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Abstract

This paper examines the uncovered interest parity puzzle in Central and Eastern European countries. Apart from investigating baseline UIP regressions, we check for structural breaks in this relationship, scrutinize deviations from the UIP, and employ different estimation methods and models augmented with various risk measures. Moreover, we offer several extensions to the common UIP testing that account for foreign-exchange interventions, the implied volatility of exchange rates, and the limited availability of data on direct measures of market expectations. The study shows that the choice of the reference currency matters for the outcome of the interest parity tests in the CEE economies. In particular, we demonstrate that inconsistencies between the results of the UIP tests vis-à-vis the euro and the US dollar that appear in CEE economies may be accounted for by the movements of the euro-dollar risk premium. This regularity has not been documented in previous studies. Additionally, we show that (a) the FX interventions in Czechia distorted the UIP, (b) the directly measured exchange rate expectations (granular survey data) in Poland do not seem to be informed by the UIP relationship, (c) the limited resilience of CEE economies to rare disasters may plausibly explain deviations from the UIP.

Keywords: interest parity puzzle; forward premium puzzle; risk premium; Fama regression; Central and Eastern Europe.

JEL Classification: F31, F41, G15.

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1 Introduction

Why do high-interest rate currencies appreciate rather than depreciate, as implied by the uncovered interest parity (UIP)? Given their stark consequences for international financial integration, central banking, and exchange rate management, as well as practical implications for cross-border investment, anomalies in the UIP have raised various controversies. Considerable progress in this area, both on empirical and theoretical fronts, has not yet led to a consensus on how prevalent or robust is the UIP (or forward premium) puzzle and which mechanisms explain its existence (see Engel, 2016). Specifically, the recent findings in this area highlight two issues concerning the puzzle. First, the empirical results on the violation of the UIP seem to differ across countries and currencies: between advanced and emerging economies or various exchange rate systems (Frankel and Poonawala, 2010; Engel and Zhu, 2019). Second, some studies indicate that the global financial crisis (GFC), the subsequent zero lower bound environment and unconventional monetary policies implemented by the major central banks may have brought lasting distortions in the interest rate parities (Bussière et al., 2022; Ismailov and Rossi, 2018). In this paper, we take up both of these threads and study the group of Central and Eastern European (CEE) economies, an appealing example of small open economies, financially integrated with the Economic and Monetary Union but retaining, for the most part, independent monetary policies and floating exchange rate regimes.

The paper re-examines the evidence of the presence of the UIP puzzle for four CEE countries – Czechia, Hungary, Poland, and Romania – as well as their aggregates from 1999 to 2019. Even though there is some research on these countries (see, e.g., Jiang et al., 2013), it is still relatively scarce, especially in comparison with abundant literature on advanced economies. Thus, not only do we review the related literature on the UIP puzzle in both advanced and CEE economies, but we also provide new and up-to-date evidence for the latter group using insights from the literature on the former group. The relationship is investigated vis-à-vis three major currencies: the euro, the US dollar, and the Swiss franc, using one-month money market interest rates. We start by providing a full set of baseline forward premium regressions for which we examine possible structural breaks and perform a decomposition of deviations from the UIP. We further explore the augmented UIP models and introduce several model specifications that include various factors which potentially account for the UIP puzzle, such as the realized volatility of the exchange rate, a volatility model of excess returns, and international risk and business cycle measures.

We arrive at several noteworthy findings. First, we demonstrate that even though the risk-based measures, such as the realized volatility of the exchange rate and the VIX, largely improve the baseline testing regressions, both from statistical and economic points of view, they do not alter the overall outcomes of our empirical models. Second, we show that the choice of the reference currency matters for the outcome of the interest parity tests in the CEE economies. The UIP puzzle is stronger for the euro and the Swiss franc than for the US dollar, a regularity that has not been documented in previous studies. Third, we provide a plausible explanation for the inconsistencies in the UIP test that relies on the role of the euro-dollar risk premium for interest parities of individual CEE versus those two major currencies. Fourth, in a series of extensions of the baseline framework, we show that: (a) the exchange rate peg of the Czech koruna from 2013 to 2017 to the euro had a significant impact on the UIP, (b) forecast and risk premium errors play leading roles in UIP deviations in Poland, (c) crash risks are priced into the CEE currencies exchange rates.

The principal contribution of this paper comes from a comparative analysis of the CEE economies, which we scrutinize under competing specifications of regression models. Importantly, given that the sample encompasses the turmoil times of the GFC, it is long enough to mitigate the so-called peso problem and the results, therefore, are not subject to a small sample bias. Moreover, CEE economies fit well 'a risky country' profile in the rare disaster hypothesis developed by Farhi and Gabaix (2016). We demonstrate how their hypothesis can be used empirically to explain the deviations from the UIP relationship in CEE economies. We also contribute to the literature in three other respects. First, we provide novel evidence of the role of the euro-dollar risk premia on the UIP between CEE economies and those two international currencies. Second, building on the insights into a decomposition of the UIP regression slope coefficient from Bussière et al. (2022), we put forward a simple way to carry out such a decomposition when constrained by the lack of data on exchange rate expectations. Such data are seldom available for emerging market economies, including CEE countries, over long periods (see, e.g., Cuestas et al., 2015). Third, as an extension to the main part of the study, we provide a country-specific insight using the datasets that have not been used in empirical studies on the forward premium puzzle: foreign exchange (FX) market interventions by the Czech National Bank, survey-based forecasts of the Polish zloty exchange rate by financial analysts, and option-implied risk reversals for CEE currencies. These datasets enable us to demonstrate that: (a) the UIP relationship can break down in the presence of FX market interventions, (b) the directly measured exchange rate expectations do not seem to be informed by the UIP relationship, and (c) the limited resilience of CEE economies to world disasters may plausibly explain deviations from the UIP.

The paper is organized as follows. The next section reviews recent studies on the UIP puzzle. In sections 3 and 4, we lay down a theoretical framework for the paper and describe the empirical models and data. Section 5 contains the empirical results, along with the discussion of our main findings. We then turn to additional case studies on Czechia, Poland, and rare disasters (section 6). The final section concludes and outlines further research in the area.

2 Related literature

There is a long research tradition of the UIP relationship in advanced economies, especially for the US dollar. Many empirical studies find little support for the UIP, revealing that higher interest rate currencies often appreciate against lower interest rate ones, the phenomenon dubbed the UIP puzzle or the Fama (1984) puzzle.

The main empirical findings in this literature can be summarised in several important observations. First, the UIP works better in emerging market economies (EMEs) compared to advanced economies (AEs) (see, e.g., Bansal and Dahlquist, 2000; Frankel and Poonawala, 2010; Gilmore and Hayashi, 2011). In an important study, Bansal and Dahlquist (2000) define the puzzle as a case in which the observed direction of an exchange rate change is opposite to the one implied by the UIP. In other words, deviations from the UIP that do not violate the direction of the implied relationship between the exchange rate and interest rates are not considered the puzzle. Using this terminology, they document that the UIP is rejected in both groups of economies, but the puzzle is not a pervasive phenomenon: it is not present in EMEs. They put forward a conjecture that there is a relationship between the level of development, average inflation, inflation volatility, and the presence of the UIP puzzle.

In more recent studies, these findings are corroborated. Frankel and Poonawala (2010) hypothesize that EMEs' currencies have more easily identified trends of depreciation than currencies of AEs. This does not remove the bias in the forward discount as a predictor of the future change in the spot exchange rate, but the bias is less severe among the former currencies than the latter. Given this finding and the observation that EMEs' currencies are probably riskier, it is suggested that a time-varying exchange risk premium may not be a proper explanation of the UIP puzzle.

In line with the previous studies Gilmore and Hayashi (2011) establish that the extent of the puzzle is smaller for EMEs than for AEs. Interestingly, using aggregate data for both groups of economies, they observe that the excess return on emerging market currencies is better explained by the interest rate differential for major currencies than by the interest rate differential between emerging market currencies. They conjecture that the excess return for individual currencies has a common global real interest factor and the interest rate differential between AEs is a better predictor of this factor.

The second key finding on the UIP condition is that it works systematically better in crisis times when both the exchange rate and interest rate volatilities are high (see, e.g., Flood and Rose, 2002; Clarida et al., 2009; Czech, 2017). Using the data that include the major currency crises in the 1990s, Flood and Rose (2002) report that the high interest rate currencies tend to depreciate, although the exchange rate changes are short of those implied by the interest rate differential. In this sense, the UIP 'works better than it used to'. The likely reason behind this improvement is that the increased exchange rate and interest rate volatilities raise 'the stakes for financial markets and central banks' and 'may provide a more statistically powerful test for the UIP hypothesis' (Flood and Rose, 2002).

A violation of the UIP obtained in Fama regressions is an artefact of the volatility regime according to Clarida et al. (2009). In line with the massive literature, they document that the high interest rate currency tends to appreciate but show that this finding holds in low volatility environments only. In high volatility states, it is the low interest rate currency that appreciates, and the change is greater than the one implied by the UIP. This is in line with the finding that as volatility grows, the available speculator capital shrinks due to higher margins and capital requirements, so traders cut back on their carry trade activities (Brunnermeier et al., 2008). Interestingly, using a similar explanation Ismailov and Rossi (2018) argue that deviations from the UIP are more likely in highly uncertain environments because investors might not be willing to take advantage of arbitrage opportunities. In turn, Kumar (2020) emphasises that these opportunities may not be exploited by arbitragers because their demand for liquidity is constrained.

Third, some papers find that the UIP holds more often in the long run than in the short run (see, e.g., Juselius, 1995; Chinn and Meredith, 2004; Lothian and Wu, 2011; Chinn and Quayyum, 2012). Using ultra long time series on two currency pairs, the French franc versus the pound sterling and the US dollar versus sterling that span the 1800-1999 Lothian and Wu (2011) demonstrate that the UIP holds over the very long haul, and the puzzle emerges only when the

1980s dominate the sample. Their conjecture is that the deviations from the UIP, including deviations over long spans of time, are due to slow adjustment of expectations to actual regime shifts as well as anticipations for extended periods of regime shifts that never materialize. Thus, their explanation is related to the well-known peso problem.

The joint hypothesis of UIP and rational expectations is found to hold better at long horizons than at short ones by Chinn and Meredith (2004). This is in line with the point raised by Juselius (1995) that over the long run, the exchange rate and interest rates cannot diverge substantially without evoking adjustment forces that tend to restore equilibrium. The findings are corroborated by Chinn and Quayyum (2012) who additionally observe that the effect is somewhat weaker in a sample that includes the close to zero lower bound interest rates in Japan and Switzerland. At the same time, they admit that the failure of the UIP is more pronounced at long horizons when the pound sterling instead of the US dollar is used as the base currency.

Empirical studies on the UIP puzzle in CEE countries often reflect broader themes present in macroeconomic research on these economies.[1] Primarily, the UIP is revoked when discussing the unique situation of the CEE economies being the EU members outside of the Eurozone, the effectiveness of their independent monetary policies, and costs and benefits of the future euro adoption in these countries. Against this backdrop, Filipozzi and Staehr (2012) estimate the UIP regressions for the CEE economies and the euro using data ranging from 1999 to 2011. Their results show that the UIP holds for Romania, there is a forward premium puzzle for Czechia and Hungary, and the outcomes for Poland are not conclusive. The study also demonstrates that low and high interest rate spread regimes and global risk factors may explain some shifts in risk premia in CEE economies, although their role is not uniform across countries.

Using slightly longer time-series, Cuestas et al. (2017) estimate a similar set of models but consider possible structural breaks in the regressions. They show that the UIP hypothesis cannot be rejected when the empirical models assume rational expectations of the exchange rate. Since this is a standard way of testing for the UIP, the results stand at odds with the majority of international empirical studies. However, when regression specifications incorporate static expectations (i.e., naive, based just on the observable exchange rate), there is more substantial evidence for the UIP anomaly in all CEE economies. It is therefore argued that the way that expectations are formed by FX market participants is crucial in interpreting the results of the UIP tests in Central Europe.

In a recent study, Ferreira and Kristoufek (2020) investigate the UIP condition for the entire EU using the so-called fractal analysis, based on the cross-correlations of daily financial data. They report that, as a rule, the UIP is less likely to hold for the non-Euro economies, but there is considerable evidence against investors' risk-neutrality and rational expectations in all EU economies. The results for the group of CEE economies are mixed, with some indication of the UIP puzzle for Hungary, Poland, and Romania. This result is attributed to the monetary autonomy of these economies and the systematic reactions of their central banks to asymmetric shocks.

