Monetary model of exchange rate determination: evidence from the Czech Republic, Hungary and Poland

Victor Shevchuk[1]

Abstract

Using a monetary model of exchange rate determination that suggests a strong link between the nominal exchange rate and a set of monetary fundamentals, exchange rate dynamics for the Czech Republic, Hungary and Poland is studied. As the cointegration relationship between exchange rate, output and the monetary fundamentals (money supply and interest rate) is found, VAR/VEC and 2SLS error-correction models are used in this context, since both approaches allow estimate short-run correlations between exchange rates and fundamentals while taking into account the existent long-run exchange rate constraints. Based on the quarterly data for the 1998–2012 period, it is found that for all countries an increase in the money supply, domestic output slowdown or stronger growth abroad are factors behind a nominal exchange rate depreciation, just as predicted by the monetary model of exchange rate. However, the effects of domestic-foreign interest rate differential are quite heterogeneous, being in line with theoretical predictions of a standard monetary model for Poland only. According to the decomposition of variance, money supply and interest rates account for 30% to 46% of the exchange rate variation in the Czech Republic, from 10% to 14% in Hungary, and from 23% to 42% in Poland.

Keywords**:** *monetary model of exchange rate, the Czech Republic, Hungary, Poland, error-correction models*

JEL Classification**:** E41; F31; F37

AMS Classification**:** 62P20

1. Introduction

Main implication of the monetary model of exchange rate determination is that the nominal exchange rate is determined by relative levels of money supply, output and interest rate [4; 10; 11]. Despite strength of its theoretical foundations, empirical evidence in favour of the monetary model had not been overwhelming, especially in the 1980s and 1990s [2; 14]. Based on the study of fiscal and monetary models of exchange rates over the 1974–1993 period for main industrial countries, Chinn [7] concluded that no model consistently and significantly outperforms a random walk, with the fiscal models outperforming monetary models in out-of-sample forecasting exercises. However, recently there are numerous studies in favour of the monetary model, including the assumption of cointegration between nominal exchange rates and monetary fundamentals [3; 5; 6]. Arguments in favour of the monetary model are found for the Central and East European (CEE) countries as well [8].

[1] Cracow University of Technology, Institute of Economy, Sociology and Philosophy, Cracow, 24 Warszawska Street, Poland, victorshevchuk@netscape.net.

A main contribution of this paper is reconsideration of a standard monetary model of exchange rate determination for the Czech Republic, Hungary and Poland. In Section 2, theoretical arguments and empirical evidence for the monetary model are discussed. Section 3 contains description of data and statistical methodology. Estimation results are presented in Section 4. Main findings are summarized in section 5.

2. Theoretical framework and empirical evidence for the monetary model to exchange rate determination

Regardless of the price specification[2], the monetary model assumes a stable real money-demand function in domestic and foreign countries (in logs):

$$m_t - p_t = \alpha_1 y_t - \beta_1 i_t, \tag{1}$$

$$m_t^* - p_t^* = \alpha_1 y_t^* - \beta_1 i_t^*, \tag{2}$$

where m_t and m_t^*, p_t and p_t^*, y_t and y_t^*, i_t and i_t^* are the domestic and foreign money supply, price level, output and nominal interest rate, respectively.

Assuming that the purchasing power parity (PPP) does hold,

$$e_t = p_t - p_t^*, \tag{3}$$

rearrangement of equations (1) and (2) for domestic and foreign price levels brings about a flexible-price monetary model:

$$e_t = m_t - m_t^* - \alpha_1(y_t - y_t^*) + \beta_1(i_t - i_t^*) + \varepsilon_t, \tag{4}$$

where e_t is the nominal exchange rate (units of domestic currency per foreign currency).

Following Frenkel [11], the nominal exchange rate is defined as:

$$e_t = m_t - m_t^* - \alpha_1(y_t - y_t^*) + \gamma_1(\pi_t^e - \pi_t^{e*}) + \varepsilon_t, \tag{5}$$

where π_t^e and π_t^{e*} are the expected domestic and foreign inflation rates, respectively.

