

# **Are the New Member States on the fast track to the EMU?**

## **An analysis of exchange rates misalignments in Central European countries<sup>\*</sup>**

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### Abstract

The paper presents the panel VEC model of the floating exchange rates in the New EU Member States. It was found 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 risk premiums, approximated by the sovereign credit default swaps. In case of the Romanian leu, the common relationship is rejected due to differences in the economic setting. The estimated misalignments exhibit some common patterns in terms of time spans and percentage values of under/overvaluation.

*JEL classification:* C33, E44

*Keywords:* equilibrium exchange rate, exchange rate misalignments, New Member States, EMU, panel VEC model

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## 1. INTRODUCTION

The new EU member states (the NMS henceforth) obligation to join the euro area raises a number of issues which have to be addressed before entering the Eurosystem. One of the key items is an assessment of the equilibrium exchange rates in order to suitably set the central values of the bands in the ERM2 mechanism and finally the parities of the currencies exchange. Therefore, the determinants of the exchange rates should be carefully identified. However, it needs to be stressed at the outset that the usefulness of the approaches to the exchange rate modeling which are typical of the large developed economies and seems to be useful in long time spans covering several tens of years, including the BEER and the FEER approaches, is limited. Instead, in this paper we use the medium-run approach based on the CHEER hypothesis complemented by the sovereign credit default risk perceived by financial investors.

A quick view at the history of twelve currencies in the NMS shows that the Estonian kroon, Lithuanian litas and Slovenian tolar entered the ERM2 mechanism in the year 2004, Cypriot pound, Latvian lats, Maltese lira and Slovak koruna joined in the next year, whereas the Bulgarian lev is pegged to the euro from the year 1999 (see also Katsimi, 2008). The currencies of Czech Republic, Hungary, Poland and Romania are (dirty) float, as the floating exchange rates of the Czech koruna, the Hungarian forint, the Polish zloty and the Romanian leu against the euro 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 +/- 15% in May 2001, and the crawling peg was repealed on 1 October 2001. Therefore, it can be assumed at best that the joint time span of the (nearly) floating exchange rate in these four 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 the empirical results based on the panel VEC model. Exchange rates misalignments are analysed in section 4. Section 5 summarizes the findings and concludes.

## 2. RISK PREMIUMS AND THE EXCHANGE RATES DETERMINATION

There is a general recognition that the balance of payment condition is a proper framework for exchange rates modeling. However, it leads to different theories about the determinants of exchange rates. On the one hand, there are approaches which highlight the both sides of the balance of payment condition. Therefore, apart from spreads of interest rates and prices driving the capital account, the real determinants of the current account are included. On the other hand, it can be stated 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 the capital account, it is natural to combine the purchasing power parity and the net interest rate differential in order to explain the exchange rate 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 the real exchange rates. This allows verifying a number of theories, like the term structure or the real interest rate parity. Kębłowski and Welfe (2010) found that the extended CHEER model is useful for the explanation of the euro/zloty exchange rate behaviour. However, the worldwide financial market turmoil at the end of 2008 year reveals that there is an additional factor which is usually omitted. The rapid change of risk aversion and capital

outflows led to similar rapid depreciation of exchange rates in case of emerging markets. Therefore, it seems that a measure of risk premium should be suitably taken into account.

The risk premium can be approximated by means of liquidity indices, which are calculated as proportions of M2 aggregate to reserves or debt to GDP, etc. (see Kelm (2011)), an equity market variables, term spreads or over-the-counter derivatives, assuming that the counterparty risk and the credit default risk are mutually independent and the market of the derivative is enough deep and liquid (see Kębłowski and Welfe (2011)). The quotations of the over-the-counter derivatives represent the level of risk perceived by financial investors, which is responsible for capital flows and therefore the exchange rate fluctuations in the medium-run.

The focus here is on credit default swaps for the governmental bonds. The quotations for the sovereign CDS started in: May 2006 for the Czech Republic, March 2002 for Hungary, November 2000 for Poland and October 2002 for Romania. In turn, the CDS index for the German bonds, which represents the euro area counterpart, is quoted from 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.