The literature also links the tests for interest parity to the long-lasting effects of economic transformation, financial liberalization, differences in income levels between 'old' and 'new' EU

[1] Long-run relationships in the CEE real exchange rates are studied by, e.g., Kębłowski (2011).

members, and a relatively fast catching-up process of the region. An example of such an analysis was provided by Jiang et al. (2013) who test for stationarity in the FX risk premium components among the CEE economies. The study uses of the money market interest rates from 1997 to 2011 and estimates a set of threshold autoregressive models. It finds that the long-run UIP holds for all four CEE countries included in our study and the euro as a base currency. However, little is said about possible short-term deviations from the parity.

An important theme in the UIP research on CEE economies is the role of exchange rate volatility and related country-specific risks in explaining forward premium anomalies. Horobet et al. (2010) perform the UIP tests augmented with the capital market and FX market volatility vis-à-vis four developed market currencies. They demonstrate that the UIP anomalies for the comparatively high-yielding CEE currencies may be related to the role that market volatility plays in asymmetric adjustments in exchange rates, particularly during episodes of swift depreciation followed by periods of sluggish appreciation. Using threshold and component GARCH models, Triandafil and Richter (2012) generally reject the UIP for the CEE economies, both when the euro and US dollar are taken as a reference currency. They explain these results with relatively high inflation rates specific to transition economies that translate to an elevated risk aversion in CEE financial markets.

The UIP puzzle in CEE economies is also studied through the lens of carry trade opportunities in their FX markets. Hoffmann (2012) shows that strong economic growth and relatively high yields have attracted substantial un-hedged portfolio investment flows to the region. He finds that the UIP is typically violated in CEE countries between 1999 and 2009 when the euro or the Swiss franc are used as funding currencies. The returns to carry trade tend to be higher when a CEE economy retains a managed floating or fixed exchange rate regime. They also depend on the global risk factors and increase when the interest rate spread against the Eurozone or Switzerland are sizeable.

Covering a broader sample of post-transition economies, Hayward and Hölscher (2014) confirm that over the period 2000-2011, the average carry trade returns in the region differ between the two regimes, which they describe as ‘moderation’ and ‘crisis’. The latter is associated mainly with the GFC: abrupt unwinding of cross-border investment, followed by low mean returns. They demonstrate that abnormal returns on CEE currencies were easier to obtain in the 2000s when the US dollar was the funding currency, while the results for the euro were mixed. For all four economies studied in this paper, carry trades funded by the Swiss franc also turned out to be profitable.

3 Theoretical background

In this section, we first explain the UIP relationship and then discuss three insights on deviations from that relationship that have testable implications. A central building block of our theoretical framework is the uncovered interest rate parity (UIP) relationship. This condition says that the differences between domestic and foreign interest rates are compensated for by the expected changes in the exchange rate. In other words, the expected excess return on any currency is zero unless there exist barriers that hinder capital flows between countries.

Let the nominal exchange rate S_t be defined as the price of domestic currency in terms of foreign currency, so its increase reflects an appreciation of domestic currency. In line with a small open economy assumption, the foreign currency is assumed to be one of the major currencies, e.g., the US dollar. The risk-free rates of return on domestic and foreign assets are i_t and i_t^*, respectively. In order to make them comparable, the former is expressed as a dollar rate of return, i.e., $i_t + s_{t+1}^e - s_t$, where $s_{t+1}^e = \ln S_{t+1}^e$ and $s_t = \ln S_t$. The UIP condition is given by:

$$s_{t+1}^e - s_t + i_t - i_t^* = 0. \tag{1}$$

Under rational expectations $S_{t+1}^e = \mathrm{E}_t S_{t+1}$, so the log rate of expected appreciation of domestic currency is $\ln(\mathrm{E}_t S_{t+1}) - \ln S_t$. Following the literature (see, e.g., Engel, 1996), it is assumed that the exchange rate is conditionally log-normally distributed, so the UIP can be restated as:

$$\mathrm{E}_t s_{t+1} - s_t + i_t - i_t^* = -0.5\mathrm{var}_t(s_{t+1}) \tag{2}$$

where the term on the right hand side is the conditional variance of the log of the exchange rate. Given that there are no systematic errors (expectations are rational), the UIP becomes:

$$\rho_{t+1} = -0.5\mathrm{var}_t(s_{t+1}) + \varepsilon_{t+1} \tag{3}$$

where $\rho_{t+1} \equiv s_{t+1} - s_t + i_t - i_t^*$ and ε_{t+1} is an i.i.d. error term. If the term with conditional variance is negligible, ρ_{t+1} can be simply interpreted as an *excess* return on domestic currency.

Equations (1)-(3) apply to risk neutral agents. More general formulation of the UIP condition can be derived from the Euler equation in the model of utility maximising agents. For example, Bekaert and Hodrick (1993) demonstrate that:

$$\rho_{t+1} = \mathrm{cov}_t(s_{t+1}, q_{t+1}) - 0.5\mathrm{var}_t(s_{t+1}) + \varepsilon_{t+1} \tag{4}$$

where $\mathrm{cov}_t(s_{t+1}, q_{t+1})$ is the conditional covariance between the exchange rate and the (log of the) intertemporal marginal rate of substitution of a dollar between period t and $t+1$, q_{t+1}. The attitudes towards risk enter this equation through the conditional covariance. For example, under risk aversion, ρ_{t+1} is positive because agents need to be paid a risk premium to hold domestic assets.

Engel (2016) explains that ρ_{t+1} is 'the object of almost all of the empirical analysis of excess returns in foreign exchange markets'. In other words, both conditional variance and covariance are rarely included in the regression tests of the UIP condition. The well-known finding in this literature is that the behaviour of the exchange rate is puzzling: '[...] when the interest rate (one country relative to another) is higher than average, the short-term deposits of the high-interest rate currency tend to earn an excess return' (Engel, 2016), so:

$$\mathrm{cov}(\mathrm{E}_t \rho_{t+1}, i_t - i_t^*) > 0. \tag{5}$$

This finding is considered the uncovered interest rate parity puzzle because, according to the UIP condition, an increase (decrease) in the interest rate differential is offset by subsequent depreciation (appreciation) of domestic currency, so no relationship between the excess return and interest rate differential should be observed.

Bussière et al. (2022) demonstrate that the relationship between the exchange rate and interest rates can be decomposed using the following equations:

$$\begin{aligned} f_t - s_t &= -(i_t - i_t^*) + \varepsilon_t^{cip} && (6)\\ f_t &= s_{t+1}^e - \varepsilon_t^{rp} && (7)\\ s_{t+1}^e &= s_{t+1} - \varepsilon_{t+1}^f && (8) \end{aligned}$$

where f_t is the (log of the) forward exchange rate at time t for delivery at time $t+1$. Equation (6) is the covered interest rate parity condition and ε_t^{cip} represents deviations due to transactions costs, measurement errors etc. (see, e.g., Thornton, 2019). Equation (7) indicates that the forward rate is equal to the market participants' expectation of the future spot rate corrected by an exchange risk premium, ε_t^{rp}. Equation (8) states that the expectations are formed rationally: the right-hand side of this equation is simply $\mathrm{E}_t s_{t+1}$ (the conditional variance is neglected). These three equations can be used to derive equation (3) (up to the neglected conditional variance term) with $\varepsilon_{t+1} = \varepsilon_t^{cip} + \varepsilon_t^{rp} + \varepsilon_{t+1}^f$.

A simple regression-based test of the UIP condition can be specified using the Fama regression:

$$\rho_{t+1} = \zeta + \beta(i_t - i_t^*) + \varepsilon_{t+1} \tag{9}$$

and the null hypothesis that both coefficients are zero.[2] Using equations (6)-(8), Bussière et al. (2022) demonstrate that the deviations from the null hypothesis can be explained by several moment conditions:

$$\operatorname{plim} \hat{\beta} = \frac{\operatorname{cov}(i_t - i_t^*, \varepsilon_t^{cip})}{\operatorname{var}(i_t - i_t^*)} + \frac{\operatorname{cov}(i_t - i_t^*, \varepsilon_t^{rp})}{\operatorname{var}(i_t - i_t^*)} + \frac{\operatorname{cov}(i_t - i_t^*, \varepsilon_{t+1}^f)}{\operatorname{var}(i_t - i_t^*)}. \tag{10}$$

Thus, a non-zero estimate of β can stem from the deviations from the covered interest rate parity, a time-varying risk premium, and departures from rational expectations. Their relative importance in CEE countries is examined in the empirical part of the paper.

The foreign exchange (FX) market interventions can affect the observed changes in the exchange rate. For example, interventions based on the 'leaning against the wind' rule would generally reduce fluctuations in the exchange rate without corresponding changes in the interest rates. If it is uncertain whether a monetary authority will continue to intervene, then the agents have to take it into account when forming their expectations about the exchange rate. A framework similar to the one used in the literature to model uncertainty about the future shift in the regime can be applied here (see, e.g., Sarno and Taylor, 2003). This framework is helpful to interpret the results obtained for Czechia in Section 5.5.

Let us assume that agents expect a monetary authority to continue (sterilized) FX interventions with probability $(1 - \lambda)$ and to shift to a policy of no interventions with probability λ.

[2] It is common in the literature to use the change in the exchange rate as a dependent variable, *not* the excess return, and treat the US dollar as a domestic currency. In such a case the null is that β equals 1.

Given that the expected exchange rate is not the same under the two regimes:

$$\mathrm{E}_t s_{t+k} = (1-\lambda)\mathrm{E}_t(s_{t+k}|M_1) + \lambda \mathrm{E}_t(s_{t+k}|M_2) \tag{11}$$

where M_1 and M_2 are old and new regimes, respectively. If the regime shift does not occur, the forecast error is:

$$s_{t+1}^{M_1} - \mathrm{E}_t s_{t+1} = \eta_{t+1} + \lambda \nabla s_{t+1} \tag{12}$$

where $s_{t+1}^{M_1}$ is the actual exchange rate at time $t+1$ in an old regime, $\eta_{t+1} = s_{t+1}^{M_1} - \mathrm{E}_t(s_{t+1}|M_1)$ is the rational expectation forecast error, i.e. the error that would be observed if agents were certain that a monetary authority would continue its interventions, and $\nabla s_{t+1} \equiv \mathrm{E}_t(s_{t+1}|M_1) - \mathrm{E}_t(s_{t+1}|M_2)$. Using (12) in (2), one obtains that:

$$s_{t+1}^{M_1} - s_t + i_t - i_t^* = -0.5\mathrm{var}_t(s_{t+1}) + \lambda \nabla s_{t+1} + \eta_{t+1}. \tag{13}$$

Three remarks seem to be relevant here. First, when there is an appreciation (depreciation) pressure that the central bank tries to mitigate by buying (selling) foreign currency, the interventions can contribute to a downward (upward) revision of the expected exchange rate (the signalling channel). This implies that an excess return on domestic currency (again neglecting the term with conditional variance) is negatively (positively) related to purchases (sales) of foreign currency by the central bank. For example, if a dummy for purchases of foreign currency is used as a proxy of ∇s_{t+1}, then the coefficient of this dummy is negative. It is because under appreciation pressure ∇s_{t+1} becomes negative.

Second, monetary authorities change the relative supplies of domestic and foreign assets (the portfolio balance channel). Thus, FX interventions affect the relative value of these assets and this – given that interventions are sterilized, and assets are not perfect substitutes – results in the alleviation of appreciation/depreciation pressure.

Third, the uncertainty about a shift in an FX intervention policy will result in a skewed forecast error distribution. The size of skewness depends on the probability of a regime shift, its likely scale, and it can prevail even after the regime shift until market participants get convinced that there will be no return to an old regime.

Additional insights into the Fama regression coefficient and its possible deviations from the UIP relationship are offered by Farhi and Gabaix (2016) in their rare disaster hypothesis (RDH). The hypothesis is well-fitted to the CEE countries' characteristics as their currencies can be considered more risky than the major currencies. According to the RDH, a country that is perceived as relatively risky has a high interest rate and weak currency because investors need to be compensated for the risk of depreciation of that currency during a potential world disaster. Thus, positive expected returns from investing in a high interest rate currency are simply compensation for bearing disaster risk.