As noticed by Hsing [12], several well-known monetary models share the assumption of equations (4) and (5) for the money supply and relative output effects, but differ in respect to

[2] Regarding different assumptions of the price setting, there are several versions of the monetary approach to exchange rate determination: (i) the flexible price monetary model, (ii) the sticky price monetary model, (iii) the sticky price monetary model augmented with relative price differential [6].

the effects of interest rate differential. According to the Frankel model, an inverse relationship between the interest rate differential and a nominal exchange rate is explained by the capital flows effects. Recent empirical studies are predominantly in support of the monetary model. While Basher and Westerlund [3] established that the monetary model emerges for industrial countries only when the presence of structural breaks and cross-country dependence has been taken into account, strong evidence for cointegration between nominal exchange rates and monetary fundamentals is found by Cerra and Saxena [5]. With the use of the autoregressive vector-error-correction model (VAR/VEC), empirical support for the monetary model is found for Mexico [13] and Turkey [6]. Crespo-Cuaresma *et al.* [8] show that the monetary model, augmented for the Balassa–Samuelson effect, provides a good description of nominal exchange rate trends in several CEE counries. For Poland, Hsing [12] established that the monetary model can explain the behaviour of the zloty/USD exchange rate reasonably well, with all the components of the model (4) having expected signs.

3. Data and statistical methodology

We used quarterly time series data for the Czech Republic, Hungary and Poland for the period 1998–2011, as provided by the online IMF *International Financial Statistics*. The data consist of a nominal exchange rate vis-à-vis US dollar (E_t), the money supply (M_t), the real GDP (Y_t), domestic interest rate on government bonds (I_t) and the LIBOR six-month rate (I_t^*), as a proxy for foreign interest rate. As all the variables are I(1) processes, it is possible to make use of the Johansen cointegration test to take into account short run dynamics of the exchange rate. As reported in Table 1, the Johansen test implies the existence from $r = 3$ (Hungary, Poland) to $r = 5$ (the Czech Republic) cointegrating vectors at 5% of confidence between exchange rate, money supply, output and interest rate for all three countries.

	Lags	LR(0)	LR(1)	LR(2)	LR(3)	LR(4)	LR(5)
Czech Republic	2	243.1*	163.8*	107.5*	70.2*	34.3*	12.6**
Hungary	2	138.3*	91.4*	51.9*	26.1**	5.9	0.4
Poland	2	150.8*	97.5*	49.6*	23.6***	8.3	2.4

Note: LR(r) denotes the likelihood ratio statistic for H_0: r cointegrating vectors against H_1: stationary VAR; * denotes rejection of the hypothesis of H_0 at the 1% level. ** at the 5% level, *** at the 10% level.

Table 1 Cointegration tests for the nominal exchange rate and its determinants.

Following the Engle–Granger two-step methodology [9], cointegration of the data containing unit roots allows estimate the long-run relationship (in levels)

$$E_t = \alpha + \beta \mathbf{X}_t + \varepsilon_t, \tag{6}$$

and then use the lagged residuals to estimate a short-run dynamics (in first differences)

$$\Delta E_t = \delta_0 + \delta \Delta \mathbf{X}_t - \gamma \varepsilon_{t-1} + \xi_t, \tag{7}$$

where $\mathbf{X}_t$ is a vector of independent variables, ε_t and ξ_t are stochastic factors.

As the Engle–Granger procedure cannot deal with cases of more than one cointegrating vectors (Table 1), an alternative procedure of VAR/VEC can be used [1]:

$$\Delta \mathbf{X}_t = \sum \delta_i \Delta \mathbf{X}_{t-i} - \gamma \varepsilon_{t-1} + \xi_t, \tag{8}$$

where γ measures feedback coefficients, which allows estimate the instantaneous correlations (very short run) between exchange rates and fundamentals while taking into account the existent long-run exchange rate equation in the estimation procedure.

