UNCOVERED INTEREST PARITY IN CENTRAL AND EASTERN EUROPE: CONVERGENCE AND THE GLOBAL FINANCIAL CRISIS[1]

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Abstract

This paper presents tests of uncovered interest parity in Croatia, the Czech Republic, Hungary, Poland and Romania; all countries in Central and Eastern Europe with floating exchange rates. Data are monthly and the trading horizon is three months. The estimations show that the UIP hypothesis is rejected for the full sample from 1999 to 2011 for all five countries. A number of reasons for the rejection were investigated. Rolling regressions show that standard versions of the UIP essentially lose all explanatory power in 2008-10, which was a period in which the global financial crisis led to instability in currency and interest markets in Central and Eastern Europe. Two indicators of global risk aversion were also found to enter significantly in the many UIP estimations. Finally, the size of the interest rates spread also seems to be of importance, at least for Poland and Romania.

Keywords: UIP, financial integration, global financial crisis, Central and Eastern Europe

JEL Classification: E43, F36, G01, G15

"*Uncovered interest rate parity remains a key assumption in international economics despite the massive body of empirical evidence against the hypothesis.*"

A. Alexius (2001, p. 505)

1. Introduction

This paper presents the results of econometric analyses testing the uncovered interest parity (UIP) hypothesis on data from Poland, the Czech Republic, Hungary, Romania and Croatia. The data sample starts in 1999 or shortly afterwards and ends in September 2011, and as such spans a period in which the countries experienced both rapid economic and financial integration and also the fallout from the global financial crisis. The UIP hypothesis is tested for a trading horizon of three months using monthly data. The five countries in the sample are the main countries in

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Central and Eastern Europe having floating or essentially floating exchange rate regimes during the sample period.[3] Poland, the Czech Republic and Hungary joined the European Union in May 2004 and Romania in January 2007, while Croatia was in the final stages of membership negotiations at the time of writing in August 2011.

The hypothesis of uncovered interest parity rests on the idea that arbitrage leads to equalisation of the return on assets or liabilities in the domestic currency and the expected return on comparable assets or liabilities in a foreign currency. Testing the UIP hypothesis may thus provide information as to whether the exchange and interest markets under consideration function so that all the gains from trade are exploited, i.e. whether the markets are efficient. In practice, however, divergence between domestic and expected foreign returns may also be due to issues such as transaction costs, different risk profiles and non-symmetric tax treatments.

This paper presents tests of the UIP hypothesis for Croatia, the Czech Republic, Hungary, Poland and Romania. Section 2 provides a survey of empirical studies of the UIP hypothesis with a particular focus on studies dealing with countries in Central and Eastern Europe (CEE). There are only a very limited number of studies that examine the UIP hypothesis for Central and East European countries, particularly studies which use data covering the EU accession and the global financial crisis. The CEE countries liberalised their capital markets and removed their remaining exchange rate restrictions before joining the EU (European Commission 2010a). Many of the countries experienced substantial capital inflows in the years immediately before and after accession to the EU, just to see a reversal of the flows in 2008-09 following the global financial crisis (Jevcak et al. 2011). It is a largely un-researched question whether these abrupt changes in capital flows have affected the relationship between exchange rates and interest rates in the CEE countries.

Testing the UIP hypothesis for the CEE countries is also important because households and firms in many countries in the region have borrowed extensively in foreign currencies, mostly the euro and the Swiss franc (Rosenberg & Tirpak 2008). In essence borrowers expect that borrowing in a foreign currency is cheaper than domestic currency borrowing, meaning they have bet that the UIP will not hold within the horizon of the loan contract. Speculators without an underlying motive of borrowing or saving have also taken positions, *carry trade*, in the currencies of the CEE countries. Rosenberg & Tirpak (2008) and Brzoza-Brzezina et al. (2010) find that the interest differential between domestic and foreign rates is an important determinant of borrowing and saving in foreign currencies in the CEE countries.[4]

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[3] The study excludes countries with fixed exchanges and countries that adopted the euro during the sample period.

[4] Batini & Dowling (2011) use a UIP framework to decompose exchange rate movements between major currencies and the US dollar into shocks stemming from US monetary policy and other sources. The sharp depreciation of most of the sample currencies against the US dollar during the global financial crisis cannot be attributed to changes in the interest rate spread, but rather to changes in the risk premia. The subsequent appreciation of many of the

This paper seeks to contribute to the empirical literature on the UIP by investigating its empirical validity in the main CEE countries that have a floating exchange rate. The paper tests the UIP hypothesis using individual regressions for each of the five CEE countries. As typically found in the literature, the UIP holds better for some countries than for others and better in some periods than in others. The paper investigates factors that may explain the variation across countries and across time, linking the findings to the different stages of convergence attained in the countries and to the global financial crisis that unfolded in 2007-2009.

The rest of the paper is organised as follows: Section 2 discusses the theoretical foundation of the UIP hypothesis. Section 3 surveys a number of empirical studies with a particular emphasis on the CEE countries. Section 4 documents the data and shows the results of unit root tests. Section 5 presents the baseline estimations using the full sample available. Section 6 contains the estimations when structural change is identified using rolling windows. Section 7 considers whether there are non-linear effects. Section 8 shows the results when different proxies of external determinants of the risk premium are included. Finally, Section 9 summarises the results.

2. The theory of uncovered interest parity

The *theory* underlying the Uncovered Interest Parity is fairly simple as it builds on the assumption of arbitrage equalising expected returns in different markets (Levi 2005, Ch. 8).

Consider the investment decision of an investor who at time *t* seeks to invest a sum for a period of *m* time units. Assuming that the interest rate is constant and equal to $i_{t,m}$ for the entire investment horizon, the gross return from investing domestically is $1+i_{t,m}$ per time unit leading to $(1+i_{t,m})^m$ compounded during the *m* periods of the investment. The sum can alternatively be exchanged at the spot exchange rate S_t and invested abroad at the interest rate $i^*_{t,m}$. The foreign denominated gross return after *m* periods is $(1+i^*_{t,m})^m / S_t$ and this sum can be exchanged into domestic currency at the exchange rate S_{t+m}.

