# MONETARY EXCHANGE RATE MODEL FOR THE CENTRAL EUROPEAN COUNTRIES – EVIDENCE FROM A PANEL APPROACH

**Marek A. Dąbrowski - Monika Papież - Sławomir Śmiech**

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**Abstract**

This paper is focused on the issue of empirical congruity of the monetary exchange rate model for the Central European countries with relatively flexible rates. We use quarterly data from 2001 until 2012. The standard flexible-price monetary model is extended to allow for differences in the relative prices of nontradeables. This correction relaxes the assumption about the law of one price – it need not hold for nontradable goods and allows for the Balassa-Samuelson effect.

Empirical analysis is divided into two steps. First, using panel cointegration techniques we find evidence of cointegration between the nominal exchange rate and fundamentals implied by the monetary exchange rate model. The signs of estimated long-term coefficients for fundamentals are in line with those implied by the model. Second, the error correction model is used to explore causality. The null hypothesis that the exchange rate fails to Granger-cause fundamentals is rejected for the relative price of nontradables. In addition, the null hypothesis on a lack of reverse causality between relative money and exchange rate is rejected as well. These findings lend support to the monetary model for the Central European countries.

**Key words:** Monetary exchange rate model, Central European countries, Panel cointegration, Granger causality

**JEL Code:** C33, E40, F31

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## Introduction and brief literature review

On the one hand, exchange rates are similar to stock prices, as their fluctuations are mainly driven by changes in expectations. Thus, one should not be surprised at the weak relationship between the exchange rate and any set of macroeconomic fundamentals. On the other hand, Obstfeld and Rogoff (Obstfeld & Rogoff, 2001) aptly pointed that “the links between the exchange rate and the real economy are much more direct than for stock prices” in a sense that it is “the single most important relative price, one that potentially feeds back immediately into a large range of transactions”. They termed these two observations the exchange rate disconnect puzzle. The widely cited finding of Meese and Rogoff (Meese & Rogoff, 1983), that a random walk model predicts exchange rates as well as any structural model in the short-run, is a manifestation of this puzzle. Another one is the observation that the exchange rate volatility differs across exchange rate regimes even though volatility of macroeconomic fundamentals does not (Baxter & Stockman, 1989, Flood & Rose, 1999). Empirical evidence on monetary approach to the exchange rate was summarised by Shone (Shone, 2002) in dismal words: “regression results are generally poor, with exogenous variables being insignificant or even having the opposite sign from what theory predicts. Monetarist models of exchange rate determination are not alone on this”.

Such a dreary picture of empirical evidence, however, is not correct. Mark and Sul (Mark & Sul, 2001), using a quarterly panel of 19 OECD countries extending from 1973Q1 to 1997Q1, found out that the nominal exchange rate was cointegrated with monetary fundamentals. Their results suggest that cross-section dependence plays an important role in the relation between the exchange rate and its fundamentals. This observation was exploited by Beckmann et al. (Beckmann, Belke & Dobnik, 2012) who first isolated common factors from idiosyncratic components and then presented evidence that there exists an international long-run relationship between the common factors of exchange rates and fundamentals[1].

The similar observation was made by Engel et al. (Engel, Mark & West, 2008) with respect to out of sample forecasts of exchange rates. Using panel techniques they were able to show that monetary models do help to forecast changes in exchange rates, i.e. “generally produces better forecast than the random walk”[2]. Moreover, they presented empirical evidence both for the hypothesis that exchange rates are useful in forecasting fundamentals and against the null that the fundamentals do not Granger-cause changes in exchange rates[3].

In a study on exchange rates in EU-10 countries[4] and Turkey Uz and Ketenci (Uz & Ketenci, 2008) used panel cointegration tests to identify the long-term relation between

[1] Some other empirical studies are shortly reviewed by Beckmann et al. (Beckmann, Belke & Dobnik, 2012). The main findings are twofold. First, exchange rates for economically advanced countries are cointegrated with monetary fundamentals and estimated coefficients are in line with the model. Second, panel cointegration techniques are more powerful if the mean reversion is weak and time series are short.

[2] They, however, admitted that they “do not have a precise economic or econometric story as to why panel estimates do so much better than country-by-country estimates”. Their conjecture was that the efficiency of panel estimation is key.

[3] The former hypothesis relies on the assumption that the unobserved fundamentals are not the key drivers of exchange rates.

