The Behaviour of Exchange Rates in the Central European Countries and Credit Default Risk Premiums

Piotr Kębłowski[*]

Abstract

We test whether the floating exchange rates of the EU New Member States against the euro are determined jointly within the panel VEC framework. We find that the exchange rates of the Czech koruna, the Polish zloty and the Hungarian forint follow the same long-run relationship, in which the real exchange rates are explained by the real interest rates parities and the spreads of the credit default risk premiums. In case of the Romanian leu, the common relationship is rejected, which is likely due to differences in the economic setting. The results confirm that the currency markets of these three countries are closely related, since the appreciation/depreciation of one currency leads to similar movements in the other currencies of the NMS. The estimated misalignments exhibit some common patterns in terms of time spans and percentage values of under/overvaluation.

[*]University of Lodz, e-mail: emfpiok@uni.lodz.pl

Piotr Kębłowski

1 Introduction

The behaviour of the floating exchange rates of the EU New Member States (NMS) during the turbulent period of the sub-prime crisis in the United States and the European sovereign debt crisis reveals that the level of credit default risk assigned to sovereign issuers by financial investors is one of the non-negligible drivers of the exchange rates. However, for the most part credit default risk is omitted in empirical models of exchange rates, or it is approximated by measures, which neglects changes in investors' risk aversion.

The usually considered risk premium proxies are mainly liquidity indices (e.g. share of debt in GDP, M2 vs. official reserves, see e.g. Kelm 2010, 2011), an equity market variable or a term structure, the last one being affected by liquidity premium. However, there are a number of risk premium measures, which are based on market quotes. On the one hand, the risk premium can be measured by credit default swaps (CDS) quoted at OTC markets (see Duffie 1999, Adler and Song 2010, Blancheton, Bordes, Maveyraud, Rous 2012 and Kębłowski and Welfe 2012), or alternatively by collateralized debt obligations (CDO) in case of emerging market sovereigns (see Duffie and Gârleanu 2001). On the other hand, sovereign credit ratings and historical default rates can be used for calculation of expected loss (see e.g. Remolona, Scatigna, Wu 2008).

In our study we use sovereign CDS premiums, since they have certain advantages as a market measure of a credit default risk. First of all, assuming that there is no counterparty risk or the risks are mutually independent, the CDS premium represents the cost of buying insurance against a default of issuer of the underlying asset. Secondly, the CDS premium seems to be helpful in empirical verification of the uncovered interest parity, since it is equal to the spread over the risk-free rate of a floating rate bond (see Duffie 1999). Finally, sovereign CDS premiums are widely used by financial investors on currency markets.

The sovereign CDS premiums as well as the floating exchange rates of the NMS exhibit some easily visible comovements between countries. Therefore, the purpose of this paper is to investigate whether the CDS premiums and the exchange rates of the NMS are cointegrated. To do so, we use the medium-run approach based on the CHEER hypothesis augmented by the sovereign credit default risk perceived by financial investors. Within the panel VEC framework we verify, whether there is a common long-run relationship explaining the behavior of the exchange rates in the NMS and if exchange rates misalignments affect each other. We also compare the evolution of the exchange rates misalignments.

We analyse the panel of four floating exchange rates against the euro, for the Czech Republic, Hungary, Poland and Romania, since the other NMS has either joined the EMU or their currencies are pegged to the euro. The floating exchange rates of the Czech koruna, the Hungarian forint, the Polish zloty and the Romanian leu were officially brought in on 26 May 1997, 26 February 2008, 12 April 2000 and 31 December 1997 respectively. According to the forint, the crawling peg mechanism

was introduced in March 1995 with a narrow fluctuation band, which was widened to $\pm 15\%$ in May 2001, and the crawling peg was repealed on 1 October 2001. Therefore, it can be assumed at best that joint time span of the (nearly) floating exchange rate for these countries starts in the year 2001.

The outline of the remainder of the paper is as follows. In Section 2 we discuss the economic background of the exchange rates determination in the NMS. Section 3 presents empirical results based on the panel VEC model. Exchange rates misalignments are analysed in Section 4. Section 5 summarizes the findings and concludes.

2   Credit default risk premiums and the exchange rates determination

Exchange rate models are mostly related to the balance of payments condition. Therefore, the determinants of both the capital account as well as the current account are considered. However, some approaches highlight real determinants of the current account, whereas others state that the balance of payment condition is satisfied mainly due to adjustments of the capital account. The latter approach decidedly seems to be particularly well suited for the case of small open economies investigated in this study.