In practice the exchange rate *m* periods ahead is unknown, so the investor will have to form expectations for this exchange rate. The variable S^e_{t+m} denotes the expectation in period t for the exchange rate in period $t+m$. A *risk-neutral investor* would be indifferent as to whether to invest in the domestically denominated asset or in the foreign denominated asset if the expected returns are identical, i.e. if uncovered interest parity holds:

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currencies may partly reflect the carry trade exploiting low US interest rates and higher interest rates in other countries. None of the CEE countries are included in the sample.

$$(1+i_{t,m})^m = (1+i^*_{t,m})^m \frac{S^e_{t+m}}{S_t} \quad (1)$$

This condition is usually log-linearised. We adopt the notation $\Delta_m \log S^e_{t+m}$ = $\log S^e_{t+m} - \log S_t$, which is approximately the relative change in the exchange rate over the *m*-period horizon of the investment. The variable $\Delta_m \log S^e_{t+m}$ is positive if the investor expects that the domestic currency will depreciate from period *t* to period *t* + *m* and negative if the investor expects that the domestic currency will appreciate. Using this notation eq. (1) becomes:

$$\frac{\Delta_m \log S^e_{t+m}}{m} = \log(1+i_{t,m}) - \log(1+i^*_{t,m}) \quad (2)$$

Using the approximations $i_{t,m} \approx \log(1+i_{t,m})$ and $i^*_{t,m} \approx (1+i^*_{t,m})$ and lowercase s_t to denote the logarithm of the exchange rate, i.e. $s_t = \log(S_t)$ and $s^e_{t+m} = \log(S^e_{t+m})$, the version of the UIP in eq. (2) can be rewritten as:

$$\frac{\Delta_m s^e_{t+m}}{m} = i_{t,m} - i^*_{t,m} \quad (3)$$

The left-hand side is the annualised average expected capital gain from the foreign currency investment. The right hand side is the spread between the domestic and foreign interest rates. The upshot is that a positive spread is consistent with the UIP hypothesis only if the spot rate is expected to depreciate in the way given in eq. (3), i.e. investment in the foreign denominated asset will only take place if the positive interest spread is compensated for by a corresponding capital gain.[5]

Eq. (3) can be tested empirically if a measure of the *expected* spot exchange rate *m* periods ahead is available, for instance from surveys or market data. A more common methodology, however, is based on the assumption of rational expectations, i.e. $\Delta_m s^e_{t+m}/m$ = $\Delta_m s_{t+m}/m + \varepsilon_{t+m}$, where $\mathrm{E}_t[\varepsilon_{t+m}] = 0$, i.e. the mathematical expectation of ε_{t+m} is zero, conditional on information in period *t*. This empirical version of the UIP is:

$$\frac{\Delta_m s_{t+m}}{m} = i_{t,m} - i^*_{t,m} + \varepsilon_{t+m} \quad (4)$$

[5] The domestic interest rate that is consistent with UIP follows directly from Eq. (3), i.e. $i_{t,m} = i^*_{t,m} + \Delta_m s^e_{t+m}/m$.

A simple empirical methodology for a test of the UIP hypothesis entails estimation of the following standard UIP regression model:

$$\frac{\Delta_m s_{t+m}}{m} = \alpha + \beta(i_{t,m} - i^*_{t,m}) + \varepsilon_{t+m} \qquad (5)$$

Eq. (5) is the model used in most estimations in the paper. The UIP corresponds to the joint null hypothesis that the constant α = 0, the slope coefficient β = 1 and $\mathrm{E}_t[\varepsilon_{t+m}] = 0$; the UIP hypothesis cannot be rejected if none of these conditions can be rejected.[6] Three comments are appropriate:

First, the assumption that $\mathrm{E}_t[\varepsilon_{t+m}] = 0$ implies that the residuals are serially uncorrelated if the investment horizon coincides with the sampling frequency. If, however, the investment horizon exceeds the investment frequency (as would be the case with, for instance, monthly data and a quarterly investment horizon), overlapping data emerge and the residual will be subject to serial correlation of order *m* – 1 even if $\mathrm{E}_t[\varepsilon_{t+m}] = 0$ is satisfied for the investment horizon (Baillie & Bollerslev 2000).

Second, the test implies essentially a joint test of several hypotheses, including the hypothesis that arbitrage equalises the expected currency gain and the interest rate differential and the hypothesis that investors have rational expectations (Alper et al. 2009). If α = 0 and β = 1 cannot be rejected (in a model with non-serially correlated residuals), it is reasonable to assume that both hypotheses are satisfied. Rejection implies that the UIP does not hold, but the underlying reason (such as absence of arbitrage trades or non-rational expectations) cannot be identified right away.

Third, the test entails the estimation of *one* coefficient of the interest spread $i_{t,m} - i^*_{t,m}$, not separate coefficients for each of the interest rates. The implicit assumption is that the investors react only to the interest rate spread, i.e. in similarly sized but opposing ways to each of the two interest rates (Mehl & Cappiello 2007). In practice, the assumption is convenient as it typically implies that the interest spread $i_{t,m} - i^*_{t,m}$ is stationary, but this may not be the case for each interest rate considered individually.

The theoretical model in eq. (3) and the empirical model in eq. (5) are based on the assumption that the investors are risk-neutral and do not require a risk premium to hold one currency or the other. This assumption is unrealistic in practice insofar as investors are risk averse. A constant risk premium can be included by allowing the

[6] Fama (1984) suggests a narrower test of the UIP hypothesis, essentially testing whether the forward rate is an unbiased estimator of the future exchange rate. The Fama regression entails that the forward premium is regressed on the future exchange rate change and a slope coefficient of one is interpreted as confirmation of the efficient market hypothesis.

constant α to differ from zero.[7] This assumption might be too restrictive if the risk premium is non-constant, but it would then be necessary to model the risk premium. The presence of a risk premium – and in particular a non-constant risk-premium – does not contradict the UIP hypothesis *per se*, but it complicates the empirical testing as it requires that the risk premium can be identified empirically.

Beyond the presence of a risk premium, it is possible to point out a number of factors which would entail that eq. (3) would not hold (Levi 2005, Ch. 8):

- Financial markets may not be fully integrated because of regulation, institutional barriers or undeveloped trading possibilities (lack of instruments). In this case, the trades needed to arbitrage different expected returns may not be available.
- Illiquidity or thin markets may lead to market inefficiency as prices may not reflect available information. Illiquidity creates more risks and complicates arbitrage trades, but this may not play a major role in currency markets with large turnovers.
- Transaction costs may make it unprofitable to execute trades that exploit small deviations from the UIP.
- Information costs may be high, in part because information is needed for expectations about exchange rate movements to be formed.
- Investors in exchange and interest markets may not have fully rational expectations. Investors may use mechanical or momentum-based trading strategies, essentially disregarding the available information.
- Liquidity preference may favour investment in domestic currency assets, as investment in foreign currency assets may be more difficult to wind down if there is a sudden need for liquidity in the domestic currency.
- The asymmetric tax treatment of interest returns and returns from capital gains (here stemming from exchange rate changes) may mean that the strict UIP hypothesis which does not take account of taxation would not hold.

3. Empirical studies

The uncovered interest parity hypothesis has been tested empirically for a long time, but better financial data have continuously expanded the possibilities for testing. We will briefly discuss the results of studies using datasets covering developed economies, emerging market economies and countries in Central and Eastern Europe.

Meese & Rogoff (1983) is an influential early study showing that the interest rate spread has essentially no predictive power for the future exchange rate movements of the US dollar when evaluated on data from the 1970s.