< Figure 1. about here >

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 counterpart. However, the most noticeable are huge increments of the CDS quotations 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 than 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).

Summarizing, it will be verified whether the four aforementioned countries share similar mechanism determining their exchange rates and the real exchange rates movements can be explained by differentials of interest rates, inflation rates and sovereign credit default swaps indices:

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

where  $q_{it} = p_{it} - pe_t - s_{it}$ ,  $p_{it}$  is the log of consumer price index for the NMS ( $i = 1, 2, \dots, I$ ,  $i = 1$  is equal to 1 for the Czech Republic, 2 - Hungary, 3 - Poland, 4 - Romania),  $pe_t$  is the log of consumer price index in the euro area, and  $s_{it}$  is the log of spot exchange rate,  $i_{it}$ ,  $ie_t$ ,  $cds_{it}$  and  $cdse_t$  denote the long-term interest rate yield and the logs of the CDS indices in the NMS and in the euro area respectively.

### 3. THE PANEL VEC MODEL

According to (1) the analysis is based on four variables ( $P = 4$ ):

$$\mathbf{y}_{it} = [q_{it} \ (i_{it} - ie_t) \ (\Delta p_{it} - \Delta pe_t) \ (cds_{it} - cdse_t)]'$$

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 and the long-term interest rates are monthly average yields on ten-year floating-rate bonds. The monthly data over the period January 2001 – April 2011 comes from Eurostat. All series are demeaned and integrated of order one (see Figure 2 and Table 1)<sup>2</sup>.

< Table 1. about here >

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<sup>2</sup> The procedures written by Tom Doan (Estima) and Johan Lyhagen (Stockholm School of Economics) were employed in calculations.

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

$$\Delta \mathbf{y}_t = \mathbf{\Pi} \mathbf{y}_{t-1} + \sum_{k=1}^K \mathbf{\Gamma}_k \Delta \mathbf{y}_{t-k} + \boldsymbol{\varepsilon}_t, \quad (2)$$

where  $\mathbf{\Pi}$  and  $\mathbf{\Gamma}_k$  ( $k=1, 2, \dots, K$ ,  $K=3$  in the study) are  $IP \times IP$  matrices of parameters,  $\boldsymbol{\varepsilon}_t = [\boldsymbol{\varepsilon}'_{1t} \boldsymbol{\varepsilon}'_{2t} \dots \boldsymbol{\varepsilon}'_{It}]'$  and  $\boldsymbol{\varepsilon}_t \sim N_{IP}(\mathbf{0}; \boldsymbol{\Omega})$ . If the variables are cointegrated then the matrix  $\mathbf{\Pi}$  has a reduced rank and can be decomposed as  $\mathbf{\Pi} = \mathbf{A} \mathbf{B}'$ , where  $\mathbf{A}$  and  $\mathbf{B}$  are  $IP \times \sum R_i$  matrices of parameters and  $R_i$  denotes the cointegration rank in the cross-section  $i$ .

The general model (2) allows for simultaneity to a large extent, since the loadings matrix  $\mathbf{A}$ ,  $\mathbf{\Gamma}_k$  and  $\boldsymbol{\Omega}$  are unrestricted:

$$\mathbf{A} = \begin{bmatrix} \mathbf{A}_{11} & \mathbf{A}_{12} & \dots & \mathbf{A}_{1I} \\ \mathbf{A}_{21} & \mathbf{A}_{22} & \dots & \mathbf{A}_{2I} \\ \vdots & \vdots & \ddots & \vdots \\ \mathbf{A}_{I1} & \mathbf{A}_{I2} & \dots & \mathbf{A}_{II} \end{bmatrix}, \quad (3)$$

whereas the matrix of cointegrating vectors  $\mathbf{B}$  has the following structure:

$$\mathbf{B} = \begin{bmatrix} \mathbf{B}_{11} & \mathbf{0} & \dots & \mathbf{0} \\ \mathbf{0} & \mathbf{B}_{22} & \dots & \mathbf{0} \\ \vdots & \vdots & \ddots & \vdots \\ \mathbf{0} & \mathbf{0} & \dots & \mathbf{B}_{II} \end{bmatrix}. \quad (4)$$

Therefore, heterogeneous long-run relationships are allowed within cross-section and each cointegration vector can affect any variable in the system (see Larsson and Lyhagen, 1999, also Groen and Kleibergen, 2001, and Larsson et al., 2001, for more restricted models). Note however, that even though the decomposition of matrix  $\mathbf{\Pi}$  into matrices  $\mathbf{A}$  and  $\mathbf{B}$  allows a non-diagonal structure of matrix  $\mathbf{\Pi}$ , the variables from different cross-sections are not permitted to cointegrate (see also Jacobson et al., 2002).