The implication for the Fama regressions is that the β coefficient should deviate from its UIP value. The reason is that it is a weighted average of two terms: a) the coefficient derived from the rare disaster model of an exchange rate with no inflation, and b) the coefficient implied by the UIP. Farhi and Gabaix (2016) demonstrate that the former is determined by the world intensity of disasters and the country's resilience to such disasters and is smaller than the coefficient implied by the UIP. Accordingly, the β coefficient can be written as:

$$\beta = \nu\beta^{NI} + (1-\nu)\beta^{UIP} \tag{14}$$

where β^{NI} is the coefficient derived from the rare disaster model setup with no inflation. The weighting parameter ν depends on the variability of inflation differential, σ_π^2, and the variability of relative resilience of a given country, σ_H^2:

$$\nu = \left[1 + \frac{\sigma_\pi^2}{c\sigma_H^2}\right]^{-1} \tag{15}$$

where c is a positive scaling constant. Using these equations, one can observe that the more variable the inflation differential, the closer the β coefficient to the coefficient implied by the UIP. Conversely, the higher the variability in the country's relative resilience, the greater the deviation of β from its UIP level. The implication is, therefore, that the relative importance of these two variabilities should be negatively associated with the deviations from the UIP condition, i.e., the higher the *ratio* of inflation differential variability to the variability in the relative country resilience, the smaller the deviation from the UIP relationship.

4 Empirical models and data

This section presents a sequence of empirical models, along with the description of the estimation procedures and data used in the paper. Coming back to the definition of the excess return on domestic currency (used in Equation 3), we define our baseline forward premium regression (Model 1) as:

$$\rho_{t+1} = \zeta + \beta(i_t - i_t^*) + \varepsilon_{t+1}, \tag{16}$$

where ε_t is i.i.d.N$(0, \sigma_\varepsilon^2)$. Following the standard Fama regression, $\beta > 0$ indicates a positive relationship between the excess returns and interest differentials, the anomaly described as the UIP puzzle.

In Model 2, the benchmark equation is augmented with the realized volatility (RV) of the exchange rate:

$$\rho_{t+1} = \zeta + \beta(i_t - i_t^*) + \gamma v_{t+1}^2 + \varepsilon_{t+1}. \tag{17}$$

The realized volatility, v_t^2, is calculated on a monthly basis using daily exchange rate returns and given as:

$$v_t^2 = \sum_{m=2}^{M_t} (s_{m,t} - s_{m-1,t})^2, \tag{18}$$

where $s_{m,t}$ is the log of the daily exchange rate and M_t is the number of trading days, changing each month.

Following the ample literature on the role of the exchange rate volatility for the UIP tests, in the next specification, we explicitly account for the potential heteroskedasticity in the baseline regression residuals. Hence, Model 3 (GARCH) is a set-up extended with a GARCH$(1, 1)$ process:

$$
\begin{aligned}
\rho_{t+1} &= \zeta + \beta(i_t - i_t^*) + \varepsilon_{t+1} && \text{(19a)}\\
\varepsilon_t &= z_t \sigma_t && \text{(19b)}\\
\sigma_t^2 &= \omega + \alpha_1 \varepsilon_{t-1}^2 + \beta_1 \sigma_{t-1}^2, && \text{(19c)}
\end{aligned}
$$

where z_t has the generalized error distribution (GED), $\omega \geq 0$, $\alpha_1 \geq 0$, $\beta_1 \geq 0$, and $\alpha_1 + \beta_1 < 1$. It must be noted here that before this form of Model 3 was selected, we examined multiple GARCH specifications, i.a., exponential, component, and GJR GARCH processes, with in-variance or in-mean GARCH terms, and various residual distributions. The results obtained under these specifications were qualitatively quite similar to those from the GARCH-GED model, but their statistical properties were inferior.

The next equation, which we call Model 4, extends the previous models with several financial and macroeconomic risk measures. In its full form, the regression is given as the following mean equation:

$$\rho_{t+1} = \zeta + \beta(i_t - i_t^*) + \gamma v_{t+1}^2 + \delta_1 \Delta vix_{t+1} + \delta_2 kilian_{t+1} + \delta_3 \Delta epu\_eu_{t+1} + \varepsilon_{t+1}, \tag{20}$$

where Δvix_{t+1}, $kilian_{t+1}$, and Δepu_eu_{t+1} denote the indicators of international financial risk (VIX), the global index of economic activity (Kilian, 2009), which approximates the overall macroeconomic risk, as well as the economic policy uncertainty in Europe, a news-based indicator put forward by Baker et al. (2016).

The inclusion of these variables in the regression may be traced back to theoretical explanations for the UIP failure. The risk-taking behaviour is likely to change along financial and business cycles in the world economy due to common drivers of volatility. Risk fluctuations, in turn, tend to shape cross-border portfolio flows which impact the risk premia that investors demand for holding certain assets. Hence, the risk measures may be treated as control variables that potentially alter the relationship between interest rate differentials and the excess return in the benchmark regression model.

The CEE countries are financially integrated with the euro area, so it is reasonable to expect that the excess returns on CEE currencies are related to movements of the USD/EUR exchange rate.[3] In the limiting case of hard pegging a currency to the euro, the excess returns on both currencies against the US dollar would follow the same path. In the case of CEE countries, which have more flexible exchange rate regimes, such a claim would be far-fetched. Notwithstanding, there is a whole range of options between movements of excess returns along the same path and their perfect independence. Accordingly, Model 5 extends the baseline specification as follows:

$$\rho_{t+1} = \zeta + \beta(i_t - i_t^*) + \theta_1 \rho_{t+1}^* + \theta_2(i_t^{US} - i_t^{EA}) + \varepsilon_{t+1}, \tag{21}$$

where ρ_{t+1}^* is an excess return on the US dollar against the euro. We include both the external excess return and interest rate differential $(i_t^{US} - i_t^{EA})$ to make the specification flexible enough to capture the dynamics of modelled variables. In the empirical part, however, we keep the

[3] Kębłowski et al. (2020) raise a similar point when modelling the real exchange rates of CEE currencies.

external interest differential in the specification only when it is statistically significant.

An important part of our empirical analysis consists of the decomposition of the UIP slope deviations from the theoretical value of zero (when the UIP holds) in Equation (16). However, it must be noted that we face a shortage of data on market expectations or forecasts of exchange rates across the CEE economies over a longer period. Because these series are needed to calculate the risk premium and expectation error terms in Equations (7) and (8), the estimated β cannot be decomposed as shown in Equation (10). Given these limitations, in this paper we suggest a simple alternative to achieve this decomposition. The procedure we introduce consists of four steps.

In the first step, we obtain the error term that may be worked out directly from the data, i.e., the CIP error. It is calculated with observable forward and spot rates, as well as domestic and foreign interest rates, as in Equation (6), $\varepsilon_t^{cip} = f_t - s_t + i_t - i_t^*$.

Next, we postulate that the risk premium can be disentangled from the full ε_{t+1} error using additional information embedded in a major risk factor that likely drives this premium. We assume here that this variable should be highly correlated with the risk premium chunk of the UIP error term and not with its remainder. The CBOE VIX, used to proxy the global risk levels and often dubbed the 'fear index', seems to be a plausible choice of such a factor. Hence, we regress the UIP error series ε_{t+1} on a constant and the log rate of growth of VIX:

$$\varepsilon_{t+1} = \alpha_0 + \alpha_1 \Delta vix_{t+1} + u_{t+1}. \tag{22}$$

In the third step, the risk premium term is retrieved as the fitted value of the OLS regression, $\hat{\varepsilon}_t^{rp} = \hat{\alpha}_0 + \hat{\alpha}_1 \Delta vix_t$. Because the residuals of the regression still contain both the forecast and CIP errors, the former one is isolated as the following difference, $\hat{\varepsilon}_{t+1}^{f} = \hat{u}_{t+1} - \varepsilon_t^{cip}$.

Finally, given the estimates of the three error terms, we may use them to break down the UIP deviations into the CIP, risk premium, and forecast components, and assess their relative importance.

The dataset that we collect covers four CEE economies: Czechia, Hungary, Poland, and Romania, for a maximum period of 1999 to 2019. The choice of the end date of the analysis is warranted by two considerations. On the one hand, the set covers relatively long series both before and after the GFC of 2007-2009, which allows us to consider probable shifts in the UIP in the post-crisis new normal state. On the other hand, the time span excludes the post-COVID period of unusually high inflation rates that might alter the relationships among exchange rates and interest rates towards the end of the sample. The spot exchange rates are in monthly frequency, defined as end-of-month daily observations. There are three reference currencies, the euro, the US dollar, and the Swiss franc, and in each case we standardize the exchange rate series relative to their mean value in 2005. The regressions are based on the one-month money market interest rates, which are again taken as the end-of-month values and transformed from annual into monthly rates.

The data on the forward rates turns out to be more problematic because their availability differs across economies. As far as the US dollar and the euro are concerned, the series cover the entire time span of the analysis with just three exceptions. The PLN/EUR, RON/EUR, and RON/USD series start later, between 2002 and 2004. The forward rates for the CEE currencies and the Swiss franc are either unavailable or cover a very short period, going back just a few years. Hence, some parts of the empirical analysis cannot be conducted for the franc.

In addition to individual data, we calculate an aggregated series for the CEE economies, which we dub V3R, as includes the three non-euro Visegrad Group countries, Czechia, Hungary, and Poland, along with Romania. The aggregate exchange rates and interest rates are computed using the geometric weighted mean and constant trade weights (sum of exports and imports) based on annual data for 1995-2019. As suggested by Engel (2016), such a setting may be superior to a pooled panel regression because it is unlikely that all four economies will be characterized by the same estimates of β.

All exchange rate and interest rate series were obtained from the Refinitiv Datastream. Trade weights used to calculate the aggregates come from the Eurostat database. The risk measures are retrieved from the datasets indicated above.

5 Results and discussion

This section presents and discusses our empirical findings. We start with the outcomes of baseline Fama regressions for the CEE economies. We then show a decomposition of their slope coefficients, investigate the UIP regressions augmented with various risk measures, and compare various specifications across the economies. Next, we focus on the role of the UIP relationship between the US and the euro area for the UIP puzzle in CEE economies.

5.1 Baseline Fama regressions

The baseline UIP regressions are estimated for four economies and their aggregate, V3R, based on weighted averages of interest rates and exchange rates in the CEE economies (Table 1). In each case, there are three reference currencies. Out of the total of 18 regressions, the estimates of β (a coefficient on $i_t - i_t^*$) are statistically significant at the 0.1 level in 10 cases. In general, the results are more supportive of the UIP when the US dollar is the reference currency and indicate deviations from the parity vis-à-vis the euro and the Swiss franc. This pattern is discernible for Poland and the V3R aggregate, although point estimates of β are also lower for the US than for the euro area for Czechia and Romania. Czechia is the only economy where the slope is insignificant for the euro area and Switzerland, which indicates that the UIP holds simultaneously against these two benchmarks. Conversely, the point estimates of β are high (above 1.5) for Hungary and statistically significant in all three specifications, pointing to a firm rejection of the UIP. The estimates of regression intercepts, in turn, are almost never significant. Two notable exceptions are the negative intercepts for Hungary when the CHF or the EUR are taken as reference currencies, making it the only economy with a non-zero constant risk premium in the UIP regressions.