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[7] If the exchange rate is expected to remain constant ($\Delta_m s^e_{t+m} / m = 0$) and α > 0, the domestic interest rate $i_{t,m}$ must exceed the foreign currency interest $i^*_{t,m}$ in order for UIP to hold.

A range of empirical studies have subsequently examined the UIP hypothesis using different currency and time samples and different econometric methods. Froot & Thaler (1990) survey 75 published estimates and conclude that the strict version of the UIP hypothesis is rejected in almost all cases. Similar conclusions have been reached in other subsequent survey papers (e.g. Engel 1996, Alexius 2001). The consistent finding that the estimated slope coefficient is far below one and often negative has been labelled the *forward premium anomaly* (Froot & Thaler 1990, Booth & Longworth 1986, Olmo & Pilbeam 2011).

Most studies are based on data with investment horizons of one month, three months or six months as such data are readily available. Studies suggest, however, that the UIP may hold better at longer investment horizons. Chinn & Meredith (2004) study the empirical validity of the UIP hypothesis for the currencies of the G7 countries using a sample from 1983 to 2000. For short investment horizons, the UIP is rejected in all cases, but when the UIP regression is estimated using 5 or 10 year horizons, the slope coefficient is always positive and in many cases not statistically different from one.[8] Qualitatively similar results are obtained by Alexius (2001) and Mehl & Cappiello (2007) although the UIP hypothesis is still rejected for some countries.

The time sample also seems to be of importance, which is unsurprising given that financial markets and regulatory schemes change over time. Lothiana & Wu (2011) use a sample of 200 years and consider the UIP hypothesis between the dollar and sterling and between the franc and sterling. They find that the slope estimate β typically is positive although far from one until 1980, but then turns negative for most periods after that. It is argued that the limited support for the UIP hypothesis is the result of expectations that ex-post are wrong for extended periods of time. Flood & Rose (2002) reach different conclusions using data from the 1990s and a broad sample of high-income and emerging economies. Estimation of standard UIP regressions leads to the conclusion that the hypothesis received more support from their data from the 1990s than from earlier data, although the overall conclusion is still negative as spelled out in the title: "Uncovered interest parity in crisis".

Baillie & Bollerslev (2000) suggest that the forward premium anomaly can, at least partly, be explained by the different time series properties of the variables in the standard UIP regression. The relative exchange rate change ($\Delta_m s_{t+m} / m$) is close to a random walk (at least at relatively high frequencies), while the interest rate spread ($i_{t,m} - i^*_{t,m}$) typically exhibits substantial persistence (but not a unit root). Baillie & Bollerslev (2000) simulate data based on these characteristics and show that the resulting slope, although centred around one, exhibits a very high variance. The upshot is that estimations with relatively few observations are likely to produce

[8] The finding that the UIP hypothesis generally holds better for long investment horizons than for short horizons can be related to the *peso problem* (Froot & Thaler 1990). In this context, the peso problem implies that adjustments of the exchange rate to the UIP may occur in discrete and infrequent steps of substantial magnitude.

coefficient estimates that are sensitive to sample changes and that may differ significantly from one even if the UIP is in fact satisfied.

It is typically found that the UIP holds better for cases where the interest rate spread is substantial and less well for cases where the interest rate spread is small. Mehl & Cappiello (2007) find that UIP relations estimated for some high-income and emerging market economies exhibit non-linearities. They estimate a smooth transition regression implying different marginal effects of the interest rate spread when the interest rate spread is small and when it is large. The upshot is that the standard linear model mixes the effects of different regimes. Using data for selected European currencies, Lothiana & Wu (2011) find more support for the UIP hypothesis in periods in which the interest rate spread is large. This result seems intuitively reasonable as factors such as risks and transaction costs may not warrant arbitrage trading if the returns from such trades are limited (Froot & Thaler 1990).

Alper et al. (2009) survey the literature on UIP testing in emerging market economies. On the one hand, the high trend inflation observed in many emerging markets facilitates the forecasting of exchange rate developments and therefore makes it more likely that the UIP hypothesis does hold. On the other hand, structural breaks and uncertainties are likely to be more pronounced in emerging markets, which would suggest that the UIP does not hold. Empirical studies confirm that UIP estimations frequently exhibit different properties for emerging markets and for high-income economies. Alper et al. (2009, p. 123) conclude that "…identifying and modelling structural breaks provide room for improvement for further research on the UIP condition for [emerging markets]". Bansal & Dahlquist (2000) provide an explicit comparison of results for high-income and emerging market economies and conclude that the UIP is more likely to hold for emerging markets than for high-income economies. Different per capita GNP, average inflation and inflation volatility are factors that may explain the different results.

Only a small number of studies have examined the empirical validity of the UIP hypothesis for countries in Central and Eastern Europe. Brasili & Sitzia (2003) estimate panel models based on CEE data in which future exchange rate changes are explained by the interest rate spread and a range of other factors that may be considered proxies of the risk premium. The spread is not statistically significant in a specification in which it enters linearly, but a non-linear transformation of the spread attains statistical significance, suggesting that non-linearities play an important role. Ho & Ariff (2009) also use a panel explaining the future exchange rate change with many variables along with the interest rate spread. A range of specifications all produce positive and statistically significant coefficients to the interest rate spread for the sample of Eastern European countries, but the coefficients vary substantially across different specifications. The use of panel data in these two studies precludes the estimation of country-specific coefficients of the interest rate spread.

Mansori (2003) compares results for the Czech Republic, Hungary and Poland from 1994 to 2002 with results for a number of West European countries. There is more support for the UIP hypothesis for the three East European countries, especially the Czech Republic and Hungary, than for the West European countries. The results for the CEE countries are however very sensitive to changes in the time sample, possibly as a result of the convergence processes underway during the period analysed. Horobet et al. (2009, 2010) estimate standard UIP regressions for eight countries, including four from Central and Eastern Europe using monthly data from 2006 to 2009. The estimated slope coefficients are positive in all cases, but neither economically nor statistically different from zero. This result seems to hold whether or not exchange market volatility is taken into account.

4. Data and unit root tests

This section provides an overview of the dataset and the main features of the series for the five sample countries, Croatia, the Czech Republic, Hungary, Poland and Romania. The samples vary across the five countries but generally span a bit more than a decade, starting in 1999 and ending in September 2011. The five countries all had floating exchange rates during this period, although Poland formally used managed devaluations until April 2000 and Hungary used different corridors until 2008.[9]

The analyses are undertaken for positions with a 3-month horizon, implying that the returns from the currency exposure and the interest rate differential are both calculated for a 3-month holding period. As discussed in the literature survey in Section 3, the results may vary with the investment horizon, but the 3-month horizon has been chosen because the 3-month money market is one of the most liquid segments of the market.

