the inflation-targeting regime started in 2001, I used a longer sample that contains the crawling narrow band regime as well. The Hungarian sample, therefore, starts in 1995. In order to ensure comparability, I used almost the same sample period for Canada, Sweden and the U.K.

As a lag selection criterion, I aimed at eliminating the autocorrelation in the residuals. Usually 3 to 7 lags were enough[2]. It more or less coincided with what the popular model selection criteria suggested.

B. IDENTIFICATION

There is a huge strand of literature on how to identify monetary policy shocks[3]. One of the most widely used assumptions is that monetary policy affects the economy only with some delay. In our case this means that the contemporaneous effect on production and prices is zero. Apparently, financial variables, such as the exchange rate, can react immediately to monetary policy. The intuition behind contemporaneous zero restrictions is that it takes time until shocks originated in financial markets or monetary policy spread onto the real economy.

Contemporaneous zero restrictions on quarterly data in small open economies can be criticized. For example, an unexpected rate hike by the Central Bank may appreciate the currency. In the presence of producer currency pricing, this may lead to a quick drop in import prices, and even in overall consumer prices. Since I use monthly data, this problem is probably less relevant.

The aim of this paper is to identify two sources of economic disturbances: monetary policy and risk premium shocks. The argument of delayed transmission applies to both; therefore, I imposed contemporaneous zero restrictions on industrial production

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2 For the exact numbers see Table A4.1 in the appendix.

3 For a review, see Christiano et al. (1998)

and consumer prices for each shock[4]. Unfortunately, this is not sufficient to disentangle them from each other. Additional assumptions are needed.

Authors using VAR to estimate the exchange rate pass-through usually apply a recursive identification scheme[5]. This approach rests on the assumption that there exists an ordering of the variables so that a shock to one variable affects the ones preceding it only with delay. These authors sometimes do robustness checks with respect to the ordering. Hahn (2003) reports no change of her estimated pass-through for the euro area after using alternative identification schemes. Ca'Zorzi et al. (2007) obtain similar estimates using various ordering for almost all transition countries, with the exception of Hungary. Faruqee (2006) found his results for the euro area and the U.S. robust to the ordering.

If both the interest rate and the exchange rate are included in the VAR, recursive identification can be justified as long as either the monetary or the risk premium shock has no impact on one of these variables within one period. We have, however, strong reasons to believe that our shocks may have contemporaneous effect on both financial variables, at least in the case of Hungary. Based on the methodology of Rigobon and Sack (2004), Rezessy (2005) showed that unexpected interest rate changes made by the Hungarian Central Bank have impact on the exchange rate even within the same day. On the other hand, minutes of the Monetary Council meetings and press releases reveal that change in the perceived risk premium did play an important role in rate setting, even within the same month.

In order to distinguish between monetary policy and risk premium disturbances, I exploit the interest parity condition. The most standard assumption when characterizing exchange rate dynamics in theoretical macro models is that expected returns are equal among various currencies. In other words, interest rate differentials reflect only the expected change in the exchange rate. This is the uncovered interest parity (UIP). Partly because of the failures of empirical UIP tests, the equation is often augmented by a residual term, which can be thought of as an excess compensation for investing in assets denominated in a risky currency:

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4   I also experimented with identification without imposing zero restrictions on prices. The impulse responses in some cases were implausible, as consumer prices jumped immediately with the risk premium shocks to an extent that was counterintuitive.

5   See McCarthy (2006), Ca'Zorzi et al. (2007), Hahn (2003), and Faruqee (2006).

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$$i_t = i_t^* + E_t s_{t+1} - s_t + \rho_t \tag{1}$$

where i and i^* denote the domestic and the foreign one-period interest rate, respectively; s stands for the logarithm of home currency price of one unit of foreign currency, hence increasing values indicate the depreciation of the domestic currency. The residual term, ρ, is the risk premium.

Within this framework, monetary policy and risk premium shocks can be represented in unexpected changes in i and ρ, respectively. Assuming constant foreign interest rate and exchange rate expectations, a higher domestic interest rate is associated with a stronger exchange rate. An increase in the risk premium increases the domestic interest rate and depreciates the currency. Note that a shock to the foreign interest rate is equivalent to a risk premium shock[6].

I derived the remaining identifying restrictions from this partial exchange rate model. I assumed that after an unexpected monetary tightening the domestic interest rate increases and the exchange rate appreciates. Regarding the risk premium shocks, an increase in the expected excess return leads to a higher interest rate and depreciates the currency. Therefore, it is the sign of correlation between the exchange rate and interest rate that makes the distinction between the two sources of financial market disturbances[7].

I imposed the sign restrictions on the first 12 months of the impulse responses. It is difficult to judge a priori whether this is too restrictive. If the restriction period is too long, plausible impulse responses might be excluded, since the interest rate and the exchange rate may endogenously respond to movements of other variables. Hence, it is reasonable to perform a consistency check of the results.

In the case of my identification scheme, the most questionable restriction may be that interest rates remain higher after an unfavorable change in risk premium throughout a whole year, even if the output or the price level falls. That would obviously be

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6   The stability of exchange rate expectations is not necessary for the identification. After an unexpected tightening it is reasonable to think that market participants expect lower prices, and consequently, a more appreciated exchange rate in the long run.

7   Among the first and most influential applications of sign restrictions were Faust (1998), Canova and de Nicolo (2002), and early versions of Uhlig (2005).

counterintuitive, considering the countercyclical behavior of monetary policy. As it turns out, in most cases the sign of the interest rate response is consistent with that of prices and output. Exceptions are Canada and Hungary, where the industrial production drops after the shock. On the other hand, prices rise in these countries due to the exchange rate depreciation, making the Central Bank face a trade-off. Therefore, it is not unlikely for interest rates to be permanently higher. One should, however, keep in mind that this kind of internal consistency is not a guarantee for not having excluded the true impulse responses if in reality prices, production, and therefore, interest rate fall within one year.

In Vonnák (2005), I demonstrated that in the case of Hungary this identification scheme provides similar results to a completely different approach. Jarociński (2006) uses very similar identification of monetary policy shocks in CEE countries. Sánchez (2007) relies entirely on sign restrictions when identifying technology, preference, monetary policy, and risk premium shocks for emerging countries.

The case of Hungary is worth some further discussion. During the crawling band exchange rate regime, which covers roughly the half of the observations, the logic of monetary policy may have been different. The preannounced devaluation rate of the intervention band served as an anchor for the exchange rate expectations, and thus monetary tightening appeared in the form of a reduction of the devaluation rate. When the exchange rate could not react to this change of the policy stance, because of the strong edge of the band, the interest rate should have been lowered according to the interest parity. The exchange rate depreciated less due to the slower pace of devaluation. Since we cannot control exchange rate expectations, monetary policy and risk premium shocks may have been observationally equivalent during this period: a lower interest rate is associated with a stronger exchange rate. Nonetheless, the similarity of the impulse responses across CEE countries suggests that the identifying restrictions seem to fit even Hungary quite well.[8]

The numerical implementation followed the algorithm of Uhlig (2005), and is presented in more detail in the appendix. To obtain posterior distribution of impulse response functions and other derived statistics, I drew jointly from the posterior

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8   Moreover, I estimated the same model for Hungary excluding the crawling band period, but including 2007 and the first half of 2008. The results were fairly similar to those obtained from the 1995–2006 sample. For the sake of comparability, in the paper I report estimates from the sample containing the crawling band regime.

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of VAR coefficients, variance-covariance matrix and the subspace of potential candidates for shock vectors. The VAR prior was a normal-Wishart distribution parameterized by the OLS estimates. The posterior distribution is constituted by drawings that fulfilled the sign restrictions. For a more detailed description see the appendix or the referred paper.

III.   THE EFFECT OF MONETARY POLICY AND RISK PREMIUM SHOCKS

The following section summarizes what we can learn by investigating impulse responses (IRFs) to structural shocks[9]. During the discussion I rely on median estimates and confidence bands obtained by connecting the corresponding percentile values of the posterior distribution for each period. Since there are some arguments that these statistics might not be the best way to present the results, in the appendix I present other reporting techniques as well. There, I conclude that the main findings of the paper are not due to an improper way of statistical reporting.

The first general observation is that in most cases the effect on industrial output and consumer prices was not significantly different from zero[10]. This means that the middle 95% of the IRF posterior distribution contains the x-axis in most cases. There might be several explanations for that. First, the effects may be indeed small. Second, our data may contain a considerable amount of noise. Third, there may be important structural breaks in the data-generating processes. And finally, our identification may be too loose to capture the structural shocks precisely, so the IRFs may show the mixed effect of several different shocks.

Another important observation is that the point estimates for the three CEE countries are fairly similar. Since results for the control group differ significantly, this is probably not an artifact of the estimation method, but rather an evidence for common features of the transition economies.

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9   See Graph A5.5 – A5.10 in the appendix.

10   The significance of interest rate and exchange rate responses are not as informative as in the case of the other two variables. The Bayesian approach and the identifying restrictions together imply that these posterior distributions do not contain the zero during the first 12 months; but this is an assumption, not a result.

An additional important feature of our results is that the exchange rate responses to monetary policy shocks do not significantly suffer from the so-called delayed overshooting phenomenon. Tipically, the literature reports that monetary contraction leads to a prolonged appreciation of the currency. This contradicts to the prediction of the model proposed by Dornbusch (1976), in which the appreciation takes place immediately, followed by gradual depreciation to the new equilibrium level. Delayed overshooting violates the UIP condition, as for several periods there is a predictable excess return in the home currency[11]. Although there exist theoretical models explaining the delayed response of the exchange rate, this remains an unpleasant feature of empirical estimates[12]. Our impulse responses, however, exhibit only a temporary deviation from the UIP, as the maximum exchange rate response is reached within six months in each country. This may be an indication that our identifying restrictions are more capable to capture the characteristics of monetary policy shocks than those typically used in the literature. Scholl and Uhlig (2005) also find that proper identification can mitigate the delayed overshooting puzzle.

The responses to monetary policy shocks are intuitive. After a surprise interest rate increase —typically of 30–40 basis points in Hungary and Poland, and of 15–20 basis points in the Czech Republic—, output seems to fall, especially in the former two countries. Consumer prices decrease too, with the trough being within the second year by -0.1 percent lower than the initial level. Since the shock is associated with an almost 1 percent appreciation, the IRFs suggest rather similar exchange rate pass-through dynamics.