According to the determinants of both sides of the balance of payments, it is natural to combine the purchasing power parity and the net interest rate differential in order to explain the exchange rate's behaviour, which is known as a CHEER approach (see MacDonald 2000). Juselius and MacDonald (2003, 2004) extended this approach by joint inclusion of short- and long-term interest rates and inflation rates, which influence spot exchange rates in the long-run. Therefore, the simple purchasing power parity is replaced by the hypothesis assuming that spot exchange rates are affected by price misalignments present in both difference between inflation rates and real exchange rates:

$$\Delta_m s^e_{i,t+m} = \phi_1 \left(\Delta_m p^e_{i,t+m} - \Delta_m p e^e_{t+m} \right) + \phi_2 q^e_{i,t+m} + \varepsilon_{it}, \tag{1}$$

where $q_{it} = p_{it} - pe_t - s_{it}$, p_{it} is the log of consumer price index for the NMS ($i = 1, 2, 3, 4$, for the Czech Republic, 2 - Hungary, 3 - Poland, 4 - Romania), pe_t denotes the log of consumer price index in the euro area, and s_{it} stand for the log of spot exchange rate against the euro, Δ_m is the difference operator – $\Delta_m x_{it} = x_{it} - x_{i,t-m}$ and m denotes maturity. An application of the aforementioned approach for the exchange rate of the zloty can be found in Kębłowski and Welfe (2010).

The uncovered interest rate parity underpinning the CHEER approach assumes that the floating rate bonds yields are risk-free interest rates. Since the sovereign credit default risks vary significantly, the basic uncovered interest rate parity needs to be

Piotr Kębłowski

augmented by the credit default premium (see Kębłowski and Welfe 2012):

$$\frac{\Delta_m s^e_{i,t+m}}{m} = \left(i^m_{i,t} - cds_{i,t}\right) - \left(ie^m_t - cdse_t\right), \tag{2}$$

where i_{it}, ie_t, cds_{it} and $cdse_t$ denote the long-term interest rate yield and the logs of the CDS contracts prices in the NMS and in the euro area respectively. Duffie (1999) showed that the credit default swap is equal to the spread over the risk-free rate of par risky floating rate bond. Thus the CDS premium allows recovering the uncovered interest parity in case of risky sovereign interest rate yields.

The hypotheses (1) and (2) enable to write the CHEER approach augmented by the credit default risk premium:

$$q_{it} = \beta_{1i}\left(i_{it} - ie_t\right) + \beta_{2i}\left(\Delta p_{it} - \Delta pe_t\right) + \beta_{3i}\left(cds_{it} - cdse_t\right) + \varepsilon_{it}. \tag{3}$$

Therefore, the real exchange rates movements are explained by differentials of interest rates, inflation rates and sovereign credit default swaps indices. We will verify whether the exchange rates in the four aforementioned countries share common mechanism determining their evolution.

The quotations of the sovereign CDS's started in: May 2006 for the Czech Republic, March 2002 for Hungary, November 2000 for Poland and October 2002 for Romania. In turn, the first CDS contracts for the euro area bonds, represented by the German bonds, were priced in March 2003. The series have been backdated to January 2001, as they are significantly correlated with the equity market variable. The series are illustrated in Figure 1.
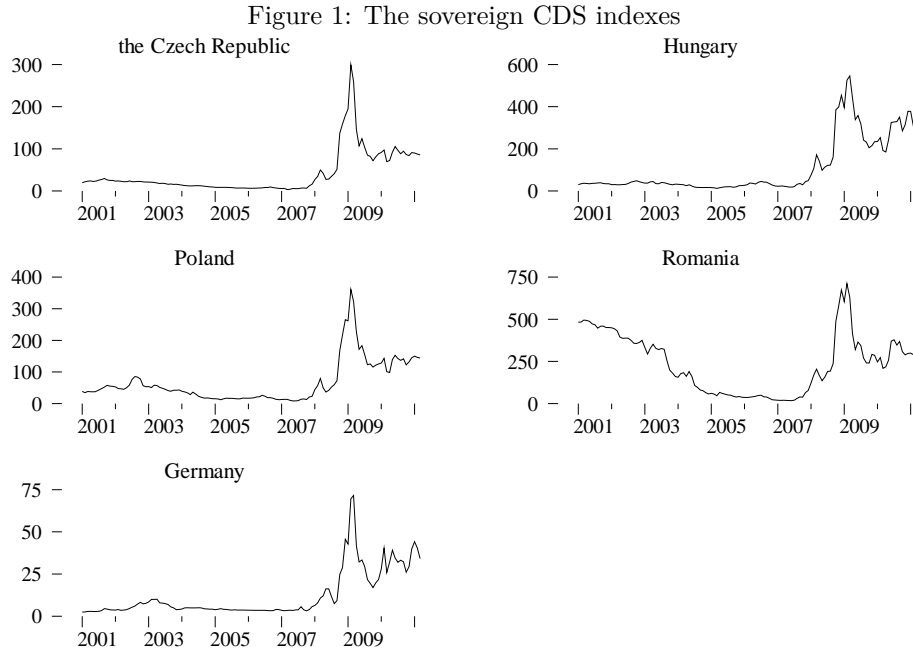
The first half of the sample reveals that up to the year 2005 the new member states bonds are perceived as less risky year by year on average, in comparison with the euro area. However, the most noticeable are huge increments of the CDS premiums for all countries at the end of the year 2008, when the financial markets turmoil begins. In fact, the first evidence of the increasing risk are visible at the beginning of the year 2008 and from that date on the new member states bonds are perceived as more risky then previously. Interestingly, the NMS and the euro area share the same shock at the end of 2008 year to a large extent (see Figure A.1).

