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Do foreign exchange forecasters believe in Uncovered Interest Parity?[☆]

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HIGHLIGHTS

- UIP is generally rejected in empirical work.
- We analyse whether Consensus Forecasts are based on UIP in a group of Central and Eastern European countries.
- We use econometric methods which account for breaks.

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ABSTRACT

Uncovered Interest Parity (UIP) is typically rejected in empirical studies, but this letter finds nevertheless that Consensus Forecasts of the exchange rate for Central and Eastern European countries are based on UIP. When structural breaks are included, the forecasts are found to deviate from UIP in 2008–09 when financial markets were under severe stress.

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1. Introduction

The hypothesis of Uncovered Interest Parity (UIP) is one of the pillars of international finance. The UIP hypothesis is derived from arbitrage principles and posits that a country with a higher interest rate than that abroad is expected to see a weakening of its currency. The UIP hypothesis is almost uniformly rejected in empirical studies; countries with higher interest rates do not generally witness a weaker currency and often the opposite is found, this being the *forward premium puzzle* (Engel, 2014).

Exchange rates are important for trade, finance, etc. and expert forecasts of future nominal exchange rates are plentiful. The literature finds that expert forecasts of the exchange rate typically are biased and do not outperform a simple random walk model (MacDonald and Marsh, 1994; MacDonald, 2002; Mitchell and Pearce, 2007). This raises the issue of how these expert forecasts are arrived at, an area where there is limited empirical evidence.

Frankel and Froot (1987) conclude that expert forecasts of the US dollar against major currencies depend on lagged forecasts, the lagged realised spot rate, and a measure of the long-term equilibrium spot rate. Schröder and Dornaub (2002) find that forecasts of the exchange rate between major economies are in large part informed by expectations of GDP developments and the interest rate differential, but the latter factor enters with different signs for different currency pairs. Haunera et al. (2014)

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consider the exchange rate between the US dollar and more than 50 other currencies. Their panel data estimations reveal that expert forecasts are informed by inflation and productivity differentials, but generally not by interest rate differentials.

There are no country-specific studies investigating expert forecasts of the exchange rate for smaller European countries, including countries from Central and Eastern Europe. Arguably the most important expert forecasts of the exchange rate are the Consensus Forecasts (CF) published by Consensus Economics. These have been available for a number of Central and Eastern European countries since 2007. Each month a number of professional forecasters provide their forecasts of the exchange rate one month ahead, and the means of these forecasts are published as the CF forecasts. This letter uses the CF forecasts to examine whether forecasters believe in the UIP hypothesis or, more precisely, to what extent expected changes in the exchange rate computed from CF forecasts depend on interest rate differentials. The letter also examines whether the forecasting behaviour changes over time by including specifications with an endogenous determination of structural breaks.

2. Full sample

The countries in the sample are the Czech Republic, Croatia, Hungary, Poland and Romania. Data are monthly from 2007:05 to 2014:10.

Eastern European Consensus Forecasts publishes forecasts for the exchange rate one month ahead in the week containing the third Monday of the month. The CF forecast is computed as a simple average of a large number of expert forecasts submitted on the third Monday of the month (or occasionally at the end of the preceding week). The CF forecast of the exchange rate of the local currency against the euro one month ahead is labelled $s^{CF}_{t,t+1}$. Consensus Forecasts also publishes the actual exchange rate as of the third Monday of the month and this spot exchange rate is labelled s_t.

The interest rate data are sourced from Ecowin. In order to ensure that the interest rate records are the latest known to the forecasters when they submit their forecasts, we use the interest rates published for the Friday before the third Monday of the month. The local-currency one-month interbank deposit interest rate is denoted i_t and the corresponding euro area (EA) interest rate i^{EA}_t. The interest rates are recalculated to denote the return *per month*.

As a first step the actual exchange rate depreciation was regressed on the CF forecast of the depreciation. The results (not reported) were in all cases a statistically insignificant slope parameter close to 0, suggesting that the CF forecasts have very little explanatory power in the present sample. This leaves the question of whether there is a pattern in the way the CF forecasts have been produced.

