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ECONOMETRIC TESTING OF UNCOVERED INTEREST RATE PARITY IN SERBIA

ABSTRACT: *This paper provides econometric evidence of the interest parity puzzle in Serbia over the period 2005–2016. Econometric findings are derived from the following techniques: long-run parameter estimation based on the autoregressive distributed lag model, impulse response function computed from the bivariate vector autoregressive model, and estimation of the two-regime Markov switching parameter model. Our results indicate that a positive interest differential corrected for country risk leads to significant dinar appreciation against the euro. The intensity of this impact is different across sub-periods of low*

exchange rate variability and high variability. Exchange rate movements are found to appreciate more strongly during lower variability episodes. Preliminary econometric investigation of four other European emerging economies documents similar findings only for Romania. Our results suggest that there is a huge incentive for short-term carry trades in Serbia, regardless of substantial risks.

KEY WORDS: *Uncovered interest rate parity, Country risk, Interest parity puzzle, ARDL bounds test, VAR model, Regime switching.*

JEL CLASSIFICATION: F31; F41; E43; C32

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1. INTRODUCTION

In recent years, financial integration has become more visible and global capital flows have become more important. On one hand a growing body of evidence suggests that financial openness in developing countries enhances growth, while on the other hand this increases a country's vulnerability during financial crisis and the possibility of speculative attacks. Emerging economies' interest rates are usually high and are potentially attractive to investors looking to make a profit. Remarkably, most of these economies are characterized by underdeveloped institutions, weak macroeconomic performance, political instability, and volatile economic conditions. These cause higher risk premiums for both currency risk and country risk. Despite the high risk, such markets are attractive to foreign investors for short-run transactions. Therefore, these countries use government restrictions on capital account transactions extensively, which are associated with complex policies regarding capital flows across national borders.

The main objective of this paper is to econometrically identify the opportunity to gain arbitrage profit in Serbia by modelling uncovered interest rate parity (UIP). Because Serbia has had very high interest differentials, even when corrected for country risk (6.23% on average in the past 10 years), it could be a potential target for speculative capital.

We consider changes in the average monthly exchange rate between the Serbian dinar and the euro (FX changes) and the average money market monthly interest rate differential for the period September 2005 to December 2016. The sample period is chosen by data availability. Because expected exchange rates are not available time series, we test the ex post UIP condition that assumes rational expectations. Some authors (Rojas-Suarez and Sotelo 2007; Ferreira 2009) emphasize the significance of country default risk (or country risk) in emerging economies and find a strong relationship between country default risk and interest rates. Hence, we incorporate country risk a priori in the model using the Emerging Market Bond Index (EMBI).

The key econometric results are derived from bounds-testing methodology within the autoregressive distributed lag (ARDL) model advanced by Pesaran et al. (2001). This provides a framework for examining the existence of a long-run relationship. This approach is suitable for both stationary and non-stationary data, and for a mixture of stationary and unit root time series. Our empirical findings support the existence of the interest parity puzzle in Serbia. We also

employ the vector autoregressive (VAR) model as an additional tool to reassess the relationship between FX and interest rate differentials. Based on the computation of the impulse response function, we explore how sensitive variables are to unexpected random shocks in the system throughout time. Finally, to account for the possibility that UIP model parameters may be time varying we use the Markov-switching (MS) model (Hamilton 1989), assuming that they change randomly across different regimes. Such a framework may link switches in parameters to sub-periods of currency depreciation and appreciation.

This study contributes to econometric literature on the Serbian economy in two ways. First, previously a priori country risk has not been considered part of the UIP model. The explanatory power of the model is increased by the inclusion of country risk, which makes statistical inference more reliable. Second, the different econometric time series techniques implemented improve the understanding of the nature of this relationship.

The rest of the paper is organized as follows. Section 2 presents a literature review. Section 3 gives an overview of the theory of interest rate parity. Section 4 describes the time series of our sample and briefly discusses the methodological framework. Econometric results for Serbia are given in Section 5. Preliminary results for some other emerging economies (Czech Republic, Hungary, Poland, and Romania) obtained by similar methodology are discussed and compared with Serbia in Section 6. Concluding remarks are offered in Section 7.

2. LITERATURE REVIEW

There are many papers in the literature that investigate the conditions for UIP, most of them referring to developed countries. UIP has become a favourite theoretical abstraction that is rejected by the data. In the literature it is recognized as an interest parity puzzle (or forward premium puzzle). The interest parity puzzle finds that over the short-term horizon (one week to a quarter) the positive interest differential is associated with an appreciating currency, rather than a depreciating currency as UIP predicts. Fama (1994) and Froot and Thaler (1990) provide early evidence of this rejection.

One of the possible explanations of the puzzle is related to the time-varying risk premium (Engel 1996). Engel (2015) finds that the relatively higher real interest rate strength of a currency is even greater than predicted by the theory of

rational expectations of future short-term nominal interest. Based on Fama regression and the vector equilibrium correction model, Engel produces evidence of risk reversal over time (higher risk in the short run, and lower in the long run) based on monthly panel data for G7 countries relative to the US dollar over the period June 1979 – October 2009. Risk reversal is explained by non-pecuniary liquidity return on assets. Ismailov and Rossi (2017) introduce a new exchange rate uncertainty index into the model and show that UIP fails in five industrialized countries (Canada, Japan, Europe, Switzerland, UK) relative to the US dollar during periods of high uncertainty, but it holds at times when uncertainty is low. They test the ex ante UIP relationship based on monthly data spanning November 1993 to January 2015, considering regression where the constant and slope parameter might be time-varying.