The model (2) with matrices  $\mathbf{A}$  and  $\mathbf{B}$  as in (3) and (4) 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 (2) with  $\mathbf{A}$ ,  $\mathbf{\Gamma}_k$  and  $\mathbf{\Omega}$  unrestricted and semi-diagonal matrix  $\mathbf{B}$ , Larsson and Lyhagen (1999) proved that the  $LR$  statistic for the common cointegration rank hypothesis, the panel counterpart of the trace test,  $\mathcal{H}_0 : rk(\mathbf{\Pi}_{ii}) = R_i \leq R$  for  $i = 1, 2, \dots, I$  vs.  $\mathcal{H}_1 : rk(\mathbf{\Pi}_{ii}) = P$  for  $i = 1, 2, \dots, I$ , converges to the convolution of the Dickey-Fuller type distribution and the  $\chi^2$  distribution:

$$LR_{IT} \stackrel{D}{=} \chi_{I(I-1)R(P-R)}^2 + \text{tr} \left( \int_0^1 d\mathbf{W}\mathbf{W}' \left[ \int_0^1 \mathbf{W}\mathbf{W}' du \right]^{-1} \int_0^1 \mathbf{W} d\mathbf{W}' \right) \text{ as } T \rightarrow \infty, \text{ for fixed } I, \quad (5)$$

where  $\mathbf{W}$  denotes the  $I(P-R)$ -dimensional Brownian motion. The results of inference on the common cointegration rank are presented in Table 2.

< Table 2. about here >

First of all, due to huge size distortion the standard  $LR$  statistic rejects all consecutive null hypotheses in both cases, which leads to the false 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 3. shows the maximum likelihood estimates of cointegrating vectors for  $R = 1$ . The results support the relationship (1), 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 semi-elasticities for interest rates and inflation rates seem to be proper for monthly data.

< Table 3. about here >

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  $\mathcal{H}_0 : \mathbf{B}_{11} = \mathbf{B}_{22} = \dots = \mathbf{B}_{II}$ :

$$\mathbf{B} = \begin{bmatrix} \mathbf{B}_{11} & \mathbf{0} & \dots & \mathbf{0} \\ \mathbf{0} & \mathbf{B}_{11} & \dots & \mathbf{0} \\ \vdots & \vdots & \ddots & \vdots \\ \mathbf{0} & \mathbf{0} & \dots & \mathbf{B}_{11} \end{bmatrix} \quad (6)$$

vs.  $\mathcal{H}_1 : \mathbf{B}_{ii} \neq \mathbf{B}_{jj}$  for some  $i, j$ . The  $LR$  statistic for the common cointegration space hypothesis converges asymptotically to the  $\chi^2$  distribution (see Larsson and Lyhagen (1999)):

$$LR_{IT} \mathcal{D} \chi_{(I-1)R(P-R)}^2 \text{ as } T \rightarrow \infty, \text{ for fixed } I. \quad (7)$$

The  $LR$  statistics of common cointegration space for  $R = 1$  are  $\chi_9^2 = 63.47 (0.00)$  for the panel of all four countries, and  $\chi_9^2 = 9.34 (0.16)$  for the case of three countries ( $p$ -values in the 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 common cointegrating vector is given as follows:

$$q_{it} = 28.15(i_{it} - ie_t) - 30.54(\Delta p_{it} - \Delta pe_t) - 0.098(cds_{it} - cdse_t) + ec_{it}, \quad (8)$$

where  $ec_{it}$  denotes weakly-stationary error correction term.

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 (8) 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.