The IRFs of the developed countries show a much more diverse picture. While in the UK the responses to the monetary policy shocks are intuitive and comparable to those in CEE countries, in the case of Canada and Sweden I encountered the phenomenon called price puzzle, that is, rising prices after a monetary contraction. As this might indicate the failure of identification of monetary policy shocks, I will use results for these two countries only in the case of risk premium shocks.

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11 Predictable excess return is the interest rate spread plus the expected appreciation over the corresponding period.

12 For instance, Bacchetta and Wincoop (2006) explain it with costs associated with processing information.

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The effect of risk premium shocks on the price level looks alike in CEE countries too: the immediate depreciation is followed by a relatively quick rise in prices. A rough comparison of exchange rate and CPI responses suggests, again, comparable exchange rate pass-through dynamics. The reaction of the industrial output, however, is ambiguous, which is in line with the findings of other authors[13].

The reaction of the monetary policy to risk premium shocks shows significant differences in CEE countries. While in the Czech Republic and Poland the interest rate responds slowly, with the peak being around 6 months after the shock, in Hungary the Central Bank reacts by a quick interest rate hike. After a negative risk premium shock causing similar depreciation of the currency, the MNB typically increased its policy rate within two months by 6-8 times more than the other two banks. This can be understood easily by taking into account that Hungary operated an intervention band for its currency during the first half of the sample period. Even in the beginning of the inflation targeting regime introduced in 2001, some exchange rate preferences were regularly announced, and the monetary policy showed high sensitivity to exchange rate shocks[14]. However, the IRF of the exchange rate does not seem to justify the Hungarian practice: although in the other two countries the interest rate policy seems to be less concerned about risk premium changes, the correction of the initial exchange rate shock has the same pace as in Hungary.

The reason for the MNB's failure in offsetting the effect of the risk premium on the exchange rate is that the size of the shocks are on average much larger than the interest rate steps. To demonstrate this, I calculated the predictable excess return (PER) after a risk premium shock:

$$PER_t = i_t - i_t^* - \Delta s_{t+1} \tag{2}$$

This is the deviation from the UIP, assuming that the dynamic effect of a risk premium shock can be perfectly foreseen. In other words, it is the excess yield a typical investor is expected to receive as a compensation for either the higher perceived risk

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13 Sánchez (2007), based on his calibrated theoretical model, could not impose a priori sign restriction on real output response to risk premium shocks. His empirical results confirmed this indeterminacy.

14 I also estimated the same model for Hungary on a short sample starting in 2001, when the exchange rate band was widened and the Central Bank began to announce inflation targets. The results show that high sensitivity to risk premium shocks remained even after the narrow band regime.

associated with having a home currency exposure or the lower propensity to take this risk. Predictable excess return can properly characterize the dynamics of an exogenous risk premium shock, as long as a representative investor is only interested in the total return, not in whether it was obtained from the foreign exchange or from the interest rate position. If monetary policy shocks cannot cause large and sustained deviations from the UIP, this measure is comparable across countries, even if their interest rate reactions are different with respect to exchange rate shocks.

According to Graph 1, the predictable excess return on a 3-month maturity T-bill after a negative risk premium shock is typically around 3% in CEE countries. The typical interest rate step after the shock is between 0.1 and 0.4%. Hence, the bulk of the excess return comes from the exchange rate appreciation following the immediate depreciation. Therefore, considerably much bigger steps would be needed in order to smooth the exchange rate.

Graph 1
Annualized Predictable Excess Returns on 3-Month Maturity after a One Standard Deviation Risk Premium Shock (Median Estimates)
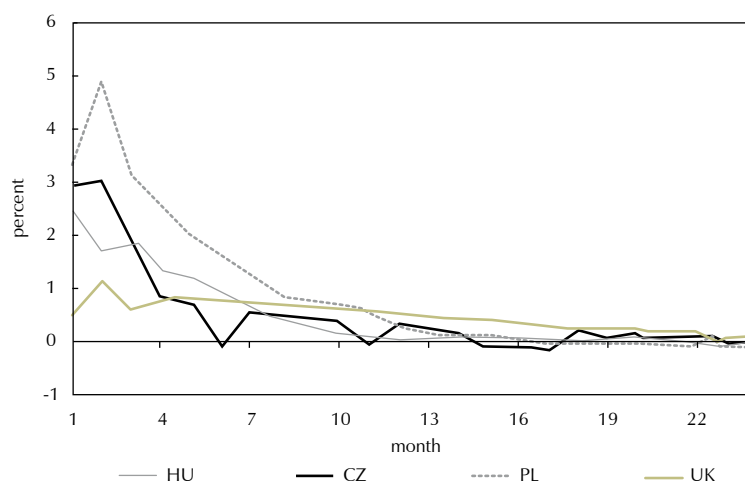
Source: Author´s calculation

Median estimates (thick line), the middle 68 percent (± 1 st. dev. for normal distribution; solid lines) and the middle 95.4 percent (± 2 st. dev. for normal distribution; dotted lines) of the posterior distribution are reported.

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IV.   THE EXCHANGE RATE PASS-THROUGH

In this section I investigate how exchange rate changes pass through to consumer prices. In particular, I will focus on the role of monetary policy. Two dimensions are of interest here: comparing pass-through estimates across different shocks, and comparing pass-through estimates across countries with possibly different monetary policies.

First, it is worth to have a look at some rough measures of exchange rate pass-through. To ensure comparability, I normalized the impulse responses by the average exchange rate response during the first year. Although the dynamics of the shocks may still differ across shocks and countries, it can be interpreted as the response of consumer prices to a shock that causes the exchange rate to depreciate by one percent in the first year.

As can be seen in Graph 2, monetary policy and risk premium shocks that depreciate the exchange rate pass through to consumer prices as well. The effect is the largest during the first two years in all countries. Although not statistically significantly, transition economies show higher exposure to these shocks as consumer prices increase more than in developed countries.
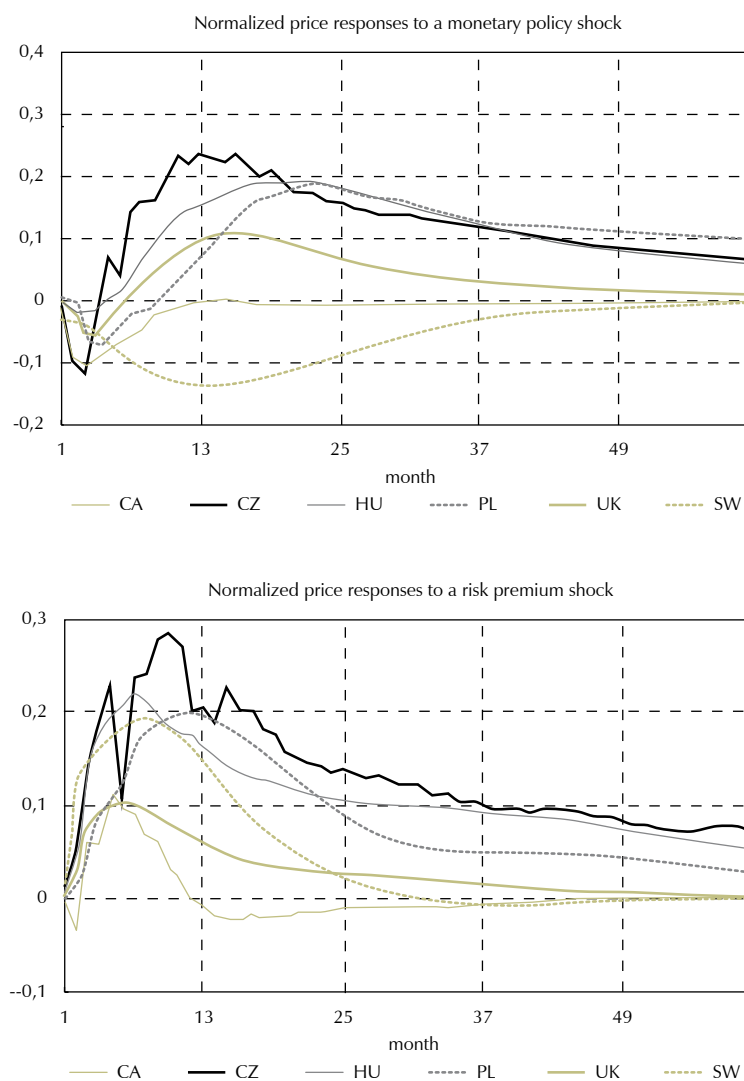
To obtain fully comparable impulse responses, it is reasonable to calculate the reaction of consumer prices to a permanent one percent depreciation of the currency for each country and each shock. Despite its conceptual drawbacks, this statistics is what is usually meant by exchange rate pass-through[15]. It is important to stress that the pass-through may be dependent on which shock caused the depreciation.

According to Graph 3, the pass-through estimates are quite close to each other in CEE countries for each shock, but slightly differ across shocks. One percent permanent depreciation of the currency due to consecutive risk premium shocks typically increases consumer prices by 0.2 and 0.3 percent by the end of the first and second year after the shock, respectively. In the case of monetary policy shocks, the corresponding values are 0.1 and 0.2. This means that risk premium shocks have a relatively large effect on the annual inflation rate in the first year, but it halves during the second and virtually disappears later. By contrast, monetary policy has a smaller,

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15   Derivation and interpretation of the exchange rate pass-through are described in detail in the appendix.

but more prolonged effect on the year-on-year consumer price index. Nevertheless, the difference between pass-through dynamics conditioned on different shocks is insignificant at usual levels.