The five countries saw increased integration with Western Europe, and in particular with the euro area, during the sample period. The reference area is therefore taken to be the euro area: the exchange rates are in units of local currency per euro and the interest rate spreads of the local interest rate are against the Euribor rate. It is noticeable that the countries considered here were at different stages of their processes of convergence with Western Europe during the sample period.[10]

[9] The Hungarian bands changed frequently before they were finally removed in February 2008. Until May 2001, the managed devaluation was based on a "daily rate of devaluation" against, in 1999, a basket (30 percent USD, 70 percent EUR) and, thereafter, the euro. The band around the central rate of the devaluation path was +/– 2.25 percent. From May to October 2001 the band around the central rate was increased to +/– 15 percent. From October 2001 the central parity was fixed at 276.1 HUF/EUR and in June 2003 to 282.36 HUF/EUR, while the band remained at +/– 15 percent.

[10] For an overview of the stages of convergence, see the European Commission (2010a, 2010b). Different indicators can be used to assess the degree of convergence of the CEE countries with Western Europe. European Commission (2010a, 2010b) asserts that the convergence process in Romania and Croatia has been slower than that in the other three CEE countries in our sample.

Most of the estimations are based on only two variables, cf. eq. (5).[11] The variable FX_CHG is the percentage change of the spot exchange rate over a 3-month period, where the exchange rate denotes units of local currency per euro at the end of month. A positive value of FX_CHG indicates a depreciation of the local currency against the euro over the 3-month period; a negative value indicates an appreciation. The variable INT_SP is the annualised interest spread between a 3-month domestic currency deposit and the 3-month Euribor.

The available sample of data varies across the countries. For Croatia, the series on the nominal exchange rate starts in November 1999, implying that the 3-month FX_CHG variable starts in February 2000. For Poland, the local 3-month interest rate is available from the beginning of 2001. Table 1 reports summary statistics of the exchange rate changes and the interest rate spreads for the five sample countries.

Table 1. Descriptive statistics for 3-month exchange rate change and 3-month interest rate spread

FX_CHG	**Mean**	**Median**	**Max.**	**Min.**	**Std. Dev.**	**Obs.**
Croatia	-0.20	-0.51	17.09	-20.97	6.46	140
Czech Republic	-2.94	-4.24	60.48	-23.00	12.15	153
Hungary	1.99	2.04	63.03	-47.54	18.54	153
Poland	2.39	-0.97	98.36	-37.77	25.06	129
Romania	9.26	6.90	76.87	-32.82	21.12	153

INT_SP	**Mean**	**Median**	**Max.**	**Min.**	**Std. Dev.**	**Obs.**
Croatia	3.30	2.74	11.05	-0.05	2.50	140
Czech Republic	0.36	0.15	5.04	-1.35	1.25	153
Hungary	6.19	5.71	12.97	2.66	2.52	153
Poland	3.70	3.27	13.03	0.66	2.62	129
Romania	22.75	13.00	145.07	2.38	26.58	153

Figure 1 depicts the nominal exchange rate of each Eastern European country against the euro from the beginning of 1999 and until December 2011. The first thing to notice is that the exchange rate dynamics vary considerably across the five sample countries. The currencies of Croatia and the Czech Republic have tended to appreciate against the euro, while the currency of Romania has tended to depreciate. The currencies of Hungary and Poland have been relatively stable with exchange rates fluctuating around a relatively constant level.

[11] The variables are calculated based on Ecowin source data.

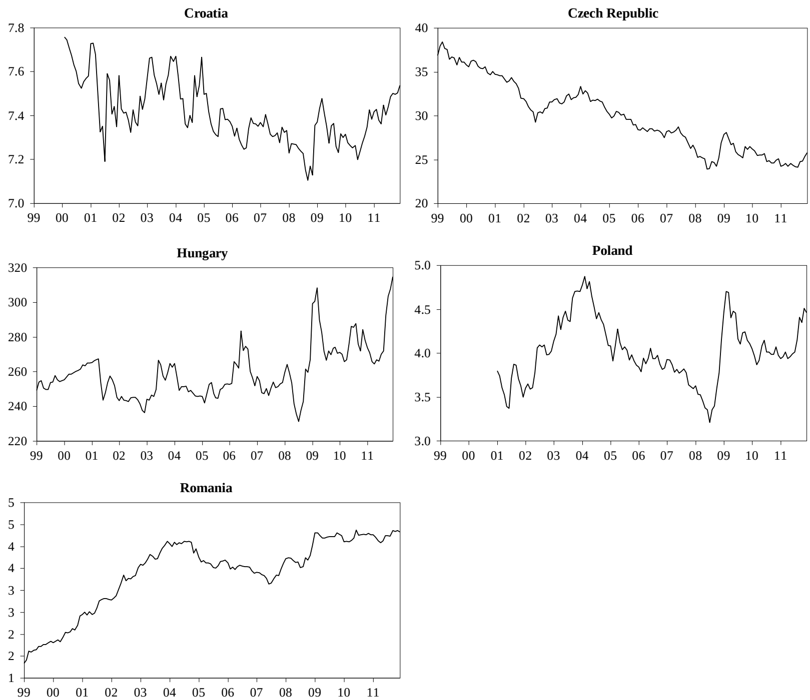

Figure 1. Nominal exchange rate of local currency against euro.

The different exchange rate development across the sample countries is the result of many factors. The process of integration into EU structures, and the associated confidence effects, has affected the exchange rate dynamics in the Central and Eastern European countries. The speed of and commitment to integration has differed across the countries.[12] The main message for our analyses is that there is no "Central and Eastern European block" with closely co-moving exchange rates; the exchange rate developments are fundamentally different across the five sample countries.

Figure 2 depicts the 3-month annualised change of the exchange rate against the euro. The series are very volatile, which suggests that, for the UIP to hold, the interest rate differential between the country and the euro area would also have to be volatile.

[12] The Romanian case is noticeable because the period from 2003 to 2005 represents a political and economic regime switch. During this period Romania joined the Council of Europe and the WTO, and became an associated member of the European Union. These steps were part of the process of stabilising the political and economic situation in the country, and helped to increase the confidence of financial markets in the Romanian economy (European Commission 2010a).

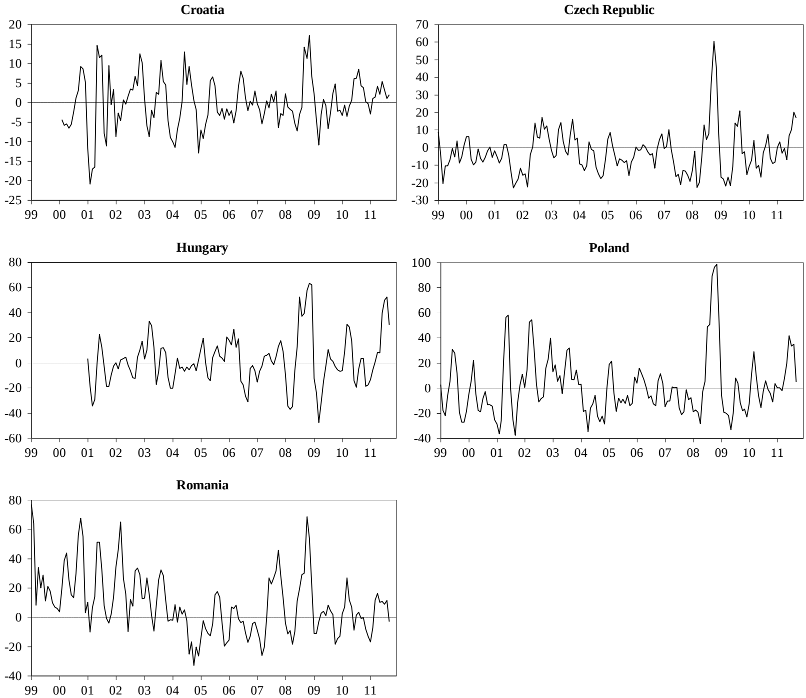

Figure 2. Annualised changes of local currency versus euro over 3-month period, %.