< Table 4. about here >

The loadings matrix in Table 4. shows the influence of deviations from the steady-states implied by (8) 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.

#### 4. EXCHANGE RATES MISALIGNMENTS

The common cointegrating vector given in (8) 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:

$$s_{it}^* = p_{it} - pe_t - 28.15(i_{it} - ie_t) + 30.54(\Delta p_{it} - \Delta pe_t) + 0.098(cds_{it} - cdse_t), \quad (9)$$

where the weakly-stationary error correction term was omitted. Assuming that the explanatory variables are in their steady-states the paths of 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 equilibrium and the actual values of nominal exchange rates are given in Figure 2.

< Figure 2. about here >

The estimated paths of equilibrium exchange rates in the NMS evolve dissimilarly in general. Nevertheless, at the beginning of the sample, i.e. till the year 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 Czech Republic and Poland there is a two year depreciatory trend from 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 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.

< Figure 3. about here >

The comparison of the equilibrium and actual values of nominal exchange 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 the year 2007 the currencies of the NMS are overvalued, till the end of 2008, when the financial crisis begins. 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 real exchange rates. Based on the BEER approach, Magyari (2008) found that the Czech koruna, the zloty and the leu were undervalued in the 2003-2004 years, whereas the forint was close to its equilibrium level in these years. The time span of 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 et al. (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.

## 5. CONCLUSIONS

The paper examined the joint determination of the 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 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 patterns 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 join the Eurosystem, it is obvious that basically this will 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 seems to be more stable than the actual values and the exchange rates misalignments exhibits 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.

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Table 1. Inference on integration order

|              | $\Delta q_{it}$           | $\Delta i_{it} - \Delta i e_t$ | $\Delta^2 p_{it} - \Delta^2 p e_t$ | $\Delta cds_{it} - \Delta cdse_t$ |
|--------------|---------------------------|--------------------------------|------------------------------------|-----------------------------------|
| $Z_\mu^a$    | -0.90 (0.81) <sup>c</sup> | -2.02 (0.98)                   | 0.49 (0.31)                        | -0.87 (0.81)                      |
| $Z_{tbar}^b$ | -7.90 (0.00)              | -16.52 (0.00)                  | -8.11 (0.00)                       | -10.98 (0.00)                     |
| $Z_{tbar}^c$ | -6.87 (0.00)              | -12.05 (0.00)                  | -7.03 (0.00)                       | -9.00 (0.00)                      |
|              | $q_{it}$                  | $i_{it} - i e_t$               | $\Delta p_{it} - \Delta p e_t$     | $cds_{it} - cdse_t$               |
| $Z_\mu$      | 72.83 (0.00)              | 4.05 (0.00)                    | 7.98 (0.00)                        | 17.34 (0.00)                      |
| $Z_{tbar}$   | -0.76 (0.23)              | -0.98 (0.16)                   | -1.14 (0.13)                       | -1.53 (0.06)                      |
| $Z_{tbar}$   | -0.69 (0.24)              | -0.92 (0.18)                   | -1.07 (0.14)                       | -1.44 (0.07)                      |

<sup>a</sup> Note: the statistic of Hadri (2000) stationarity test,  $Z_\mu \overset{D}{\sim} N(0;1)$  as  $T \rightarrow \infty$ ,  $I \rightarrow \infty$  and  $I/T \rightarrow 0$ .

<sup>b, c</sup> Note: the statistics of Im, Pesaran, Shin (2003) unit root test,  $Z_\mu \overset{D}{\sim} N(0;1)$  as  $T \rightarrow \infty$ ,  $I \rightarrow \infty$  and  $I/T \rightarrow const$ .

<sup>c</sup> Note:  $p$ -values in the brackets.

Table 2. Inference on common cointegration rank

| $H_0$   | 4 countries <sup>a</sup> |              |            | 3 countries <sup>b</sup> |              |            |
|---------|--------------------------|--------------|------------|--------------------------|--------------|------------|
|         | $LR$                     | $LR^{BC\ c}$ | 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      |

<sup>a</sup> Note: the Czech Republic, Hungary, Poland, Romania.

<sup>b</sup> Note: the Czech Republic, Hungary, Poland.

