Monetary Policy in Transition: Structural Econometric Modeling and Policy Simulations

Wojciech S. Maliszewski
Centre for Social and Economic Research CASE
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Abstract

This paper estimates Bayesian structural VAR models for the Czech Republic and Poland, allowing for changes in parameters between two monetary policy arrangements. The five-variables structural VAR is successful in identifying monetary policy shocks and their effects on the Czech and Polish economy. The model is also capable of detecting changes in monetary policy reaction functions consistent with introduction of floating exchange rate systems and switching to short-term interest rates as policy instruments. The results point to a dominant role of exchange rate in the monetary transmission mechanism.

I. Introduction

Conducting monetary policy in a transition environment is a difficult task. Available time series are short and often unreliable, there is an ongoing structural adjustment in the real sector of the economy, and even more dramatic developments in the financial sector. It is therefore not surprising that transition economies opted for exchange rate anchors at the beginning of the reform process. Fixing the exchange rate was relatively easy to implement and effective in taming inflationary expectations, but became difficult to sustain when inflation reached a moderate-level plateau. Exit from the peg was either volatile as in the Czech Republic or carried on in an orderly manner as in the case of Poland. In the latter the crawling peg system gradually evolved into a target zone with a decreasing rate of crawl of the central parity and a widening width of the band. At the end of the decade both Poland and the Czech Republic conducted their monetary policies in the framework of direct inflation targeting and with floating exchange rates. Although inflation targeting requires a thorough knowledge about a monetary transmission mechanism, understanding of this mechanism in transition countries remains weak, not only because of structural changes in the real and financial sectors, but also because of a short history of monetary policy under the floating exchange rate system. Researchers quickly documented that the highly uncertain relationship between monetary policy instruments and inflation may hamper implementation of the framework (Christoffersen et al. [2001] and Gottschalk and Moore [2001]).

This paper presents structural VAR models for analysis of the effects of monetary policy in Poland and the Czech Republic. The identifying restrictions are supplemented by Bayesian priors, which turned out to be useful in estimation of the models with the short time series available. To account for policy changes I apply the methodology recently proposed by Sims and Zha [2002] and allow for changes in some parameters of the model. The changes are restricted to parameters of the policy reaction functions, leaving coefficients of other equations constant across regimes. This specification may seem to be restrictive, but it captures well the policy changes and it is parsimonious enough to produce reasonable results in the small samples under investigation[1].

The remainder of the paper is organized as follows: the first section describes monetary and exchange rate policies in Poland and the Czech Republic. Section two discusses estimation strategy. Sections three presents estimation results. In section four I run counterfactual policy simulations, replacing policy reaction function parameters in the flexible exchange rate period by their pegged exchange rate counterparts. The last section concludes. Data sources and technical details of the estimation strategy are relegated to appendices.

I. MONETARY AND EXCHANGE RATE POLICY IN POLAND AND THE CZECH REPUBLIC

A. Poland

Poland experienced a period of near-hyper inflation after price liberalization in 1989-90. One of the main elements of the implemented stabilization package was temporary fixing

[1] The most recent policy regime has been introduced less than four years ago.

of the zloty against the dollar. The inflation was quickly reduced, but remained high and exchange rate appreciated in real terms. In 1991 the exchange rate was devalued, the dollar was replaced by a basket of five currencies and subsequently a crawling peg system replaced the peg. The National Bank of Poland (NBP) faced difficulties in conducting monetary policy in the early stabilization period, when the money-market was non-existing and the policy relied on credit ceilings. Open market operations have become the main policy instruments only since 1993. In the 1991-94 post-stabilization period, financing of fiscal deficits made the monetary policy relatively lax, but the NBP real interest rates remained at a positive level and inflation was declining in line with the exchange rate crawling peg.

In 1995 an increase in capital inflow to the country – driven by an increasing credibility of economic reforms and an ongoing privatization process – became a source of excessive monetary expansion. In order to increase the risk faced by portfolio investors and tame the capital inflow, the crawling peg system was replaced by a ±7 percent band in May 1995. As anticipated, zloty appreciated after introduction of the new system, but the new regime did not bring any significant increase in volatility since exchange rate was persistently close to the strong bound of the zone. The capital inflow slowed down only after interest rate cuts and revaluation of the central parity at the end of 1995. The revaluation, however, did not increase the exchange rate volatility because the NBP conducted a policy of keeping average deviations of the exchange rate from parity within ±2.5 per cent zone.

In 1996 a fast expansion of credit to non-financial institutions became the main source of money creation. Monetary policy reacted sluggishly to this rapid credit growth and significant interest rate increases came only at the end of 1996 and in 1997. Policy tightening was facilitated by an increased risk on international financial markets after the Asian crisis in October 1997, giving the NBP more scope for independent monetary policy actions.

In January 1998, the newly established Monetary Policy Council (MPC) adopted inflation targeting as a monetary policy strategy. The MPC considered a break from the previous exchange rate policy a necessary step in implementation of the inflation targeting framework (NBP 1998). In February 1998 the exchange rate band was widened to ±10 per cent and the exchange rate was allowed to move more freely within the zone. In October 1998 the target zone was widened to ±12.5 per cent and in April 2000 zloty started floating.

Throughout 1998 the reference rate was reduced from 24 percent in February to 15.5 percent in December. The sharp reduction was a response to economic slowdown, declining inflation, and a large portfolio capital inflows at the beginning of the year. Central bank interest rates remained at a stable level for the first half of 1999 but a higher inflation - overshooting the target in 1999 - prompted monetary authorities to reverse the cuts starting from the second half of 1999. In 2000 interest rates were further increased and the cautious period of interest rate cuts started only from the beginning of 2001, when inflation was declining and economic activity plunged into recession. Despite the relaxation of monetary policy, inflation remained below the target in 2001 and 2002.