The second explanation is based on exchange rate expectations. The vast majority of the papers assume rational expectations. Juselius and Assenmacher (2017) emphasize that most of the puzzle vanishes if an uncertainty premium (proxied by the persistent PPP gap) in the foreign exchange market is introduced, which supports imperfect knowledge-based expectations instead of rational expectations. This evidence is the result of applying the cointegrated VAR model to Swiss-US parity for 1973–2014 monthly data. Ter Ellen et al. (2013) test heterogeneous expectations using a dataset of survey expectations for four exchange rates ($\$/\text{€}$, $\$/\text{£}$, $\text{¥}/\text{€}$, $\text{¥}/\text{\$}$) on foreign exchange markets for the period January 2003 – February 2008, and show that in the short run (1 and 3 months), positive interest differentials lead to a carry trades strategy, while in the long run (12 months) the expectations are in line with UIP theory. By applying the structural VAR model in three small open economies (Australia, Canada, the UK) and using money market quarterly data (1992 – 2007), Felcser and Vonnak (2014) show that an unexpected shock in the interest rate affects exchange rate appreciation and carry-trade movements. Brunnermeier et al. (2008) arrive at the same results. Also, a potential reason for the puzzle could be the presence of a different group of investors in the market and their interaction (Froot and Frankel 1990; Bacchetta and van Wincoop 2010).

McCallum (1994) considers central bank intervention regarding the UIP condition for three exchange rates ($\$/\text{DM}$, $\$/\text{£}$, and $\text{\$}/\text{¥}$). It is argued that empirical applications of UIP (OLS estimates) give a negative slope estimate due to monetary authorities reacting to the movement in the exchange rate by setting a policy interest rate. In this way, all items (interest differential and expected depreciation) in the UIP condition are determined by the central bank.

Additionally, Chinn and Meredith (2004) expand McCallum's model by including intervention output and inflation for G7 countries in the function of the central bank. They conclude that deviations in UIP are primarily the result of the central bank reaction to shocks in exchange rates in the short run (1 month). Similarly, Lothian and Wu (2012), using French francs and US dollars to the pound sterling, show that in periods of exceptionally high inflation (late 1970s–1980s) the regression slope estimates become negative.

Lothian (2016) assesses the UIP relationship over an extremely long time horizon (two centuries) using panel estimation for 16 developed countries (Australia, Belgium, Canada, Denmark, Finland, France, Germany, Italy, Japan, Netherlands, New Zealand, Norway, Spain, Sweden, Switzerland, United Kingdom) and based on annual long-term bond yields. The results confirm the UIP condition that it holds for a longer time horizon, in line with previous literature (Alexius 2001; Chin 2006).

The empirical evidence for emerging economies is less frequent. Since the UIP condition holds under the assumption of full capital mobility, low transaction costs, and risk neutrality (identical assets in terms of default risk), it is expected that all these characteristics contribute to larger UIP deviations in emerging economies than in developed economies. However, empirical findings do not support this. Some authors find that UIP rather holds for countries with higher inflation rates and higher volatility because high inflation rates are followed by large exchange rate depreciation (pass-through effect) and high interest differentials (Bansal and Dahlquist 2000; Bacchetta and Wincoop 2006; Flood and Rose 2002). Ferreira (2004) shows that monetary authorities' interventions determine both expected exchange rates and the interest differential, which induce a simultaneity bias on the interest differential and expected exchange rate for observed emerging markets. The model proposed by McCallum (1994) is used to test this for five emerging markets (Argentina, Brasil, Chile, Mexico and Turkey) from 1995–2004. Cavoli and Rajan (2006) analyse this phenomenon for East Asian countries from January 1990 to May 1997 and show that large capital inflows in those countries have a negligible effect on UIP deviation due to the partial capital mobility that leads monetary authorities to sterilize substantial parts of that inflow. Also, Ito and Chinn (2007), based on panel estimates for 21 industrial countries (1984–2006), find that capital account openness and greater financial development decrease the wide range of deviations, while greater inflation volatility and higher per capita income increase the size of UIP deviations.

Only a few papers investigate the UIP condition for Central and East European (CEE) countries. In some of them these countries are included as part of a wider group of emerging markets, so that the overall results are very general. Sarmidi and Salleh (2011) put Latin America countries, Asian emerging markets, and CEE countries into the panel and report that the UIP condition holds better for a longer time horizon (12 months) than for shorter periods (1 and 3 months). However, for all time horizons they find exchange rate depreciation, rather than appreciation as appears in the puzzle. The dynamic nature of the UIP condition has also been examined. By comparing the monthly data of 18 industrialized economies and 25 emerging markets (including CEE countries) and following the linear regression approach, Burnside (2014) confirms more UIP failure for industrialized countries than for emerging economies. Although carry trade profitability is demonstrated for both groups, the time varying risk premium only explains the returns in industrialized countries. In emerging markets the returns mainly address the high interest rate differentials. Vasilyev et al. (2017) show that UIP holds better for Russia than for other emerging economies for 2001–2014 monthly data, which is explained by accumulation of foreign exchange reserves. Additionally, Burnside's (2014) results are confirmed, based on panel data for 10 advanced and 15 emerging economies.