Since most of the reported regressions display serial correlation of residuals, as confirmed by the Ljung-Box statistics with one and ten lags (Q_1 and Q_{10}), we calculate the HAC robust standard errors to correct for a possible bias in the estimates of β. It must be noted, however, that potential misspecification of the baseline UIP regression may also be reflected in the correlation of squared residuals, as indicated by further residual diagnostics (Q_1^2 and Q_{10}^2). We come back

Table 1: Baseline UIP regression results for CEE economies

		const	$i_t - i_t^*$	Q_1	Q_{10}	Q_1^2	Q_{10}^2
Czechia	EA	0.141	1.068	0.360	8.059	4.891	44.155
		[0.203]	[0.353]	[0.548]	[0.623]	[0.027]	[0.000]
	US	0.144	-0.382	0.164	11.357	0.082	39.960
		[0.519]	[0.858]	[0.685]	[0.330]	[0.775]	[0.000]
	CH	-0.051	1.280	3.079	13.620	0.492	4.893
		[0.842]	[0.436]	[0.079]	[0.191]	[0.513]	[0.898]
Hungary	EA	-0.312	1.527	0.314	28.801	1.661	35.714
		[0.053]	[0.000]	[0.575]	[0.001]	[0.197]	[0.000]
	US	-0.320	1.591	0.133	12.588	0.123	47.774
		[0.224]	[0.024]	[0.715]	[0.248]	[0.725]	[0.000]
	CH	-0.507	1.510	0.356	21.019	6.131	53.263
		[0.059]	[0.001]	[0.551]	[0.021]	[0.013]	[0.000]
Poland	EA	-0.180	1.530	9.639	17.824	26.035	56.652
		[0.472]	[0.006]	[0.002]	[0.058]	[0.000]	[0.000]
	US	0.129	0.503	0.675	11.749	0.074	34.986
		[0.674]	[0.398]	[0.411]	[0.302]	[0.785]	[0.000]
	CH	-0.282	1.281	0.147	7.480	1.650	37.731
		[0.353]	[0.032]	[0.702]	[0.679]	[0.199]	[0.000]
Romania	EA	-0.010	0.511	0.396	9.746	19.094	29.305
		[0.947]	[0.000]	[0.529]	[0.463]	[0.000]	[0.001]
	US	0.092	0.383	0.945	9.730	3.575	25.710
		[0.719]	[0.018]	[0.331]	[0.464]	[0.059]	[0.004]
	CH	-0.143	0.527	0.620	6.902	5.517	10.012
		[0.510]	[0.000]	[0.431]	[0.735]	[0.019]	[0.493]
V3R	EA	-0.052	1.009	12.636	24.904	40.025	78.369
		[0.715]	[0.000]	[0.000]	[0.006]	[0.000]	[0.000]
	US	0.117	0.432	0.398	12.839	0.015	41.841
		[0.655]	[0.395]	[0.528]	[0.233]	[0.902]	[0.000]
	CH	-0.194	0.969	0.008	7.294	5.597	29.854
		[0.388]	[0.004]	[0.928]	[0.697]	[0.018]	[0.000]

Notes: dependent variable: ρ_{t+1}, see Equation (16); V3R is the aggregate for Czechia, Hungary, Poland, and Romania, based on trade weights; estimation period: 1999:01 - 2019:12; p-values calculated with HAC robust standard errors in brackets; Q_1, Q_{10} and Q_1^2, Q_{10}^2 indicate the Ljung-Box statistics for the first 1 and 10 autocorrelations of standardized residuals and squared standardized residuals.

to these issues in Sections 5.3– 5.5, where we consider augmented UIP regression models.

Another potential caveat in the baseline regressions is related to breakpoints that may appear in the relationship between interest rate differential and excess returns. Appendix discusses this issue in detail and provides the results of the Chow breakpoint test in the entire set of the UIP regressions. Overall, we find weak evidence for the role of breakpoints in tested relationships. For example, there are no significant breaks that could be associated with the GFC or the zero-lower bound in the US and the EMU. Hence, we proceed with the next steps of the analysis using models without structural breaks.

So far, our main finding from the baseline Fama regressions is that the UIP puzzle appears in most of the tested currency pairs. It means, for example, that a change in the interest rate differential between the CEE economy and the Eurozone or Switzerland is a significant predictor of the excess returns on their domestic currencies. This result is consistent with the majority of the literature, which shows that the UIP puzzle occurs more often for economies maintaining floating exchange rate arrangements, low capital controls, and low inflation rates (e.g., Engel and

Zhu, 2019), such as the CEE economies in most of the period that we investigate. However, we also find a considerable variation in our results. The primary one is the difference in β estimates for the euro and the Swiss franc, on the one hand, and the US dollar, on the other. Although there are no failures of the UIP for the USD in Hungarian in Romanian cases, this would suggest the UIP holds more often in this case than for the two European currencies.

5.2 Decomposition of the UIP slope coefficients

Additionally to the joint tests of the forward premium puzzle, we further isolate the error terms in the UIP equations using market forward rates and the VIX as an international risk measure. This allows us to decompose β into three sources of deviations from the UIP hypothesis, which states that $\beta = 0$. Figure 1 depicts these deviations. The forecast error turns out to be the most important driver of the departure from the UIP across the CEE economies. For almost all currency pairs, it displays a positive impact on point estimates of β, indicating a substantial and persistent bias in the expectations of the FX market participants. In as many as five out of eight cases, this error raises the slope parameter value by more than one. The three exceptions are the CZ-US and PL-EA pairs, where the contribution of the forecast error is smaller than the risk premium error (negative in the Czech case), and the PL-US pair, for which this contribution is around 0.5.

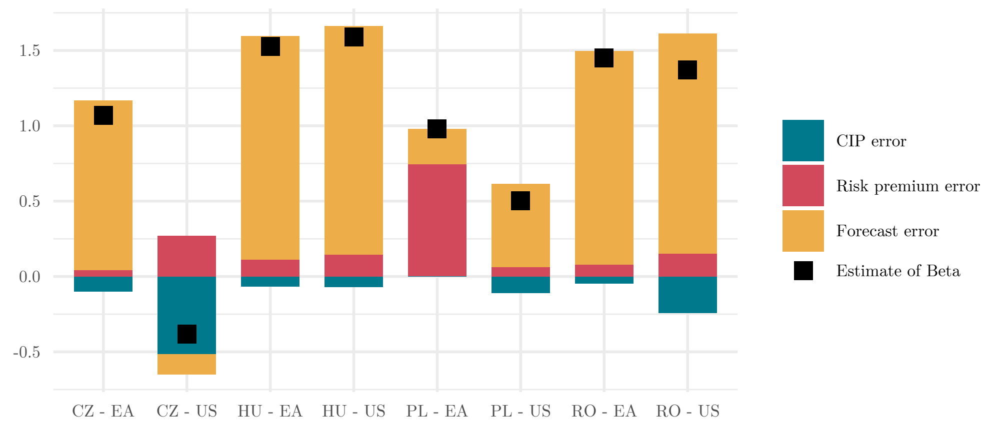

Figure 1: Decomposition of the slope parameter in baseline UIP regressions for CEE economies

Notes: see Equation (10); due to data availability, the sample for PL – EA covers the period 2003:01 - 2019:12, while samples for RO – EA and RO – US, the period 2004:03 - 2019:12.

Much like the forecast error, the risk premium component positively influences β estimates. In general, a time-varying risk premium appears to produce a relatively minor distortion to the UIP condition. In Czechia, Hungary, and Romania, the contribution of the risk premium error is larger for the US dollar than for the euro. This may indicate a more significant role of the dollar as a global risk factor and the fact that its value is related to shifts in risk-taking attitudes of international investors (Bruno and Shin, 2015). The same, however, is not true for the Polish zloty and the euro. This currency pair again stands as an outlier and exhibits a comparatively large risk premium error.

The third component, the CIP error, turns out to be of much lesser importance than the two remaining factors. Its impact on β estimates is slightly larger only for Romania and the

US, as well as for Czechia and the US relationships. The results do not indicate a considerable covered interest parity failure that could appear, for example, due to a rise in risk related to individual financial institutions. It is despite the fact that our calculations are based on domestic interest rates rather than assets that are comparable in terms of all characteristics, such as the Euro-currency deposits (see Sarno and Taylor, 2003), which are not fully available for the CEE economies. A limited role of the CIP error speaks to the small transaction costs attributed to CEE currencies. This, in turn, may be attributed to open financial accounts and the absence of capital controls in those economies.

5.3 UIP regressions augmented with risk measures: the CEE economies aggregate

To provide a more detailed picture of the UIP relationships in CEE economies, we estimate a set of augmented regression models. Because there are numerous regression specifications for all CEE economies, their aggregate, and three reference currencies (see Section 4), at this point, we compare four regression models estimated for the V3R group before going to a summary of the rest of the results. It must be noted that due to the decidedly similar outcomes that we obtain for the euro area and Switzerland, we do not report the latter case in this and the following subsection. The set of regressions calculated for the Swiss franc reveals a similar pattern of mostly positive values on the estimated UIP coefficient.[4]

Table 2 shows the results for the euro area. Typically, for this reference currency, significant and positive estimates of β in the baseline regression imply the existence of the UIP puzzle. Further specifications, augmented with GARCH components and risk measures, corroborate the initial conclusions regarding the UIP puzzle. They indicate that the risk aversion of market participants is indeed time-varying, as indicated by statically significant estimates of coefficients on the realized volatility of the exchange rate and VIX. What is more, the point estimates on v_t^2 in Model 2 are very close to -0.5, which corresponds to their anticipated value (see Equation 2). The same is true for Kilian's measure of global economic activity. This confirms the previous sections' results, particularly the significant impact of various risk factors on currency excess returns.

At the same time, the inclusion of risk measures improves the overall performance of the regression models by fixing the autocorrelation of the residuals that we presented before. The specifications in which the GARCH error terms are introduced tend to alleviate the autocorrelation of squared residuals. Their results show that the extension of the UIP regressions is justified from the point of view of the time series properties used in the study, and the UIP relationship likely contains a non-constant volatility term. Interestingly, even though the extensions to the baseline regression seem to be substantiated, the estimates of β do not change. The estimates are always statistically significant and fall within a range from 0.945 to 1.358.

Table 3 presents the detailed regression results for the V3R group vis-à-vis the US. Unlike in the case of the euro area, the results show the lack of a significant positive relationship between interest rate differentials between the entire CEE economies and the US, $i_t - i_t^*$, and the excess return against the dollar, ρ_{t+1}, although the point estimates of β somewhat increase

[4]Those results, along with all the empirical results not reported in the paper, are available upon request.

Table 2: Model estimates for the V3R group vis-à-vis the euro area

	const	$i_t - i_t^*$	v_{t+1}^2	Δvix_{t+1}	$kilian_{t+1}$	Δepu_{t+1}	Q_1	Q_{10}	Q_1^2	Q_{10}^2
M1: Baseline	-0.052	1.009					12.636	24.904	40.025	78.369
	[0.715]	[0.000]					[0.000]	[0.006]	[0.000]	[0.000]
M2: RV	0.218	1.186	-0.499				4.660	13.457	10.275	98.739
	[0.068]	[0.000]	[0.001]				[0.031]	[0.199]	[0.001]	[0.000]
M3: GARCH	0.011	0.945					3.629	15.756	0.056	10.500
	[0.929]	[0.001]					[0.057]	[0.107]	[0.813]	[0.398]
M4: GARCH+risk	0.152	1.358	-0.506	-0.095	0.003	-0.002	-0.002	0.030	10.949	0.928
	[0.224]	[0.000]	[0.000]	[0.000]	[0.003]	[0.204]	[0.173]	[0.862]	[0.361]	[0.335]

Notes: dependent variable: ρ_{t+1}, see Equations (16)-(20); p-values in brackets; the OLS estimation (M1, M2) uses HAC robust standard errors; regressions augmented with GARCH term (M3, M4) assume the generalized error distribution (GED) of the conditional errors; Q_1, Q_{10} and Q_1^2, Q_{10}^2 indicate the Ljung-Box statistics for the first 1 and 10 autocorrelations of standardized residuals and squared standardized residuals.

in models M2–M4 relative to M1. The estimated signs on risk measure coefficients align with our expectations, but the coefficient on the realized volatility is markedly higher than -0.5. An increase in the global risk measures is negatively related to the excess returns on CEE currency, even though the UIP holds in this case for the USD. As indicated by regression diagnostics, the overall misspecification of the baseline regression seems smaller than in the euro case. In particular, there is no notable autocorrelation in the baseline regression residuals. Still, the risk measures and the GARCH functions improve the overall performance of the models.

Table 3: Model estimates for the V3R group vis-à-vis the US

	const	$i_t - i_t^*$	v_{t+1}^2	Δvix_{t+1}	$kilian_{t+1}$	Δepu_{t+1}	Q_1	Q_{10}	Q_1^2	Q_{10}^2
M1: Baseline	0.117	0.432					0.398	12.839	0.015	41.841
	[0.655]	[0.395]					[0.528]	[0.233]	[0.902]	[0.000]
M2: RV	0.564	0.534	-0.321				0.010	11.633	0.288	52.978
	[0.036]	[0.302]	[0.009]				[0.920]	[0.310]	[0.591]	[0.000]
M3: GARCH	0.151	0.597					0.845	6.855	0.909	15.651
	[0.547]	[0.346]					[0.358]	[0.739]	[0.341]	[0.110]
M4: GARCH+risk	0.289	0.596	-0.217	-0.163	0.005	-0.005	0.045	4.069	2.638	20.954
	[0.278]	[0.360]	[0.000]	[0.000]	[0.039]	[0.264]	[0.131]	[0.832]	[0.944]	[0.104]

Notes: see Table 2.