3   The panel VEC model

According to (3) the analysis is based on four variables ($P = 4$):

$$y_{it} = \left[q_{it}\ (i_{it} - ie_t)\ (\Delta p_{it} - \Delta pe_t)\ (cds_{it} - cdse_t)\right]'$$

in four cross-sections ($I = 4$) for the full model. Since Romania seems to be a heterogeneous entity in the panel (lack of long-term inflation targeting for example), the small model with 3 countries is also considered. The prices are deseasonalized using the Census X-11 method. The long-term interest rates are monthly average

Figure 1: The sovereign CDS indexes

yields on ten-year floating-rate bonds. The monthly data cover the period January 2001 - April 2011 and come from Eurostat. All series are demeaned. The panel stationarity test as well as the unit root tests indicate that the data seem to be generated by the process integrated of order one (see Figure 1 and Table 1). An analysis of double unit roots in the DGP of prices can be found in Majsterek and Welfe 2012.

Let $y_t = [y'_{1t}\, y'_{2t}\, \ldots\, y'_{It}]'$ denotes an IP-dimensional vector of variables in the cross-sections in time t. The panel vector error correction model can be written as:

$$\Delta y_t = \Pi y_{t-1} + \sum_{k=1}^{K} \Gamma_k \Delta y_{t-k} + \varepsilon_t, \tag{4}$$

where Π and Γ_k $(k = 1, 2, \ldots, K)$ are $IP \times IP$ matrices of parameters, $\varepsilon_t = [\varepsilon'_{1t}\, \varepsilon'_{2t} \ldots \varepsilon'_{It}]'$ and $\varepsilon_t \sim N(0, \Omega)$. The lag order is set to $K = 3$, according to indications of the information criteria and the LM test of autocorrelation for the VEC models estimated for individual countries in a preliminary analysis. If the variables are cointegrated then the matrix Π has a reduced rank and can be decomposed as $\Pi = AB'$, where A and B are $IP \times \sum R_i$ matrices of parameters and R_i denotes the cointegration rank in the cross-section i.

Piotr Kębłowski

Table 1: Inference on integration order

	Δq_{it}	$\Delta i_{it} - \Delta ie_t$	$\Delta^2 p_{it} - \Delta^2 pe_t$	$\Delta cds_{it} - \Delta cdse_t$
Z_μ[a]	-0.90[d]	-2.02	0.49	-0.87
	(0.81)	(0.98)	(0.31)	(0.81)
Z_{tbar}[b]	-7.90	-16.52	-8.11	-10.98
	(0.00)	(0.00)	(0.00)	(0.00)
$Z_{\underset{tbar}{\sim}}$[c]	-6.87	-12.05	-7.03	-9.00
	(0.00)	(0.00)	(0.00)	(0.00)
	q_{it}	$i_{it} - ie_t$	$\Delta p_{it} - \Delta pe_t$	$cds_{it} - cdse_t$
Z_μ	72.83	4.05	7.98	17.34
	(0.00)	(0.00)	(0.00)	(0.00)
Z_{tbar}	-0.76	-0.98	-1.14	-1.53
	(0.23)	(0.16)	(0.13)	(0.06)
$Z_{\underset{tbar}{\sim}}$	-0.69	-0.92	-1.07	-1.44
	(0.24)	(0.18)	(0.14)	(0.07)

[a] *Note*: the statistic of Hadri (2000) stationarity test, $Z_\mu \xrightarrow{d} N(0,1)$ as $T \to \infty$, $I \to \infty$ and $I/T \to 0$.
[b,c] *Note*: the statistic of Im, Pesaran, Shin (2003) unit root test, $Z_{tbar} \xrightarrow{d} N(0,1)$ as $T \to \infty$, $I \to \infty$ and $I/T \to const$.
[d] *Note*: p-values in brackets.

The general model (4) allows for simultaneity to a large extent, since the loadings matrix A, Γ_k and Ω are unrestricted:

$$A = \begin{bmatrix} A_{11} & A_{12} & \cdots & A_{1I} \\ A_{21} & A_{22} & \cdots & A_{2I} \\ \vdots & \vdots & \ddots & \vdots \\ A_{I1} & A_{I2} & \cdots & A_{II} \end{bmatrix}, \tag{5}$$

whereas the matrix of cointegrating vectors B has the following structure:

$$A = \begin{bmatrix} B_{11} & 0 & \cdots & 0 \\ 0 & B_{22} & \cdots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & \cdots & B_{II} \end{bmatrix}. \tag{6}$$

Therefore, heterogeneous long-run relationships are allowed within cross-sections and each cointegration vector can affect any variable in the system (see Larsson and Lyhagen 2007, also Groen and Kleibergen 2003, and Larsson, Lyhagen, Loethgren 2001, for more restricted models, and a review in Kębłowski (2009)). Note, however, that even though the decomposition of the matrix Π into matrices A and B allows a non-diagonal structure of the matrix Π, the variables from different cross-sections are not permitted to cointegrate, which seems to be reasonable assumption here. If the assumption of a block-diagonal structure of the matrix B cannot be maintained, e.g. in case of the simple purchasing power parity testing, then specific solutions can

be used (see Jacobson, Lyhagen, Larsson, Nessen 2002, Banerjee, Marcellino, Osbat 2004, 2005).