Triandufal and Richter (2012) reject the UIP condition for five CEE countries (Bulgaria, Czech Republic, Hungary, Poland, and Romania) using the GARCH model based on money market data over the 1997–2011 period. Filipozzi and Staehr (2012) apply rolling regressions to five CEE countries (Bulgaria, Czech Republic, Hungary, Poland, and Romania) using three-month data and get negative estimates of slope parameters for all countries except Romania.

Jiang et al. (2013) and Cuestas et al. (2015) explore whether UIP deviation is a stationary process for each country using money market monthly data for interest rates and ex post exchange rates. Based on the threshold autoregressive model, Jiang et al. (2013) conclude for seven CEE countries (Bulgaria, Croatia, Czech Republic, Hungary, Poland, Romania, Russia) that UIP deviation is a non-linear stationary process, which confirms the theory; however, the unit root was found for Belarus, Latvia, and Macedonia. Cuestas et al. (2015) apply several unit root tests (with and without structural break analysis) for a similar sample and confirm stationarity with constant terms in most economies.

3. UNCOVERED INTEREST RATE PARITY (UIP): OVERVIEW

The relationship between interest rates and the forward exchange rate is defined as covered interest rate parity (Levi 2005, Ch. 8). In a world without restrictions on capital flows or transaction costs and with profit-oriented market players, in equilibrium Covered Interest rate Parity (CIP) would hold:

$$\frac{1 + i_t}{1 + i_t^*} = \frac{F_t}{S_t} \quad (1)$$

where i_t and i_t^* denote the interest rate between time t and $t+1$ on domestic and foreign assets respectively, S_t is the spot exchange rate at time t (domestic currency price of foreign currency), and F_t is forward exchange rate (domestic currency price of foreign exchange delivered at $t+1$). This would hold due to investors (maximizing profit agents), who always monitor this relationship, trying to exploit mispricing between these four key variables and gaining riskless profit. It applies to countries at the same level of development. The CIP does not hold for foreign exchange rates between an emerging market economy and a developed economy due to the country default risk premium (country risk). For example, the interest differential between the dinar interest rate in Serbia and the euro interest rate in Germany does not equal the forward exchange rate. It must be corrected for country risk. Actually the appropriate forward exchange rate in Serbia is determined by the interest differential between the interest rate on RSD bonds and the interest rate on EUR bonds, both issued by the Serbian government.

Assuming that the forward exchange rate equals the expected future exchange rate ($F_t = S_{t+1}^e$), we end up with the Uncovered Interest rate Parity (UIP) condition that must hold in equilibrium:

$$\frac{1 + i_t}{1 + i_t^*} = \frac{S_{t+1}^e}{S_t} \quad (2)$$

where S_{t+1}^e is expected exchange rate at time $t+1$. Contrary to CIP, UIP does not consider investors who lock in future exchange rates on the forward exchange market, but investors who buy foreign assets and expose themselves to

the risk that the money earned could be squeezed out by unexpected movements in the exchange rates.

Taking the logarithm of (2), we get:

$$s_{t+1}^e - s_t = i_t - i_t^* \tag{3}$$

where the small case letters denote logs.

There is another way to express the interest rate differential, which is important for emerging market economies (f_t denotes log value of F_t):

$$i_t - i_t^* = [(i_t - i_t^*) - (f_t - s_t)] + (f_t - s_{t+1}^e) + (s_{t+1}^e - s_t) \tag{4}$$

If we rearrange the relationship in (4), we get a modified UIP equation:

$$(s_{t+1}^e - s_t) = (i_t - i_t^*) - \underbrace{[(i_t - i_t^*) - (f_t - s_t)]}_{\text{Country risk}} - \underbrace{(f_t - s_{t+1}^e)}_{\text{Currency risk}} \tag{5}$$

We may observe that in (5) the interest rate differential is corrected by two terms, denoted respectively as country risk and currency (exchange rate) risk.

As mentioned above, if we consider two developed countries, then CIP should always hold. However, in emerging markets there is a resounding risk associated with the issuer of the asset. This risk is known as country risk.

The risk associated with variability of the currency itself in terms of another currency is recognized as currency risk. Fama (1984) highlights that the forward exchange rate contract must contain a premium ($F_t = S_{t+1}^e + RP_t$, where RP_t is a risk premium). Throughout the literature it is often identified as the main cause of UIP rejection, primarily for developed countries.

Given that the time series of expected exchange rates are unavailable in emerging economies, the ex post UIP condition is tested:

$$(s_{t+1} - s_t) = (i_t - i_t^*) - [(i_t - i_t^*) - (f_t - s_t)] - (f_t - s_{t+1}) \tag{6}$$

The substitution of ex ante exchange rates with ex post exchange rates is due to the assumption of rational expectations.

Finally, the empirical model considered in this paper takes the following form:

$$s_{t+1} - s_t = \alpha + \beta \cdot (i_t - i_t^* - CR_t) + \xi_t \quad (7)$$

where α and β are parameters, ξ_t is an error term, and CR_t is the country risk.

Currency risk is omitted from the model because time series of the expected FX movements and forward rates are unavailable. However, country risk is included and subtracted from the interest differentials. If it is positive, then the domestic country is higher risk than the foreign country, and vice versa. Most previous studies have not included the country risk.

The following hypothesis is commonly tested in the empirical literature: $H_0: \alpha = 0, \beta = 1$. As already mentioned, parameter β is often estimated to be negative, which is defined in the literature as the interest parity puzzle or forward premium puzzle.