B. The Czech Republic

In Czechoslovakia, the macroeconomic disequilibrium inherited from the Soviet-style system was less severe than in Poland and price liberalization produced a smaller effect on inflation. At the early stage of transition, the Czech National Bank (CNB)[2] controlled money supply through credit ceilings, but in 1992 the CNB switched to indirect monetary policy instruments. The koruna was initially pegged to a basket of five currencies and since 1993 to the basket of DM and USD. Inflation fell to a low, one-digit level.

From 1994 the fixed exchange rate regime combined with liberalization of the foreign exchange market, improved rating of the Czech Republic and ongoing privatization process led to a massive capital inflow, a pressure on further real appreciation and a growing current account deficit. In 1997, a growing external disequilibrium changed market sentiments towards koruna. A speculative attack, triggered by the Asian crisis, led to abandoning the peg in May 1997. The CNB adopted measures to avoid a sharp drop in the domestic currency value. The koruna was well inside the band when exit from the previous exchange rate regime was announced, leaving some scope for depreciation even within the previous,

[2] At that time the National Bank of Czechoslovakia.

narrow band. The CNB maintained a tight liquidity control after the regime change, driving up money market rates and preventing excessive depreciation. Pressure on koruna eased after June 1997 and the central bank was gradually reducing interest rates, closely monitoring developments on the foreign exchange market. Real interest rates converged to the pre-crisis level in August 1997, completing the process of interest landing after the hike triggered by the crisis (Smidkova et al. 1999).

Inflation targeting, a new regime in the Czech monetary policy, was introduced at the beginning of 1998. The CNB initially targeted net inflation rate (excluding changes in administrative prices and changes in direct taxes). Economic slowdown, brought about by the currency crisis and the tight monetary policy in its aftermath, exerted downward pressure on inflation. The 1998 target was missed by a large margin: the net inflation was 1.7 percent against the 6 ± 2 per cent target range. In mid-1998 the CNB started reducing interest rates and, in a series of cuts, the rate was lowered from 15 percent in July 1998 to 5.6 percent in December 1999. Policy relaxation did not prevent the net inflation rate to drop below the 4.5 ± 0.5 percent target in 1999. Interest rates continued falling at a slower rate in 2000 and 2001, reaching 5 percent in May 2001. Low interest rates induced net portfolio outflows, but capital inflows related to foreign direct investments put a pressure on koruna appreciation and in 2000 the currency appreciated in nominal terms. The net inflation in 2000 was again below the target (4.5 ± 1 percent). In 2001, the CNB announced the target band for a headline inflation, declining from 3-5 percent in January 2002 to 2-4 percent in December 2005. Further appreciation of koruna, despite record-low interest rates and interventions in the foreign exchange market, kept monetary conditions tight and hampered implementation of the inflation targeting framework

II. SPECIFICATION AND IDENTIFICATION OF THE MODEL

The empirical model adopted in this study is a version of a small, open-economy structural vector autoregression (SVAR)[3]. Similar models have been successfully estimated for Canada by Cushman and Zha [1997] and for other OECD countries by Kim and Roubini [1995]. Due to data limitations (short time series), the size of the model is kept at the minimum and consequently the adopted identification scheme is simpler than in the other studies. The basic model is described by the following system of simultaneous equations:

$$x_t' A_0 = z_t' A_+ + \varepsilon_t, t = 1...T \qquad (1)$$

where $x_t = [\, y_t, p_t, i_t, e_t, i_t^* \,]$ is a $n \times 1$ vector of endogenous variables (y_t is output, p_t denotes prices, i_t is short-term interest rate, e_t is exchange rate and i_t^* is a measure of returns on foreign assets), z_t is a $m \times 1$ vector of lagged endogenous variables and exogenous variables specified below, and ε_t is a $n \times 1$ vector of structural disturbances with the following distribution:

$$\varepsilon_t | Y_t \sim N(0,1) \qquad (2)$$

Five variables in the x_t vector can be grouped into two blocks: output and prices form a ”real sector” block, and short-term interest rate, exchange rate, and return on foreign assets form a ”financial sector” block. Identification of shocks in the system described by equations 1 and 2 requires imposing a sufficient number of restrictions on the A_0 matrix. Table 1 presents the imposed restrictions.

Table 1: Restrictions on coefficients of the A_0 matrix.

[3] There have been only limited attempts of estimating structural VAR for transition economies going beyond a simple recursive identification scheme. A notable exception is Dibogloo and Kutan [2000], who estimate structural VARs for Poland and Hungary.

	Real sector bloc		Financial sector bloc		
			Policy reaction function	Information equation	Foreign financial sector
p	a_0			a_5	
y	a_1	a_2		a_6	
i			a_3	a_7	
e			a_4	a_8	
i^*				a_9	a_{10}

The first column of the table lists endogenous variables. The top rows divide the table into two blocs and – when economically interpretable – into behavioral equations defined by columns with non-empty cells.

Identification of shocks in the real sector block is achieved by assuming that financial sector variables affect real sector variables only with a lag. This restriction makes the A_0 matrix bloc triangular. The real sector bloc itself is assumed to be triangular, which is a normalization. I do not attempt to give any economic interpretation to shocks and equations in the real sector bloc and the normalization does not affect identification of shocks in the financial sector bloc (Zha [1999]).

Identification of shocks in the financial sector is achieved by assuming that the foreign interest rate is not affected by domestic variables and that a time-lag in data collection and dissemination does not allow real sector variables to contemporaneously enter the policy reaction function. Additional restriction is provided by the assumption that the foreign interest rate does not contemporaneously affect variables in the policy reaction function.

The policy reaction equation may be interpreted as a function describing the behavior of an implicit monetary condition index – a linear combination of short term interest rate and exchange rate – by lagged endogenous and exogenous variables. The last restriction originates from Gerlach and Smets [1996] and Smets [1996], who show theoretically that it is optimal for monetary authorities targeting the monetary condition index not to react to foreign financial markets' shocks. The weights in the monetary condition index can be normalized to sum up to one and a successful identification should give a higher weight to exchange rate in the target zone regime and a higher weight to interest rate in the floating regime. The information equation reflects the fact that – if markets are efficient – financial variables quickly respond to shocks in all other sectors. The restrictions identifying the financial sector bloc do not depend on any particular policy regime and can be used under both pegged and floating exchange rate arrangements.