The model (4) with matrices A and B as in (5) and (6) enables to test the hypothesis of common cointegration rank, which implies that the number of long-run relationships in each cross-section is equal – $\forall_i R_i = R$. For the model (4) with A, Γ_k and Ω unrestricted and a block-diagonal matrix B, Larsson and Lyhagen (2007) proved that the LR statistic for the common cointegration rank hypothesis, the panel counterpart of the trace test, $\mathcal{H}_0$: $rk(\Pi_{ii}) = R_i \leq R$ for $i = 1, 2, \ldots, I$ vs. $\mathcal{H}_1$: $rk(\Pi_{ii}) = P$ for $i = 1, 2, \ldots, I$, converges to the convolution of the Dickey-Fuller type distribution and the χ^2 distribution:

$$LR_{IT} \xrightarrow{d} \chi^2_{I(I-1)R(P-R)} + tr\left(\int_0^1 dWW'\left[\int_0^1 WW'du\right]^{-1}\int_0^1 WdW'\right) \quad (7)$$
$$\text{as } T \longrightarrow \infty, \text{for fixed } I,$$

where W denotes the $I(P-R)$-dimensional Brownian motion and the symbol $\xrightarrow{d}$ stands for convergence in distribution.

The results of inference on the common cointegration rank are presented in Table 2. First of all, due to huge size distortion the standard LR statistic rejects all consecutive null hypotheses in both cases, i.e. with and without Romania in the panel, which leads to the conclusion that the variables are jointly stationary, contrary to the results in Table 1. However the Bartlett corrected test clearly indicates that there is one cointegrating vector in the system. Therefore, we assume that $R = 1$.

Table 2: Inference on common cointegration rank

$\mathcal{H}_0$	4 countries[a]			3 countries[b]		
	LR	LR^{BC}[c]	crit. val.	LR	LR^{BC}[c]	crit. val.
R=0	1829.85	617.69	530.75	1105.07	414.25	307.15
R=1	511.39	303.16	345.61	263.93	181.17	197.46
R=2	310.76	162.16	194.89	143.96	88.63	111.05
R=3	135.57	61.11	82.56	75.73	40.22	47.18

[a] *Note*: the Czech Republic, Hungary, Poland, Romania.
[b] *Note*: the Czech Republic, Hungary, Poland.
[c] *Note*: Bartlett corrected LR statistic.

Table 3. shows the maximum likelihood estimates of cointegrating vectors for $R = 1$. The results support the relationship (3), since the coefficients have sound values and are similar for each country, with the exception of Romania. The results for the Czech Republic, for example, shows that an 10% increase of the domestic sovereign credit default swaps index *ceteris paribus* leads to a depreciation of the Czech koruna of about 1%, which seems to be in line with the historical observations. Similarly, the obtained semi-elasticities for interest rates and inflation rates seem to be proper for monthly data.

Table 3: The estimates of the cointegrating vectors

	q_{it}	$i_{it} - ie_t$	$\Delta p_{it} - \Delta pe_t$	$cds_{it} - cdse_t$
the Czech Republic	1	-21.40 [a] (20.5)	32.84 (21.7)	0.102 (8.8)
Hungary	1	-38.20 (5.6)	25.92 (6.3)	0.061 (4.6)
Poland	1	-43.77 (15.2)	25.48 (9.7)	0.087 (3.3)
Romania	1	-26.43 (5.0)	11.22 (2.9)	0.194 (10.3)

[a] *Note*: absolute values of t-ratios in brackets, standard errors bootstrapped with 1000 replications.

Acceptance of the common cointegration rank and similarity of the estimated relationships for the NMS enables to tests whether the countries share the same cointegration space $\mathcal{H}_0$: $B_{11} = B_{22} = \cdots = B_{II}$:

$$
A = \begin{bmatrix}
B_{11} & 0 & \cdots & 0 \\
0 & B_{11} & \cdots & 0 \\
\vdots & \vdots & \ddots & \vdots \\
0 & 0 & \cdots & B_{11}
\end{bmatrix}, \tag{8}
$$

vs. $\mathcal{H}_1$: $B_{ii} \neq B_{jj}$ for some i, j. The LR statistic for the common cointegration space hypothesis converges asymptotically to the χ^2 distribution (see Larsson and Lyhagen 2007):

$$
LR_{IT} \xrightarrow{d} \chi^2_{I(I-1)R(P-R)} \qquad \text{as } T \longrightarrow \infty, \text{ for fixed } I. \tag{9}
$$