Graph 2
Response of Consumer Prices to Unexpected Exchange Rate Depreciations
(Median Estimates)

Source: Author´s calculation

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Graph 3
The Pass-Through of One Percent Nominal Exchange Rate Depreciation to The
Consumer Prices in CEE Countries, Conditioned on Monetary Policy (Mp) and Risk
Premium (Rp) Shocks
(Median Estimates; Extremely Oscillating Responses Are Omitted from the Chart)

Price level

Annual inflation rate
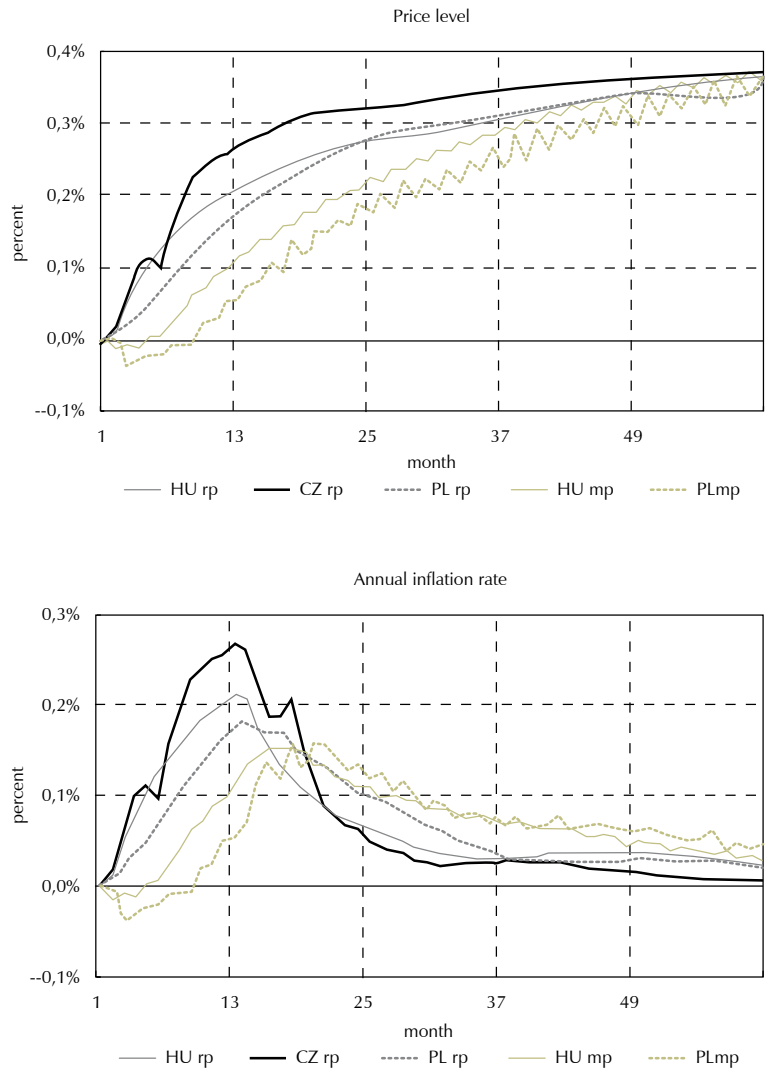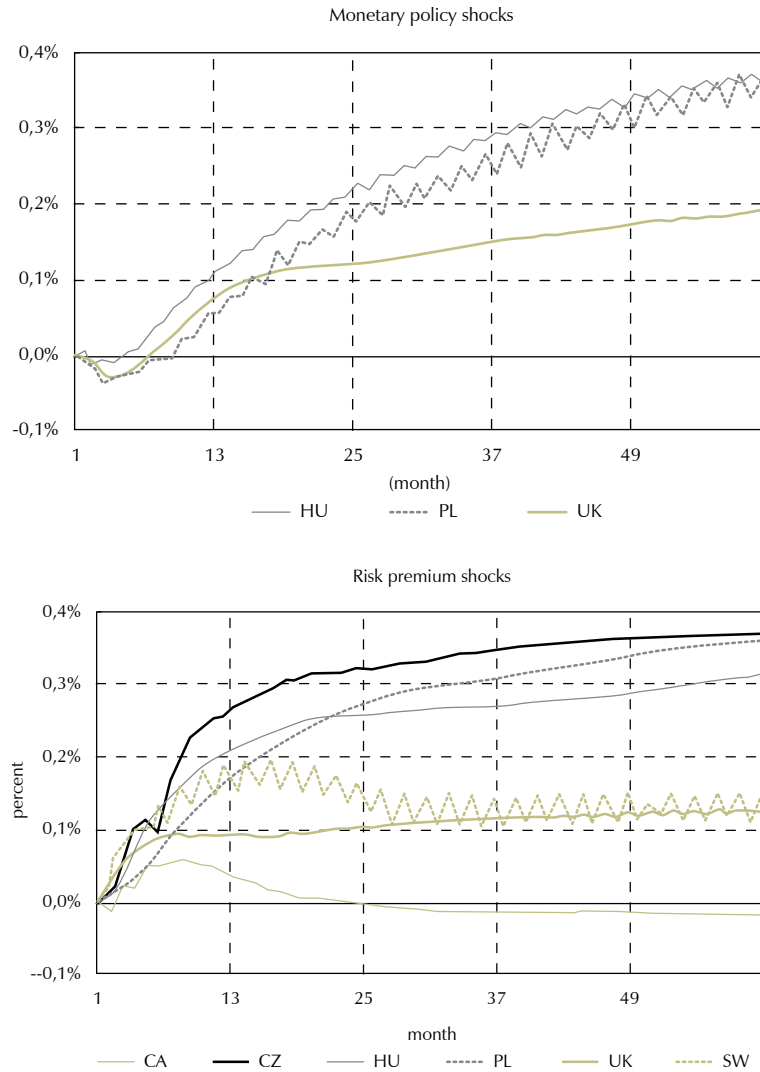

Source: Author´s calculation

The similarity of exchange rate pass-through in CEE countries is in line with our earlier findings drawn from investigating pure and normalized impulse responses. It also supports our earlier finding that the bigger weight Hungarian monetary policy gave to the exchange rate stability did not foster price stability. Although the interest rate steps of the Magyar Nemzeti Bank to the same risk premium shocks were larger, the exchange rate and the consumer prices reacted just like in the other two countries. The interest rate change was not enough to influence significantly the exchange rate; hence, the exchange rate pass-through should not differ from those of the other countries, other things being equal.

The comparison of pass-through dynamics conditioned on different types of shocks may give an impression about how price-setting works. As we noted in the previous section and according to the estimates, monetary policy has a more gradual and more protracted effect on the exchange rate than risk premium shocks do. If price-setters were rational and forward-looking, one would expect lower pass-through in the latter case, as changing prices may be costly and the temporary feature of these shocks limits the potential benefits of repricing. On the contrary, monetary policy generates more permanent revaluation of the currency, so it may pay off more to reoptimize prices in such cases.

The results do not indicate shock-dependent pass-through in CEE countries. Although there is some difference in the short-run pass-through dynamics, it is not significant enough, and it even has a counterintuitive direction. Based on the reaction of the exchange rate to these shocks, we would expect a faster reaction of consumer prices to the exchange rate movements generated by monetary policy. However, the estimates show a slightly faster pass-through in the case of risk premium shocks.

In Graph 4 we can observe again that the exchange rate pass-through is higher in transition countries, independently of the nature of the shock. Monetary policy induced depreciation has the same effect on prices in both the CEE region and in the UK during the first year, but later it seems to slow down in the UK, while continuing further in the transition economies. Moreover, the pass-through of exchange rate movements to consumer prices caused by risk premium shocks is much higher in transition economies. The role of the exchange rate in price-setting seems to be lower in Canada, Sweden and the UK. This result reinforces again that our findings for CEE countries are not pure artifact of the methodology; these countries do share some common features regarding the exchange rate pass-through.

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Graph 4
The Pass-Through of One Percent Nominal Exchange Rate Depreciation to the Consumer Prices Conditioned on Monetary Policy and Risk Premium Shocks in CEE and Developed Countries
(Median Estimates; Extremely Oscillating Responses Are Omitted from the Chart)

Monetary policy shocks

Risk premium shocks

Source: Author´s calculation

It is important to note here that by exchange rate pass-through we do not exclusively mean the strength of the exchange rate channel. In a general macro model where all the relevant variables are endogenous it is not feasible to investigate partial causal chains. Hence, our measure of exchange rate pass-through inevitably incorporates the effect of the interest rate and of other variables on prices. The difference between exchange rate pass-through dynamics, as we defined, may be therefore the consequence of different strength of other channels, such as the interest rate channel. However, since almost the same pattern can be observed in the case of monetary policy and risk premium shocks, where the interest rate goes in the opposite direction, we may feel some confidence that what we obtained is primarily about the exchange rate channel.

According to Taylor (2000), the reason for the difference in pass-through may be the credibility of the monetary policy, and the inflationary environment. While central banks in CEE countries had — during the sample period— the main objective of bringing down the inflation to the level corresponding to price stability, in Canada, Sweden and the UK the low inflation was already a fact. As Taylor argues, the observed decline in the pass-through in developed countries can be the consequence of the higher nominal stability of the nineties.

Our results are in line with Gagnon and Ihrig (2004). They estimate long-run pass-through coefficients for several developed countries controlling also for the change in the level of inflation. They found that low inflation brought about a lower rate of exchange rate pass-through in these countries. For Canada, Sweden and the UK and using a sample period almost like ours as the "low inflation era", they estimated virtually no long-run effect.

V.   CONCLUSION

In this paper I investigated the effect of monetary policy, investors' risk assessment and risk preference on the exchange rate, interest rate, industrial production and consumer prices in the Czech Republic, Hungary and Poland. The main focus was on the role monetary policy plays in the pass-through of exchange rate changes.

A VAR model was used to capture the dynamic interaction among our variables. I identified monetary policy and risk premium shocks by imposing restrictions on impulse responses. For identification, I assumed that neither of these shocks has impact on industrial production and consumer prices within one month. I also assumed that

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an interest rate increase caused by an unexpected monetary tightening is coupled with an appreciation of the currency. In the case of the risk premium shock, a higher interest rate comes with a weaker exchange rate. I estimated the same model to Canada, Sweden and the UK in order to identify features that are specific to the transition economies. Roughly ten years of data was used, starting in the middle of the nineties.

I found that the transmission of monetary policy is fairly similar in CEE countries. After a monetary contraction, consumer prices and industrial production fall in line with intuition.

Concerning risk premium shocks, the case of Hungary is somewhat special. Among the three transition countries, Hungarian monetary policy has been less tolerant towards exchange rate movements caused by financial markets. Its policy assigned a larger weight to exchange rate stability than in the other two countries. To achieve this stability, the interest rate responded faster and sharper to risk premium shocks. However, the results did not confirm that this policy resulted in higher price stability. The reason is that the interest rate responses proved to be still too small compared to the shock, and therefore could not dampen significantly the exchange rate response.

Based on the impulse responses to the identified shocks, I constructed the dynamic response of consumer prices to a hypothetical, permanent, one percent depreciation of the currency. I derived this measure of exchange rate pass-through with respect to both shocks. The implied pass-through rates were almost identical in CEE countries, independently of the nature of the shock. They were, however, higher than in the countries of the control group of developed countries. I explained the difference by the credibility of the monetary policy and the average inflation rate. While in Hungary, Poland, and to a lesser extent, in the Czech Republic the sample period could be characterized by "converging to the price stability", in Canada, Sweden and the UK it was the era of low and stable inflation. In the latter countries anchored inflation expectations may have been the main reason behind the low exchange rate pass-through.