4. Estimation results

As implied by the Engle–Granger two-step methodology, the long-run and short-run 2SLS estimates of the exchange rate vis-à-vis US dollar are as follows (uppercase letters are for the levels, and lowercase ones are for the first differences):

a) the Czech Republic

$$\begin{array}{llllll} E_t = & 1.259E_{t-1} & -0.667E_{t-2} & -0.355M_t & -0.106Y_t & +0.842Y_t^* \\ & (8.35^*) & (-3.87^*) & (-1.78^{***}) & (-0.17) & (1.73^{***}) \\ & +0.305I_t & -0.015I_t^*; & & & \\ & (1.77^{***}) & (-0.88) & & R^2 = 0.94 & ADF = -2.98^{**} \end{array} \tag{9a}$$

$$\begin{array}{lllll} e_t = & 0.622m_t & -3.219y_t & +1.099y_t^* & \\ & (1.75^{***}) & (-4.72^*) & (1.22) & \\ & -0.597i_t & +0.097i_t^* & +0.141\varepsilon_{t-1}; & \\ & (-1.62) & (3.20^*) & (0.831) & R^2 = 0.35 \quad ADF = -3.02^* \end{array} \tag{9b}$$

b) Hungary

$$\begin{array}{llllll} E_t = & 1.174E_{t-1} & -0.288E_{t-2} & -0.053M_t & -0.249Y_t & +0.525Y_t^* \\ & (9.70^*) & (-2.58^{**}) & (-0.97) & (-1.17) & (2.66^{**}) \\ & -0.064I_t & +0.004I_t^*; & & & \\ & (-1.68^{***}) & (0.76) & & R^2 = 0.92 & ADF = -3.52^{**} \end{array} \tag{10a}$$

$$
\begin{aligned}
e_t = \; & -0.182 e_{t-2} && +0.629 m_t && -2.502 y_t && +1.463 y^*_{t-1} \\
& (-1.72^{***}) && (2.37^{**}) && (-2.73^{*}) && (1.47) \\
& -0.104 i_t && +0.057 i^*_t && +0.251 \varepsilon_{t-1}; && \\
& (-1.73^{***}) && (1.81^{***}) && (1.83^{***}) && R^2 = 0.20 \quad ADF = -3.84^{*}
\end{aligned} \tag{10b}
$$

c) Poland

$$
\begin{aligned}
E_t = \; & 0.669 E_{t-1} && +0.548 M_t && -1.768 Y_t && +0.306 Y^*_t \\
& (7.60^{*}) && (3.02^{*}) && (-3.56^{*}) && (1.83^{**}) \\
& +0.033 I_t && -0.038 I^*_t && +0.077 crisis; && \\
& (1.43) && (-3.01^{*}) && (2.51^{**}) && R^2 = 0.90 \quad ADF = -3.08^{**}
\end{aligned} \tag{11a}
$$

$$
\begin{aligned}
e_t = \; & 0.931 e_{t-1} && +0.786 m_t && -2.233 y_t && +1.494 y^*_{t-1} \\
& (5.64^{*}) && (2.90^{*}) && (-3.47^{*}) && (1.95^{***}) \\
& +0.094 i_t && -0.021 i^*_t && -1.061 \varepsilon_{t-1}, && \\
& (1.34) && (-0.74) && (-2.87^{***}) && R^2 = 0.43 \quad ADF = -3.64^{*}
\end{aligned} \tag{11b}
$$

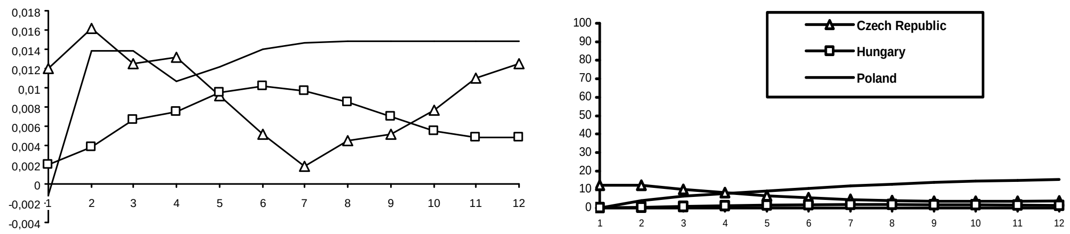
where dummy *crisis* is included to capture the effects of the 2008─2009 financial crisis[3].