The LR statistics of the common cointegration space for $R = 1$ are $\chi^2_9 = 63.47(0.00)$ for the panel of all four countries, and $\chi^2_9 = 9.34(0.16)$ for the case of three countries (p-values in brackets). Hence, the test clearly rejects the null hypothesis of common cointegration space for the model with four countries (including Romania), whereas it is accepted for the smaller one, covering the Czech Republic, Hungary and Poland. The point estimate of the common cointegrating vector is given as follows:

$$
q_{it} = \underset{(5.5)}{28.15} \left(i_{it} - ie_t\right) + \underset{(4.2)}{30.54} \left(\Delta p_{it} - \Delta pe_t\right) + \underset{(3.0)}{0.098} \left(cds_{it} - cdse_t\right) + ec_{it}, \tag{10}
$$

where ec_{it} denotes weakly-stationary error correction term (absolute values of t-ratios in brackets, standard errors bootstrapped with 1000 replications).

The long-run semi-elasticity of the exchange rate with respect to the interest rates differential is considerably lower then 120 predicted by the uncovered interest rate parity and it is comparable to the estimated semi-elasticities provided by conventional models, such as the monetary approach. It is noteworthy that the semi-elasticities of interest rates and inflation rates are almost the same with respect to their modulus. Therefore, the relationship (10) indicate that the real exchange rates of domestic

currencies versus euro in the Czech Republic, Hungary and Poland are driven by the real interest rates parities and the spreads of the risk premiums, with the euro area as the point of reference. Therefore, an 10% increase of the sovereign CDS index *ceteris paribus* leads to nearly 1% depreciation of domestic currency, whereas one percentage point rise in domestic annual long-term interest rate (or one percentage point fall in annual inflation rate) *ceteris paribus* leads to about 2,5% appreciation of the NMS currencies.

The loadings matrix in Table 4. shows the influence of deviations from the steady-states implied by (10) on the system's variable. It is interesting to note that the disequilibriums in the other countries usually push the real exchange rates outside their steady-states. Hence, the currency market of these three countries are closely related in such a manner that appreciation/depreciation of one currency leads to similar movements in the other currencies of the NMS in the short-run.

Table 4: The estimate of the loadings matrix

	$\widehat{\alpha}_1$	$\widehat{\alpha}_2$	$\widehat{\alpha}_3$
Δq_{1t}	-0.13 (3.7)	0.03 (2.0)	0.01 (2.5)
$\Delta i_{1t} - \Delta ie_t$	-0.10 (2.6)	0.00 (1.1)	-0.02 (0.4)
$\Delta^2 p_{1t} - \Delta^2 pe_t$	0.01 (1.8)	0.00 (1.0)	-0.01 (0.9)
$\Delta cds_{1t} - \Delta cdse_t$	-0.22 (0.1)	-0.06 (0.5)	-0.03 (0.2)
Δq_{2t}	0.07 (3.9)	-0.11 (3.1)	-0.01 (1.7)
$\Delta i_{2t} - \Delta ie_t$	-0.01 (0.2)	-0.09 (2.5)	-0.01 (0.4)
$\Delta^2 p_{2t} - \Delta^2 pe_t$	0.01 (0.6)	0.00 (0.9)	0.00 (1.2)
$\Delta cds_{2t} - \Delta cdse_t$	0.00 (0.3)	-0.43 (4.2)	-0.06 (1.9)
Δq_{3t}	0.08 (1.8)	0.02 (2.2)	-0.14 (2.8)
$\Delta i_{3t} - \Delta ie_t$	-0.01 (1.3)	0.01 (1.7)	-0.07 (2.4)
$\Delta^2 p_{3t} - \Delta^2 pe_t$	0.00 (0.1)	0.00 (0.7)	0.01 (1.6)
$\Delta cds_{3t} - \Delta cdse_t$	-0.13 (1.1)	-0.03 (0.5)	-0.41 (2.0)

[a] *Note*: absolute values of t-ratios in brackets, standard errors bootstrapped with 1000 replications.
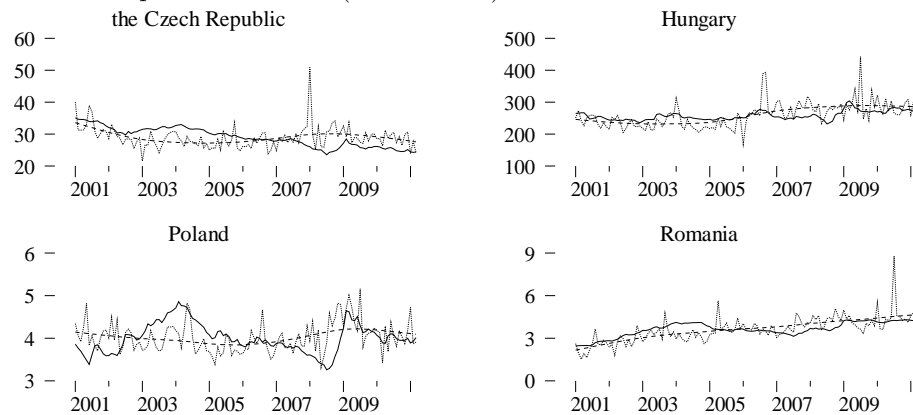
Piotr Kębłowski