A change in the monetary policy strategy reflected in a switch to inflation targeting makes the fixed-coefficient VAR model potentially unsuitable for analysis of the whole sample. I approach this problem by allowing for a limited time-variation in coefficients of the model, closely following the methodology proposed by Sims and Zha [2002] and similar to Del Negro and Obiols-Homs [2001]. In the specification with time-varying coefficients, equation 1 is replaced by:

$$x_t^{'} A_0(s_t) = z_t^{'} A_+(s_t) + \varepsilon_t, t = 1...T \qquad (1')$$

where s_t is a state of the economy, which is assumed to be observable.

If all elements of matrices A_0 and A_+ vary across states, the number of parameters in the system described by equation 1' can be very large, making it unsuitable for short transition series. In what follows, I restrict time-variation in the system to parameters of the policy reaction function, both in the A_0 and in the A_+ matrix. I set the number of states to two, corresponding to the pegged and floating exchange rate regime. If a change in the monetary policy reaction function is the main factor behind parameters' variation, a model of this form will be correctly specified. I do not test this specification against a more flexible alternative, but it seems that the policy break was the major change in the estimation period, which for Poland runs from May 1995 (introduction of the target zone) until December 2002. Estimation period for the Czech Republic starts in March 1996 (*de facto* introduction of the target zone) and ends in June 2002. In the second half of 2002 Czech monetary authorities heavily intervened on the foreign exchange rate market. The new policy constituted a break from the floating exchange rate period, but there is not enough observations in the last regime to estimate a three-state model accounting for a change. I set the date of policy change to January 1998 for the Czech Republic (corresponding to introduction of the inflation targeting framework) and February 1998 for Poland (corresponding to widening of the target zone to ±10 percent by the Monetary Policy Council, after announcing a switch to inflation targeting).

III. ESTIMATION AND RESULTS

In empirical specification, the variables in the financial sector bloc are short term interest rates (1 month WIBOR for Poland and 1 month PRIBOR for the Czech Republic), exchange rate indices defined as a weighted average of USD and DM rates to domestic currencies (with a dollar weight of 35 percent for the Czech Republic and 50 percent and Poland), and the emerging market bond index EMBI+. The interest rates are best proxies for short term policy rates in the sample. The dollar weight in the exchange rate index for the Czech Republic is the official basket weight in the March 1996 – June 1997 period. For Poland the weight is an approximation, since the official basket included four non-dollar currencies, later replaced by the Euro. The EMBI+ index is used as a measure of return on foreign assets. Advanced transition economies are regarded as emerging markets and capital flows to these countries are affected by developments in other emerging markets to much higher extent than by changes in developed economies' interest rates (Habib [2002])[4]. Real sector variables are consumer price and industrial production indices. Industrial production and consumer price indices are seasonally adjusted and in logs, exchange rate and EMBI+ indices are in logs, and money market rates are not transformed. A dummy equal to one in May and June 1997 and zero elsewhere is introduced in the model for the Czech Republic. Introduction of the dummy reflects the assumption that large shocks related to a sudden capital outflow in this period are not drawn from the same distribution as other shocks in the

[4] Estimation of the models with 1 month LIBOR interest rates instead of the EMBI+ index yields similar, but less precise estimates.

model. Following Begg [1998], I assume that the capital outflow was driven by a confidence crisis, which was contemporaneously exogenous to other variables in the model. A description of variables is given in appendix A. Appendix B discusses the Bayesian estimation technique.

As discussed in Sims (2000), estimation of the system like the one analyzed here with flat priors attributes a high share of variation to a deterministic component of the model. Sims' (2000) solution to the problem is to use priors pushing the system into a non-stationary region. Although my base priors for coefficients of the A+ matrix are unit root priors (Minnesota priors), I do not add other priors advocated by Sims (2000) and Sims and Zha (2002) to enforce the non-stationarity of the system. According to Sims (2000), the use of these priors should not be automatic but should rather depend on investigator's prior knowledge about the behavior of the system. Powerful price liberalization shocks and deep recessions experienced in transition countries make the initial conditions for the system different from its subsequent behavior. Attributing a significant share of variation to the deterministic component seems to be a reasonable strategy in this case[5].

The lag length in the VAR is set to six, which is a balance between limiting degrees of freedom and allowing for a reach dynamics in the system. All calculations are conducted in OX (Doornik [1999]) and programs replicating the results are available on request.

Coefficients of the A_0 matrix

Tables 2 and 3 lists modes, 5th and 95th percentiles of the posterior distribution of coefficients in the A_0 matrix for two regimes for the Czech Republic and Poland.

[5] In other words, the estimated model leans toward a hypothesis that the gradual disinflation path experienced by transition economies could had been expected at the start of the process.

Table 2: Coefficients of the A_0 matrix for the Czech Republic.

	Regime 1			Regime 2		
	5th percentile	Mode	95th percentile	5th percentile	Mode	95th percentile
a_0	34.58	38.85	43.50	34.58	38.85	43.50
a_1	-120.30	-65.91	-11.04	-120.30	-65.91	-11.04
a_2	296.01	329.73	366.60	296.01	329.73	366.60
a_3	30.62	92.82	136.58	37.84	104.44	132.59
a_4	-144.21	-113.46	-80.52	-51.70	-29.82	-13.28
a_5	-16.21	-9.74	-3.62	-16.21	-9.74	-3.62
a_6	-84.69	-35.46	16.67	-84.69	-35.46	16.67
a_7	60.81	93.22	130.22	60.81	93.22	130.22
a_8	20.54	50.06	63.22	20.54	50.06	63.22
a_9	3.82	11.94	17.22	3.82	11.94	17.22
a_{10}	25.95	28.55	31.45	25.95	28.55	31.45

Table 3: Coefficients of the A_0 matrix for Poland.