The differences between the results for the V3R that we and the two reference currencies must be considered surprising for at least two reasons. First, various risk measures do not account for the failure of the UIP when the euro is the benchmark currency, but they tend to improve the baseline specification. Second, there is a consistent difference between the euro and the US dollar, with the former indicating the UIP puzzle. As we turn to the summary of the results, we ask whether the same conclusions can be made for other currency pairs.

5.4 UIP regressions augmented with risk measures: a summary

This subsection reports a summary of β estimates in the baseline and extended regression models. The bar plot in Figure 2 depicts 90-percent confidence intervals for the CEE economies and the euro area. Almost all intervals are located in the positive territory and do not include zero, even in the most expanded form. The exception to this pattern appears in the Czech case, in which the confidence band are particularly wide and shift above zero in Model 3. This indicates an enduring UIP puzzle for all other economies, including the V3R aggregate. The inclusion of the realized volatility and the global risk proxies, as well as the GARCH errors, does not substantially change the estimates of β. On average, the CEE currencies appreciate when their domestic interest rates are higher than in the euro area, which violates the UIP. When we omit the singular Czech case, the confidence intervals for β shift more for Hungary and Poland than for Romania and the V3R. As we see from Models 1 to 4, the confidence bands tend to become slightly narrower, indicating more precise estimates of β, but in some cases, such as the V3R, they remain roughly the same width.

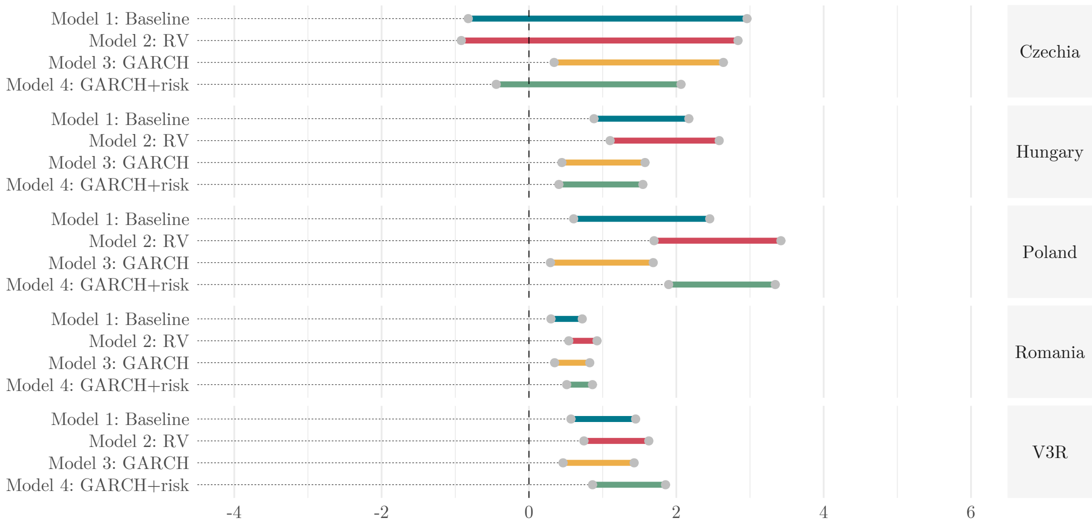

Figure 2: 90-percent confidence intervals of β in the UIP regressions for the CEE economies and the euro area

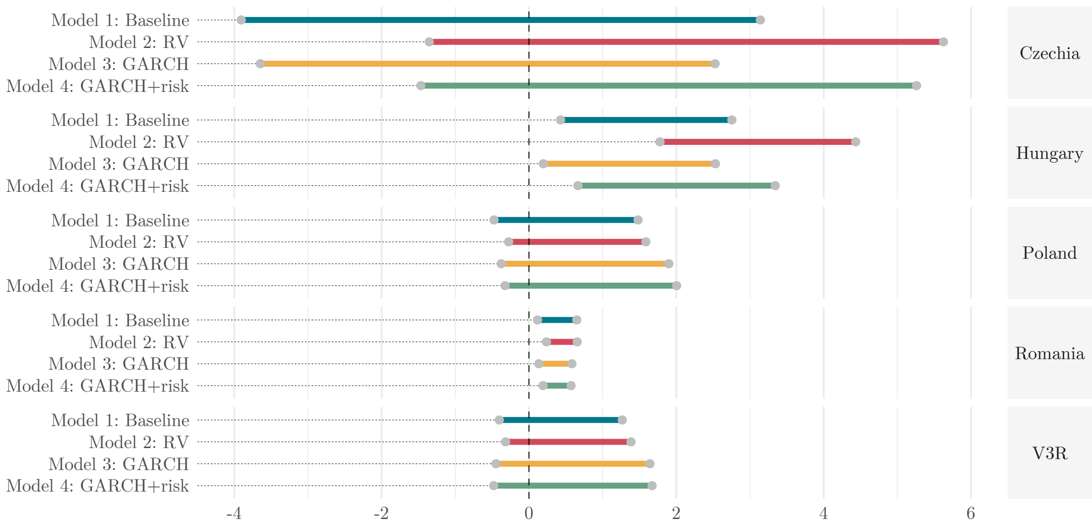

Figure 3: 90-percent confidence intervals of β in the UIP regressions for the CEE economies and the US

The bar plots constructed for the regressions that use the US dollar as the reference currency reveal further differences in the distribution of β in the Fama regression across economies and their aggregates (Figure 3). Even though the confidence bands are even wider than for the euro area, we find fairly strong evidence that the UIP vis-à-vis the US holds in Czechia and Poland. It translates into similar results for the V3R aggregate, in which those two economies weigh the most. On the contrary, an indication of deviations from the UIP is found for Hungary and Romania. However, for both economies, the confidence intervals exclude zero by a minimal margin. In the Hungarian case, the UIP puzzle becomes even more visible when β is estimated in a model with realized volatility and risk measures.

The summary of the results we obtain for the CEE economies shows that the estimated UIP slopes differ across countries, but there is little variation within economies and the augmented specifications produce comparable results for a given economy or the aggregate of the countries. Even though subsequent specifications, the ones that control for various risk measures or include GARCH errors, may be justified, both from economic and statistical points of view, they do not alter our previous conclusions on the UIP puzzle.

5.5 Accounting for differences between the UIP regressions vis-à-vis the euro area and the US

One of the noteworthy features of our results is the inconsistency between the UIP test outcomes for the euro area and the US. In particular, the evidence against the UIP is stronger in the former case. The existence of the UIP puzzle for the euro may be considered surprising since all the CEE economies exhibit strong economic and financial ties with the EMU, specifically with Germany. However, when we recognize that the empirical studies reviewed in Section 2 consistently indicate UIP deviations when the dollar is the reference currency rather than confirm the UIP, it is weaker evidence of the puzzle for the dollar that raises more questions. One compelling rationale for the lack of the UIP puzzle for the dollar was put forward in a recent study by Engel et al. (2021). This explanation is related to a major shift in the US monetary policy after the GFC. In this period, the Fed cut the interest rates almost to zero and turned to forward guidance, quantitative easing, and other non-standard programs. As it started to react to economic and financial developments, including inflation rates, with unconventional tools, the interest rate ceased to be its primary policy instrument. Hence, interest rates in the US fluctuated very little, undermining their predictive power for the dollar exchange rates.

As described in Section 4, in this study, we investigate an explanation for the differences between the results obtained with the euro and the dollar as reference currencies based on a 'triangular' relationship among CEEs, the euro area, and the US. Specifically, we expand the baseline regression specification with the movements in the USD/EUR exchange rate. The model with the excess return on the US dollar against the euro and the external interest rate differential (Model 5) can potentially be better suited to explain excess returns on CEE currencies (see Equation 21). Given the good statistical properties of the GARCH-GED model found in the previous subsections, we stick to such a framework when expanding the models.

Table 4 reports the estimation results for Model 5. They can be summarised in two points. First, as expected, the coefficients on external excess returns (ρ_{t+1}^{US-EA}) are highly significant in all CEE countries. In contrast, the external interest rate differential ($i_t^{US}-i_t^{EA}$), which we include in the regressions only when it is significant at the 0.1 level, seems less important. Moreover, the external excess return is correctly signed, i.e., when it increases, the excess returns on CEE currencies tend to decrease. Interestingly, the transmission effect is much stronger for the excess return against the US dollar than the euro. In the former case, the coefficients are close to -1, while for the euro, the lowest estimate is -0.156 (Poland). This indicates that the CEE excess returns against the dollar mimic, to a large extent, those of the euro against the US. In the latter case, the coefficients are small and negative, which signals that the sensitivity of CEE currencies to global factors is greater than that of the euro. These findings are reasonable, given that the trade and financial linkages between CEE countries and the euro area are tighter than between CEE countries and the US.

Table 4: UIP regressions in CEE economies and dollar-euro excess returns

	$const$	$i_t - i_t^*$	ρ_{t+1}^{US-EA}	$i_t^{US} - i_t^{EA}$	Q_1	Q_{10}	Q_1^2	Q_{10}^2
Euro area								
Czechia	0.098	1.476	-0.092		0.001	4.340	0.151	2.757
	[0.151]	[0.038]	[0.000]		[0.975]	[0.987]	[0.698]	[0.987]
Hungary	0.002	1.181	-0.045	-1.538	0.538	16.184	0.225	4.816
	[0.990]	[0.000]	[0.152]	[0.044]	[0.463]	[0.903]	[0.635]	[0.903]
Poland	0.020	1.192	-0.156		2.684	12.803	1.465	7.567
	[0.911]	[0.008]	[0.000]		[0.101]	[0.671]	[0.226]	[0.671]
Romania	-0.012	0.607	-0.025		0.360	7.837	0.008	9.594
	[0.863]	[0.000]	[0.321]		[0.548]	[0.477]	[0.931]	[0.477]
V3R	0.012	1.044	-0.119		3.388	16.217	0.836	6.446
	[0.914]	[0.000]	[0.000]		[0.066]	[0.777]	[0.361]	[0.778]
United States								
Czechia	0.198	0.005	-1.070		1.303	4.725	0.024	4.941
	[0.014]	[0.994]	[0.000]		[0.254]	[0.895]	[0.878]	[0.895]
Hungary	0.089	1.039	-1.049		0.107	15.468	0.877	8.506
	[0.493]	[0.001]	[0.000]		[0.744]	[0.580]	[0.349]	[0.580]
Poland	-0.080	1.201	-1.121	2.456	0.508	12.920	1.305	11.577
	[0.672]	[0.005]	[0.000]	[0.048]	[0.476]	[0.314]	[0.253]	[0.314]
Romania	-0.104	0.598	-0.934	1.244	1.026	7.245	2.215	7.592
	[0.318]	[0.000]	[0.000]	[0.078]	[0.311]	[0.669]	[0.137]	[0.669]
V3R	-0.022	0.983	-1.048	1.556	0.471	13.609	1.084	10.714
	[0.861]	[0.001]	[0.000]	[0.062]	[0.493]	[0.380]	[0.298]	[0.381]

Notes: dependent variable: ρ_{t+1}, see Equation (21); regressions augmented with GARCH error terms and the generalized error distribution (GED) of conditional errors; Q_1, Q_{10} and Q_1^2, Q_{10}^2 indicate the Ljung-Box statistics for the first 1 and 10 autocorrelations of standardized residuals and squared standardized residuals; the interest rate different $i_t^{US} - i_t^{EA}$ is included only when significant at the 0.1 level.

Second, the regression results offer some additional insights into the UIP puzzle. Under a simple GARCH framework, we found that the results for Czechia, Poland, and the V3R aggregate were incompatible across alternative reference currencies, i.e., we observed the puzzle for the euro but not for the US dollar (see Model 3 in subsection 5.4). Once we augment the regressions with the external excess return, the results do not change qualitatively for the excess return against the euro but do change vis-à-vis the US dollar. In the former case, the puzzle still holds for all currencies, whereas in the latter case, the puzzle appears in Poland and the V3R aggregate. Our interpretation is that the inconsistency of the findings on the UIP puzzle across alternative reference currencies stems from the conventional omitted variable problem. Czechia is the only country in which the incompatibility remains and calls for an additional explanation. It is discussed in detail as a first extension to the main part of the paper in Section 6.1.