4. DATA DESCRIPTION AND OVERVIEW OF THE METHODOLOGY

This study is based on average money market interest rates and average monthly changes in exchange rates, since financial institutions undertake much of the arbitrage rather than individual investors. The interest rates of Serbian government bonds could also be used, but due to the short time series and low liquidity the results might be spurious. As a measure of interest rates we take the money market rate BELIBOR in Serbia, and EURIBOR for EMU. The exchange rate is represented by the average monthly price of the euro in units of dinars. This data is taken from the National Bank of Serbia website and the Eurostat database. We opt for the Emerging Market Bond Index (EMBI) as a proxy for country risk. EMBI is spread between international government bonds issued by the Serbian government in dollars and US Treasury Bonds. This is not the most appropriate measure, given that debt in the index is more than one year to maturity, while we consider the UIP condition over a one-month horizon. However, this incorporates the relatively high volatility of the Serbian economy in the modelling. The data is obtained from Bloomberg. The sample period is

defined by availability of EMBI index data and covers the period from September 2005 to December 2016.

Exchange rate movements, the interest rate differential, the EMBI index, and the interest rate differential corrected by the EMBI are shown in Figures 1–3.

Figure 1. Monthly Exchange Rate Movements (RSD/EUR)

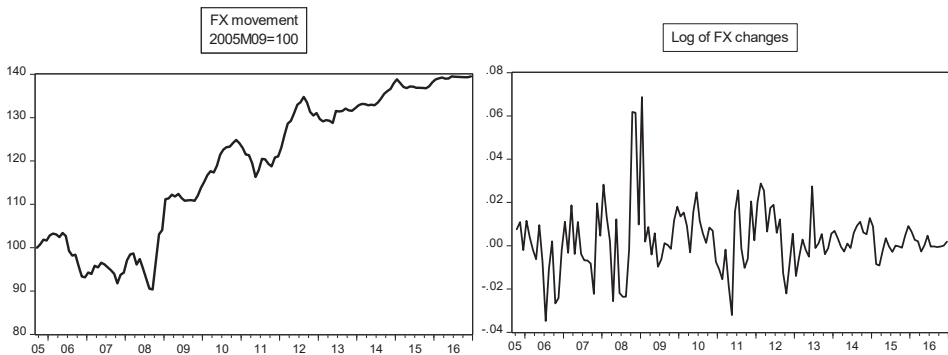
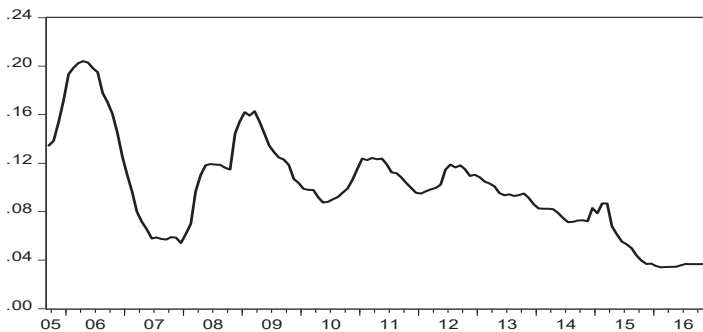


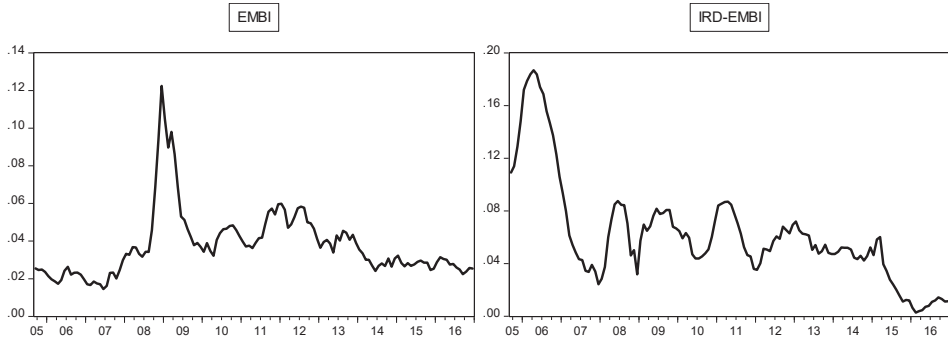
Figure 1 shows a distinct upward trend of the nominal exchange rate, especially after the financial crisis in 2008 when the Serbian dinar experienced sharp depreciation (about 25% in the six months from July 2008 to February 2009).

Figure 2. Interest Rate Differential (BELIBOR 1M – EURIBOR 1M)



The highest interest rate (Fig. 2) is recorded at the beginning of 2006 (over 20%). This is the result of the Serbian monetary authority's tendency to curb inflation through changing the policy rate (the reference rate), directly affecting money market interest, since these rates appear in interbank markets. The inflation rate in 2005 was 16.1%.

Figure 3. EMBI Index and Interest Rate Differential (IRD) corrected by EMBI



The EMBI index (Fig. 3) also reached a peak of above 12% during the financial crisis.

Basic descriptive statistics are presented in Table 1. We can observe that Serbia had very attractive interest rates during the sample period, even corrected for country risk, and could have been a potential target for carry trade investors. The average interest rate differential corrected for country risk was 6.23%, ranging between 0.27% and as much as 18.68%.