	Regime 1			Regime 2		
	5th percentile	Mode	95th percentile	5th percentile	Mode	95th percentile
a_0	34.41	39.03	44.13	34.41	39.03	44.13
a_1	-199.03	-133.22	-70.64	-199.03	-133.22	-70.64
a_2	307.37	346.08	388.34	307.37	346.08	388.34
a_3	17.61	30.07	36.46	287.99	325.49	383.64
a_4	-167.33	-121.29	-109.52	-28.01	-9.21	-2.90
a_5	-1.45	4.80	10.68	-1.45	4.80	10.68
a_6	-70.55	-14.82	49.08	-70.55	-14.82	49.08
a_7	42.42	51.72	66.16	42.42	51.72	66.16
a_8	59.26	68.90	77.00	59.26	68.90	77.00
a_9	4.93	9.55	14.47	4.93	9.55	14.47
a_{10}	25.65	28.51	31.56	25.65	28.51	31.56

Coefficients of the policy reaction function a_3 and a_4 validate our identification restrictions. For the fixed exchange rate regime in the Czech Republic the mode of the weight attached to exchange rate in the monetary condition index is much higher than the weight attached to the interest rate. For the floating exchange rate regime, the reverse is true. Both weights are precisely determined.

For the target zone regime in Poland, the mode of the interest rate weight in the monetary condition index, a_3, is positive and lower than the mode of the exchange rate weight a_4. In the floating exchange rate regime the reverse is true.

Dynamic responses to policy shocks in two regimes

Responses of consumer price index, industrial production, money market rate and exchange rate to a monetary policy shock are plotted on figures 1 and 2 for the Czech Republic and Poland respectively. Graphs in left columns of figures 1 and 2 show the results for the first regime. Right columns on these graphs show the results for the second regime. The graphs report posterior modes and 5th and 95th percentiles of the posterior distribution.

In both regimes the policy shock leads to a simultaneous appreciation of exchange rate and increase in interest rate, but the path of response is different, reflecting a change in operation of the monetary policy in these periods. The response of macroeconomic variables is similar in both regimes. Output decreases relatively quickly after the shock, prices take longer to react but decline permanently. Industrial production drops to the lowest level 10 to 12 months after the shock and recovers afterwards. Impulse responses of the CPI indicate that prices fully fall about 20 months after a monetary policy tightening and the effects of the shock are permanent.

In Poland, the reaction of financial variables has expected sign in both regimes. Dynamic responses of macroeconomic variables are again consistent with theoretical presumptions and other VAR studies: output decreases quickly and prices react later but more persistently. Industrial production drops to the lowest level 5 to 10 month after the shock and recovers afterwards. Prices fall to the lowest level about 15-20 months after a monetary policy tightening and the effects of the shock are persistent. In the inflation targeting period responses of production block variables are qualitatively similar but weaker than in the first regime, especially in case of the CPI response.

IV. Counterfactual Simulation of the 1998 Policy Change

In this section I attempt to estimate the effects of the 1998 change in policy framework on the behavior of policy instruments, inflation and output for the Czech Republic and Poland. The aim of these simulations is to validate the adopted identification procedure and estimation results. Following Sims and Zha (2002) the counterfactual history is modeled as follows. For each draw from the parameters' density I save a sequence of unit variance structural shocks. In the next step I generate a one-step to n-step ahead forecast (where n is the last period of the simulation) starting from February 1998 and using the saved structural shocks and parameters from the previous policy regime. The counterfactual paths for year-on-year changes in industrial production, year-on-year changes in CPI, short-term interest rate and the exchange rate index are plotted on figures 3 and 4 for the Czech Republic and Poland.

For Poland, the simulated path for the exchange rate is remarkably different from the actual one. Lower than actual depreciation of the currency in 1999 results in lower inflation rate in 2000. Lower appreciation in 2001, combined with lower interest rates in this year, lead in turn to significantly higher inflation rate in 2002. The difference between the actual and the simulated path for industrial production follows a similar pattern: the simulated industrial production is lower than the actual in 2000 and higher at the end of the sample. The results of simulations give a quantitatively reasonable interpretation of the effect of changes in operating procedures of monetary policy on macroeconomic aggregates.

For the Czech Republic, the differences between the actual and the simulated path are more dramatic. The results suggest that managing the nominal exchange rate along a stable, before-the-crisis path would result in a significantly lower inflation rate and would require significantly higher nominal interest rates. These results validate the policy of the Czech National Bank following the crisis.

V. CONCLUSIONS

In this paper I attempt to shed some light on the effects of monetary policy on output and inflation and evaluate the effects of changes in the way monetary policy is conducted in the Czech Republic and Poland after introduction of inflation targeting framework in 1998. The five-variables structural VAR methodology adopted in the study is successful in identifying monetary policy shocks and their effects on the Czech and Polish economies. The time-varying model is capable of detecting changes in the policy reaction function consistent with introduction of the floating exchange rate system and switching to short-term interest rate as the main policy instrument. The results indicate the dominant role of exchange rate in the monetary transmission mechanism. The typical unexpected monetary tightening during the fixed exchange rate and the target zone period leads to a persistent appreciation of the exchange rate and a temporary increase in the short term interest rate. In the inflation targeting framework the monetary policy tightening is reflected in significant and persistent increase in the short-term interest rate, also leading to an exchange rate appreciation. Responses of prices and output are consistent with macroeconomic theory and other VAR studies: both prices and output decline after a contractionary monetary policy shock, but the response of output is faster than the response of prices.

Figure 1. Czech Republic: Impulse responses to a monetary policy shock in two regimes: 1993.7-1998.1 (left column) and 1998.2-2002.4 (right column) with 95% confidence interval.