<sup>c</sup> Note: Bartlett corrected  $LR$  statistic.

Table 3. Cointegrating vectors

|                    | $q_{it}$ | $i_{it} - ie_t$ | $\Delta p_{it} - \Delta pe_t$ | $cds_{it} - cdse_t$ |
|--------------------|----------|-----------------|-------------------------------|---------------------|
| the Czech Republic | 1        | -21.40          | 32.84                         | 0.102               |
| Hungary            | 1        | -38.20          | 25.92                         | 0.061               |
| Poland             | 1        | -43.77          | 25.48                         | 0.087               |
| Romania            | 1        | -26.43          | 11.22                         | 0.194               |

Table 4. Loadings matrix

|                                    | $\hat{\alpha}_1$ | $\hat{\alpha}_2$ | $\hat{\alpha}_3$ |
|------------------------------------|------------------|------------------|------------------|
| $\Delta q_{1t}$                    | -0.13            | 0.03             | 0.01             |
| $\Delta i_{1t} - \Delta i e_t$     | -0.10            | 0.00             | -0.02            |
| $\Delta^2 p_{1t} - \Delta^2 p e_t$ | 0.01             | 0.00             | -0.01            |
| $\Delta cds_{1t} - \Delta cdse_t$  | -0.22            | -0.06            | -0.03            |
| $\Delta q_{2t}$                    | 0.07             | -0.11            | -0.01            |
| $\Delta i_{2t} - \Delta i e_t$     | -0.01            | -0.09            | -0.01            |
| $\Delta^2 p_{2t} - \Delta^2 p e_t$ | 0.01             | 0.00             | 0.00             |
| $\Delta cds_{2t} - \Delta cdse_t$  | 0.00             | -0.43            | -0.06            |
| $\Delta q_{3t}$                    | 0.08             | 0.02             | -0.14            |
| $\Delta i_{3t} - \Delta i e_t$     | -0.01            | 0.01             | -0.07            |
| $\Delta^2 p_{3t} - \Delta^2 p e_t$ | 0.00             | 0.00             | 0.01             |
| $\Delta cds_{3t} - \Delta cdse_t$  | -0.13            | -0.03            | -0.41            |