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APPENDIX 1
Generating draws from the posterior distribution

Following Uhlig (2005), if the maximum likelihood estimator of the coefficient and the residual covariance matrix of the stacked VAR

$$Y = XB + u \tag{A1.1}$$

are denoted by $\hat{B}$ and $\hat{\Sigma}$, respectively, and with a "weak" normal-Wishart prior, the posterior distribution will be normal-Wishart as well, with the parameters $\hat{B}, \; \hat{\sigma}, \; T$ and $X'X$.

The posterior distribution of the responses to the structural shocks and other derived statistics of interest are jointly determined by the VAR posterior and the posterior of the A matrix connecting the reduced form residuals (u_t) with the structural innovations (ε_t):

$$u_t = A\varepsilon_t \tag{A1.2}$$

Therefore, in order to simulate from the posterior distribution of the impulse responses, one has to generate draws from the joint posterior of $(B, \; \Sigma, \; A)$. The posterior distribution of A conditioned on a particular VAR realization $(\tilde{B}, \tilde{\Sigma})$ can be characterized as follows: (1) the elements relating to the output and consumer prices of the columns of monetary and risk premium shocks are zero due to the contemporaneous restrictions, and (2) the impulse responses fulfill the sign restrictions. Actually, we are interested in only two columns of A, corresponding to monetary policy and risk premium shocks. Let us denote that half of A by A^{mr}. It can be shown that if we draw θs uniformly from the $[0, 2\pi]$ interval, simulation from the posterior of A^{mr} can be done by calculating the impulse responses to the shocks:

$$\tilde{A}^{mr} = \tilde{C} \begin{bmatrix} 0 & 0 \\ 0 & 0 \\ \cos\theta & -\sin\theta \\ \sin\theta & cos\theta \end{bmatrix} \tag{A1.3}$$

and keeping draws that satisfy the sign restrictions, where $\tilde{C}$ is the lower triangular matrix from the Cholesky decomposition of $\tilde{\Sigma}$. The reported posterior distributions in the paper have been made of at least 1000 draws in each case.

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APPENDIX 2
Calculation of the exchange rate pass-through

Comparing exchange rate pass-through dynamics across countries and different shocks is not straightforward. The problem is that the exchange rate itself is an endogenous variable having its own specific response to each shock. Exchange rate pass-through usually means the dynamic response of some price variables to a change in the exchange rate. How can one compare these responses if the exchange rate itself responds differently to different shocks?

The way other authors using VAR calculated exchange rate pass-through is not always clearly presented. Faruqee (2006) reports the price response divided by the exchange rate IRF. Others, like Hahn (2003), McCarthy (2006), and Ca'Zorzi et al. (2007) normalize the CPI impulse response by the size of the immediate exchange rate response. Darvas (2001), in a slightly different but simultaneous modeling framework, reports long-run price responses in a similar way. However, the comparison of these statistics is still questionable, as they may differ if the exchange rate IRFs have different shape.

My alternative approach was to create a sequence of the same type of shocks, but of different size, in order to achieve a permanent one percent depreciation of the exchange rate. I then calculated the accumulated response of consumer prices to these series of shocks. The comparison now makes sense, as these responses correspond to exactly the same exchange rate movement.

Probably the main objection to this approach could be a conceptual one. While this calculation is technically feasible, it may contradict to the logic of impulse response analysis. Impulse responses are the reactions of certain variables to a certain unexpected event that could not have been forecasted from the lags of the variables in the VAR. Impulse responses are valid as long as the shock is really unexpected. If, for example, the monetary policy tries to surprise the economic agents in a way that it generates a permanent 1 percent depreciation of the currency, agents may realize this "new" policy rule. If they learn the modified behavior of the Central Bank, they can forecast it, and their reaction will change. Our experiment, however, assumes that for several periods private agents are again and again surprised by the fact that the interest rate is still not where it should be if the monetary policy followed its rule. Since in the long run this may be —but not necessarily is— an unrealistic assumption, we restrict ourselves to formulate firm statements only for the short-run pass-through.

Even though the method looks straightforward to implement, I had some problems in calculating the pass-through conditioned on monetary policy shocks, especially in the case of the Czech Republic. The posterior distributions became extremely wide, compared to those of the original IRFs. Moreover, even the median response exhibited a disturbing oscillation.

The reason for this is the delayed overshooting, that is, it takes some months for the exchange rate to reach its maximum deviation after the shock. The standardization was based on achieving an immediate and permanent one percent shift in the exchange rate by giving a series of the same shock to the system. Therefore, a large initial monetary policy shock is needed to generate the required appreciation or depreciation. But this initial shock has an even bigger impact on the exchange rate during the next few periods. Since it would cause too much appreciation or depreciation, the monetary policy has to surprise the markets again with a shock, but now to the opposite direction. Since the earlier steps have a large cumulative effect on the exchange rate, bigger and bigger shocks with opposite signs are needed to stabilize the exchange rate. This results in an exploding oscillation in the path of other variables, like prices, which is of primary interest here.

The policy implication of this seemingly technical finding is that in the presence of delayed overshooting, tight exchange rate targeting may be difficult to implement if it is not the announced rule known by the public. The econometric implication is that estimating a response to a disturbance that is not observed in the sample requires some artificial transformations that question the validity of the results.

Since the delayed overshooting, and therefore the oscillation, is most pronounced in the Czech impulse responses, I omitted them when discussing exchange rate pass-through related to monetary policy shocks. A similar oscillation in pass-through dynamics in the Hungarian and Polish case is also present, but the magnitude is not that disturbing. Fortunately, in the case of risk premium shocks, exchange rate jumps immediately, so the modified responses of prices can be derived and compared without any difficulties.

As a control group I use first of all estimates for the UK because its IRFs for both shocks are plausible. Since the identification of monetary policy shocks was not convincing for Canada and Sweden they are considered as a benchmark only in the case of risk premium shocks.

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APPENDIX 3
Alternative methods for selecting the mean responses

Fry and Pagan (2007) criticized the practice that reports posterior distribution of impulse responses by connecting median estimates for each period. Their main argument was that, typically, the pointwise median response does not belong to any plausible shock; its points usually lie on different impulse responses. Hence, deriving conclusions about the shape of the response might be misleading. The second problem arises when one identifies more than on shock simultaneously. In this case, median responses usually belong to shocks that are not orthogonal, which renders their comparison and the use of forecast error variance decomposition invalid.

They proposed an algorithm to choose impulse responses by minimizing their squared deviation from the first period median. Using Peersman's (2005) model and data set, they demonstrated that the difference between their proposed median and the pointwise median can be substantial, affecting even the main conclusions of that study.

Liu (2007) generalized Fry and Pagan's median selection rule by minimizing the squared deviation of the responses from the pointwise median over several periods. Obviously, this approach tends to select a median response closer to the pointwise median than that of Fray and Pagan's, since the latter uses information of only the first period posterior distribution.

One may, however, raise some arguments against using these median selection algorithms. First, their methods can pick only the mean, and it cannot be extended to provide confidence bands or percentiles. Second, especially with VARs with several lags, impulse responses may oscillate excessively, which is obviously the result of imprecise estimation. While pointwise medians sometimes tend to smooth out these counterintuitive oscillating, the Fry-Pagan or Liu median usually inherits this unappealing feature. Third, there is some inconsistency between opposing the use pointwise median on one hand, and penalizing the deviation from it when choosing the mean response, on the other. Finally, choosing "median shocks" based on the distribution of the impulse responses of the VAR's variables may produce outlier results when one considers other statistics that are not linear or even not monotonous function of the original IRFs.

In order to assess whether our results depend on focusing on the popular pointwise percentiles, I calculate alternative mean responses as well. The first candidate was

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Liu's approach. I used it instead of the Fry-Pagan mean because it uses more information than the latter. Practically, I chose the pair of orthogonal monetary policy and risk premium shocks that produced impulse responses as close to the pointwise medians as possible for 60 months, or formally:

$$\min_i \sum_{t=1}^{60} \left(y_t^{m(i)} - y_t^{m,med}\right)^2 + \left(p_t^{m(i)} - p_t^{m,med}\right)^2 + \left(i_t^{m(i)} - i_t^{m,med}\right)^2 + \left(s_t^{m(i)} - s_t^{m,med}\right)^2 +$$

$$\left(y_t^{r(i)} - y_t^{r,med}\right)^2 + \left(p_t^{r(i)} - p_t^{r,med}\right)^2 + \left(i_t^{r(i)} - i_t^{r,med}\right)^2 + \left(s_t^{r(i)} - s_t^{r,med}\right)^2 \tag{A3.1}$$

where i counts the successful draws, $m(i)$ and $r(i)$ denote the response to the ith monetary policy and to the risk premium shock. Superscript med stands for the pointwise median.

I also propose another mean that does not use the pointwise median as a benchmark. I define a distance between two draws (ith and jth), similarly to the sum in the previous expression, by replacing the median impulse responses with the IRFs of the second (jth) draw. In other words, I calculate the unweighted sum of the squared differences of impulse responses for each variable to each shock over 60 periods, that is, five years[16]. Then, I choose the draw that has the minimum sum of distances from all the other draws. This algorithm minimizes the expected deviation when all draws are equally likely and the distance measure is the squared difference – analogously what the arithmetic mean does in the one dimensional case, formally:

$$\min_i \sum_j \sum_{t=1}^{60} \left(y_t^{m(i)} - y_t^{m(j)}\right)^2 + \left(p_t^{m(i)} - p_t^{m(j)}\right)^2 + \left(i_t^{m(i)} - i_t^{m(j)}\right)^2 + \left(s_t^{m(i)} - s_t^{m(j)}\right)^2 +$$

$$\left(y_t^{r(i)} - y_t^{r(j)}\right)^2 + \left(p_t^{r(i)} - p_t^{r(j)}\right)^2 + \left(i_t^{r(i)} - i_t^{r(j)}\right)^2 + \left(s_t^{r(i)} - s_t^{r(j)}\right)^2 \tag{A3.2}$$

I will call this mean "Euclidean", referring to the sum of squares distance definition.

In most cases these alternative means were within the middle 68% of the posterior distributions of the impulse responses. As mentioned earlier, it does not necessary imply that the same is true for nonlinear transformations of the IRFs. To check the

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16 Should we have a preference regarding the most relevant shocks, variables or horizons, a proper weight vector can be used.

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robustness of our main findings, I also calculated the corresponding price responses
to the permanent exchange rate shock, as defined earlier. Graph A3.1 presents the
exchange rate pass-through for both shocks, using both alternative mean measures.