Table 1. Descriptive statistics of considered time series

	RSD/EUR	$i_t - i_t^*$	EMBI (CR)	$i_t - i_t^* - CR_t$
Mean	102.36	10.00%	3.77%	6.23%
Minimum	76.42	3.41%	1.45%	0.27%
Maximum	123.42	20.41%	12.24%	18.68%
Std. Deviation	15.52	4.12%	1.78%	3.98%

To test the validity of the UIP model the following econometric approach is employed. First, we implement several unit root tests to find out if the time series are stationary. Specification (7) is a valid regression equation only if both the variables on either side of the equation are stationary or non-stationary, but cointegrated. In the second step we apply bounds testing methodology based on the ARDL model suggested by Pesaran and Shin (1999) and Pesaran et al. (2001). The approach is flexible so that the long-run relationship (level relationship) among time series of a different order of integration can be considered. Also, different short-term dynamics across variables are allowed. If

a long-run relationship exists, then the ARDL model has an adequate equilibrium correction form (ECM) that provides information about long- and short-run adjustments.

The baseline ARDL(k+1,q+1) model can be reformulated as an unrestricted conditional ECM specification:

$$\Delta y_t = \gamma_0 + \sum_{i=1}^k \gamma_i \Delta y_{t-i} + \sum_{j=0}^q \delta_j \Delta x_{t-j} + \theta_0 y_{t-1} + \theta_1 x_{t-1} + e_t \quad (8)$$

where y_t denotes the FX changes and x_t represents the interest rate differential corrected for country risk. Δ is the first difference operator and e_t is an error term.

After verifying the assumptions of this model (stability and approximation of an error term by the Gaussian white noise process), the bounds testing procedure is applied. First, as in the conventional ECM model, the F-test is used to test for the absence of a level relationship between the variables ($H_0: \theta_0 = \theta_1 = 0$). A null hypothesis rejection implies the existence of a level relationship between FX changes and interest rate differentials corrected for country risk. As the distribution of the test statistic is non-standard and critical values are not available for a mixture of I(0) and I(1) time series, Pesaran et al. (2001) provide lower and upper asymptotic critical value bounds for the distribution of the F-statistic for a different number of variables. The constant term is restricted to be a part of the level relationship. Corresponding critical values at the 5% significance level are given in Table 2.

Table 2. Asymptotic critical value bounds for F-statistics
(restricted intercept and no trend, Pesaran et al. 2001)

Bound	F-test
I(0)	3.62
I(1)	4.16

If the computed F-statistic is below the lower bound critical value, it implies that all variables are I(0) and hence there is no level relationship. If the F-statistic is above the upper bound critical value a level relationship exists. If the F-statistic is in between critical values the test is inconclusive. Assuming that the bounds

test implies long-run existence, it can be extracted from equation (8) given that $\beta = -(\theta_1/\theta_0)$ holds for the slope parameter.

The third step estimates the standard VAR model. It consists of our two key time series, but one purely exogenous variable is also included (monetary intervention). This model estimates the impulse response function so that the impact of unanticipated random shocks on FX changes and interest rate differentials can be traced over time.

Finally, the MS approach is implemented to allow for changes in the parameters of the UIP model, assuming that intercept, slope, and error-term variability differ between two regimes. The specification takes the following form (similar versions are used, for example, in Engel and Hamilton 1990; Bekaert and Hodrick 1993; and Ichiue and Koyama 2011):

$$y_t = (\alpha_0 + \alpha_1 P_t) + (\beta_0 + \beta_1 P_t) \cdot x_t + (h_0 + h_1 P_t) \varepsilon_t \tag{9}$$

P_t is the unobserved random variable that follows a first-order Markov chain defined by transition probabilities between two states. The full matrix of transition probabilities for two states reads as follows:

State at $t+1$	Condition at t	
	$P_t=0$	$P_t=1$
$P_{t+1}=0$	$q=p_{0/0}$	$f=p_{0/1}$
$P_{t+1}=1$	$p_{1/0}$	$p_{1/1}$

Note: Probabilities in each column sum to 1.

Shifts of the economy from state 0 to state 1 are governed by the introduced random variable P_t . Under this specification we have two different regimes: regime 0 (i.e., $P_t = 0$) and regime 1 (i.e., $P_t = 1$). Index 0 is associated with the parameters of regime 0. The parameters α_1, β_1, h_1 capture the changes in the intercept, slope parameter, and random term variability respectively during regime 1 relative to regime 0. Such a specification enables a more detailed analysis of UIP relationship (7) because it may detect a different reaction of FX changes that could also depend on the level and variability of FX changes.

5. EMPIRICAL RESULTS

To find out if a unit root is present in the time series we have applied the following unit root tests: Augmented Dickey-Fuller (ADF), Kwiatkowski-Phillips-Schmidt-Shin (KPSS), Phillips-Perron (PP), and Elliott-Rothenberg-Stock (ERS). Results are shown in Table 3. We are able to clearly reject the unit root null hypothesis for $s_{t+1}-s_t$, but not for $i_t - i_t^* - CR_t$. Namely, while the ADF and the ERS tests indicate that the interest differential is a stationary series, the application of the KPSS and the PP tests implies a unit root presence.

Table 3. Results of unit root testing

Variable	Level			
	ADF	KPSS	PP	ERS
$s_{t+1}-s_t$	-7.88	0.13	-7.88	-7.84
$i_t - i_t^* - CR_t$	-4.27	0.20	-2.45	-4.30

Notes: The number of lags included to take care of the autocorrelation is 0 in FX changes and 3 in interest rate differential corrected for country risk. It is determined by the SC criterion. Constant and trend are used as deterministic components. The 5% critical values are: -3.44, 0.146 and -3.00 for ADF (PP), KPSS, and ERS tests respectively.