[4] New member states that joined the European Union on 1 May 2004.

nominal exchange rate and fundamentals suggested by the monetary model. They found out that there is "strong evidence of the cointegration" and in addition all fundamentals were correctly signed (except price index in some tests) and significant. It seems, however, that their approach is deficient in three respects. First, they introduced interest rate differential and price differential to their estimated equation in an *ad hoc* manner, i.e. their modification was not founded on the theory. Second, the span of time included in the sample, i.e. 1993Q1 2005Q4, covered periods of relatively high inflation in the EU-10 countries. Thus, empirical results could be mainly driven by inflation differential[5]. Third, their sample covered countries with fixed exchange rate regimes, e.g. the Baltic States or Cyprus. In such a regime the exchange rate is determined by administrative decision and there is no need to explain its fluctuations[6].

The paper is organised as follows. In the next section we provide a concise description of the model adopted. Data and empirical results are presented in section 3. The last section concludes.

Our objective is to test whether monetary approach to exchange rate is a plausible one for Central European countries. Our sample covers a time span extending from 2001Q4 to 2012Q4 which has two advantages. First it excludes the early years of economic transformation in Central European countries, thus our results are not driven by high inflation differentials that were characteristic for early phases of transition. Second, it covers both the years of fast economic expansion and the recent global financial crisis (and other crisis episodes as well), so it is balanced in terms of the business cycles. Moreover, we are careful in adding relative prices of nontradeables to the cointegrating relation. In fact, we believe that such a correction is key for the type of economies we study: all of them are classified as emerging markets and grow at a relatively fast pace (in comparison to the euro area).

## 2 Monetary model of the exchange rate

Equilibria in domestic and foreign money markets, with foreign variables denoted with asterisks, are given by:

$$m_t = p_t + \kappa y_t - \lambda i_t , \qquad (1)$$

[5] See for instance their table 2 in which only prices are significant in all alternative methods used to estimate cointegration coefficients. The classic fundamentals (i.e. monetary aggregate and output) in turn are insignificant according to four methods out of five. Moreover, as pointed by Flood and Rose (Flood & Rose, 1999) monetary approach to the exchange rate "performs fairly well when inflation is high".

[6] The interesting issue is whether the parity was at the level corresponding to the equilibrium exchange rate. This, however, is a question of a different type.

$$m_t^* = p_t^* + \kappa^* y_t^* - \lambda^* i_t^*, \qquad (2)$$

where $p$ stands for the price level, $m$ is the level of money supply, $y$ denotes an income, and $i$ is the interest rate. All variables except interest rates are in logarithms. Parameters $\kappa$ and $\lambda$ are income elasticity and interest rate semi-elasticity of money demand. Two core elements on which the model rests are that the uncovered interest rate parity and the absolute purchasing power parity hold continuously (Sarno & Taylor, 2002). Though we stick to the former, the latter is relaxed. We narrow it down to the price of tradeables, $p_t^T$, so:

$$s_t = -(p_t^T - p_t^{T*}), \qquad (3)$$

where $s$ is expressed as the foreign currency price of domestic currency. Assuming that the general price level encompasses both tradeables and nontradables, i.e.

$$p_t = (1-\alpha) p_t^T + \alpha p_t^N, \qquad (4)$$

and taking into account equations (1)-(3) one can arrive at:

$$s_t = \kappa(y_t - y_t^*) - (m_t - m_t^*) - \lambda(i_t - i_t^*) + \alpha[(p_t^N - p_t^T) - (p_t^{N*} - p_t^{T*})], \qquad (5)$$

where for simplicity domestic and foreign parameters are assumed to be the same. This approach allows for the Balassa-Samuelson effect, i.e. if the domestic relative price of nontradeables increases faster than the foreign relative price then currency appreciates.

According to the uncovered interest rate parity the interest rate differential ($i_t - i_t^*$) is equal to the expected depreciation of the domestic currency ($-\Delta s_{t+1}^e$). Using this in (5) the rational expectations solution is[7]:

$$s_t = \frac{1}{1+\lambda} \sum_{j=0}^{\infty} \left( \frac{\lambda}{1+\lambda} \right)^j \mathrm{E}_t v_{t+j}, \qquad (6)$$

where fundamentals are encapsulated in :

$$v_{t+j} \equiv \kappa(y_t - y_t^*) - (m_t - m_t^*) + \alpha[(p_t^N - p_t^T) - (p_t^{N*} - p_t^{T*})].$$

[7] Generally equation (5) augmented with the UIP condition has multiple rational expectations solutions. For details on rational bubble terms see Blanchard and Fischer (Blanchard & Fischer, 1989).