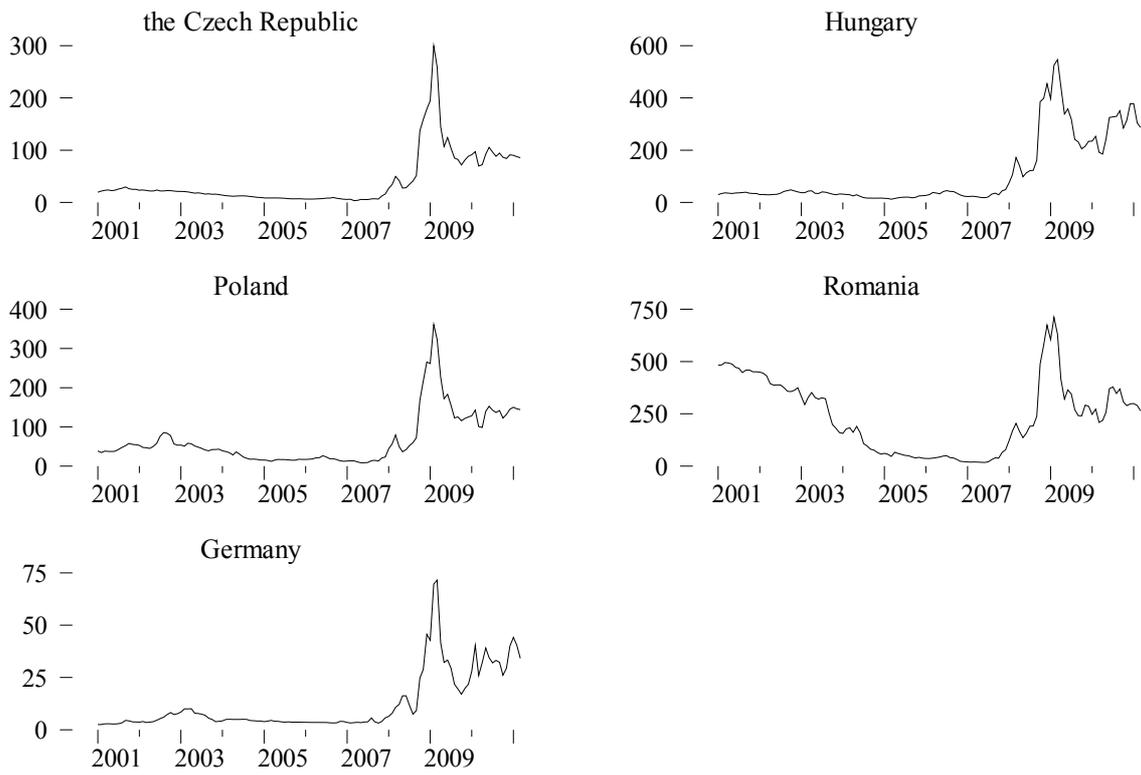


Figure 1. The sovereign CDS indexes

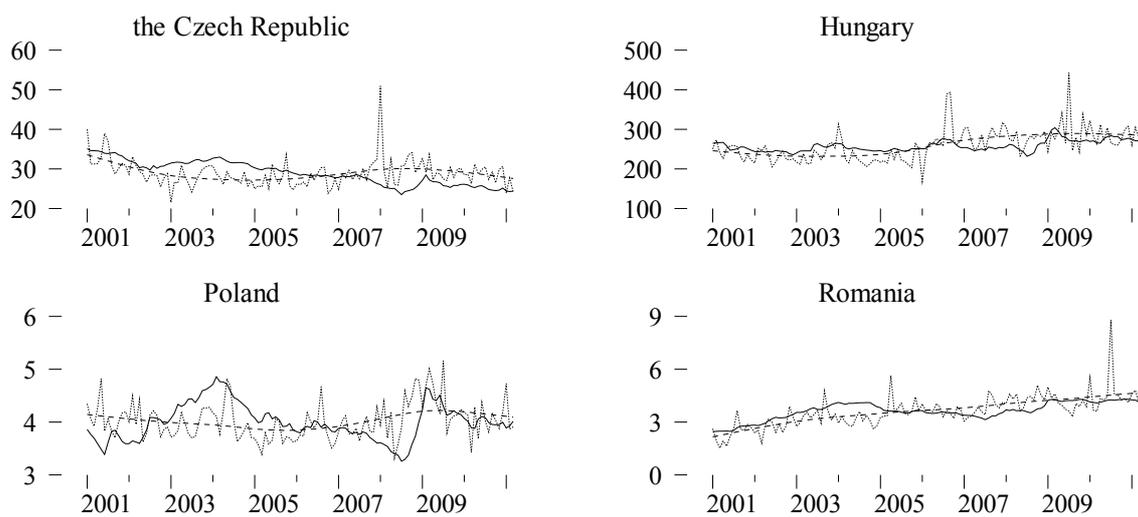


Figure 2. Spot exchange rates (solid lines), equilibrium values  $s_{it}^*$  (dotted lines) and HP filtered equilibrium values (dashed lines)