Figure 3 reports the spread between the local 3-month interbank interest rate and the 3-month Euribor. The volatility of the interest rates spread is much smaller than the volatility of the foreign exchange rate changes on the same horizon.

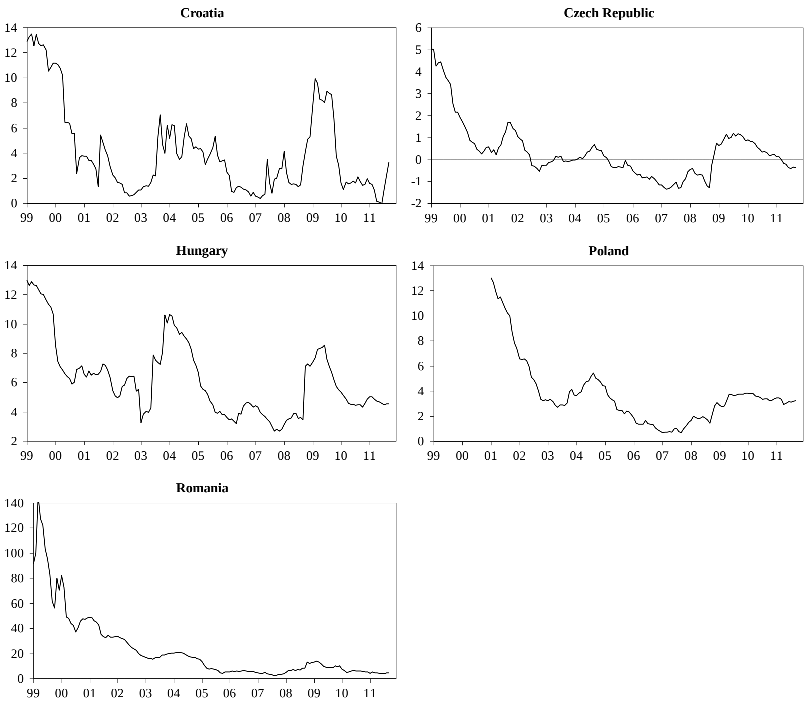

Figure 3. Annualised interest rate spreads on 3-month deposits, %.

The time series properties of the exchange rate changes and the interest rate spreads have been examined by means of Augmented Dickey-Fuller tests. Given that the variables are either changes in percentage terms (for currency pairs) or spreads (interest rates), the test is performed at the level of the variables and an intercept, but no time trend, is included in the estimations. The number of lags used is chosen by means of the Schwartz selection criterion. The results are reported in Table 2. The hypothesis of a unit root can be rejected in all cases; the series are I(0) for all five sample countries.

Table 2. Augmented Dickey-Fuller unit root tests

FX_CHG	1% C.V.	5% C.V.	10% C.V.	Statistic	Prob.	Process
Croatia	-3.479	-2.883	-2.578	-7.831	0.000	I(0)
Czech Republic	-3.475	-2.881	-2.577	-5.225	0.000	I(0)
Hungary	-3.475	-2.881	-2.577	-6.969	0.000	I(0)
Poland	-3.482	-2.884	-2.579	-5.161	0.000	I(0)
Romania	-3.475	-2.881	-2.577	-4.495	0.000	I(0)
INT_SP	**1% C.V.**	**5% C.V.**	**10% C.V.**	**Statistic**	**Prob.**	**Process**
Croatia	-3.477	-2.882	-2.578	-3.476	0.010	I(0)
Czech Republic	-3.474	-2.880	-2.577	-3.767	0.004	I(0)
Hungary	-3.473	-2.880	-2.577	-2.745	0.069	I(0)
Poland	-3.482	-2.884	-2.579	-4.352	0.001	I(0)
Romania	-3.477	-2.882	-2.578	-3.963	0.002	I(0)

Note: C.V. denotes critical value.

5. Uncovered interest parity

We start by rewriting eq. (5) using our empirical notation in which a bracket after the variable name is used to indicate a time shift (in month) of the variable:

$$FX\_CHG(3) = \alpha + \beta \cdot INT\_SP + \varepsilon(3) \tag{6}$$

Eq. (6) is estimated for each country individually using OLS. The results are reported in Table 3. The choice of a 3-month investment horizon but monthly data leads to first- and second order-autocorrelation of the residuals. We therefore report Newey-West robust standard errors. The strict version of the UIP holds if $\alpha = 0$ and $\beta = 1$ and the residuals do not exhibit serial correlation of the third or a higher order. The table reports the F-statistics for the Wald test of the joint hypothesis $\alpha = 0$ and $\beta = 1$. Examination of the residuals reveals the existence of autocorrelation of first and sometimes second order, but never of higher orders.

The estimation results reveal that the coefficients of determination, R^2, of all the regressions are extremely low. This is not surprising in light of Figures 2 and 3 and is found in all tests of the UIP hypothesis (Flood 1996). The foreign exchange return is much more volatile than the interest rate spread, which limits the ability of the interest rate spread to explain the foreign exchange change.

Table 3. UIP estimation results (OLS)

	$\hat{\alpha}$	$\hat{\beta}$	F-stat	R^2	Sample	Obs.
Croatia	1.401 (0.888)	-0.486** (0.210)	31.660 [0.000]	0.035	2000:02-2011:09	140
Czech Republic	-2.447 (1.718)	-1.380 (0.972)	9.492 [0.000]	0.020	1999:01-2011:09	153
Hungary	9.546* (5.706)	-1.220* (0.711)	10.120 [0.000]	0.028	1999:01-2011:09	153
Poland	3.658 (6.479)	-0.342 (1.319)	0.642 [0.528]	0.001	2001:01-2011:09	123
Romania	2.023 (3.290)	0.308*** (0.087)	47.944 [0.000]	0.148	1999:01-2011:09	153

Note: Newey-West standard errors are shown in round brackets. Superscripts ***, **, * denote that the coefficient estimate is statistically different from 0 at the 1, 5 and 10% level of significance respectively. The null hypothesis of the F-test is that $\alpha = 0$ and $\beta = 1$; the *p*-value is shown in square brackets.