Sarno and Taylor (Sarno & Taylor, 2002) pointed out that after subtracting $v_t$ both sides of (6) and some manipulation it is possible to show that:

$$s_t - v_t = \sum_{j=1}^{\infty}\left(\frac{\lambda}{1+\lambda}\right)^j \mathrm{E}_t \Delta v_{t+j} \,. \qquad (7)$$

If fundamentals are non-stationary processes $I(1)$ then the right-hand side of (7) has to be integrated of order 0. Therefore, the left-hand side is stationary as well and the cointegration between the exchange rate (assuming it is non-stationary) and its fundamentals exists. The cointegrating vector takes the following form:

$$s_t - \kappa(y_t - y_t^*) + (m_t - m_t^*) - \alpha[(p_t^N - p_t^T) - (p_t^{N*} - p_t^{T*})] \,. \qquad (8)$$

## 3 Data and empirical results

We use the (balanced) panel quarterly data extending from the last quarter 2001 to the last quarter 2012 for eight countries: the Czech Republic, Hungary, Moldova, Poland, Romania, Serbia, Turkey, and Ukraine. Nominal exchange rates are end-of-quarter observations from the IMF (the CEIC database was used). They are expressed in euros per national currency, so an increase means an appreciation of national currency. Our measure of money supply is an aggregate M2 from the IMF/CEIC. Real GDP are used to measure output. Data are from the national sources for Moldova and Serbia, form the IMF/CEIC for Turkey, and Ukraine and form the Eurostat for other countries. General prices are measured with GDP deflators that were collected from the same sources as GDPs. Producer price indices from the IMF/CEIC are used for the price of tradeables. All variables are expressed as indices with their average values in 2005 equal 100, seasonally adjusted and specified in natural logarithms.

The empirical part was carried out using a three-step procedure. At first a non-stationarity of individual variables was assessed. Four panel unit root test were carried out: LLC test, IPS test, F-ADF test, and Hadri test (Levin, Lin & Chu, 2002, Im, Pesaran & Shin, 2003). Though the results are not perfectly clear-cut, the evidence tilts the balance in favour of non-stationarity of exchange rates and their fundamentals[8].

Then we test whether there is a long-run cointegrating relationship between the variables. At this step the set of Pedroni panel cointegration tests (Pedroni, 2000) were carried out. In the final step (after confirming existence of cointegration relationship) a panel vector

[8] Detailed test results are available from the authors upon request.

error correction model was estimated so short and long-run causal relations between the variables were investigated.

Table 1 presents both the within and between dimension panel cointegration test statistics. The majority of tests we used reject the null of no cointegrating relationship.

**Tab. 1: Pedroni panel cointegration tests** (null hypothesis: no cointegration)

| **Within dimension test statistics** | | | **Between dimension test statistics** | |
|---|---|---|---|---|
| Panel v-Statistic | 1.5925* | | Group rho-Statistic | -1.0910 |
| Panel rho-Statistic | -1.4138* | | Group PP-Statistic | -3.0576*** |
| Panel PP-Statistic | -2.5218*** | | Group ADF-Statistic | -3.1040*** |
| Panel ADF-Statistic | -2.6361*** | | | |

Notes: ***, **, * indicate statistical significance at 1, 5 and 10 percent level of significance, respectively

Since evidence is found that there is a long-run relationship between the variables, the cointegrating vector was estimated. The fully modified ordinary least squares (FMOLS) methods developed by Pedroni (2000, 2001) were used. The FMOLS estimates are as follows:

$$s = 4.586 - \underset{(-4.717)^{***}}{0.300}\,\hat{m} + \underset{(2.649)^{***}}{0.502}\,\hat{y} + \underset{(6.776)^{***}}{1.349}\left(\hat{p} - \hat{p}^{T}\right), \qquad (9)$$

where all variables with hats are expressed as differences between the relevant home and foreign levels, e.g. $\hat{m} \equiv m - m^{*}$, and the numbers in parentheses denote the values of t-statistics.

All variables are in natural logarithms, so the estimated coefficients can be interpreted as long-run elasticities. The elasticities of both standard monetary fundamentals, i.e. money supply and income, turned out to be highly significant. Moreover, the signs of estimated coefficients for fundamentals are in line with those implied by the model. More precisely, our results show that a one per cent increases in the money supply relative to the euro area triggers depreciation against the euro by 0.3 per cent in the long-run. A rise in the relative output has a stronger absolute long-term impact on the exchange rate: it brings about an appreciation by 0.5 per cent. The finding that the income elasticity is greater than the money supply elasticity is in accordance with results of Beckmann et al. (Beckmann, Belke & Dobnik, 2012)[9].