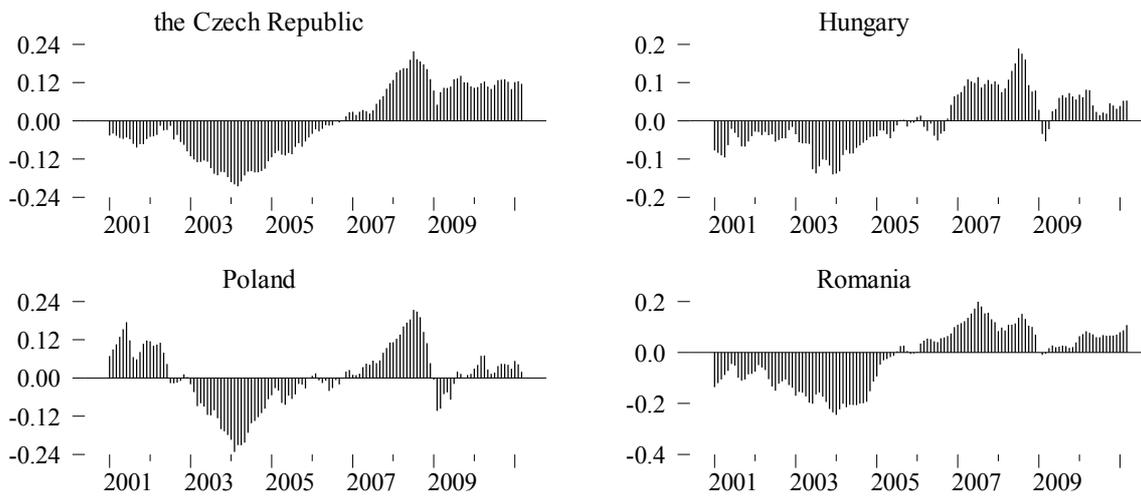


Figure 3. Exchange rates misalignments (percentage)