For all countries, the positive short-run coefficient of the domestic money supply and the asymmetric coefficients of the domestic and foreign output are as expected. However, the long-run coefficients on M_t and Y_t are consistent with the monetary model only for Poland. The coefficients on M_t have perverse negative values for the Czech Republic and Hungary, while the negative coefficients on Y_t lack statistical significance. The long-run coefficients on I_t are in accordance with the Frankel model for the Czech Republic and Hungary, while the estimates for Poland support the Bilson model (equation (4)). The long-run effects of LIBOR are consistent with the Bilson model for Poland, being neutral in respect to the nominal exchange rate for the Czech Republic and Hungary. However, in the latter case the short-run coefficients on r^*_t are in favour of the Frankel model. The error-correction negative coefficient on ε_{t-1} is statistically significant for Poland, but in two other countries it is small and positive. A dummy for the 2008─2009 financial crisis suggests the long-run depreciation of the nominal exchange rate only in Poland.

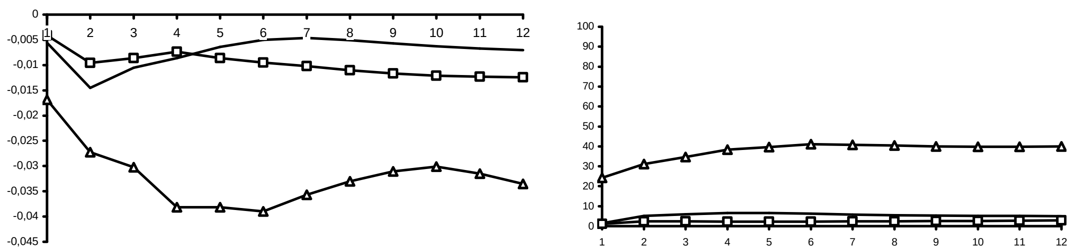
To check out the robustness of our results, the VAR/VEC methodology is implemented. The VAR/VEC system consists of six variables at a quarterly frequency, has two lags (a lag-

[3] The coefficient of determination R^2 ranges from 0.90 to 0.94 for the estimates of long-run coefficients, and from 0.20 to 0.43 for the estimates of short-run coefficients. The ADF test suggests stationarity of residuals at 5% statistical significance for all regression equations. Domestic and foreign output, as well as interest rates, are separately included, assuming different income and interest rate elasticities for industrial and CEE economies.

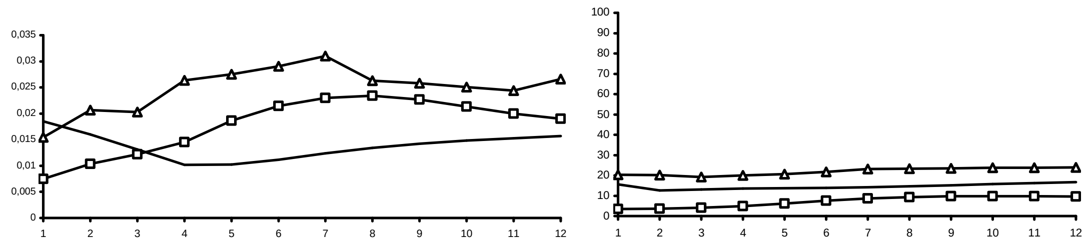
length is chosen according to the Akaike Information Criterion (AIC)), no constant or a time trend, dummy for control of the 2008–2009 financial crisis, and uses the logarithm for all variables (in levels). The impulse responses of the exchange rate to the shocks and the estimated variance decompositions are presented in Fig. 1 (the responses of other variables are not reported). In response to the money supply shock, the exchange rate depreciates in all countries (Fig. 1c). Innovations to the money supply have a positive impact on the exchange rate and the responses start to increase over the two-year horizon for the Czech Republic and Hungary. The proportion of variance explained by the money supply shock ranges from 10% (Hungary) to 15% (Poland) and 20% (the Czech Republic).