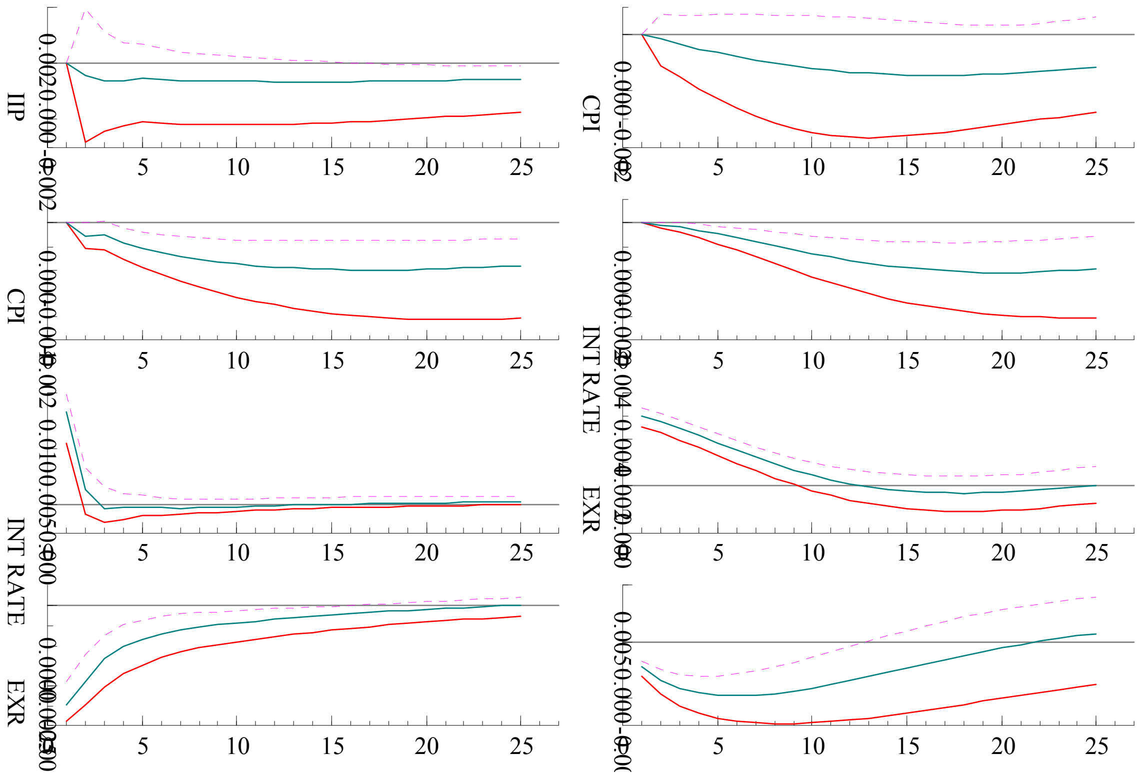


Figure 2. Poland: Impulse responses to a monetary policy shock in two regimes: 1993.1-1998.1 (left column) and 1998.2-2002.4 (right column) with 95% confidence interval.

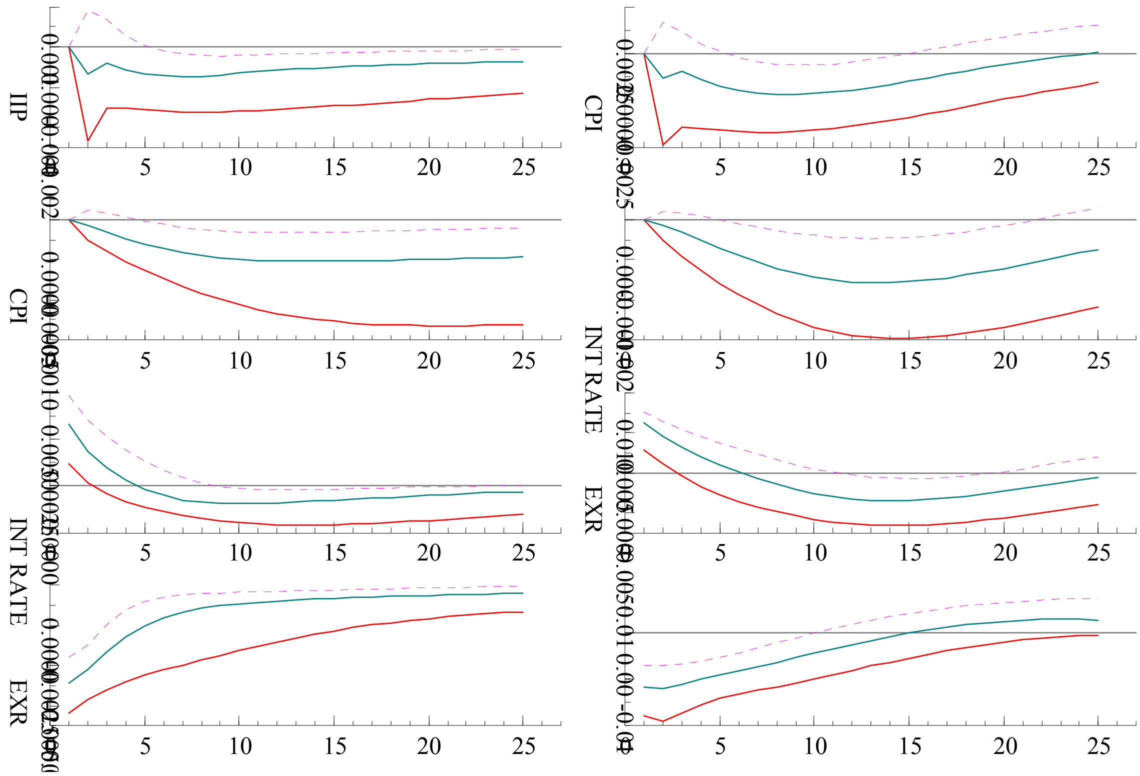


Figure 3. Czech Republic: Counterfactual simulation of policy rule from the crawling peg regime operating in the floating exchange rate regime.