Extremely oscillating and exploding lines were omitted from the charts.

Although there are some changes compared to the pointwise medians, the main find-
ing of the paper, namely, that transition countries' inflation is more exposed to ex-
change rate shocks than that of developed economies, remains valid.

Graph A3.1
The Pass-Through of One Percent Nominal Exchange Rate Depreciation to the Con-
sumer Prices Comparison of Alternative Means
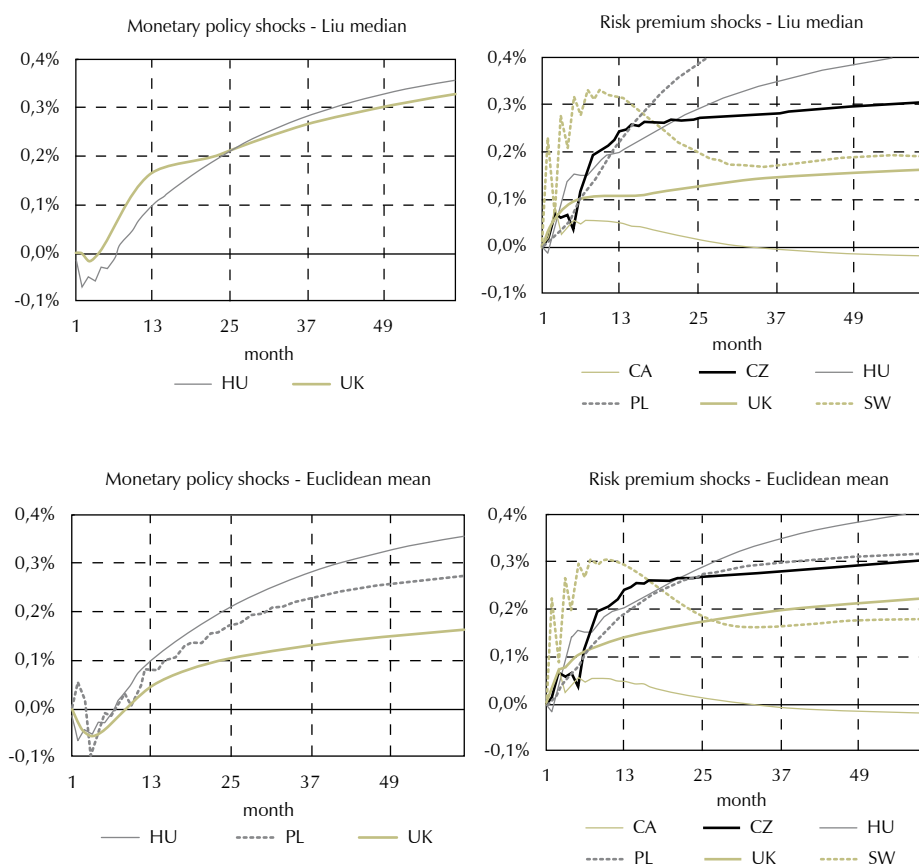
Source: Author´s calculation

APPENDIX 4
Data and VAR specification

The series are from the IMF's IFS database. The CPI series were in some cases seasonally adjusted with Demetra because the IFS data contained seasonality.
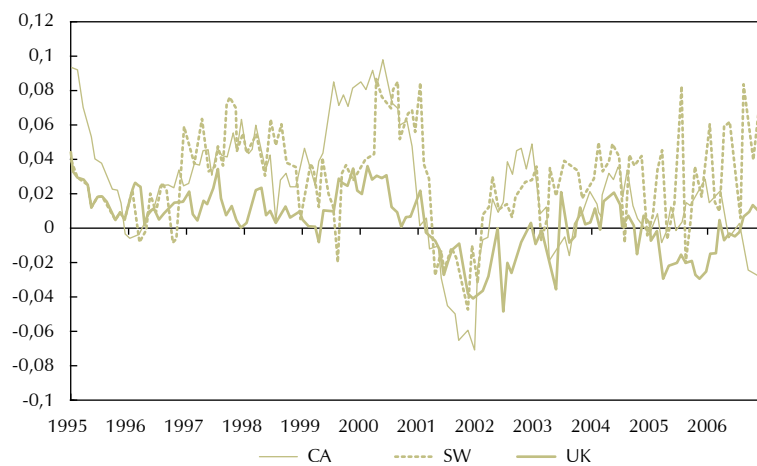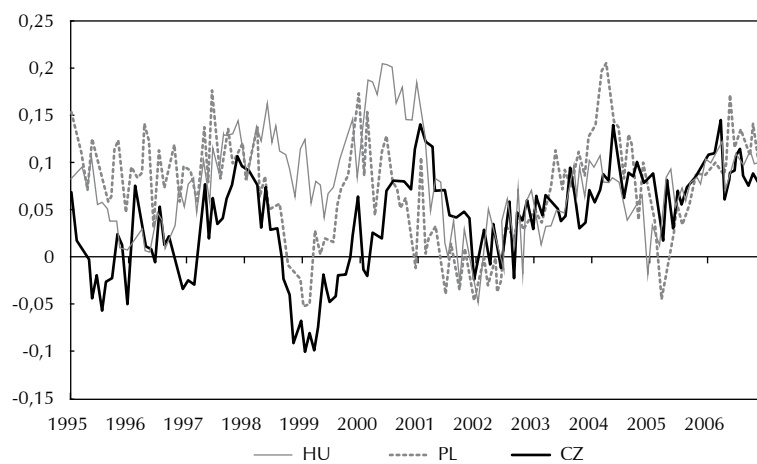
Table A4.1
IFS code of the data series and VAR specifications

Country	Industrial production	Consumer prices	Short term interes rate	Exchange rate	Estimation sample	Lags
Czecz Rep.	93566..CZF...	93564...ZF...	93560B..ZF...	935..NECZF...	M1:1998 M12:2006	7
Hungary	94466..CZF...	94464...ZF...	94460C..ZF...	944..NECZF...	M1:1995 M12:2006	4
Poland	96466..BZF...	96464...ZF...	96460C..ZF...	964..NECZF...	M1:1997 M12:2006	4
Canada	15666..CZF...	15664...ZF...	15660C..ZF...	156..NECZF...	M1:1995 M12:2006	5
Sweden	14466..CZF...	14464...ZF...	14460C..ZF...	144..NECZF...	M6:1995 M7:2006	3
U.K.	11266..CZF...	11264...ZF...	11260C..ZF...	112..NECZF...	M1:1995 M12:2006	3
Germany	13466..CZF...	13464...ZF...	13460C..ZF...	134..NECZF...		
U.S.	11166..CZF...	11164...ZF...	11160C..ZF...	111..NECZF...		

Source: Author´s calculation

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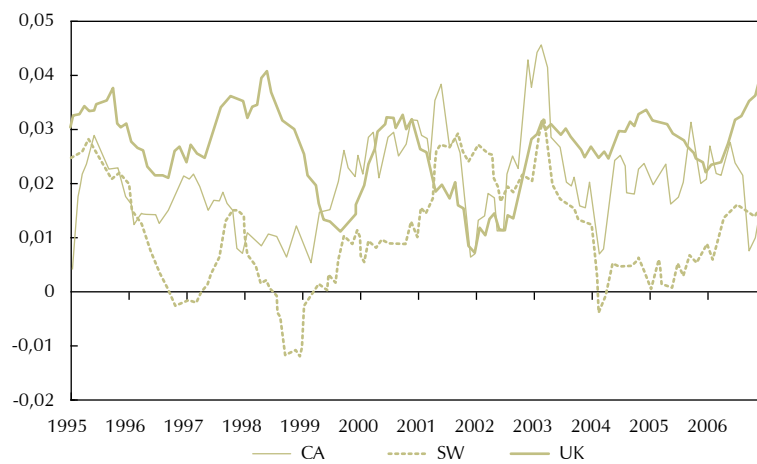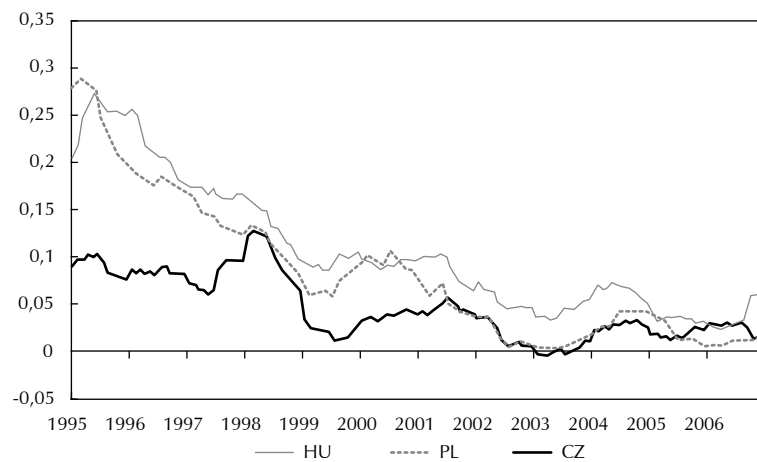in the Czech Republic, Hungary and Poland
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APPENDIX 5

Graph A5.1
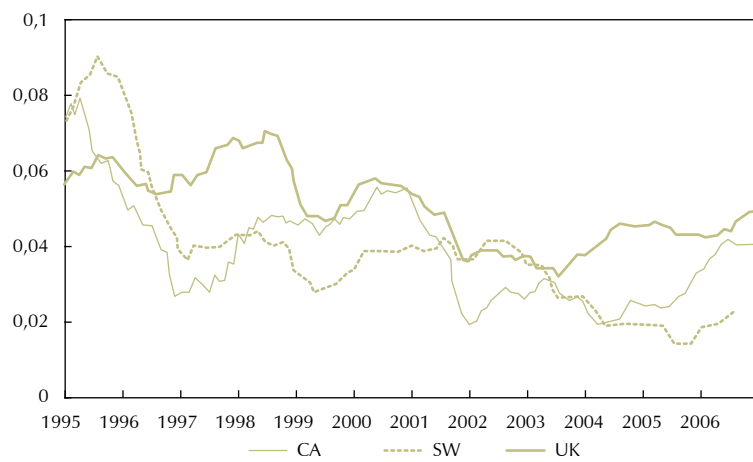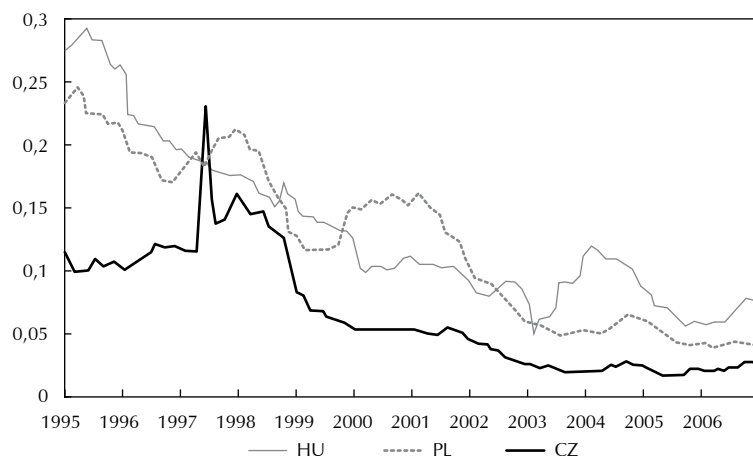Year-on-Year Growth of Industrial Production (Log-Differences)