Next, we proceed with the selection of the appropriate ARDL model. Using the criterion of the lowest value of the Akaike information criterion (AIC) starting with a maximum of 12 lags, the optimal lags in the ARDL model for FX changes and the interest differential are chosen to be 5 and 1 respectively (ARDL(5,1)). From here, we formulate the conditional ECM model, but without imposing any restriction on their coefficients.

The results of the bounds tests clearly indicate the presence of a level relationship (Table 4), with estimates given in Table 5.

Table 4. Bounds Testing Results

Computed value of F-statistic	The 5% critical value	
	I(0)	I(1)
10.42	3.62	4.16

Table 5. Coefficient estimates of level equation

Coefficient	Estimate	Standard error
α	0.007**	0.0031
β	- 1.354**	0.5566

Note: The form of level equation is presented in equation (7). The symbol ** denotes significance at the 0.05 level.

Estimates of both α and β are significant at the 5% significance level. The constant α is close to zero (0.007) and the β coefficient is negative (-1.354). Finally, we end up with the usual ECM in Table 6:

Table 6. Equilibrium correction form of the ARDL(5,1) UIP equation

Regressor	Coefficient	Standard error	p-value
z_{t-1}	- 0.607	0.1077	NA
Δy_{t-1}	- 0.206	0.1093	0.062
Δy_{t-2}	- 0.047	0.0981	0.634
Δy_{t-3}	0.020	0.0834	0.809
Δy_{t-4}	0.171	0.0676	0.013
Δx_t	- 6.391	1.7298	0.000

$$\bar{R}^2 = 0.599, \hat{\sigma} = 0.0110, AIC = -6.1143$$

$$\text{Box-Ljung } Q(8)=6.48(0.59), Q(12)=10.15(0.60),$$

$$\text{Jarque-Bera } JB = 0.25(0.88), \text{ White } WH=10.12(0.34)$$

Notes: This regression is based on the conditional ECM model given in equation (8) using the ARDL(5,1) specification where the dependent variable y_t is defined as $s_{t+1} - s_t$ and x_t as $i_t - i_t^* - CR_t$. z_{t-1} is an equilibrium correction term lagged one period that is based on estimates in Table 5. Additionally, two dummy variables are included: impulse dummy to take only non-zero value 1 for October 2008, and transitory dummy with non-zero values -1 and 1 for December 2008 and January 2009 respectively. The model performs well statistically, as confirmed by several misspecification tests.

The same approach has been employed for the system of variables that do not include information about country risk. However, the results indicate an insignificant level influence of interest rate differentials on FX changes, while the estimated unconditional ECM model has lower explanatory power than the corresponding model in Table 6.

The obtained results correspond to other findings in the literature on the interest parity puzzle (forward premium puzzle) over a one-month horizon. A positive interest rate differential corrected for country risk affects dinar appreciation against the euro, not depreciation as UIP predicts. In fact, a high interest rate differential corrected for country risk leaves space for excessive returns in dinars compared to the euro. In spite of individual investors not being able to reach these returns because capital controls were imposed, foreign financial institutions were permitted to perform carry trades. This affected the dinar appreciation.

Our result is in line with the findings in the literature (Burnside 2014; Ter Ellen et al. 2013; Felcser and Vonnak 2014; Brunnermeier et al. 2008) that in the short-run, positive interest differentials lead to a carry trade strategy. Although these findings are mostly based on data from developed countries, our evidence points to the huge incentive for short-run carry trades in Serbia, regardless of the resounding risks. Burnside (2014) emphasizes that a time-varying risk premium only explains the returns in industrialized countries, whereas in emerging markets the returns are mainly the result of the high interest rate differentials.

In the next stage of our analysis the dynamic relationship is revealed using the VAR model of FX changes and interest rate differentials that includes exchange rate intervention as an exogenous variable (data is from the National Bank of Serbia website). Due to a lack of reliable data on exchange rate intervention, the analysis is based on a sample that starts in January 2007.

A VAR model of order 4 is chosen following the general-to-specific approach. The model performs well statistically, as confirmed in Table 7. The model is stable because all roots of the companion matrix are less than one in the modulus (Table 8).

The results of the Granger causality test show there is only one-way significant causality, running from interest differentials corrected for country risk toward

FX changes, which is in line with the UIP equation. Namely, the null hypothesis that the interest rate differential corrected for country risk does not Granger-cause FX differential is strongly rejected (24.89 with p-value = 0.0). Granger causality in the opposite direction is not detected (6.90 with p-value = 0.23).

Table 7. Multivariate test statistics

Test for	Value	D. of freedom	p-value
Autocorrelation of order 6	13.88	8	0.08
Autocorrelation of order 12	40.52	32	0.15
Normality	6.62	4	0.16
Heteroskedasticity	99.02	90	0.24

Notes: Autocorrelation is tested by the multivariate version of the Box-Ljung test. Normality is assessed by the multivariate version of the Doornik-Hansen test. Heteroskedasticity is tested by the multivariate type of the White test. The estimated VAR model contains three dummy variables that are included to take care of several outliers. Two of them are already defined in the Note to Table 6. The third one is an impulse dummy variable designed to take only non-zero value 1 for 2008M3.