4   Exchange rates misalignments

The point estimate of the common cointegrating vector given in (10) can be easily used to calculate equilibrium exchange rates and deviations of exchange rates from their steady-states. A straightforward transformation gives the following relationship between spot exchange rates and their determinants:

$$s_{it}^* = p_{it} - pe_t - 28.15 \left(i_{it} - ie_t\right) + 30.54 \left(\Delta p_{it} - \Delta pe_t\right) + 0.098 \left(cds_{it} - cdse_t\right), \quad (11)$$

where the weakly-stationary error correction term was omitted. Assuming that the explanatory variables are in their steady-states the paths of the equilibrium exchange rates are given by s_{it}^*. Due to high variance of the monthly inflation rates, the short-term fluctuations of s_{it}^* were attenuated by means of the Hodrick-Prescott filter. The point estimates of equilibrium exchange rates and the actual values of nominal exchange rates are given in Figure 2. The estimated paths of equilibrium exchange
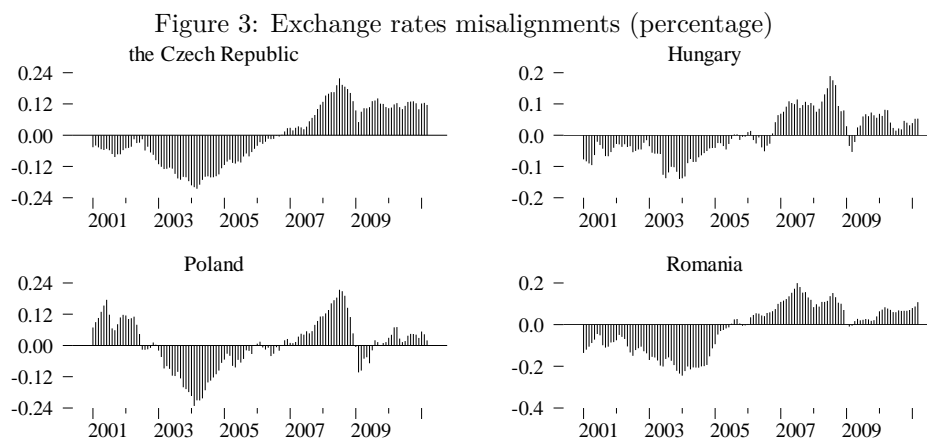
Figure 2: Spot exchange rates (solid lines), equilibrium values s_{it}^* (dotted lines) and HP filtered equilibrium values (dashed lines)

rates in the NMS evolve dissimilarly in general. Nevertheless, at the beginning of the sample, i.e. till 2005, there is an appreciatory trend in the equilibrium exchange rates of the Czech Republic, Hungary and Poland, which is due to the perspective of entering the EU. Moreover, in case of the Czech Republic and Poland there is a two year depreciatory trend starting in the middle of 2007, which seems to be related to the financial crisis in the subsequent years and increments of the risk premiums. In case of Hungary, a shift in mean of the equilibrium exchange rate is observed at the beginning of 2006, which should be attributed to the deterioration of the Hungarian economy performance and the following austerity package. The equilibrium exchange rate of the Romanian currency exhibits a depreciatory trend in the whole sample, which is mostly due to the high domestic inflation rate. Note however, that the

common cointegration vector was rejected in case of the panel including Romania. Therefore, the last result should be reexamined in a larger sample and/or separate study. The comparison of the equilibrium and actual values of nominal exchange

Figure 3: Exchange rates misalignments (percentage)

rates leads to exchange rate misalignments, see Figure 3. There are some easily visible common tendencies in their evolution. Firstly, all currencies of the NMS are undervalued in the years 2003-2004, i.e. before the EU enlargement in the next year. The maximum undervaluation took place at the turn of 2003 and 2004, reaching 21% percent for the Czech koruna, 14% for the forint, 23% for the zloty and almost 25% for the leu. Secondly, from 2007 till the end of 2008, when the financial crisis begins, the currencies of the NMS are overvalued. With the exception of Romania, the maximum overvaluation of about 20% took place on July 2008. Finally, in the last years the exchange rates in Poland and Hungary remains close to its equilibrium levels, whereas the currencies of the Czech Republic and Romania are overvalued, with respect to the estimated equilibrium levels.