Source: Author´s calculation

Graph A5.2
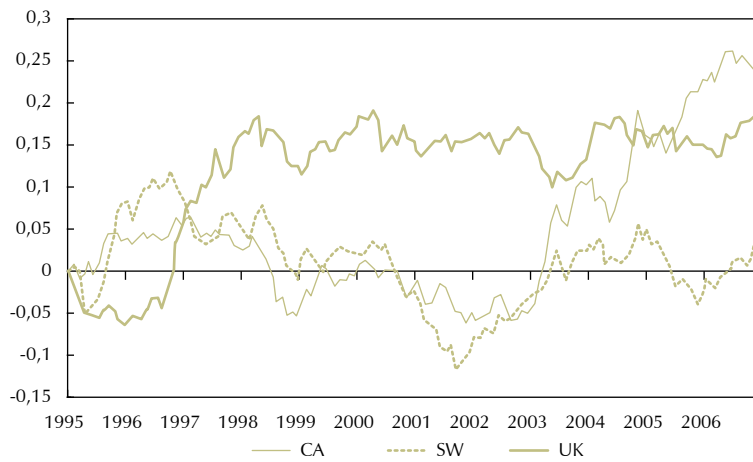Year-on-Year Growth of Consumer Prices (Log-Differences)

Source: Author´s calculation

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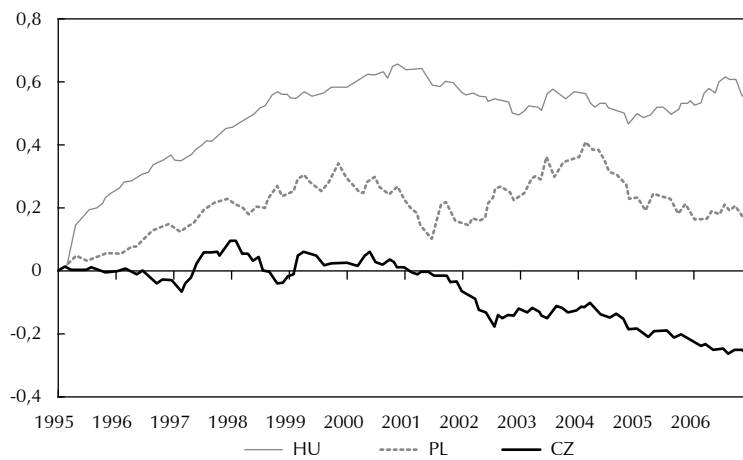
Graph A5.3
Short-Term Interest Rates

Source: Author´s calculation

Graph A5.4
Nominal Effective Exchange Rate (Logarithmic Deviation from the 1995:M1 Level)

Source: Author´s calculation

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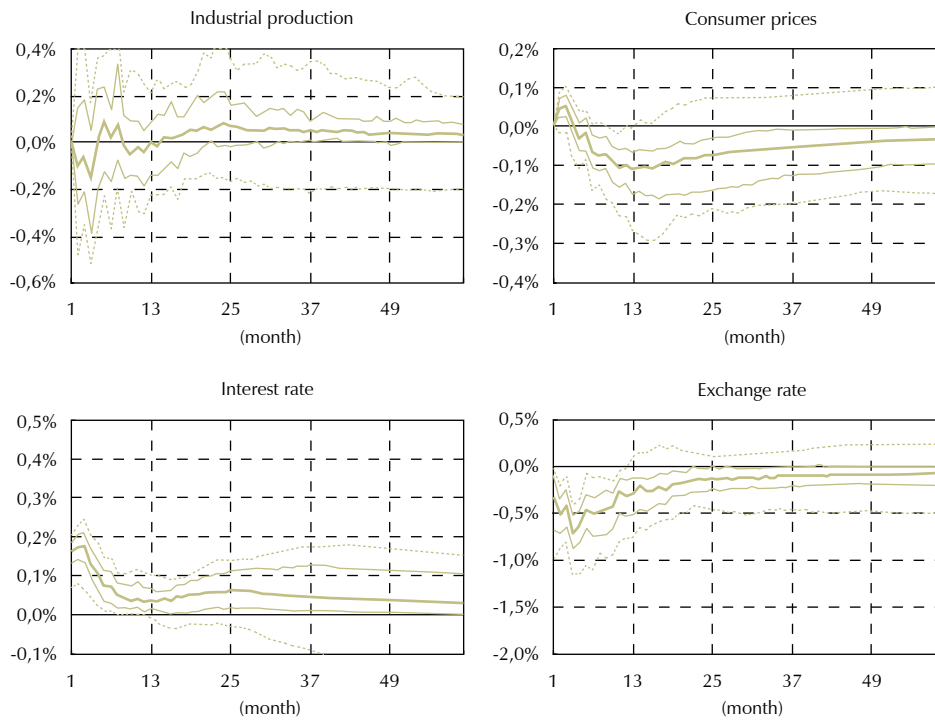in the Czech Republic, Hungary and Poland
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Graph A5.5
Impulse Responses to a One Standard Deviation Shock in the Czech Republic

A.   Monetary policy shock

Industrial production

Consumer prices

Interest rate

Exchange rate

Graph A5.5
Impulse Responses to a One Standard Deviation Shock in the Czech Republic
(Continued)

B.   Risk premium shock

Industrial production

Consumer prices

Interest rate

Exchange rate

Source: Author´s calculation

Median estimates (thick line), the middle 68 percent (± 1 st. dev. for normal distribution; solid lines) and the middle 95.4 percent (± 2 st. dev. for normal distribution; dotted lines) of the posterior distribution are reported.

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Graph A5.6
Impulse Responses to a One Standard Deviation Shock in Hungary

A.   Monetary policy shock

Industrial production

Consumer prices

Interest rate

Exchange rate

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Graph A5.6
Impulse Responses to a One Standard Deviation Shock in Hungary (Continued)
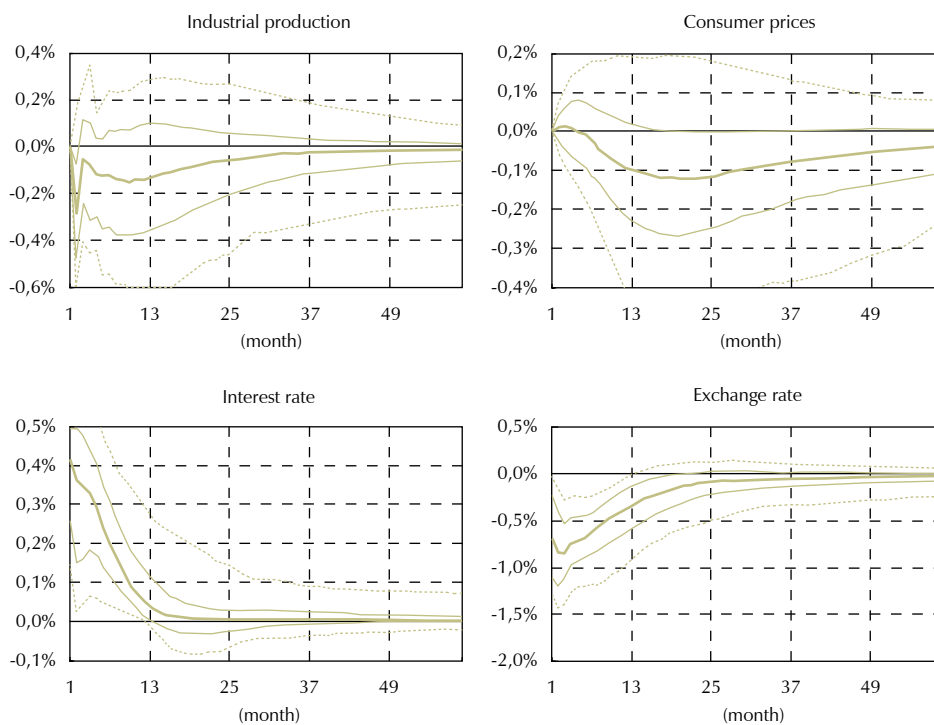
B.   Risk premium shock