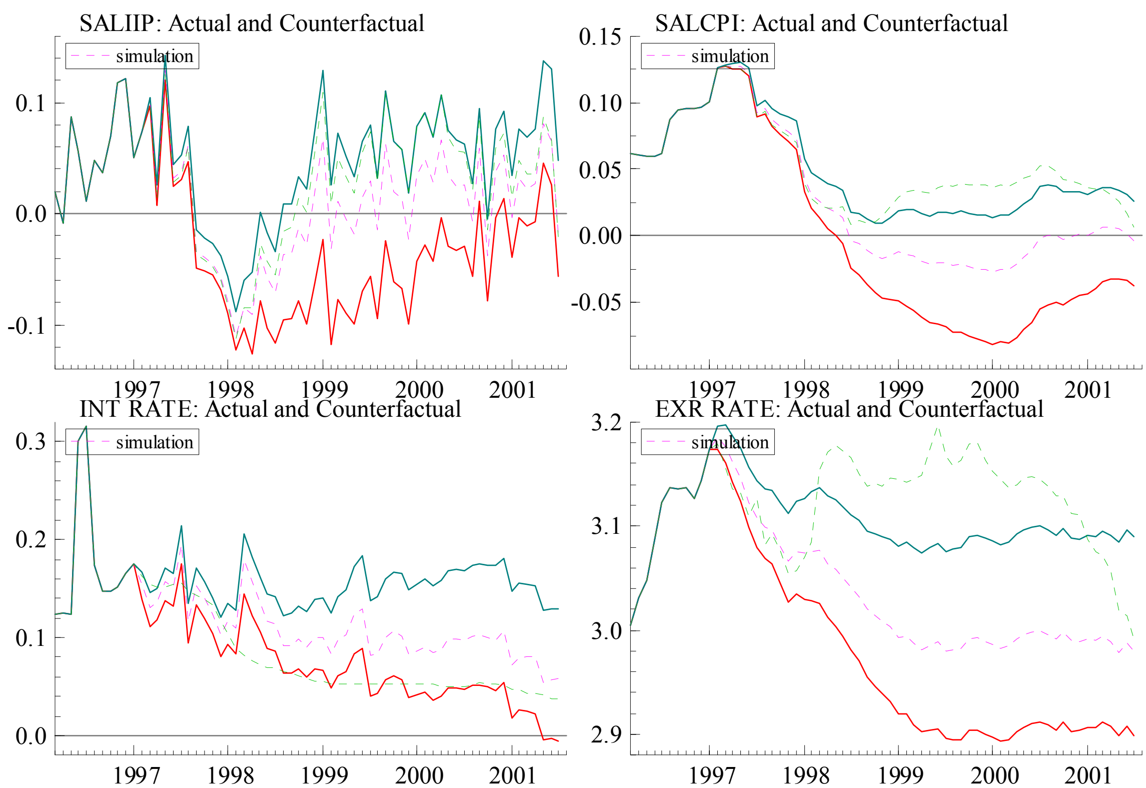
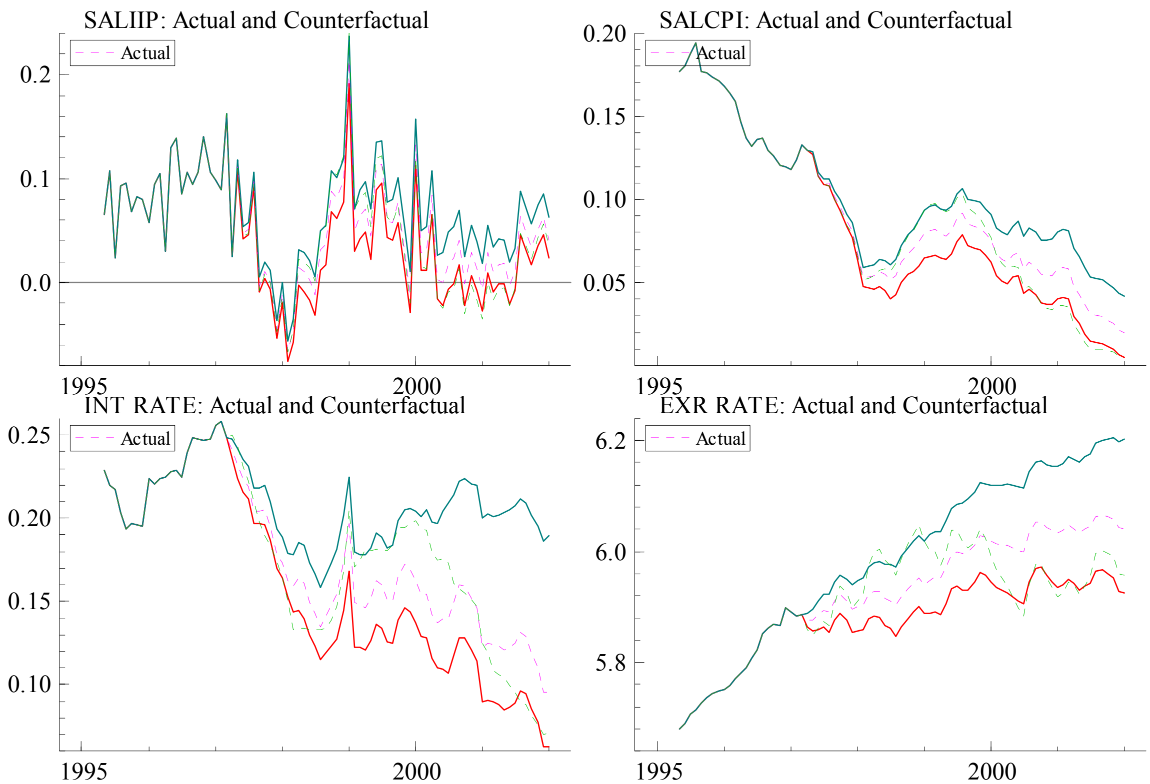


Figure 4. Poland: Counterfactual simulation of policy rule from the crawling peg regime operating in the floating exchange rate regime.



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Appendix A. Data sources and transformations

Money market rates:

Czech Republic: 1 month PRIBOR - Czech National Bank

Poland: 1 month WIBOR - Datastream

Exchange rate indices:

Czech Republic: (0.35 log(CZK/USD) + 0.65 log(CZK/DM)) - Czech National Bank

Poland: (0.5 log(PLN/USD) + 0.5 log(PLN/DM)) - National Bank of Poland

Consumer Price Indices CPI (log of X11 seasonally adjusted series)

Czech Republic: Czech National Bank

Poland - Central Statistical Office

Index of Industrial Production IIP (log of X11 seasonally adjusted series)

Czech Republic: Czech National Bank

Poland: Central Statistical Office

Appendix B. Bayesian estimation

The estimation method for the multi-regime model in this paper is adapted from Del Negro and Obiols-Homs (2001) (who estimate a similar model for Mexico) and Sims and Zha (2002). The following description follows closely Sims and Zha but is simplified due to our assumption that dates of regime changes are known.