The estimated paths of the equilibrium values and misalignments in the NMS can be compared with the results based on other approaches, even though usually there are different price indices or currency baskets employed in calculations of the real exchange rates. Based on the BEER approach, Magyari (2008) found that the Czech koruna, the zloty and the leu were undervalued in the years 2003-2004, whereas the forint was close to its equilibrium level in these years. The time span of the undervaluation of the aforementioned currencies is consistent with our results, the last conclusion is different. The application of the FEER approach in Rubaszek and Rawdanowicz (2009) leads to the conclusion that the zloty was undervalued in the 2003-2004 years, whereas the forint was overvalued in this period. The former conclusion agrees with our findings, the latter differs. Frait, Komarek, Melecky (2006) states in turn that in 2003 (the last year in the sample) the currencies of the Czech Republic, Hungary

and Poland started to be undervalued, which is similar to our results. Estimates of the exchange rate misalignments in the NMS for the preceding years can be found in Égert and Lahrèche-Révil (2003) and Rahn (2003).

5   Conclusions

The paper examined the joint determination of the floating exchange rates in the New Member States, based on the panel VEC framework. Our results indicate that the exchange rates of the Czech koruna, the Hungarian forint and the Polish zloty vs. euro follow the same long-run relationship, in which the real exchange rates are explained by the real interest rates parities and the spreads of the credit default risk premiums, approximated by the credit default swaps. In case of Romania, which seems to be a heterogeneous entity in the panel of the NMS with the floating exchange rates, the common long-run relationship is rejected.

We have found that the equilibrium exchange rates in these countries evolve dissimilarly in the long-run, even though there are some common tendencies in the short-run, resulting from the EU enlargement or the subprime crisis. On the other hand, the deviations of the actual exchange rates from its equilibrium levels show some common patters with respect to time-spans and values of exchange rate misalignments. Firstly, the currencies of the NMS were undervalued in the 2003-2004 years. Secondly, the overvaluation took place between 2007 and 2009. Finally, the forint and the zloty seem to be close to its steady-state levels after the beginning of the financial crisis, whereas the Czech koruna and the leu continue their overvaluation after a short interval.

According to the perspectives of entering the NMS currencies the ERM2 and joining the Eurosystem, it is obvious that this will basically rely on fulfillment of the strict convergence criteria and a political decision. Nevertheless, it is also clear that entering the ERM2 mechanism depends on that how probable is to meet its restrictions. Our results show that the equilibrium exchange rates of the NMS seem to be more stable than the actual values and the exchange rates misalignments exhibit common patters. Therefore, these countries are enabled to jointly enter the EMU. However, high values of the exchange rates misalignments at the periods of under/overvaluation, reaching over 20%, hinder this process essentially.

Acknowledgements

References

[1] Aixala J., Fabro G., (2009), Economic freedom, civil liberties, political rights and growth: a causality analysis, *Spanish Economic Review* 11, 165-178.

[2] Adler M, Song J., (2010), The bahaviour of emerging markets sovereigns' credit default swaps premiums and bond yield spreads, *International Journal of Finance and Economics* 15, 31-58.

[3] Banerjee A., Marcellino M., Osbat C., (2004), Some cautions of the use of panel methods for integrated series of macroeconomic data, *Econometrics Journal* 7, 322-340.

[4] Banerjee A., Marcellino M., Osbat C., (2005), Testing for PPP: Should we use panel methods?, *Empirical Economics* 30, 77-91.

[5] Blancheton B., Bordes C., Maveyraud S., Rous P., (2012), Risk of liquidity and contagion of the crisis no the United States, United Kingdom and euro zone money markets, *International Journal of Finance and Economics* 17, 124-146.

[6] Duffie D., (1999), Credit swap valuation, *Financial Analysts Journal* 55, 73-87.

[7] Duffie D., Gârleanu N., (2001), Risk and Valuation of Collateralized Debt Obligations, *Financial Analysts Journal* 57, 41-59.

[8] Égert B., Lahrèche-Révil A., (2003), Estimating the Fundamental Equilibrium Exchange Rate of Central and Eastern European Countries The EMU Enlargement Perspective, *CEPII research center Working Papers* 2003-05.

[9] Frait J., Komarek L., Melecký M., (2006), The real exchange rate misalignments in the five central european countries, *Warwick Economic Research Papers* 739.

[10] Groen J.J.J., Kleibergen F., (2003), Likelihood-based cointegration analysis in panels of vector error correction models, Journal of Business & Economic Statistics 21, 295-318.

[11] Hadri K., (2000), Testing for stationarity in heterogeneous panel data, *Econometrics Journal* 3, 148-161.

[12] Im K.S., Pesaran M.H., Shin Y., (2003), Testing for unit roots in heterogeneous panels, *Journal of Econometrics* 115, 53-74.

[13] Jacobson T., Lyhagen J., Larsson R., Nessén M., (2002), Inflation, exchange rates and PPP in a multivariate panel cointegration model, *Sveriges Riksbank Working Paper Series* 145.

[14] Juselius K., MacDonald R., (2003), International parity relationships between Germany and the United States: a joint modeling approach, available at: `http://www.econ.ku.dk/okokj/Papers/ronnie3.pdf`.

[15] Juselius K., MacDonald R., (2004), International parity relationships between the USA and Japan, *Japan and the World Economy* 16, 17-34.