The estimated slope coefficients in Table 3 are different from 1 at the 1% level of significance for all five sample countries. For all countries except Romania, the coefficients are also negative, which is in accordance with the *forward premium anomaly* found in many other studies (cf. Section 3). For Romania, the estimated coefficient is positive and significantly different from zero (but also significantly different from one). This would be consistent with the finding that the UIP hypothesis is more likely to hold when the interest rates spread is large (Froot & Thaler 1990, Mehl & Cappiello 2007, Lothiana & Wu 2011). It follows from Figure 3 that the spread between the Romanian 3-months interest rate and the 3-months Euribor rate was in the double digits until 2005 and also afterwards remained much higher than for the other sample countries. The large interest spread reflects that Romania has experienced a more prolonged convergence process the other sample countries.

The estimated constant terms are, with the exception of the Czech Republic, positive, but statistically significantly different from 0 only for one country. As already noted, this coefficient should indicate the presence of either a risk premium or barriers to entry. While it is probable that barriers to entry or other parts of the regulatory landscape do not change very often, previous research and anecdotal evidence (again, from the recent financial crisis) indicates that the risk premium varies across time and economic cycles, and therefore to model them as a constant would be to impose a tight constraint on the model.[13]

[13] The residuals generally exhibit some heteroskedasticity. To assess the impact, we estimated eq. (6) using a GARCH specification. Although the GARCH coefficients are statistically significant in many cases, the effects on the estimated α and β and the explanatory power of the regressions are modest.

The F-statistics reported in Table 3 shows that Poland is the only country for which the null hypothesis cannot be rejected. The Polish case is predicated by the fact that the standard errors of the two coefficient estimates are very high for this country. For all other countries in the sample, the joint hypothesis that α and β take values in accordance with the UIP is rejected.

6. Uncovered interest parity across time

The test of the UIP in Section 5 is undertaken on the entire available time sample from the turn of the century to September 2011. The recent global financial crisis has, however, provoked very sharp reactions in *inter alia* foreign exchange and interest markets. Eastern European countries largely escaped the first part of the crisis (the "sub-prime" phase from summer 2007), but the default of Lehman Brothers in September 2008 affected the region greatly. This is also shown by Figures 1 and 3, in which sudden depreciations of the currencies against the euro and a jump in the spreads between local interest rates and the Euribor are evident.

In order to shed further light on the impact on the UIP of the global financial crisis, and more generally to shed light on the time dimension, we undertake rolling windows estimations with samples of monthly observations for five years. The estimations are based on eq. (6), i.e. the simple linear version of the UIP. Figure 4 shows the coefficient of determination, while Figures 5 and 6 show the estimated constants and slope coefficients for the five countries. For all three figures, the date reported on the horizontal axis indicates the *end* of the sample.

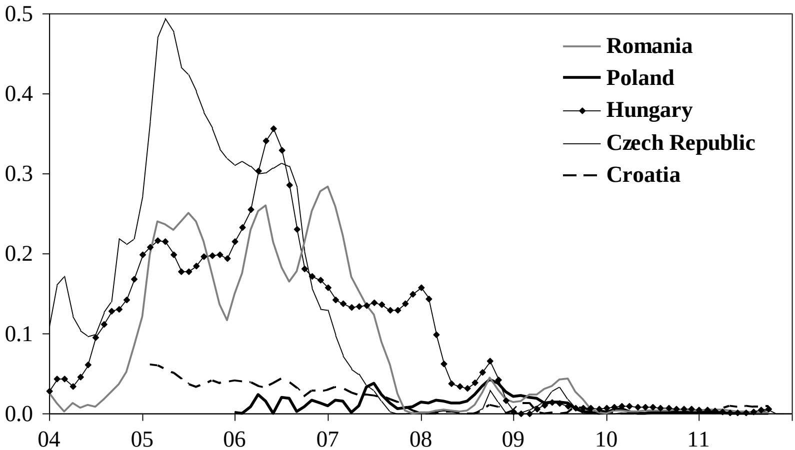

Figure 4. Coefficient of determination, 5-year rolling windows.

Figure 4 reveals that the explanatory power of the regressions is always very low for Poland and Croatia, but relatively high before the crisis for the three other countries.

This could be an indication that Poland and Croatia may have been more "closed" or insulated from external influences than the other three countries in the sample (Jevcak et al. 2011). Moreover, when the windows consist largely of the period around the global financial crisis, the simple UIP specification (without crisis indicators and with fixed coefficients) basically has no explanatory power for the five sample countries.

Further insights into developments before and after the global financial crisis hit the region can be gained from Figures 5 and 6. The coefficient estimate and +/– 2 times the Newey-West standard errors are depicted in each figure. The estimated constants and slopes for all the sample countries display extreme variation. This could be due to the relatively short span of the sample (five years for each rolling regression), or to an inherent instability in the relation between interest rate spreads and currency returns (Baillie & Bollerslev 2000).

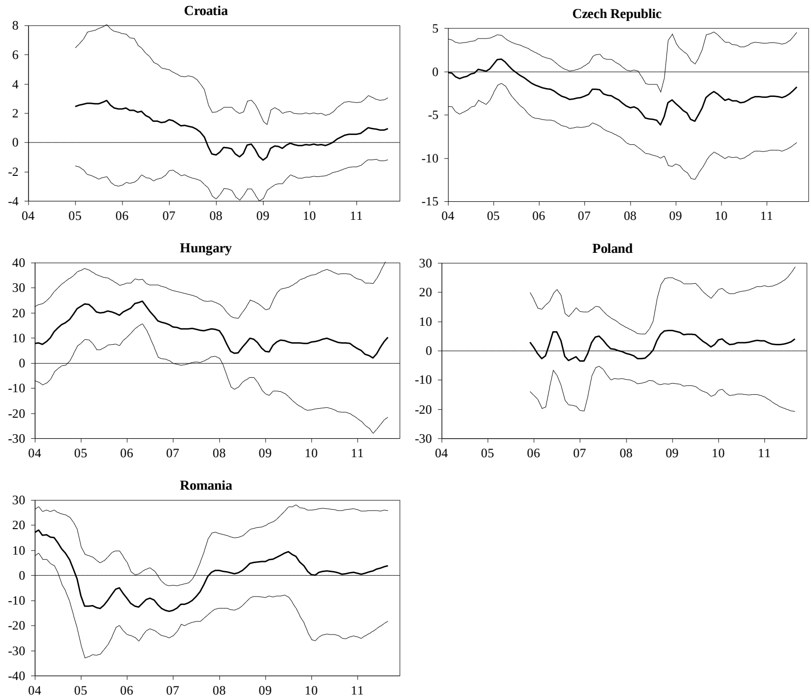

Figure 5. Estimated constants, 5-year rolling windows.