# APPENDIX

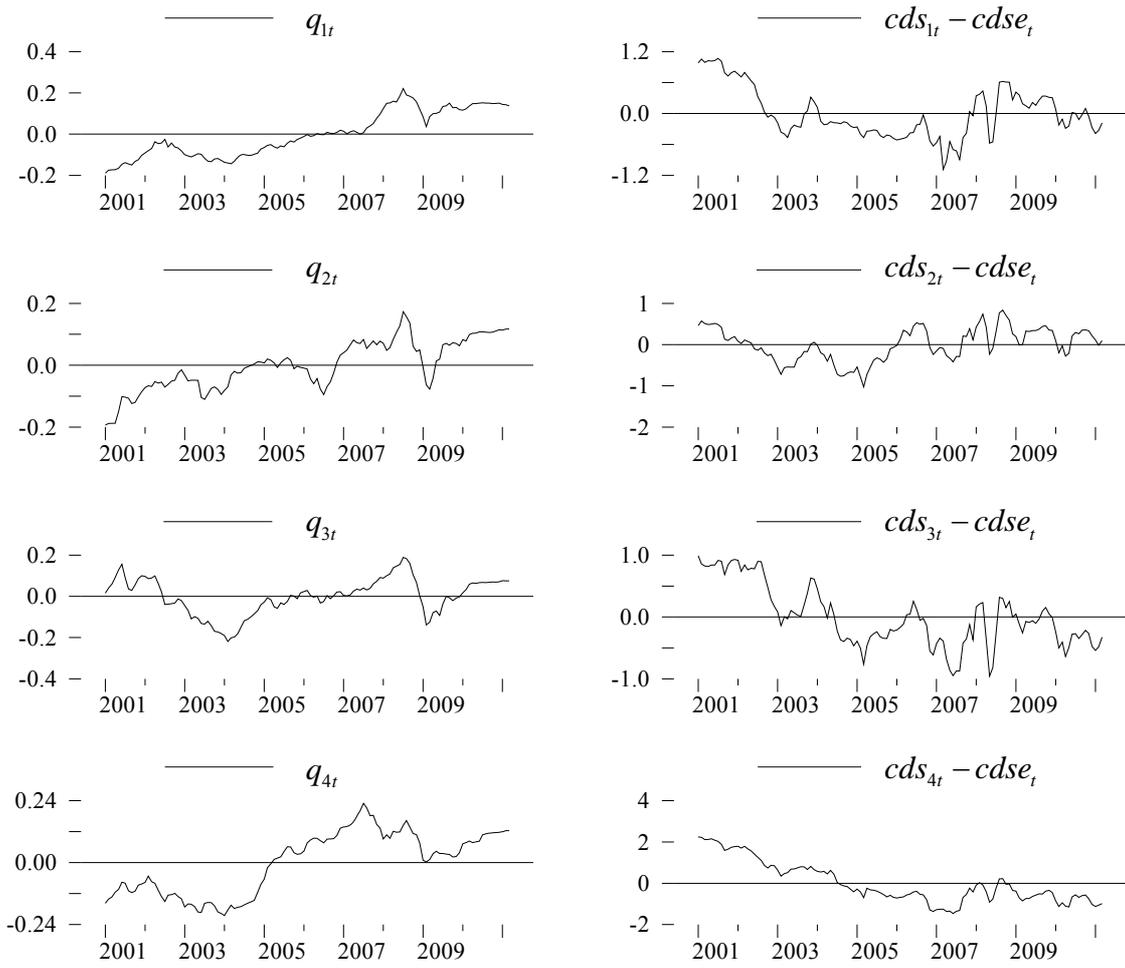


Figure A.1. The monthly data

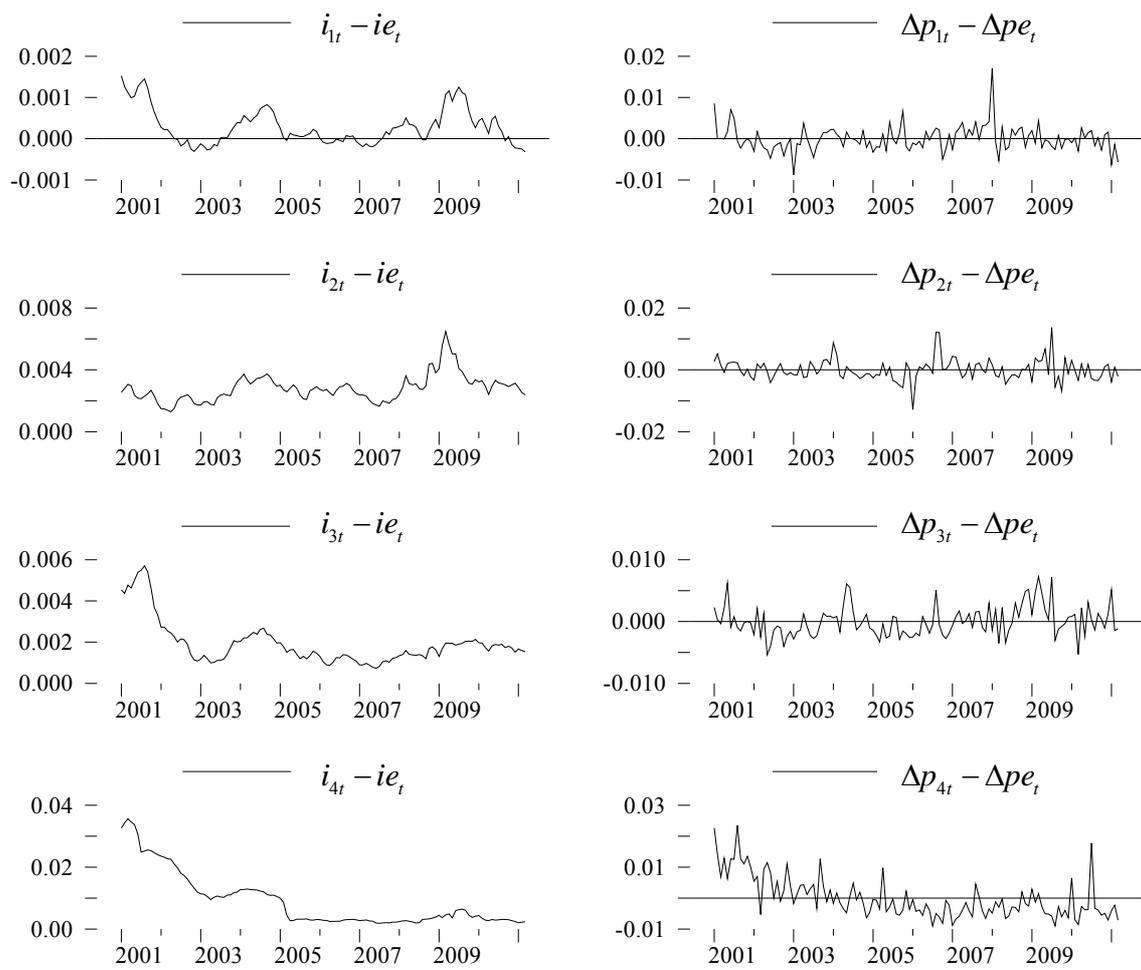


Figure A.1. cont.