Following Sims and Zha, I begin by rewriting the model 1 in the following way

$$y_t' A_0^{(k)} = x_t' [D^{(k)} + \bar{S} A_0^{(k)}] + \varepsilon_t' \qquad (1)$$

where

$$\bar{S} = \begin{bmatrix} I_{nxn} \\ 0_{(m-n)xn} \end{bmatrix} \qquad (2)$$

I specify priors for non-zero coefficients in A_0 (for regimes k=1,2) in the same way as in Sims and Zha (2002), assuming a joint Gaussian prior distribution with independent individual elements with mean zero and standard deviation set to $\lambda_0 / \hat{\delta}_i$ for the *i*'th raw of $A_0^{(k)}$, where λ_0 is a hyperparameter and $\hat{\delta}_i$ is a standard deviation from an estimated autoregressive process for the *i*'th variable. For $D^{(k)}$, I specify a joint Gaussian prior distribution with independent individual elements with mean zero and standard deviation of the coefficient on lag *l* of variable *j* in each equation given by $\frac{\lambda_0 \lambda_1}{\hat{\delta}_j l^{\lambda_3}}$, where λ_1, λ_3 are hyperparameters and $\hat{\delta}_j$ is a standard deviation from an estimated autoregressive process for the *j*'th variable. This formulation of the prior is equivalent to specifying a random walk prior for reduced form VAR coefficients.

Denoting the *j*'th column of $A_0^{(k)}$ by $a_{0,j}^{(k)}$ and the *j*'th column of $D^{(k)}$ by $d_j^{(k)}$ for $j = 1, \ldots, n$ and $k = 1, 2$, I write the proposed prior distributions for elements of $A_0^{(k)}$ in equation *j* as:

$$a_{0,j}^{(k)} \sim N(0, H_{0j}), k = 1,2, j = 1,...,n$$
$$d_{0,j}^{(k)} \sim N(0, H_{+j}), k = 1,2, j = 1,...,n \qquad (3)\ (4)$$

I also follow Sims and Zha incorporating the two sets of "dummy observations" pushing the model towards the non-stationarity region but not excluding cointegration among the series. Given very short sample under investigation, the effect of these dummies on dynamic properties of the model is very strong and the model generates reasonable impulse responses only when variance of these priors is made substantially greater than in the Sims and Zha's application. In the final specification I scale the Sims and Zha's dummy observations by a factor of 0.01. As discussed in the main text, this also reflects our beliefs that a large part of variation in the series can be explained by deterministic part of the model due to a slowly decaying effect of initial conditions.

I further rewrite the model by stacking $a_{0,j}^{(k)}$ and $d_j^{(k)}$ for $k = 1;\ 2$ into:

$$a_{0,j} = \begin{bmatrix} a_{0,j}^{(1)} \\ a_{0,j}^{(2)} \end{bmatrix}, d_j = \begin{bmatrix} d_j^{(1)} \\ d_j^{(2)} \end{bmatrix}$$

In order to avoid overparametrization of the model, the coefficients in a_j and d_j must be restricted. In our model I impose restrictions that $d_j^{(k)}$ are the same for all *k*'s and $a_{0,j}^{(k)}$ are the same for *k* = 1 and 2 in the non-policy block of the model. I then follow Waggoner and Zha by expressing coefficients in $a_{0,j}$ and $d_{0,j}$ in terms of "free" coefficients by writing these restrictions as:

$$Q_j a_{0,j} = 0, R_j d_j = 0, j = 1,...,n \qquad (5)$$

where Q_j is a nk x nk matrix with rank q_j and R_j is a mk x mk matrix with rank r_j. Denoting by U_j a nk x qj matrix such that the columns of U_j form an orthonormal basis for the null space of Q_j and denoting by V_j a mk x rj matrix such that the columns of V_j form an orthonormal space for the null space of R_j , I can express a_j and d_j in terms of a q_j x nk vector b_j and a r_j x mk vector g_j satisfying the following relationships:

$$a_{0,j} = U_j b_j, d_j = V_j g_j, j = 1,...,n \tag{6}$$

Vectors b_j and g_j contain ”free” parameters of the model and the original parameters in a_j and d_j can be recovered by linear transformations U_j and V_j.

The prior for coefficients in b_j and g_j are obtained by combining equations 3, 4 and 6:

$$b_j \sim N(0, \overline{H}_{0j}), j = 1,...,n \tag{7}$$
$$g_j \sim N(0, \overline{H}_{+j}), j = 1,...,n \tag{8}$$

where

$$\overline{H}_{0j} = (U_j^{'}(I \otimes H_{0j}^{-1})U_j)^{-1}, j = 1,...,n$$
$$\overline{H}_{+j} = (V_j^{'}(I \otimes H_{0j}^{-1})V_j)^{-1}, j = 1,...,n$$