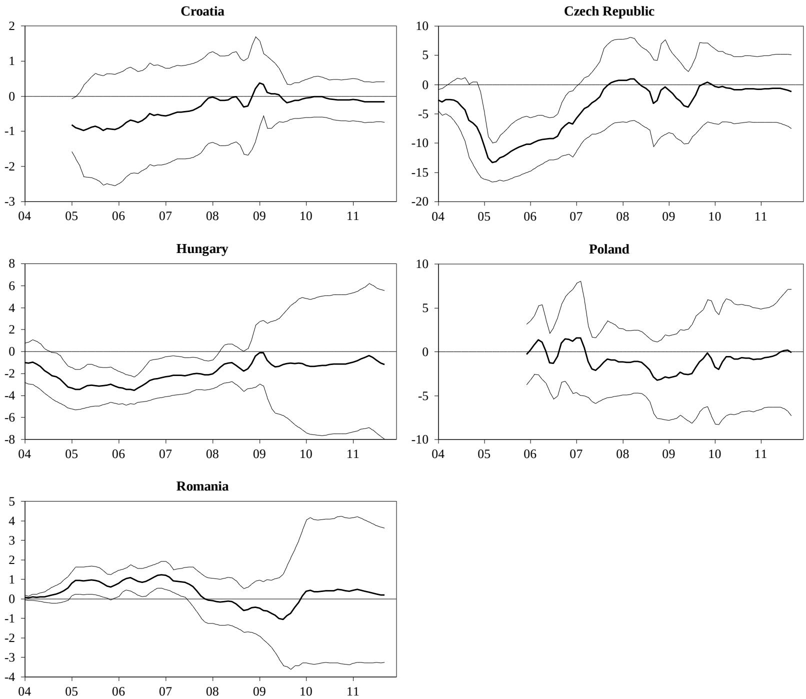

Figure 6. Estimated slope coefficients, 5-year rolling windows.

The UIP specifications exhibit some explanatory power for the Czech Republic, Hungary and Romania in the pre-crisis period. For the Czech Republic the constant was close to zero and the slope was negative. The absolute value of the slope estimate is extremely large when the period 2000-2001 is included in the sample; this was a period in which the Czech koruna appreciated rapidly. For Hungary the slope estimate is also negative (below -1), while the constant is positive. For Romania the slope is positive and the constant is negative. Moreover, the slope is close to one for all of the period before 2007 but turned negative later. This suggests that the UIP was satisfied in the transition period when the interest spread was very high, but not in later periods when the spread was reduced.

The conclusion from the estimations in Sections 5 and 6 is that the UIP has limited empirical validity in the sample of CEE countries. Still, there are noticeable differences across the sample countries and across different time samples. The rest of the paper examines a number of possible reasons for these findings. Transaction costs may limit arbitrage when the interest rate spread is small (Section 7) and the risk premium may be time-varying (Section 8).

7. Non-linearities

The size of the interest rate spread may affect whether or not the UIP hypothesis is supported. Transaction and information costs are likely to keep investors from exploiting deviations from the UIP when the interest rate spread is small, but not when the spread is high (Froot & Thaler 1990). The conjecture has some empirical support (Mehl & Cappiella 2007, Lothiana & Wu 2011).

The extreme volatility of the FX_CHG variable has made us pursue a simple and robust way to model the presence of different regimes for different levels of interest rate spreads. We separate the interest spread into two series. Taking the average spread over the sample for each country, two series of interest rate spreads are computed: the variable INT_SP_LO equals the spread when the spread is lower than the average, and zero otherwise; the variable INT_SP_HI equals the spread when the spread is higher than the average, and zero otherwise. Both spread variables are included in the UIP specification:

$$\text{FX\_CHG}(3) = \alpha + \beta^{LO}\cdot\text{INT\_SP\_LO} + \beta^{HI}\cdot\text{INT\_SP\_HI} + \varepsilon(3) \tag{7}$$

The results of the regressions are reported in Table 4. The results are as expected for Poland and Romania; the slope coefficients for high interest rate spreads are in both cases positive and statistically different from zero, while the coefficients for low spreads are statistically insignificant. The results are inconclusive for the other three countries; the slope coefficients of the high interest rate spreads are negative and the coefficients are generally estimated imprecisely. Overall, Table 4 provides some support to the hypothesis that the UIP should hold better when the interest rate spread is large than when it is low, at least for Poland and Romania.

Table 4. UIP estimation results, high and low interest rate spread variables

	$\hat{\alpha}$	$\hat{\beta}^{LO}$	$\hat{\beta}^{HI}$	F-stat	R^2	Sample	Obs.
Croatia	2.181 (1.355)	-0.969 (0.729)	-0.553** (0.225)	21.459 [0.000]	0.041	2000:02- 2011:09	140
Czech Republic	-1.905 (1.680)	0.107 (3.328)	-1.743* (0.955)	6.195 [0.000]	0.023	1999:01- 2011:09	153
Hungary	8.979 (9.084)	-1.073 (1.894)	-1.163 (0.934)	7.543 [0.000]	0.028	1999:01- 2011:09	153
Poland	1.445 (5.111)	0.221 (0.497)	0.464 (0.221)	4.936 [0.002]	0.080	2001:01- 2011:09	129
Romania	5.790 (4.523)	-0.113 (0.462)	0.266*** (0.089)	34.744 [0.000]	0.156	1999:01- 2011:09	153

Notes: OLS estimation. Newey-West standard errors are shown in round brackets. Superscripts ***, **, * denote that the coefficient estimate is statistically different from 0 at the 1, 5 and 10% level of significance respectively. The null hypothesis of the F-test is that $\alpha = 0$, $\beta^L = 1$and $\beta^H = 1$; the *p*-value is shown in square brackets.

We have also implemented two other specifications of the non-linear relation from the interest spread to the foreign exchange rate change (results not shown). One approach was the smooth transition model of Granger & Teräsvirta (1993), but we generally had problems estimating the non-linear relation. Another approach was to use a Taylor order approximation up to the third order of the Granger & Teräsvirta model and then to estimate coefficients to all the included powers. In many cases the estimated coefficients attained implausible sign and size and the R^2 of the regressions did not change from the base case (results not shown). In conclusion, non-linearities seem to play only a minor role for the UIP estimations, i.e. transaction and information costs are unlikely to be behind the weak support of the UIP for the CEE countries.

8. Risk aversion and financial instability

A possible explanation for the low explanatory power of the UIP estimations is that the risk premium is in fact not constant. We include different proxies of the risk premium.