6 Extensions

This section contains three extensions to our main results using additional data. In the Czech case, we consider the Czech National Bank's (CNB) foreign exchange interventions conducted while keeping the exchange rate pegged to the euro. For Poland, we augment the UIP regressions with survey-based expectations of the PLN/EUR and PLN/USD exchange rates. The third extension investigates the rare disaster hypothesis with option-implied risk reversals of CEE currencies.

6.1 Czechia: central bank foreign exchange interventions

Of the four CEE economies we investigate in this paper, Czechia is the only one that steered away from a standard interest rate policy by introducing an explicit form of the exchange rate control. Starting in 2008, the CNB lowered interest rates by roughly four percentage points, and in late 2012 its main policy reached 0.05%. At the same time, it introduced forward guidance to communicate that its monetary policy would remain expansionary. However, the Czech economy was still experiencing stark consequences of the GFC and the euro area debt crisis, most notably the deflationary tendencies. Searching for extraordinary policy options but unable to bring interest rates below zero and reluctant to start quantitative easing, the CNB decided to alleviate the zero lower bound by deploying an additional instrument, the exchange rate control (see Bruha and Tonner, 2018). In November 2013, the bank announced a one-sided peg of the Czech koruna to the euro. It aimed to prevent appreciation of the koruna below the EUR/CZK 27 level in order to boost the national economy and bring the inflation rate closer to the 2% target. While maintaining this target, the CNB carried out large-scale foreign exchange interventions. The exchange rate commitment was terminated in April 2017.

To account for the unique features of Czech monetary policy post-2012 and their possible implication for the validity of the UIP tests, we run a set of regressions that encompass foreign exchange interventions of the CNB. The regressions are estimated for the entire period of 1999-2019 and they build on Model 5 estimated with GARCH-GED specification and the euro-dollar excess return. However, they contain two additional dummy variables. The first one, which we denote as 'positive' foreign exchange interventions, takes the value of 1 in a month when the CNB conducted net purchases of the euro-denominated assets and zero otherwise. The second dummy, a proxy for 'negative' interventions, is defined in an analogous manner. Data on FX interventions are sourced from the CNB website.

The outcome of the extended regression models is presented in Table 5. The table reports the results of two specifications for each reference currency. As discussed before, in the first of them, without additional variables, the estimated slope coefficient is relatively high and significant for the euro area and close to zero and insignificant for the US, indicating the UIP puzzle in the former case. The additional regression specifications (the second row for each currency), however, produce considerably higher, significant estimates of coefficients on the interest rate differential. The point estimate of β jumps from 1.476 to 2.655 for the euro and – quite remarkably – from 0.005 to 2.376 for the dollar. These models consistently demonstrate that net sales ($negfxi_t$) rather than net purchases ($posfxi_t$) of euro-denominated assets by the CNB had a substantial impact on the excess returns of the Czech koruna. This shows that once we account for the central bank's FX interventions, the UIP puzzle emerges in the instance of Czechia vis-à-vis the dollar and becomes stronger when the euro is the reference currency. Hence, the Czech case, a clear outlier in previous tests, now matches the results for other CEE economies. This suggests that the period when the CNB kept its currency pegged to the euro is related to a substantial distortion in the UIP condition. There are at least three plausible explanations for this finding.

Table 5: The case of Czechia: central bank foreign exchange interventions

	$const$	$i_t - i_t^*$	ρ_{t+1}^{US-EA}	$i_t^{US} - i_t^{EA}$	$posfx_t$	$negfx_t$	Q_1	Q_{10}	Q_1^2	Q_{10}^2
EA	0.098	1.476	-0.092				0.001	4.340	0.151	2.757
	[0.151]	[0.038]	[0.000]				[0.975]	[0.987]	[0.698]	[0.987]
	-0.027	2.655	-0.057		-0.003	0.482	0.010	4.516	0.131	2.169
	[0.789]	[0.002]	[0.012]		[0.979]	[0.002]	[0.921]	[0.995]	[0.718]	[0.995]
US	0.198	0.005	-1.070				1.303	4.725	0.024	4.941
	[0.014]	[0.994]	[0.000]				[0.254]	[0.895]	[0.878]	[0.895]
	-0.074	2.376	-1.070	2.482	0.062	0.470	2.025	5.781	0.000	5.169
	[0.610]	[0.051]	[0.000]	[0.039]	[0.761]	[0.018]	[0.155]	[0.880]	[0.992]	[0.880]

Notes: dependent variable: ρ_{t+1}, see Table 4; indicator variables $posfx_t$ and $negfx_t$ take the value of "1" in months when the Czech National Bank conducted net purchases and net sales of euro-denominated assets, respectively, and "0" when there were no foreign-exchange interventions.

First, foreign exchange interventions may have a direct impact on the returns of foreign versus domestic assets and capital flows to the Czech economy, possibly changing the default or liquidity risk of assets denominated in the Czech koruna. In fact, this is confirmed by Frait and Mora (2020), who point out to shifts in capital flows, the yield curve of Czech securities, and the liquidity of the domestic banking sector during this period.

Second, by setting an exchange rate peg to the euro, the CNB explicitly changed its reaction function to the inflation rate and output gap developments in the Czech economy, which the bank actively communicated. After more than a decade of direct inflation targeting and a floating exchange rate regime, the CNB veered from its regular monetary policy strategy, as it regarded the exchange rate commitment as an instrument to exit the zero-lower bound. A set of simulations conducted by Bruha and Tonner (2018) indicates that the CNB's policy positively impacted the inflation rate in the country. In particular, it prevented the core inflation rate from falling below zero (see also Caselli, 2017).

Third, the koruna-euro exchange rate floor brought substantial operational adjustments in CNB's monetary policy and the bank's balance sheet. Together with the central bank's reaction function, its specific operational procedures, such as the type of open market operations it conducts, may have an impact on the UIP condition (see Backus et al., 2010). Given that the CNB's exchange rate commitment was one-sided, it did not have to intervene in the FX market in a systematic manner. Large, sterilized operations were needed when the peg was introduced and then after 2015, especially in 2017, when the ECB launched the quantitative easing policies (Czech National Bank, 2017). Such operations, in turn, may have diminished the relative riskiness of the koruna vis-à-vis the major currencies and influenced the UIP.

6.2 Poland: survey-based exchange rate forecasts

The second extension to the baseline results consists of a case study we conduct for Poland using professional exchange rate forecasts to approximate market expectations. The empirical foundation of this exercise comes from a market participants survey conducted by "Rzeczpospolita", a major Polish nationwide daily newspaper specializing in economic and legal issues. On a monthly basis, analysts working at the largest financial institutions located in Poland – banks, brokers, or financial groups – are asked to submit their predictions on macroeconomic and financial variables, among them the PLN/EUR and the PLN/USD exchange rates. The earliest entry in this survey comes from December 2013, with forecasts for January 2014, and the database ends in December 2019. The panel of respondents changes over time, with a maximum number of 41 financial analysts. This dataset provides us with a rare opportunity to re-run the UIP regression for Poland, in a shorter timespan but using a reliable proxy of actual market expectations and without imposing additional assumptions on the UIP error decomposition, as in Section 5.2.

To keep the empirical results comparable with the previous parts of the paper, we introduce some adjustments to the time series used in the regressions. The surveyed financial analysts are asked to forecast the *average* level of the official exchange rates quoted by the National Bank of Poland for the following month. Hence, in this part of the paper, we also use the averages of these particular exchange rate quotations instead of the end-of-month market entries. Based on individual survey responses, we calculate the median of disaggregated forecasts and use it as a measure of market expectations. Next, we estimate the basic UIP model, the one without any additional explanatory variables, and decompose the regression residuals using forward rates and expectations, according to Equations (6) to (8).

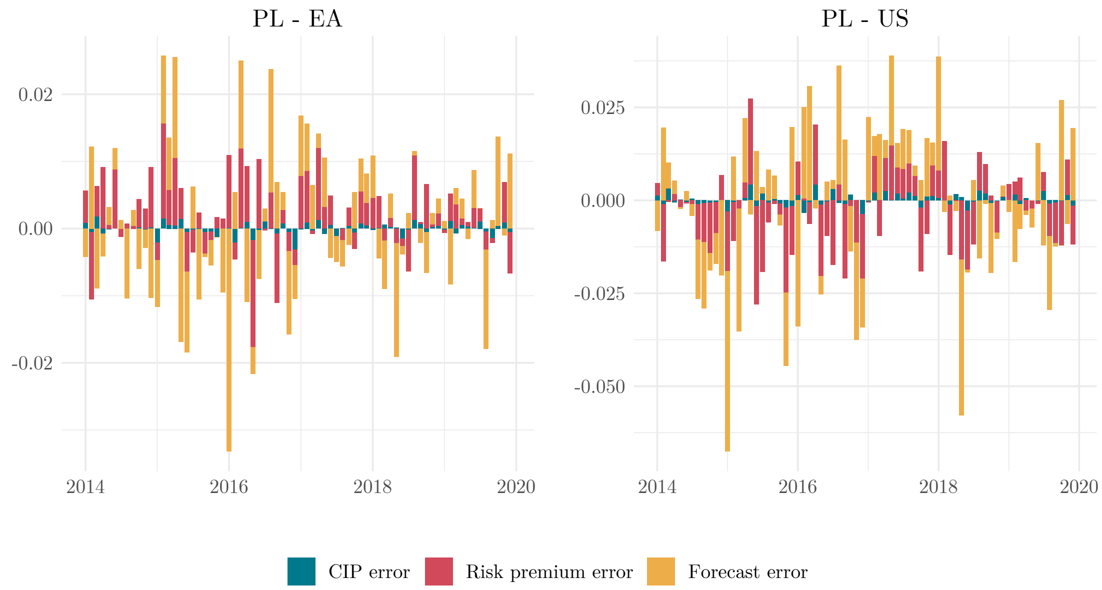

Figure 4: UIP error decomposition for Poland using survey data on exchange rate forecasts

Figure 4 plots the contribution of CIP, risk premium and forecast errors to the overall UIP regression error. The average forecast error value is slightly negative for the euro and positive for the US dollar exchange rates. The converse is true concerning the risk premium error. The CIP errors are at least an order of magnitude smaller than the remaining components. For both reference currencies, the risk premium error is highly variable, but the periods of positive or negative errors seem to be clustered and follow each other in longer 'swings', especially for the US. For example, over 2015 and 2016, there are only six months when the risk premium error of the US regression was positive, while in 2017 alone, the error was greater than zero for eight months. The forecast error tends to periodically escalate for both currency pairs, possibly in periods of elevated economic or political uncertainty.

We may next use the respective errors to explain deviations of the estimated UIP slope coefficients from the theoretical value of $\beta = 0$. The results of this decomposition are depicted in Figure 5. Point estimates of β stand at 5.445 for the euro area and -2.484 for the US, but neither are significant at the 10% significance level. The forecast error turns out to be dominant for the euro, whereas the combined CIP and risk premium errors contribute only to 0.792 (or 14%) of the departure from the UIP. For the US reference, however, the risk premium plays a leading role. More than 80% of deviation from zero may be explained by this error component. Unlike for the EUR, all three components drag the estimate of β downwards, into the negative territory. Hence, we once again obtain a distinct picture of the UIP relationship for each major foreign currencies.

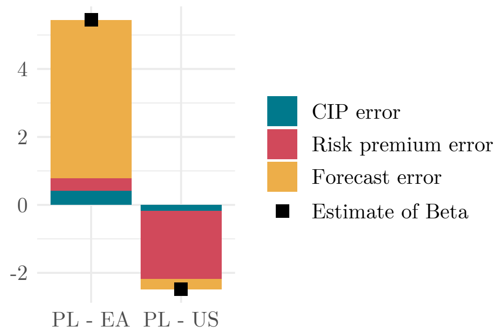

Figure 5: Decomposition of β estimates for Poland using survey data on exchange rate forecasts

Additionally, as β estimates are insignificant, the regression outcomes imply that survey-based exchange rate expectations on PLN/EUR and PLN/USD are not informed by the UIP relationship, at least between 2014 to 2019. Professional forecasters in large financial institutions in Poland seem not to be guided by interest rate differentials when formulating one-month-ahead exchange rate forecasts. In this regard, our results stand at odds with Cuestas et al. (2015), who find some evidence that forecasters 'believe' in the UIP hypothesis in the Polish zloty-euro market. However, their study covers only the euro exchange rates of CEE currencies between 2007 and 2014 and employs a different exchange rate forecast data source. Hence, our findings suggest that post-2014, some macroeconomic or financial factors other than the UIP relationship may have a more substantial impact on the market forecasts of the zloty, both vis-à-vis the euro and the dollar.