Our results lend support to the modification of the monetary approach we suggest for the emerging market economies: the long-run elasticity of price level relative to the euro area

[9] See also other studies quated by these Authors.

was highly significant and correctly signed. Thus, a one per cent increase in the relative price of nontradeables results in a nominal appreciation against the euro of 1.3 per cent. This can be interpreted as a manifestation of the Balassa-Samuelson effect. The positive coefficient means that a certain fraction of real appreciation (caused by a faster growth of the domestic relative price of nontradeables in comparison to the foreign relative price) is channelled via nominal appreciation.

The final step is to estimate a panel vector error correction model in order to infer the Granger causal relations between the variables. The Granger causality test is based on the model with a dynamic error correction term. Table 2 presents *p*-value for Granger causality test.

**Tab. 2: Panel Granger causality test results**

| Cause variable | Effect variable | F-statistic | *p*-value |
|---|---|---|---|
| $\Delta s$ | $\Delta\hat{m}$ | 2.101 | 0.1482 |
| $\Delta s$ | $\Delta\hat{y}$ | 2.084 | 0.1498 |
| $\Delta s$ | $\Delta\left(\hat{p}-\hat{p}^T\right)$ | 10.246 | 0.0015 |
| $\Delta\hat{m}$ | $\Delta s$ | 3.332 | 0.0688 |
| $\Delta\hat{y}$ | $\Delta s$ | 0.004 | 0.9506 |
| $\Delta\left(\hat{p}-\hat{p}^T\right)$ | $\Delta s$ | 0.063 | 0.8014 |

The rational expectations solution of the monetary model (see equation (6)) implies that the exchange rate is mainly determined by expectations of future fundamentals. In other words, the exchange rate responds to news about fundamentals. Thus, as explained by Engel et al. (Engel, Mark & West, 2008), changes in exchange rates might be useful in predicting the fundamentals (under the assumption that the unobservable fundamentals are not the primary drivers of exchange rates).

The null hypothesis that the exchange rate fails to Granger-cause fundamentals is strongly rejected for the relative price of nontradables. The *p*-values for the relative money supply and relative output are above the significance levels conventionally adopted, which means that the exchange rates Granger-cause neither relative money supply nor relative income.

There is also some evidence of reverse causality, the one running from the relative money supply to exchange rate. This is a bit surprising since as pointed by Engel et al. (Engel, Mark & West, 2008) “the models have little power to forecast exchange rate changes"[10]. This seems to lend support to the view that the exchange rate is under the relatively strong influence of monetary policy in countries analysed. We conjecture that the exchange rate may possibly be considered an implicit target variable (at least to a certain extent) for the central banks. This requires further research.

## Conclusion

The objective of this study was to test whether monetary model of exchange rate determination provides a reasonable framework for the exchange rate movements in the Central European countries and Turkey. We found evidence of cointegration between exchange rates and standard monetary fundamentals plus relative prices of nontradeables. The latter seem to be not only statistically significant but also important from an economic point of view. After all countries we include into the sample are emerging markets with growth rates higher relative to those observed in the euro area and therefore subjects to the Balassa-Samuelson effect. We were able to show that a rise in a relative price of nontradeables results in a long-term nominal appreciation of national currency[11]. Granger-causality tests we used together with the evidence on cointegration lend support to the monetary model for the Central European countries.

Taking into account the generally observed relative efficiency of panel cointegration analysis and our empirical results we share the view of Beckmann et al. (Beckmann, Belke & Dobnik, 2012) that exchange rates share common stochastic trends across countries. These can stem not only from an international business cycle but also from the tight economic and financial links of all the Central European countries and Turkey with the euro area. Applied to these countries their methodology (based on the direct analysis of cross-section dependence) could turn out to be fruitful. This is the issue we plan to investigate in further research.

[10] The reverse causality was also found by Engel et al. (Engel, Mark & West, 2008) for economically advanced OECD countries.

[11] This of course is not a necessary consequence of the BS effect which implies a real appreciation in fast-growing economies. Our finding is not so much about the BS effect itself but rather about the structure of real appreciation.

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**Contact**

Marek A. Dąbrowski
Cracow University of Economics, Department of Macroeconomics,
Rakowicka 27 st., 31-510 Cracow, Poland
marek.dabrowski@uek.krakow.pl

Monika Papież
Cracow University of Economics, Department of Statistics,
Rakowicka 27 st., 31-510 Cracow, Poland
papiezm@uek.krakow.pl

Sławomir Śmiech
Cracow University of Economics, Department of Statistics,
Rakowicka 27 st., 31-510 Cracow, Poland
smiechs@uek.krakow.pl