The UIP hypothesis posits that the expected exchange rate depreciation equals the interest rate differential plus a risk premium that may be constant or time varying. A test of the hypothesis using CF forecasts, a constant risk premium and a one month horizon can be based on this specification:

$$\frac{s^{CF}_{t,t+1} - s_t}{s_t} = \alpha + \beta(i_t - i^{EA}_t) + \varepsilon_t. \tag{1}$$

The left-most term is the CF forecast of the rate of depreciation of the exchange rate. To the right, α denotes the risk premium in percentage points per month. A negative value of α signifies that the investors expect or demand a higher return for investments in the country considered than for investments in the EA. This may result from illiquid financial markets or other sources of risks associated with investment in the country. The parameter β captures the effect of the interest rate spread on the CF forecast of the exchange rate depreciation; an estimate for β of around 1 would suggest that the CF forecast has been informed by the UIP hypothesis. Finally, ε_t is an error term.

Table 1
Estimation of Eq. (1) for the full sample.

	α	β
Czech Rep.	-0.05	2.94^{**}
	(0.17)	(1.22)
$Adj.\ R^2 = 0.050,\ DW = 1.71,\ White(p\text{-}value) = 0.130$		
Croatia	-0.05	0.90^{***}
	(0.05)	(0.19)
$Adj.\ R^2 = 0.180,\ DW = 2.00,\ White(p\text{-}value) = 0.094$		
Hungary	-0.22	1.29
	(0.50)	(1.16)
$Adj.\ R^2 = 0.002,\ DW = 1.43,\ White(p\text{-}value) = 0.158$		
Poland	-1.37^{***}	3.45^{*}
	(0.47)	(1.92)
$Adj.\ R^2 = 0.024,\ DW = 1.70,\ White(p\text{-}value) = 0.674$		
Romania	-0.51^{***}	1.16^{***}
	(0.17)	(0.29)
$Adj.\ R^2 = 0.693,\ DW = 1.44,\ White(p\text{-}value) = 0.693$		

Standard errors appear in brackets. DW is the Durbin–Watson statistic and White reports the *p*-value for the White test for heteroskedasticity.
* Denote statistical significance at the 10% level.
** Denote statistical significance at the 5% level.
*** Denote statistical significance at the 1% level.

Table 1 shows the results for the full sample 2007:05–2014:10. The constant is negative in all cases, although only statistically significant for Poland and Romania, suggesting that the forecasters include a risk premium in their forecasts. The estimated parameter of the interest rate differential is in all cases positive and fairly close to 1 with the possible exceptions of the Czech Republic and Poland, for which the slope parameters are higher than 1 although not statistically different from 1.

The estimation results in Table 1 suggest that the CF forecasts may have been based on the UIP hypothesis. The explanatory power of the models for the full sample is relatively low for some of the countries and the estimations for Hungary and Romania may suffer from mild autocorrelation. These issues motivate the use of a more sophisticated modelling strategy allowing for structural breaks.

3. Structural breaks

The sample period covers the global financial crisis, several debt crises and substantial financial instability. These events may have led to structural breaks and (1) is therefore altered to allow for up to two endogenously determined structural breaks in the parameters:

$$\begin{aligned}\frac{s^{CF}_{t,t+1} - s_t}{s_t} &= \alpha_1 I(t < T_1) + \beta_1 I(t < T_1)\,(i_t - i^{EA}_t) \\ &\quad + \alpha_2 I(T_1 \le t < T_2) \\ &\quad + \beta_2 I(T_1 \le t < T_2)\,(i_t - i^{EA}_t) \\ &\quad + \alpha_3 I(t \ge T_2) + \beta_3 I(t \ge T_2)\,(i_t - i^{EA}_t) + \varepsilon_t.\end{aligned} \tag{2}$$

The indicator function $I(\cdot)$ takes the value 1 when the condition in the bracket holds. Bai and Perron (1998) present a test for obtaining the number of breaks endogenously and discuss the properties of the estimators. They propose the use of the Bayesian information criteria (BIC), the Liu et al. (1997) modified Schwarz information criteria (LWZ), and two F-tests to establish the number of breaks. Bai and Perron (2003) discuss key practical issues.

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