6.3 CEE currencies and the rare disaster hypothesis

The final extension builds on the rare disaster hypothesis (RDH) put forward by Farhi and Gabaix (2016). As demonstrated in Section 3, the coefficient in the Fama-type regressions can deviate from the value implied by the UIP because they may be understood as a weighted average of the coefficient derived from the RDH model with no inflation and the UIP coefficient. The deviation, in turn, is driven by the variability of inflation differential and the variability of the relative country resilience.

Before we proceed to analyze these variabilities, it is worthwhile to check if financial market participants indeed consider the CEE currencies riskier than the major currencies, as required by the RDH. Following Farhi and Gabaix (2016), we use notions in option theory to quantify the currency riskiness. The options that protect against a crash of the exchange rate of currency i against currency j are out-of-the-money puts on currency i. In turn, out-of-the-money calls on currency i protect against a crash of the exchange rate of j against i. If currency i is riskier than currency j, then the implied volatility of out-of-the-money puts is higher than the implied volatility of out-of-the-money calls. This pattern is known as a 'smirk'. The risk reversal (RR) measure is defined as a difference between the implied volatility of an out-of-the-money put and the implied volatility of an out-of-the-money call, and it can be used as an indicator of currency riskiness. The higher the RR of currency i, the riskier this currency is.[5]

The availability of data on the implied volatility of currency options is not uniform across CEE countries. Relatively long time series can be obtained only for the Czech, Hungarian, and Polish currencies vis-à-vis the euro. Thus, the focus of this extension is on these three currency pairs. The end-of-month daily data on implied volatilities of put and call options in the period spanning from February 2006 to December 2019 are retrieved from Refinitiv Datastream. In Figure 6, the RRs for three CEE currencies against the euro are depicted. The RRs for the Hungarian and Polish currencies are positive, which confirms that these currencies are considered riskier than the euro. The same, in principle, holds for the Czech koruna, although there are two periods when the risk of appreciation is perceived as higher than that of depreciation. The first one is a very short episode in the summer of 2008, abruptly ended by an outbreak of the global financial crisis. The second one is longer and lasts from September 2015 to March 2017. It overlapped with the large-scale foreign exchange interventions carried out by the CNB to suppress the tendency of its currency to appreciate in response to quantitative easing deployed by the ECB and continued favourable developments in the domestic economy (see point 6.1 and Czech National Bank, 2017). The RR of the Czech currency increased in April 2017 when the exchange rate floor vis-à-vis the euro was abandoned. Overall, in line with intuition, CEE currencies can be considered riskier than the euro.

To provide more insights on the estimated UIP coefficients, we proceed in four steps. First, the baseline Fama regressions are re-estimated on the 2006-2019 sample. The β coefficients obtained for this sample are reported in Table 5 in column (1). Second, the variability of inflation differential is calculated using the CPI inflation rates for CEE countries and the HICP inflation rate for the euro area (data are from the International Financial Statistics database). The results are reported in column (2). Third, the variability in the relative country resilience is proxied with the variability of the RR.[6] The results are tabulated in column (3). In Section 3, it is explained that what matters for the deviation of the β coefficient from the level implied

[5] To pin down the degree of moneyness, the delta of an option equalled to +/-0.25 is used. The delta is a derivative of the option price with respect to the spot price of a currency. For details see, e.g., Garman and Kohlhagen (1983) and Farhi and Gabaix (2016).

[6] In the calibrated model of Farhi and Gabaix (2016), the latter is a multiple of the former, and the scaling constant is 1.57^2.

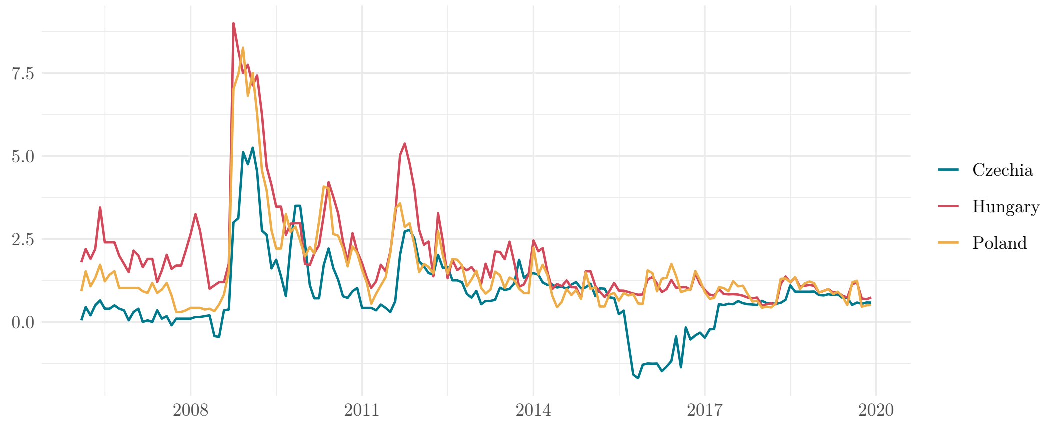

Figure 6: Risk reversals of CEE currencies vis-à-vis the euro area

Notes: Following the convention adopted in the Refinitiv Datastream, the RRs of CEE currencies are calculated as differences in the implied volatility of 25-delta out-of-the-money call options on the euro and the implied volatility of 25-delta out-of-the-money put options on the euro.

by the UIP is the importance of variability in inflation differential relative to variability in the relative country resilience. Thus, in the fourth step, the ratio of these variabilities is calculated. It is reported in column (4). For ease of interpretation, the weight on the coefficient implied by the UIP is also tabulated in column (5). This measure conveys the same information as the ratio of variabilities in column (4) but is easier to apply to the UIP puzzle analysis.

Table 6: Inflation variability, risk reversals, and the UIP puzzle for CEE currencies and the euro

	Estimated β coefficient	Inflation rate variability	Risk reversal variability	Ratio of inflation and RR vars	Implied weight on the UIP coefficient
	(1)	(2)	(3)	(4)	(5)
Sample A: February 2006 – December 2019					
Czechia	0.821	0.899	1.257	0.716	0.165
Hungary	1.360	3.501	2.307	1.517	0.295
Poland	0.046	1.539	1.857	0.829	0.186
Sample B: January 2010 – December 2019					
Czechia	2.626	0.487	0.761	0.641	0.150
Hungary	1.173	1.814	0.881	2.058	0.361
Poland	1.474	0.940	0.553	1.700	0.319

Notes: The implied weight, $1-\nu$, is implicitly defined in Equation (15). See the main text for further explanations.

The observed association between the estimated β coefficient and the ratio of variabilities is only roughly in line with the negative association implied by the model. The same holds for the association between the β coefficient and the weight on the coefficient implied by the UIP. It is because both the coefficient in the regression for the Hungarian forint and the weight are the largest. When read without taking into account the results for Hungary, the findings for the other two currencies are in line with the RDH: the estimated coefficient for Czechia deviates from the zero-level implied by the UIP more than the coefficient for Poland (0.821 vs 0.046), and at the same time the weight of the coefficient implied by the UIP relationship is smaller in Czechia than in Poland (0.165 vs 0.186).

One can make a fair point that the differences between CEE countries are not sharp since the estimated β is not statistically significant for any country. Following this concern, we repeated the whole exercise on the sample starting in January 2010. The results are reported in Table 6 in the panel entitled Sample B. The β coefficient is found to be statistically significant for Czechia but not for the other two countries, which makes the comparative analysis more justified.[7] In line with the RDH, the weight of the coefficient implied by the UIP is smaller in Czechia (0.150) than in the other two CEE economies (above 0.300).

Our findings, although preliminary, demonstrate the importance of currency resilience to rare disasters in driving the deviations from the UIP relationship. As such, they are consistent with the evidence that the returns on currencies perceived as risky are determined by rare disaster risks documented for the dollar-based exchange rates of the BRICS countries by Gupta et al. (2019). Moreover, the results fit the conjecture put forward by Ismailov and Rossi (2018) that the relationship between exchange rates and interest rate differentials is blurred in periods of high uncertainty that can be potentially linked to rare disasters. In a related context of exchange rate predictability, Bak and Park (2022) use implied volatility measures of at-the-money currency options for major advanced economies and demonstrate that models with the estimated risk premium track the actual exchange rate changes more closely than those based on the Fama regression framework.

To sum up, even though our findings lend some support to the relevance of the RDH to the CEE currencies, they should be interpreted with caution. Taking into account that the time series employed are not too long, we consider the results as tentative and requiring further research.

7 Conclusions

This paper aimed to examine the evidence of the UIP puzzle in four CEE economies: Czechia, Hungary, Poland, and Romania. Starting with baseline Fama regressions, we investigated the forward premium anomaly using the euro, the US dollar, and the Swiss franc as reference currencies, covering the period from 1999 to 2019. After checking for structural breaks in baseline regressions and providing the error decomposition of deviations from the UIP, we put forward a sequence of UIP models augmented with various risk measures and estimated with GARCH error terms. We also offered several extensions to the common UIP tests, in particular models augmented with the US–euro area risk premium. Additionally, we considered the impact of exchange rate interventions in Czechia and market expectations in Poland on the interest rate parities. We also investigated the role of the FX implied volatility measures in explaining the 'riskiness' of CEE currencies with the rare disasters hypothesis.

First, we demonstrate that even though the risk-based measures, such as the realized volatility of the exchange rate and the VIX, essentially improve the baseline UIP regressions, they do not alter the overall conclusions on the puzzle's existence (or lack thereof). Second, we show that the choice of the reference currency matters for the outcome of the interest parity tests in the CEE economies. The UIP puzzle is stronger for the euro and the Swiss franc than for the US dollar.

[7] The samples starting in 2007, 2008, and 2009 were also tested. The coefficients were insignificant in the first two samples. In the third sample, the coefficient was significant for Czechia and Hungary and insignificant for Poland. At the same time, however, the coefficients for Hungary and Poland were very close to each other. The difference in their statistical significance was due to the precision of estimation rather than economic considerations. Thus, these results are not reported, albeit they are available upon request.

Third, we provide a plausible explanation for the inconsistencies in the UIP test that relies on the role of the euro-dollar risk premium for interest parities of CEE economies versus those two major currencies but especially the US dollar. Additionally, we show that the Czech koruna exchange rate peg to the euro from 2013 to 2017 and related FX market interventions had a significant impact on the UIP. A case study on Poland confirms that forecast and risk premium errors play leading roles in UIP deviations, while the directly measured exchange rate expectations do not seem to be informed by the UIP relationship. The option-implied risk reversals for the CEE currencies and the euro indicate that crash risks are priced into their exchange rates. Hence, the limited resilience of CEE economies to world disasters may plausibly explain departures from the UIP that we observe for their currencies.

Some limitations of this study should be acknowledged. A major one comes from the availability of the datasets we use to investigate forward rates, market expectations, FX options, and risk reversals. Moreover, even though we explained the differences in the outcomes of the UIP tests against the US dollar and the euro, the result is only partially satisfactory. On the one hand, we made the results consistent across base currencies and with the literature, but on the other hand, our approach turned out to lend support to the UIP puzzle in all CEE countries. Thus, the puzzle calls for a deeper explanation. Finally, given that the exchange rate can be a part of a policy reaction function, the proper treatment of potential endogeneity could require further adjustment in the approach to testing the UIP employed in this paper.

Promising areas for an extension of this study include an investigation into a possibly asymmetric relationship between interest rates and risk premium, including higher conditional moments of exchange rate returns, currency crashes (Brunnermeier et al., 2008), and rare disasters (Gupta et al., 2019). Given the role of the global risk fluctuations for CEE economies, further investigation into a 'triangular' relationship among the euro-dollar exchange rate, international risk factors (e.g., oil prices), and the CEE exchange rates posits an interesting alternative to a single exchange rate framework (Kębłowski et al., 2020). Empirical puzzles related to the UIP anomaly, such as the one that connects deviations from UIP to levels of exchange rates (Engel, 2016), are also worth exploring for currencies other than the US dollar.