Table 8. Roots of the companion matrix, in modulus

Root 1	Root 2	Root 3	Root 4	Root 5	Root 6	Root 7	Root 8
0.91	0.73	0.73	0.67	0.67	0.62	0.56	0.56

The findings of the Granger-causality test indicate how ordering for the Cholesky decomposition of the residual covariance matrix should be chosen. Therefore, the impulse response functions are computed based on the ordering: FX changes – interest rate differential. The estimates from ordinary impulse response functions are depicted in Figure 4, whereas accumulated estimates are plotted in Figure 5. In this way we trace the reaction of the given variable from the system to one standard deviation shock in the other variable. These estimates are dynamic such that the variable reaction can be followed through different months.

Unexpected shocks to interest rate differentials corrected for country risk (in units of one standard deviation) cause dinar appreciation with one lag, which continues in the following four months. The estimates appear to be precise,

given the narrow confidence interval bands. On the other hand, unanticipated shocks in FX changes seem to be negligible for interest rate differential responses. As confidence bounds are relatively wide, these estimates are insufficiently reliable.

The Serbian economy experienced relatively high inflation rates and inflation volatility over the first half of the period considered (10% on average), which forced monetary authorities to increase policy rates to moderate demand. The policy rates referred to reference rates, which directly affect money market interest rates. Hence, high inflation rates were one of the main sources of the unexpected shocks in the interest differential. Moreover, the very attractive policy rates influenced financial institutions (mainly commercial banks) to keep money in the central bank, enabling the central bank to withdraw liquidity from the markets. Consequently, the reduced amount of Serbian dinars in the market system induced dinar appreciation.

Figure 4. Estimation of Ordinary Impulse Response Function based on Cholesky One-standard Deviation Shock

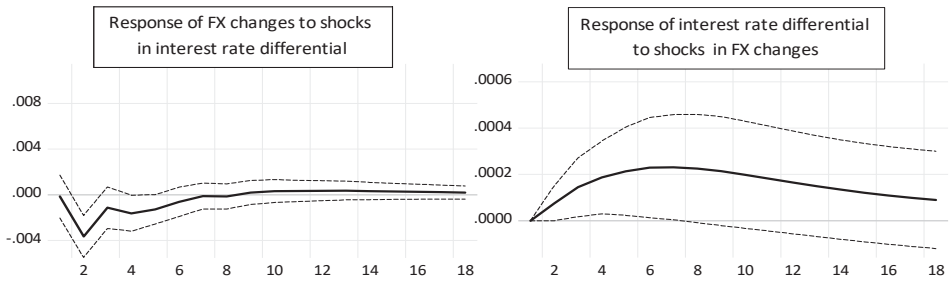
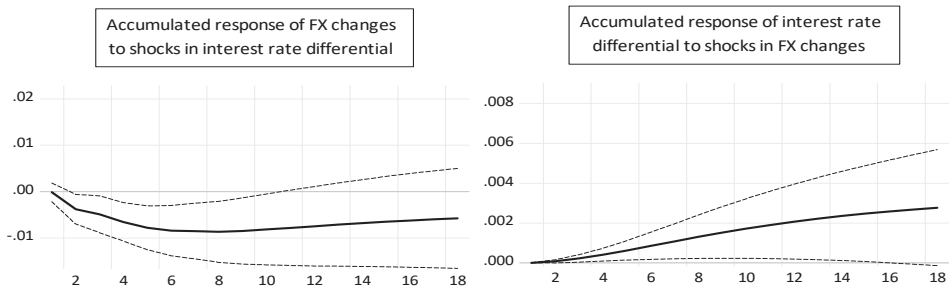


Figure 5. Estimation of Accumulated Impulse Response Function from Figure 4



Note to Figures 4–5: Dotted lines represent the 95% confidence interval bands calculated by Monte-Carlo simulations with 1000 replications.

In the final stage of our econometric research an MS model is estimated (Table 9). Two different regimes are detected as being significant. Regime 0 is characterized by the lower intercept and variability of FX changes, whereas the impact of the interest rate differential is estimated as -2.67 . Regime 1 is found to be associated with higher intercept and variability of FX changes. The slope parameter that measures the reaction of FX changes to interest rate differential in regime 1 is estimated as -0.86 . Therefore, we may argue that regime switches in the relationship between FX changes and interest rate differentials significantly determine FX changes.

A visual inspection of the regimes in Figure 6 shows that regime 0 is closely related to sub-periods of currency appreciation and/or relatively stable FX level. Regime 1 is detected for sub-periods of currency depreciation.

The probability q of remaining in the currency appreciation regime while being in that regime is 0.74. The probability f of switching from the currency depreciation regime to the currency appreciation regime is small and equal to 0.15, indicating that the probability of staying in the currency depreciation regime is relatively high, 0.85.

The economy remains in the currency depreciation regime 64% of the time, while the remaining 36% is associated with the currency appreciation/stability regime. The average duration of the currency depreciation regime is 8.6 months, whereas the average duration of the opposite regime is shorter, 4.8 months.

The parameters of the model appear to be mutually dependent to some extent. Namely, the slope parameter is bigger during currency depreciation (-0.86) when relatively higher variability in FX changes is observed. The slope parameter drops significantly, by about three times, with the appreciation trend when lower variability in FX changes is identified.

Therefore, over sub-periods of the dinar's lower variability, stronger dinar appreciation occurred as a result of higher interest rate differentials. However, higher-variability sub-periods are characterized by much lower appreciation. There is a clear incentive for carry-trade activity during periods of less uncertainty (relatively stable dinar) with high interest rate differentials. However, during periods of higher variability and persistent depreciation these activities are substantially reduced.