We start by including the VIX index as a proxy of the risk premium. The VIX index is an implied volatility index calculated from option prices on the S&P500 equity index and is often seen as a main indicator of risk aversion in global financial markets. A higher value of the VIX index is tantamount to larger financial uncertainty. We include VIX as an additional explanatory factor in the empirical UIP specification:

$$FX\_CHG(3) = \alpha + \beta \cdot INT\_SP + \gamma \cdot VIX + \varepsilon(3) \quad (8)$$

The results are reported in Table 5. While the R^2 of the estimations do not improve markedly, the coefficient of VIX is positive for all the countries and also statistically significant for Croatia and Romania. More financial instability in global financial markets puts *ceteris paribus* depreciation pressure on the local currency. The slope coefficients stay largely unchanged, while the constants change sign for three countries, becoming (with the exception of Hungary) negative, but mostly not significant. This suggests that when global risk aversion is taken into account, the time-invariant remaining part captured by the constant loses its explanatory power.

Table 5. UIP estimation results, including VIX

	$\hat{\alpha}$	$\hat{\beta}$	$\hat{\gamma}$	F-stat	R^2	Sample	Obs.
Croatia	-2.125 (1.912)	-0.65^{***} (0.230)	0.185^{**} (0.094)	25.946 [0.000]	0.096	2000:02- 2011:09	140
Czech Republic	-13.048^{**} (6.133)	-2.176^{*} (1.232)	0.488 (0.320)	3.746 [0.026]	0.131	1999:01- 2011:09	153
Hungary	2.585 (8.785)	-1.439^{*} (0.757)	0.373 (0.430)	5.580 [0.005]	0.056	1999:01- 2011:09	153
Poland	-11.250 (9.412)	-0.755 (1.488)	0.748 (0.614)	1.156 [0.318]	0.075	2001:01- 2011:09	129
Romania	-11.151^{*} (6.385)	0.271^{***} (0.082)	0.639^{**} (0.294)	40.687 [0.000]	0.210	1999:01- 2011:09	153

Notes: OLS estimation. Newey-West standard errors are shown in round brackets. Superscripts ***, **, * denote that the coefficient estimate is statistically different from 0 at the 1, 5 and 10% level of significance respectively. The null hypothesis of the F-test is that α = 0 and β = 1; the *p*-value is shown in square brackets.

An alternative measure of risk aversion, less global and more linked to European foreign exchange markets, may be based on other currency pairs in the region. As a rough measure of the external risk aversion affecting currency markets in Europe, we use the 3-month return of the Swedish *krona* against the euro. Sweden had a floating exchange rate throughout the sample period and the exchange rate is likely be affected by currency market pressures. The estimated equation is the following, where SWE_FX_CHG denotes the annualised 3-month depreciation of the Swedish krona against the euro:

$$\text{FX\_CHG}(3) = \alpha + \beta\cdot\text{INT\_SP} + \delta\cdot\text{SWE\_FX\_CHG}(3) + \varepsilon(3) \tag{9}$$

The results are reported in Table 6. The R^2 are higher and the coefficients of the Swedish krona return are always statistically significant (with the exception of the results for Croatia) and have positive signs. It seems that including the currency pressure on the Swedish krona gives the same overall result as was given when the VIX variable were included, but in an arguably stronger way. Unlike in the equation with VIX, the constants become insignificant, with the exception of the one for the Czech Republic, where the constant is still significant and negative.

Table 6. UIP estimation results, including change in Swedish krona foreign exchange rate

	$\hat{\alpha}$	$\hat{\beta}$	$\hat{\delta}$	F-stat	R^2	Sample	Obs.
Croatia	1.239 (0.915)	-0.462* (0.221)	0.094 (0.071)	30.880 [0.000]	0.064	2000:02- 2011:09	140
Czech Republic	-3.005** (1.449)	0.147 (0.846)	0.484*** (0.173)	5.042 [0.008]	0.211	1999:01- 2011:09	153
Hungary	6.714 (5.655)	-0.787 (0.754)	0.601*** (0.211)	6.992 [0.001]	0.161	1999:01- 2011:09	153
Poland	3.317 (5.075)	-0.304 (1.168)	1.199*** (0.312)	0.736 [0.481]	0.310	2001:01- 2011:09	129
Romania	1.679 (2.758)	0.324*** (0.068)	0.807*** (0.129)	77.248 [0.000]	0.334	1999:01- 2011:09	153

Notes: OLS estimation. Newey-West standard errors are shown in round brackets. Superscripts ***, **, * denote that the coefficient estimate is statistically different from 0 at the 1, 5 and 10% level of significance respectively. The null hypothesis of the F-test is that $\alpha = 0$ and $\beta = 1$; the *p*-value is shown in square brackets.

Concluding this section, the two indicators of risk aversion in international financial markets seem to exhibit substantial explanatory power. The estimated coefficients attain the expected sign and are statistically significant in many cases. The addition of these risk aversion measures, however, does not change the conclusions about the estimated slope coefficient, but has, as expected, an impact on the constant term, which becomes statistically insignificant.[14]

9. Summary

This paper presented the results of empirical tests of uncovered interest parity in Croatia, the Czech Republic, Hungary, Poland and Romania during the first decade of the 21st century. The objective was to examine whether the UIP would obtain empirical support in this particular sample, and to ascertain to which extent the convergence process and the global financial crisis have affected the UIP relation.

We proceeded from simple estimations of the link between the return on 3-month exposure to local currencies against the euro and the spread between local interest rates and Euribor. The stability of the estimated parameters was analysed using rolling windows. The analysis examined the importance of a number of issues that may affect the results. Estimations took into account the possibility of different regimes depending on the size of the interest rate spread. Various indicators of risk and risk aversion were included, chiefly to capture the effect of the global financial crisis. The main results are summarised below.

[14] For the Czech Republic, Hungary and Poland we tried to use the Exchange Market Pressure (EMP) index in Filipozzi & Harkmann (2010). The coefficients of the EMP index were not statistically significant (not reported).

The basic model used to test the UIP in the CEE countries gave a result in line with most of the previous literature, namely that the UIP relation cannot be supported in general. The forward premium anomaly is confirmed in the present sample of Central and Eastern European countries; the estimated slope coefficient is negative in all cases except Romania.

Rolling window regressions showed that the coefficient estimates generally are unstable and depend on the choice of sample. The rolling regressions also cast some light on the effect of global financial crisis on the UIP relations in the five CEE countries. At least for the Czech Republic, Hungary and Romania, there is a clear change after the crisis as the explanatory power of the UIP regressions drops dramatically after 2007.

Transaction and information costs do not seem to affect the UIP estimations in ways which can be clearly discerned through the inclusion of non-linearities in the UIP relation. It is clear, however, that the importance of the interest rate spread varies between low and high interest rate spread regimes, but the picture is not uniform across the sample countries. For Poland and Romania, the slope coefficient is positive when the interest rate spread is large, although the estimate is still statistically different from one.

There is substantial evidence suggesting that the risk premium is not constant. Both the global volatility index VIX and the movements in the Swedish exchange rate seem to exhibit substantial explanatory power although not symmetrically across all five countries. This suggests that global risk factors have considerable impact on the liquidity of financial markets and the arbitrage processes underlying the UIP in the five countries from Central and Eastern Europe.

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