Source: Author´s calculation

Median estimates (thick line), the middle 68 percent (± 1 st. dev. for normal distribution; solid lines) and the middle 95.4 percent (± 2 st. dev. for normal distribution; dotted lines) of the posterior distribution are reported.
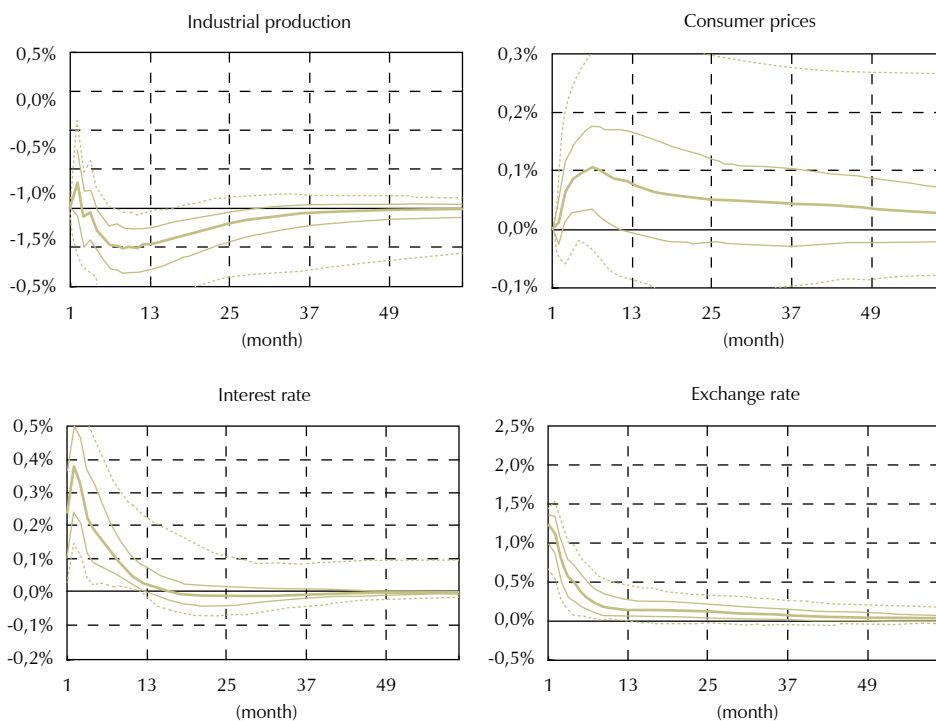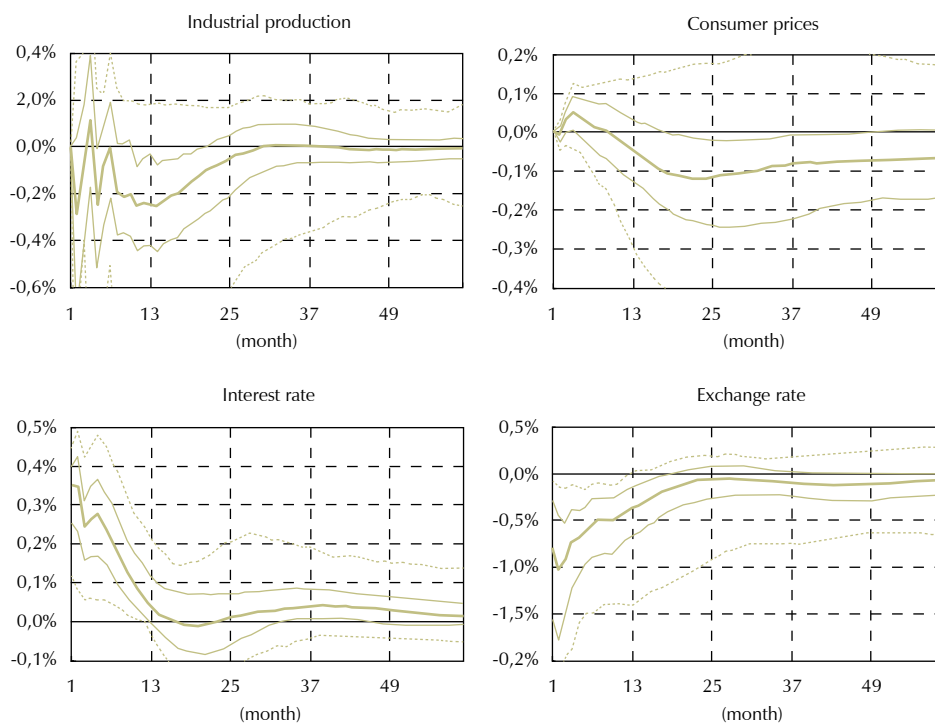
Graph A5.7
Impulse Responses to a One Standard Deviation Shock in Poland

A.   Monetary policy shock

Industrial production

Consumer prices

Interest rate

Exchange rate

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Graph A5.7
Impulse Responses to a One Standard Deviation Shock in Poland (Continued)

B.   Risk premium shock

Industrial production

Consumer prices

Interest rate

Exchange rate

Source: Author´s calculation

Median estimates (thick line), the middle 68 percent (± 1 st. dev. for normal distribution; solid lines) and the middle 95.4 percent (± 2 st. dev. for normal distribution; dotted lines) of the posterior distribution are reported.
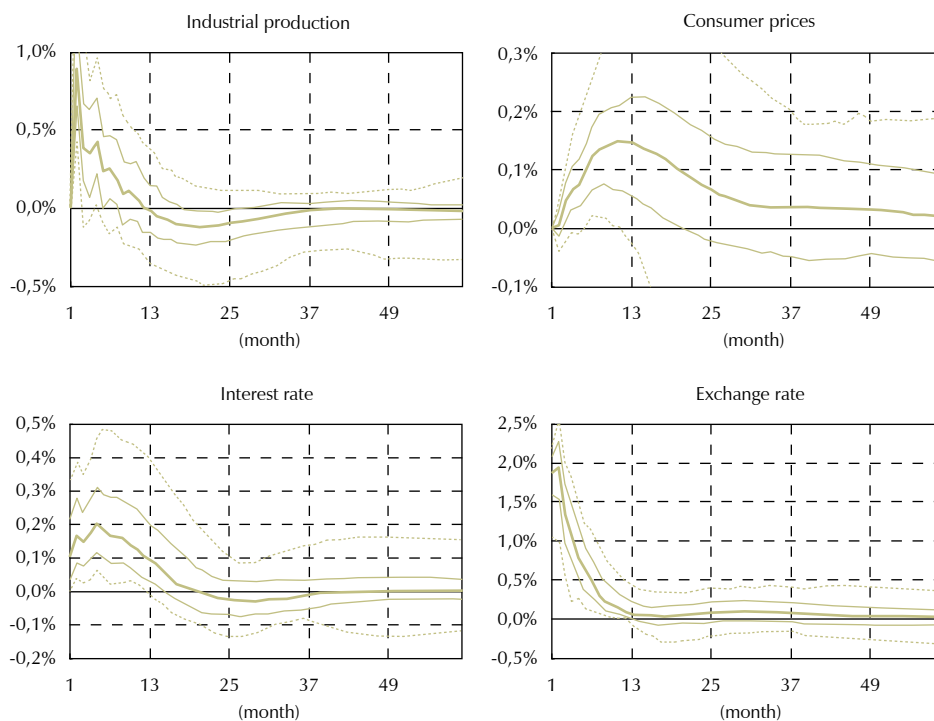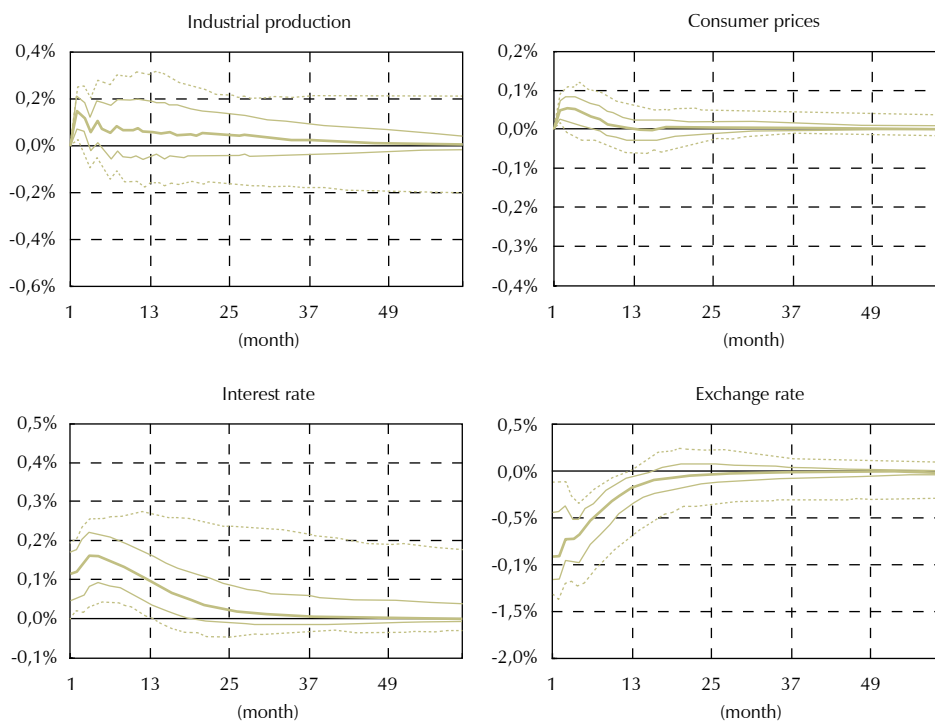
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Graph A5.8
Impulse Responses to a One Standard Deviation Shock in Canada

A.   Monetary policy shock

Industrial production

(month)

Consumer prices

(month)

Interest rate

(month)

Exchange rate

(month)

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Graph A5.8
Impulse Responses to a One Standard Deviation Shock in Canada (Continued)