[16] Kelm R., (2010), The Exchange Rate and Two Price Inflations in Poland in the Period 1999-2009. Do Globalization and Balassa-Samuelson Effect Matter?, *Central European Journal of Economic Modelling and Econometrics* 2, 315-349.

[17] Kelm R., (2011), The Polish zloty/euro exchange rate under free float: an econometric investigation, available at: `http://www.cerge-ei.cz/pdf/gdn/RRCIX_70_paper_01.pdf`.

[18] Kębłowski P., (2009) Modelling Integrated Panel Data: An Overview, [in:] Welfe W., Lang P. (eds.) *Knowledge Based Economies*, Frankfurt am Main.

[19] Kębłowski P., Welfe A., (2010), Estimation of the equilibrium exchange rate: the CHEER approach, *Journal of International Money and Finance* 29, 1385-1397.

[20] Kębłowski P., Welfe A., (2012), A risk-driven approach to exchange-rate modelling, *Economic Modelling* 29, 1473-1482.

[21] Larsson R., Lyhagen J., (2007), Inference in Panel Cointegration Models with long panels, *Journal of Business and Economic Statistics* 25, 473-483.

[22] Larsson R., Lyhagen J., Löthgren M., (2001), Likelihood-based cointegration tests in heterogeneous panels, *Econometrics Journal* 4, 109-142.

[23] MacDonald R., (2000), Concepts to calculate equilibrium exchange rates: an overview. *Deutsche Bundesbank Discussion Paper* 3, Frankfurt o/M.

[24] Majsterek M., Welfe A., (2012), Price-wage nexus and the role of a tax system, *Economic Change and Restructuring* 45, 121-133.

[25] Magyari I., (2008), Exploring the correlation between the exchange rate misalignment and economic growth in the CEE countries, Advances in Economic and Financial Reseach – *DOFIN Working Paper* Series 16.

[26] Rahn J., (2003), Bilateral equilibrium exchange rates of EU accession countries against the euro, *BOFIT Discussion Papers* 11.

[27] Remolona E.M., Scatigna M., Wu E., (2008), A ratings-based approach to measuring sovereign risk, *International Journal of Finance and Economics* 13, 26-39.

[28] Rubaszek M., Rawdanowicz Ł., (2009), Economic convergence and the fundamental exchange rate in Central and Eastern Europe, *International Review of Financial Analysis* 18, 277-284.
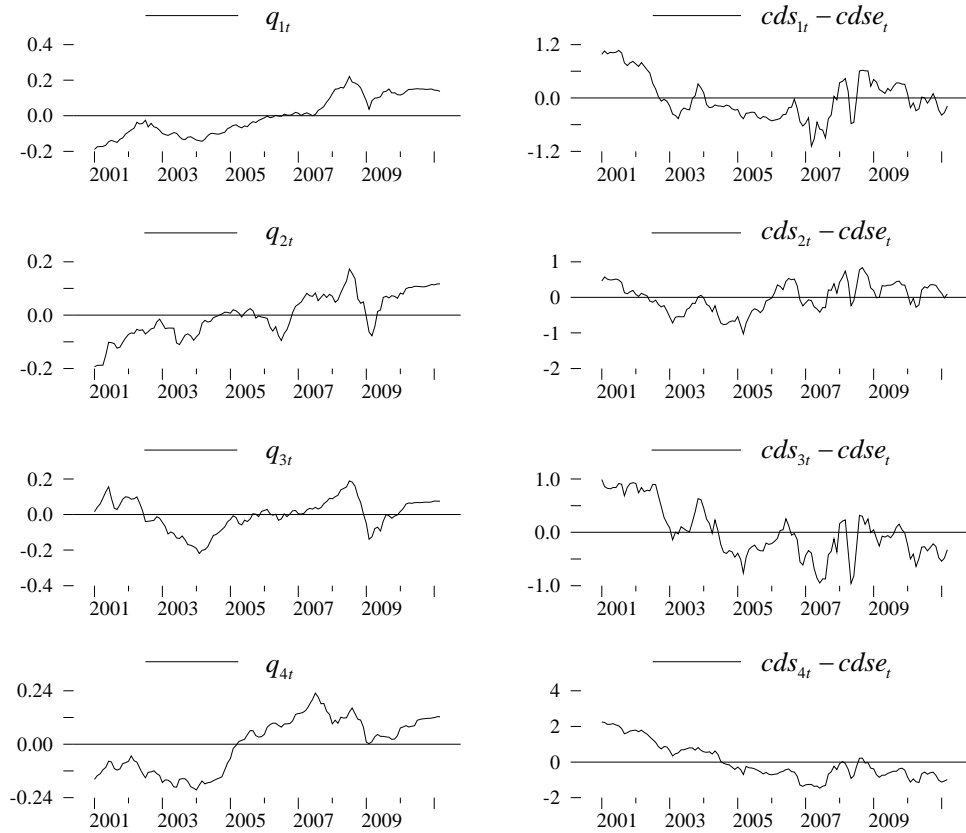
Piotr Kębłowski

Appendix

Figure 1: The monthly data

Figure 2: The monthly data – cont.