Introducing notation:

$$Y^{(k)} = \begin{bmatrix} y^{'}_{t_1^{(k)}} \\ . \\ . \\ y^{'}_{t_{T_k}^{(k)}} \end{bmatrix},\ X^{(k)} = \begin{bmatrix} x^{'}_{t_1^{(k)}} \\ . \\ . \\ x^{'}_{t_{T_k}^{(k)}} \end{bmatrix}, k=1,2$$

where T_k is the total number of observations in regime *k;* and $t_1^{(k)}$ and $t_{T_k}^{(k)}$ are respectively the first and the last observation from regime *k*, the likelihood function expressed in terms of original parameters is proportional to:

$$L \propto \prod_{k=1}^{2} \left|\det(A_0^{(k)}\right|^{T_k} \prod_{k=1}^{2} \exp\left\{-\frac{1}{2}\mathrm{trace}[Y^{(k)}A_0^{(k)} - X^{(k)}(D^{(k)} + \bar{S}A_0^{(k)})^{'}[Y^{(k)}A_0^{(k)} - X^{(k)}(D^{(k)} + \bar{S}A_0^{(k)})]\right\} =$$

$$\prod_{k=1}^{2} \left|\det(A_0^{(k)}\right|^{T_k} \prod_{k=1}^{2} \exp\left\{-\frac{1}{2}[Y^{(k)}a_{0,j}^{(k)} - X^{(k)}(d_j^{(k)} + \bar{S}a_{0,j}^{(k)})^{'}[Y^{(k)}a_{0,j}^{(k)} - X^{(k)}(d_j^{(k)} + \bar{S}a_{0,j}^{(k)})]\right\} =$$

$$\prod_{k=1}^{2} \left|\det(A_0^{(k)}\right|^{T_k} \prod_{k=1}^{2} \exp\left\{-\frac{1}{2}[a_{0,j}^{'(k)}Y^{(k)'}Y^{(k)}a_{0,j}^{(k)} - 2(d_j^{(k)} + \bar{S}a_{0,j}^{(k)})^{'}X^{(k)'}Y^{(k)}a_{0,j}^{(k)} + (d_j^{(k)} + \bar{S}a_{0,j}^{(k)})'X^{(k)'}X^{(k)}(d_j^{(k)} + \bar{S}a_{0,j}^{(k)}\right.$$

The above expression can be simplified by introducing the following notation:

$$\widetilde{\Delta}_0^{-1} = \text{diag}\left\{[Y^{(k)'}Y^{(k)} - 2\bar{S}'X^{(k)'}Y^{(k)} + \bar{S}'X^{(k)'}X^{(k)}\bar{S}]_{k=1}^{2}\right\}$$
$$\widetilde{\Delta}_{+0} = \text{diag}\left\{[X^{(k)'}Y^{(k)} - X^{(k)'}X^{(k)}\bar{S}]_{k=1}^{2}\right\}$$

$$\widetilde{\Delta}_+^{-1} = \text{diag}\left\{[X^{(k)'}X^{(k)}]_{k=1}^{2}\right\}$$

where $\text{diag}\left\{[X^{(k)'}X^{(k)}]_{k=1}^{2}\right\}$ is a matrix $X^{(1)'}X^{(1)}$ and $X^{(2)'}X^{(2)}$ on the diagonal.

Using the newly introduced symbols, the likelihood function is proportional to:

$$L \propto \prod_{k=1}^{2}\left|\det(A_0^{(k)}\right|^{T_k}\prod_{k=1}^{2}\exp\left\{-\frac{1}{2}[a'_{0,j}\widetilde{\Delta}_0^{-1}a_{0,j} - 2d'_j\widetilde{\Delta}_{+0}a_{0,j} + d'_j\widetilde{\Delta}_+^{-1}d_j]\right\}$$
(9)

Equation 9 can be re-written in terms of free parameters *bj* and *gj* implicitly defined in equation 6:

$$L \propto \prod_{k=1}^{2}\left|\det(A_0^{(k)}\right|^{T_k}\prod_{k=1}^{2}\exp\left\{-\frac{1}{2}[b'_jU'_j\widetilde{\Delta}_0^{-1}U_jb_j - 2g'_jV'_j\widetilde{\Delta}_{+0}U_jb_j + g'_jV'_j\widetilde{\Delta}_+^{-1}V_jg_j]\right\} \quad (10)$$

Combining equation 10 with priors from equation 8 and 7, and completing the squares in *gj* gives the following conditional posterior distribution for *gj* and the marginal density kernel for *b*:

$$\pi(g_j|Y_T, b) = N(\widetilde{g}_j, (zV'_j\widetilde{\Delta}_+^{-1}V_j + \overline{H}_{+j}^{-1})^{-1}) \quad (11)$$

where

$$\pi(b|Y_t, g_j)$$
$$\propto \prod_{k=1}^{2}\left|\det(A_0^{(k)}\right|^{T_k}\prod_{k=1}^{2}\exp\left\{-\frac{1}{2}[b'_j(U'_j\widetilde{\Delta}_0^{-1}U_j + \overline{H}_{0j}^{-1})b_j - (V'_j\widetilde{\Delta}_{+0}U_j)'(V'_j\widetilde{\Delta}_+^{-1}V_j + \overline{H}_{+j}^{-1})^{-1}(V'_j\widetilde{\Delta}_{+0}U_j)]b_j\right\}$$
(12)

where $\widetilde{g}_j = (V'_j\widetilde{\Delta}_+^{-1}V_j + \overline{H}_{+j}^{-1})^{-1}(V'_j\widetilde{\Delta}_{+0}U_j)]b_j$

Alternatively, combining equation 10 with priors from equation 7 gives the following conditional density kernel for *b*:

$$\pi(b|Y_t, g_j)$$
$$\propto \prod_{k=1}^{2} \left|\det(A_0^{(k)}\right|^{T_k} \prod_{k=1}^{2} \exp\left\{-\frac{1}{2}[b_j'(U'_j \widetilde{\Delta}_0^{-1} U_j + \overline{H}_{0j}^{-1})b_j - 2g'_j V_j' \widetilde{\Delta}_{+0} U_j b_j)\right\}$$

(13)

The posterior density of parameters of the model can be estimated by Gibbs sampling, combining direct draws from distribution in equation 11 with Metropolis-Hastings step for drawing from distribution in equation 13. Alternatively, modal estimates of *b* coefficients can be obtained by maximizing the expression in equation 12 with respect to *b* and modal estimates of *g*'s can be obtained by substituting estimated *b*'s in the expression for the mean of normal distribution in equation 11. Posterior density of *b* can then be approximated by a multivariate normal density with variance-covariance matrix equal to the inverse of the Hessian of equation 12 (as a function of *b*) evaluated at the maximum. This first approach is adopted here.