B.   Risk premium shock

Industrial production

Consumer prices

Interest rate

Exchange rate

Source: Author´s calculation

Median estimates (thick line), the middle 68 percent (± 1 st. dev. for normal distribution; solid lines) and the middle 95.4 percent (± 2 st. dev. for normal distribution; dotted lines) of the posterior distribution are reported.
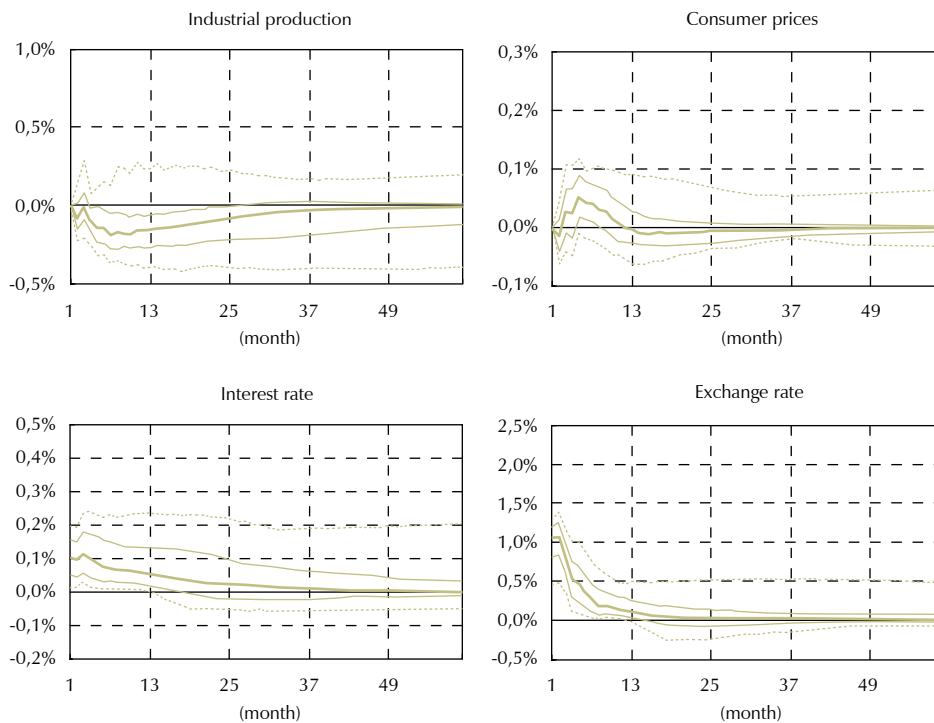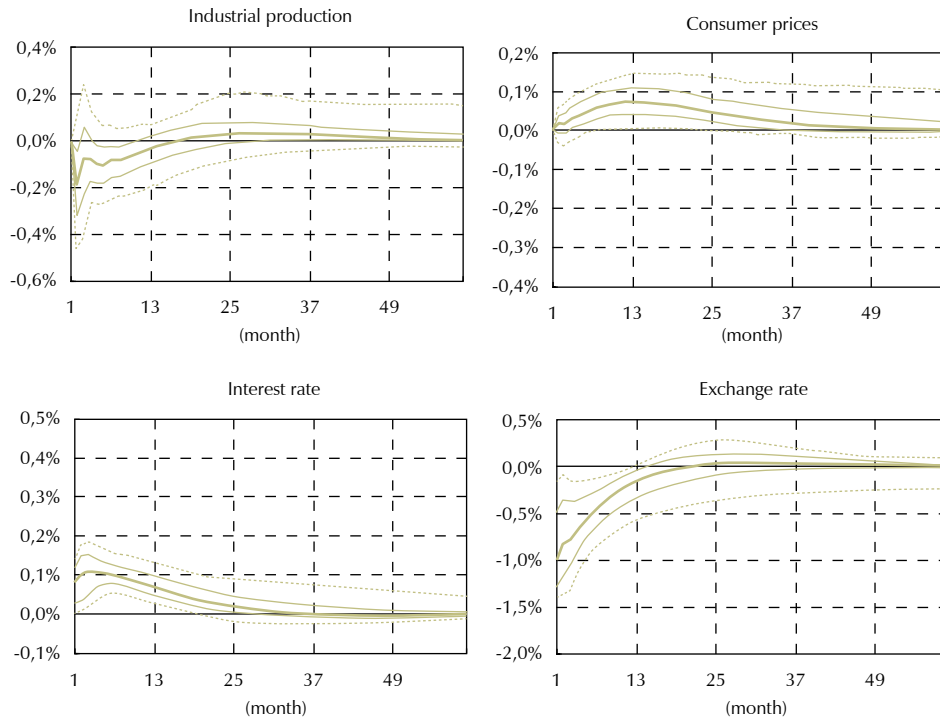
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Graph A5.9
Impulse Responses to a One Standard Deviation Shock in Sweden

A.   Monetary policy shock

Industrial production

Consumer prices

Interest rate

Exchange rate

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Graph A5.9
Impulse Responses to a One Standard Deviation Shock in Sweden (Continued)

B.   Risk premium shock

Industrial production

Consumer prices

Interest rate

Exchange rate

Source: Author´s calculation

Median estimates (thick line), the middle 68 percent (± 1 st. dev. for normal distribution; solid lines) and the middle 95.4 percent (± 2 st. dev. for normal distribution; dotted lines) of the posterior distribution are reported.
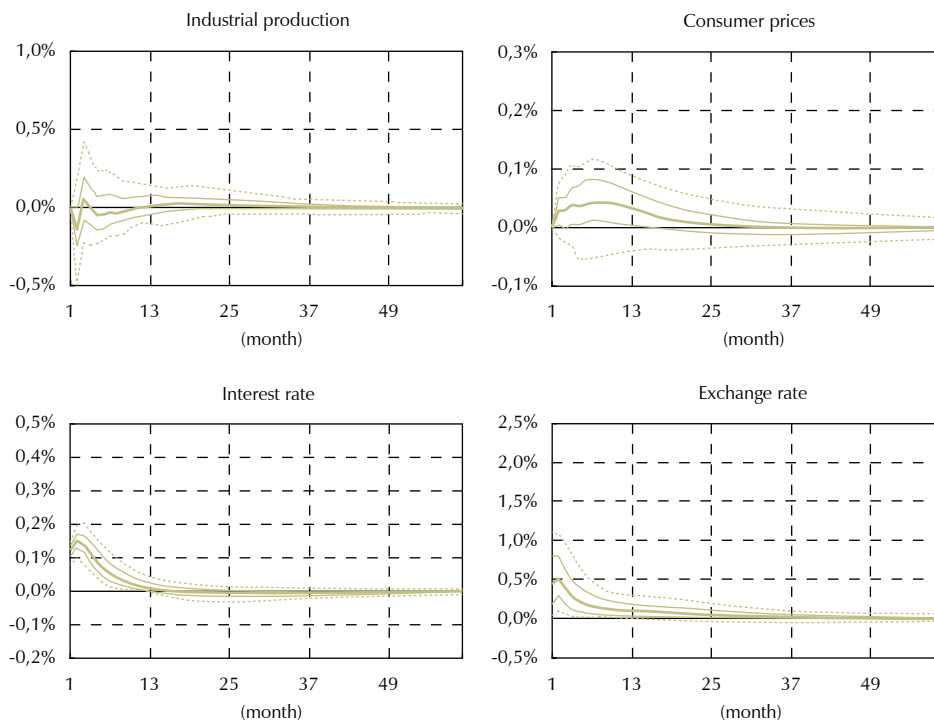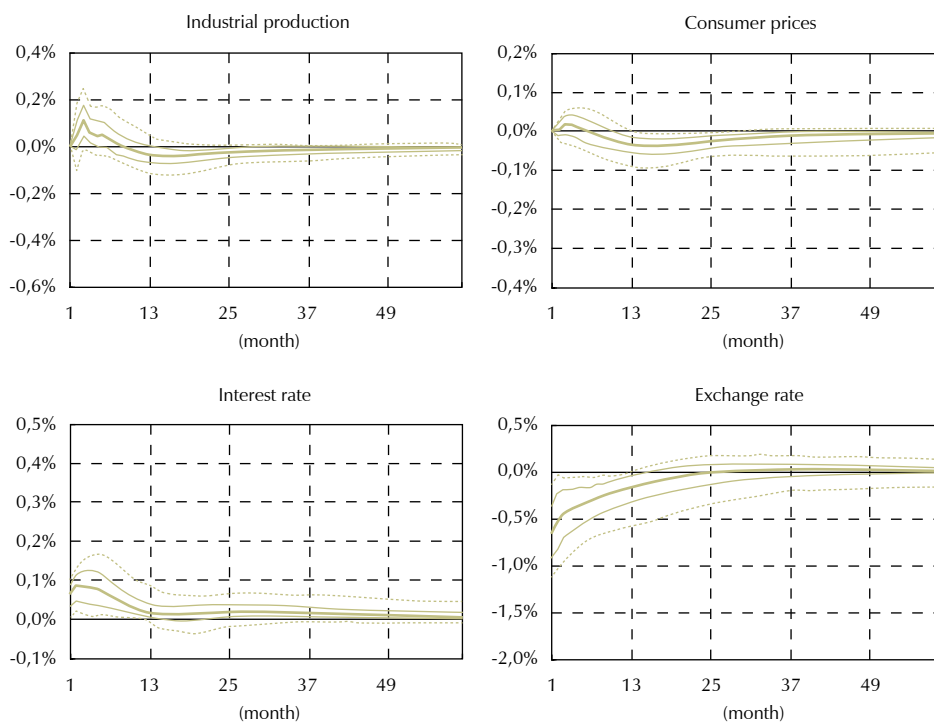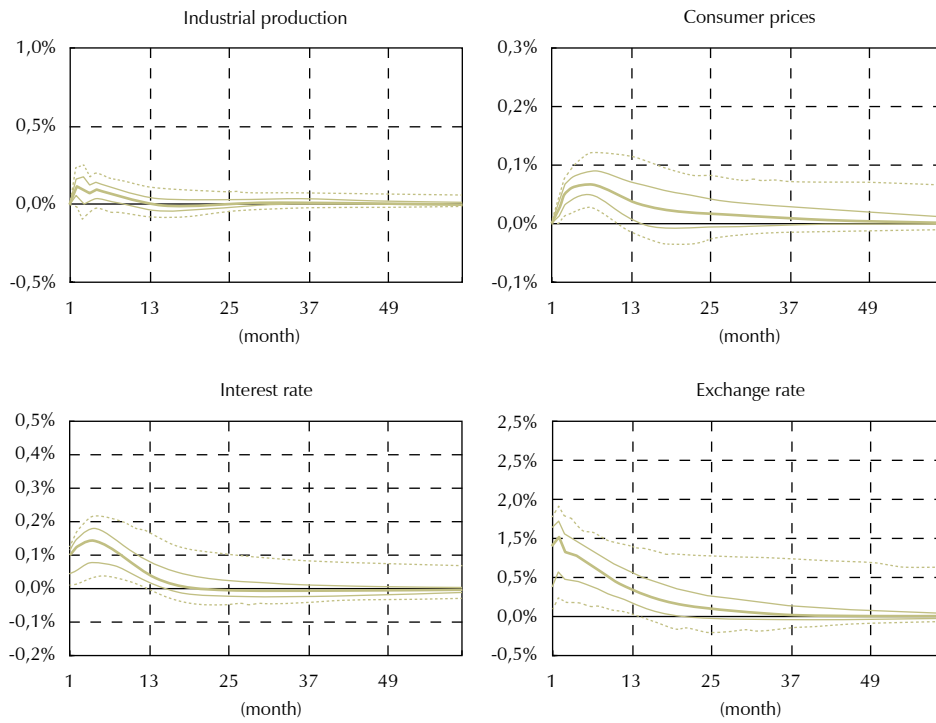
350 | Risk premium shocks, monetary policy and exchange rate pass-through
in the Czech Republic, Hungary and Poland
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Graph A5.10
Impulse Responses to a One Standard Deviation Shock in the UK

A.   Monetary policy shock

Industrial production

Consumer prices

Interest rate

Exchange rate

Ensayos sobre POLÍTICA ECONÓMICA, Vol. 28, Núm. 61, edición especial ciclos económicos globales, crisis financiera y sus efectos en las economías emergentes

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Graph A5.10
Impulse Responses to a One Standard Deviation Shock in the UK (Continued)

B.   Risk premium shock

Industrial production

Consumer prices

Interest rate

Exchange rate

Source: Author´s calculation

Median estimates (thick line), the middle 68 percent (± 1 st. dev. for normal distribution; solid lines) and the middle 95.4 percent (± 2 st. dev. for normal distribution; dotted lines) of the posterior distribution are reported.