a) foreign output;



b) domestic output;



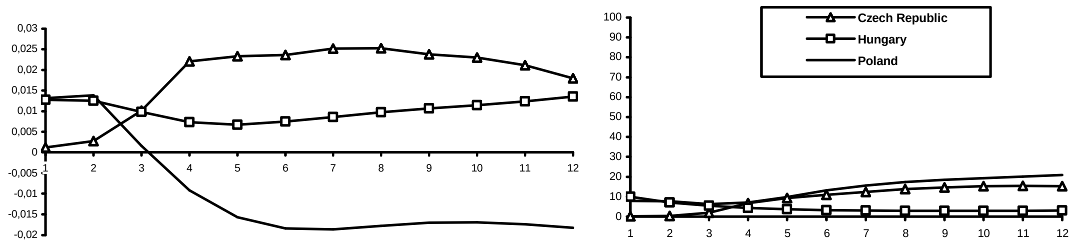
c) money supply;

Note: impulse responses and variance decomposition are on the left and right graphs, respectively.

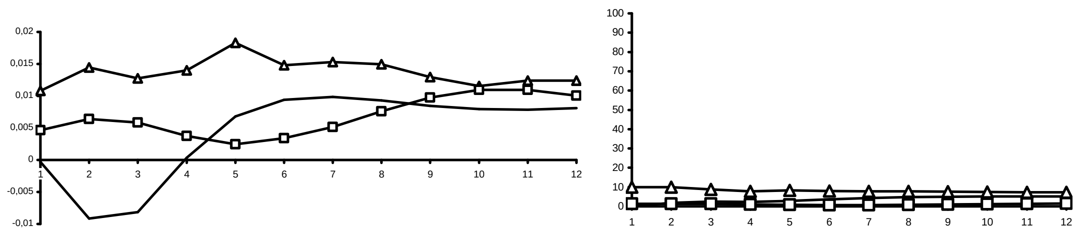
Fig. 1a. Effects of VAR shocks on the exchange rate.

As predicted by the monetary model, the domestic output is the factor behind an exchange rate appreciation (Fig. 1b), while the opposite relationship does hold for the foreign output (Fig. 1a).

Domestic output contributes as much as 40% to the variation in the exchange rate at horizons longer than 5 quarters in the Czech Republic, while accounting for less than 6% of the variation in the exchange rate in Hungary and Poland. Gradually, foreign output accounts for 15% of the exchange rate variation in Poland, while its impact fades away from 11% in the first period to less than 3% at the twelve-quarter horizon in the Czech Republic. The proportion of foreign output in the variation of exchange rate is marginal in Hungary. Except Poland, the nominal exchange rate depreciates in response to a positive LIBOR shock (Fig. 1d). The same outcome does hold for a domestic interest rate shock, though with an inverse relationship between I_t and E_t on impact for Poland (Fig. 1d). The LIBOR explains up to 18% and 20% of the variation of exchange rate in the Czech Republic and Poland, respectively, but it is of much less importance in Hungary. The combined effect of money supply and interest rates ranges at different time horizons from 30% to 46% in the Czech Republic, from 10% to 14% in Hungary, and from 23% to 42% in Poland. Domestic and foreign output account from 36% to 47% of the variance of the exchange rate in the Czech Republic, from 1% to 4% in Hungary, and from 1% to 20% in Poland.



a) foreign interest rate;



b) domestic interest rate;

Fig. 1b. Effects of VAR shocks on the exchange rate (continued).

5. Conclusions

In general, our empirical results are in favour of long-run and short-run relationships between the nominal exchange rate, money supply and macro-economy, which are consistent with the monetary model of exchange rate determination, as the increase in the money supply and negative growth differential are factors behind the nominal exchange rate depreciation (combined evidence squares best with the monetary model implications for Poland). A quite heterogeneous impact of the domestic and foreign interest rates could be explained by country-specific effects of capital flows.

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