References

Backus, D., Gavazzoni, F., Telmer, C., and Zin, S. (2010). Monetary Policy and the Uncovered Interest Parity Puzzle. *NBER Working Papers*, 16218.

Bak, Y. and Park, C. (2022). Exchange rate predictability, risk premiums, and predictive system. *Economic Modelling*, 116:106024.

Baker, S. R., Bloom, N., and Davis, S. J. (2016). Measuring economic policy uncertainty. *Quarterly Journal of Economics*, 131(4):1593–1636.

Bansal, R. and Dahlquist, M. (2000). The forward premium puzzle: Different tales from developed and emerging economies. *Journal of International Economics*, 51(1):115–144.

Bekaert, G. and Hodrick, R. J. (1993). On biases in the measurement of foreign exchange risk premiums. *Journal of International Money and Finance*, 12(2):115–138.

Bruha, J. and Tonner, J. (2018). An exchange rate floor as an instrument of monetary policy: An ex-post assessment of the Czech experience. *Czech Journal of Economics and Finance (Finance a Uver)*, 68(6):537–549.

Brunnermeier, M. K., Nagel, S., and Pedersen, L. H. (2008). Carry trades and currency crashes. *NBER Macroeconomics Annual*, 23:313–347.

Bruno, V. and Shin, H. S. (2015). Capital flows and the risk-taking channel of monetary policy. *Journal of Monetary Economics*, 71:119–132.

Bussière, M., Chinn, M., Ferrara, L., and Heipertz, J. (2022). The New Fama Puzzle. *IMF Economic Review*, 70(3):451–486.

Caselli, F. (2017). Did the Exchange Rate Floor Prevent Deflation in the Czech Republic? *Review of Economics and Institutions*, 8(2):31.

Chinn, M. D. and Meredith, G. (2004). Monetary policy and long-horizon uncovered interest parity. *IMF Staff Papers*, 51(3):409–430.

Chinn, M. D. and Quayyum, S. (2012). Long Horizon Uncovered Interest Parity Re-Assessed. *NBER Working Papers*, 18482.

Clarida, R., Davis, J., and Pedersen, N. (2009). Currency carry trade regimes: Beyond the Fama regression. *Journal of International Money and Finance*, 28(8):1375–1389.

Cuestas, J. C., Filipozzi, F., and Staehr, K. (2015). Do foreign exchange forecasters believe in Uncovered Interest Parity? *Economics Letters*, 133:92–95.

Cuestas, J. C., Filipozzi, F., and Staehr, K. (2017). Uncovered interest parity in Central and Eastern Europe: Expectations and structural breaks. *Review of International Economics*, 25(4):695–710.

Czech, K. (2017). Testing uncovered interest parity in the PLN/JPY foreign exchange market: a Markov-switching approach. *Quantitative Methods in Economics*, 18(1):18–26.

Czech National Bank (2017). Inflation Report – II/2017. Technical report, Czech National Bank.

Engel, C. (1996). The forward discount anomaly and the risk premium: A survey of recent evidence. *Journal of Empirical Finance*, 3(2):123–192.

Engel, C. (2016). Exchange rates, interest rates, and the risk premium. *American Economic Review*, 106(2):436–474.

Engel, C., Kazakova, E., Wang, M., and Xiang, N. (2021). A Reconsideration of the Failure of Uncovered Interest Parity for the U.S. Dollar. *NBER Working Papers*, 28420.

Engel, C. and Zhu, F. (2019). Exchange rate puzzles: evidence from rigidly fixed nominal exchange rate systems. *BIS Working Papers*, 805.

Fama, E. F. (1984). Forward and spot exchange rates. *Journal of Monetary Economics*, 14(3):319–338.

Farhi, E. and Gabaix, X. (2016). Rare disasters and exchange rates. *Quarterly Journal of Economics*, 131(1):1–52.

Ferreira, P. and Kristoufek, L. (2020). Uncovered interest rate parity through the lens of fractal methods: Evidence from the European Union. *Physica A: Statistical Mechanics and its Applications*, 553:124257.

Filipozzi, F. and Staehr, K. (2012). Uncovered Interest Parity in Central and Eastern Europe: Convergence and the Global Financial Crisis. *Estonian Discussions on Economic Policy*, 20(1).

Flood, R. P. and Rose, A. K. (2002). Uncovered interest parity in crisis. *IMF Staff Papers*, 49(2):252–266.

Frait, J. and Mora, M. (2020). From float to currency floor and back to float: the Czech National Bank's temporary exchange rate commitment Jan. *BIS Papers*, 113:121–131.

Frankel, J. and Poonawala, J. (2010). The forward market in emerging currencies: Less biased than in major currencies. *Journal of International Money and Finance*, 29(3):585–598.

Garman, M. B. and Kohlhagen, S. W. (1983). Foreign currency option values. *Journal of International Money and Finance*, 2(3):231–237.

Gilmore, S. and Hayashi, F. (2011). Emerging market currency excess returns. *American Economic Journal: Macroeconomics*, 3(4):85–111.

Gupta, R., Suleman, T., and Wohar, M. E. (2019). Exchange rate returns and volatility: the role of time-varying rare disaster risks. *European Journal of Finance*, 25(2):190–203.

Hayward, R. and Hölscher, J. (2014). UIP, the Carry Trade and Minsky's Financial Instability Hypothesis in the CEE and CIS. In *Poland and the Eurozone*, pages 111–138. Palgrave Macmillan UK.

Hoffmann, A. (2012). Determinants of carry trades in Central and Eastern Europe. *Applied Financial Economics*, 22(18):1479–1490.

Horobet, A., Dumitrescu, S., and Dumitrescu, D. G. (2010). Implications of Volatility for Uncovered Interest Parity Testing. In *Money, Banking and Financial Markets in Central and Eastern Europe*, pages 253–275. Palgrave Macmillan UK.

Ismailov, A. and Rossi, B. (2018). Uncertainty and deviations from uncovered interest rate parity. *Journal of International Money and Finance*, 88:242–259.

Jiang, C., Li, X. L., Chang, H. L., and Su, C. W. (2013). Uncovered interest parity and risk premium convergence in Central and Eastern European countries. *Economic Modelling*, 33:204–208.

Juselius, K. (1995). Do purchasing power parity and uncovered interest rate parity hold in the long run? An example of likelihood inference in a multivariate time-series model. *Journal of Econometrics*, 69(1):211–240.

Kębłowski, P. (2011). The Behaviour of Exchange Rates in the Central European Countries and Credit Default Risk Premiums. *Central European Journal of Economic Modelling and Econometrics*, 3(4):221–236.

Kębłowski, P., Leszkiewicz-Kędzior, K., and Welfe, A. (2020). Real Exchange Rates, Oil Price Spillover Effects, and Tripolarity. *Eastern European Economics*, 58(5):415–435.

Kilian, L. (2009). Not all oil price shocks are alike: Disentangling demand and supply shocks in the crude oil market. *American Economic Review*, 99(3):1053–1069.

Kumar, V. (2020). Liquidity shocks: A new solution to the forward premium puzzle. *Economic Modelling*, 91:445–454.

Lothian, J. R. and Wu, L. (2011). Uncovered interest-rate parity over the past two centuries. *Journal of International Money and Finance*, 30(3):448–473.

Sarno, L. and Taylor, M. P. (2003). *The economics of exchange rates*. Cambridge University Press.

Thornton, D. L. (2019). Resolving the unbiasedness and forward premium puzzles. *Scottish Journal of Political Economy*, 66(1):5–27.

Triandafil, C. M. and Richter, C. (2012). Testing the UIP Theory in the CEE Countries. Evidence from the GARCH models. *INFER Working Papers*, 8.

Appendix

Appendix A: potential structural breaks in the UIP regressions

In the baseline specification of the Fama regression (Section 5.1), we test whether the UIP holds on average between 1999 and 2019 using monthly data. However, as shown by Bussière et al. (2022), among others, a case may be made for a structural break in this relationship that appears around the GFC (2007-2009). In particular, the UIP may hold strongly during the crisis when investors forcefully adjust their positions, while persistent deviations from the UIP occur during 'normal times' (Brunnermeier et al., 2008; Clarida et al., 2009). In our case, the results may be influenced, for example, by a reversal in capital flows to and from the CEE economies around the GFC (Cuestas et al., 2017). Conceivably, such breaks make the estimates of regression parameters unstable over time. Due to its sensitivity to extreme observations, the slope of our regressions may also switch signs, ranging from large positive to large negative values in some subperiods (Thornton, 2019). What is more, for most of the post-crisis period the interest rates in the US remained near zero, and the adjustment of the interest rate differential part of the testing equation could be driven by unconventional monetary policy shocks in the US (e.g., signalling or portfolio balance channels of the quantitative easing). In the euro area and Switzerland, the policy rates also reached zero or negative values. The European Central Bank (ECB) actively engaged in its non-standard measures post-2012, while the Swiss National Bank (SNB) kept the franc pegged to EUR from 2011 to 2015.

To take into account possible structural breaks in the UIP regressions, we run the Chow breakpoint tests using a grid search for each of the baseline regressions. The resulting F-statistics, along with the probable breakpoint dates, are presented in Table A.1. We find some evidence for significant structural breaks in only three out of 18 tested cases. To some surprise, even though some of the break dates appear around the GFC (Poland–US, V3R–US, and Czechia–Switzerland), the last one alone is significant and only at the 0.1 level. Hence, we do not find considerable support for the conjecture that the GFC had a considerable impact on the UIP regression slopes in the CEE economies.

Table A.1: Structural breaks in the UIP regressions - the Chow test

	EA			US			CH		
	Date	F-stat	p-value	Date	F-stat	p-value	Date	F-stat	p-value
Czechia	2002-06	6.354	0.367	2011-04	5.285	0.515	2008-07	10.596	0.075
Hungary	2016-06	3.103	0.874	2002-02	3.820	0.758	2015-06	6.260	0.379
Poland	2016-03	12.585	0.033	2008-07	6.349	0.368	2015-01	16.054	0.007
Romania	2003-12	7.948	0.210	2014-06	9.762	0.104	2015-01	6.191	0.388
V3R	2004-01	5.436	0.492	2008-07	8.830	0.150	2016-06	6.435	0.358

Notes: the Chow parameter instability test based on the baseline regression model (see Equation 16); estimation period: 1999:12 - 2019:01; first and last 15% of observations are trimmed during the breakpoint search.

At the same time, the Chow tests indicate breaks in the latter part of the sample for Poland. The breakpoint in the regression for the Swiss franc is detected in January 2015, around the time when the SNB discontinued its currency peg to the EUR. This decision culminated in an abrupt shock to the CEE FX markets. Poland is the only case in which we also detect a significant breakpoint for the EUR in March 2016. A possible explanation could be that around this period the monetary policies in Poland and the Eurozone started to become more divergent. While the ECB was engaging in further unconventional policies, including quantitative easing, the National Bank of Poland kept the interest rates unchanged for several years. In general, however we find that structural breaks do not play an important role in the baseline UIP regressions in four CEE economies and their aggregate, V3R.

Appendix B: UIP tests for the euro area and the US

Table A.2 shows the results of the UIP regressions for the euro area and the US using two specifications: the baseline Model 1 and Model 3 estimated with GARCH-GED error terms.

Table A.2: UIP regression results for the US and the euro area

	$const$	$i_t - i_t^*$	Q_1	Q_{10}	Q_1^2	Q_{10}^2
M1: Baseline	-0.040	2.229	0.141	10.857	1.170	41.764
	[0.852]	[0.185]	[0.707]	[0.000]	[0.279]	[0.000]
M3: GARCH	-0.156	3.061	1.739	10.924	0.338	13.396
	[0.380]	[0.031]	[0.187]	[0.202]	[0.561]	[0.202]

Notes: dependent variable: ρ_{t+1}, see Equation (16); estimation period: 1999:01 - 2019:12; p-values in brackets; the OLS estimation (M1) uses HAC robust standard errors; regression augmented with GARCH (M3) assumes the generalized error distribution (GED) of the conditional errors; Q_1, Q_{10} and Q_1^2, Q_{10}^2 indicate the Ljung-Box statistics for the first 1 and 10 autocorrelations of standardized residuals and squared standardized residuals.