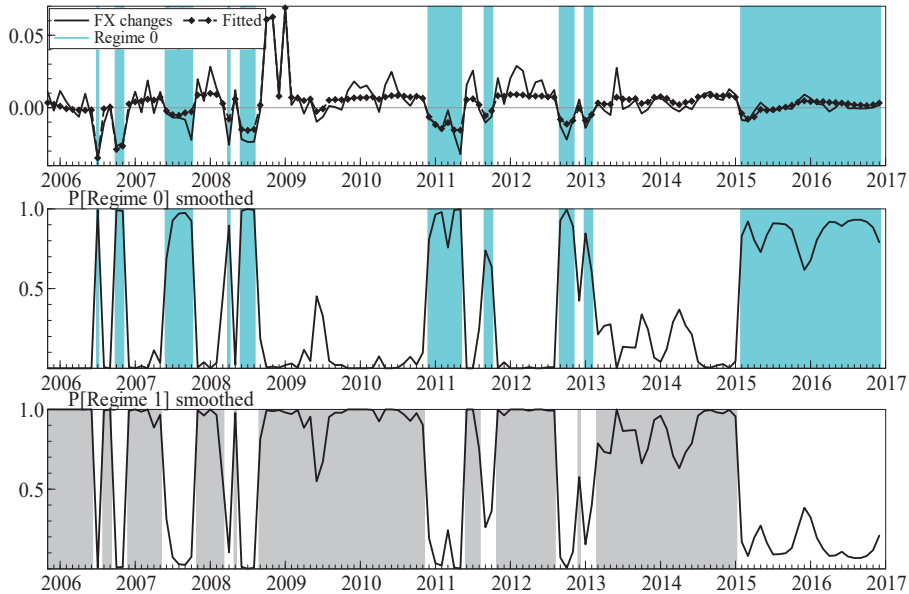
This result strongly supports a carry trade presence and overlaps with some of the previous literature. For example, Ichiue and Koyama (2011), also using the regime-switching approach, find that low volatility influences low-interest-rate currencies to depreciate, which is the cause of the puzzling relationship. This evidence is derived for developed countries (Japan, Great Britain, Switzerland, and Germany) and we would expect even greater appreciation for emerging markets, due to higher interest rates. This is what we have identified for Serbia.

Table 9. Estimated MS version of UIP model

Parameter	α_0	$(\alpha_0 + \alpha_1)$	β_0	$(\beta_0 + \beta_1)$	h_0	$(h_0 + h_1)$	q	f
Estimate	0.004	0.012	-2.672	-0.862	0.007	0.009	0.74	0.15
t-ratio	2.38	3.03	-10.0	-2.20	4.56	11.6	5.51	2.52
p-value	0.02	0.00	0.00	0.03	0.00	0.00	0.00	0.01

Linearity test LR = 22.73(0.00) Box-Ljung Q(8)=12.89(0.12), Q(12)=18.21(0.11), Jarque-Bera JB =0.86(0.65), ARCH(12) = 1.38(0.19)

Notes: Estimation is based on the BFGS algorithm. Robust standard errors are used to calculate t-ratios. The model additionally contains two impulse dummy variables defined similarly to dummies previously introduced. The first one has only two non-zero values of 1 for the following months: 2008 M10 and M11. The second one takes 1 for 2009M1 and 0 otherwise. The model performs statistically well, as shown by several misspecification tests computed for one-step prediction error. The value of the linearity test strongly confirms the presence of non-linearity captured by MS specification.

Figure 6. Estimated regimes from MS version of UIP model

6. PRELIMINARY ECONOMETRIC RESULTS FOR FOUR EUROPEAN EMERGING ECONOMIES

The bounds testing procedure is applied for four CEE countries with mainly floating exchange rates, the Czech Republic, Hungary, Poland, and Romania. Prior to this we take into account descriptive statistics of the interest rate differential corrected for country risk; the results are shown in Table 10. All countries have positive interest differentials over the whole period considered, except that in the Czech Republic, untypically for an emerging economy, the average interest differential is negative or close to 0. The result for the other three countries is consistent with the general trend of positive interest rate differentials in emerging economies. Similar to Serbia, Romania has the highest interest differentials corrected for country risk, peaking in 2003 and 2004 (16.42%), which correspond to the high inflation rates during that time (15.3% and 11.9% respectively). Results of the econometric investigation are summarized in Table 11.

Table 10. Descriptive statistics of interest rate differentials corrected for country risk for CEE countries

	Czech Republic	Hungary	Poland	Romania
Mean	- 0.55%	2.68%	1.40%	3.53%
Minimum	- 2.34%	-0.98%	-0.77%	-0.78%
Maximum	0.45%	10.32%	4.14%	16.42%
Std. Dev.	0.55%	2.79%	1.09%	4.37%

Table 11. Unit root test, Bounds test and Level equation for CEE countries

Country	Period	ADF for level		ARDL	F stat.	Level equation	
		$s_{t+1}-s_t$	$i_t - i_t^* - CR_t$			α	β
Czech	2004M1 – 2016M12	-9.327	-3.666	(2,4)	31.83	-0.00	2.94
Hungary	2003M1 – 2016M12	-9.772	-3.024	(9,0)	20.64	0.00	-0.53
Poland	2003M1 – 2016M12	-8.755	-2.090	(2,0)	38.49	-0.00	-0.21
Romania	2004M1 – 2016M12	-8.579	-2.624	(1,0)	35.55	0.00	-0.77**

Notes: The Augmented Dickey Fuller test is applied to examine stationarity. The number of lags included to take care of the autocorrelation is 0 in FX changes for all countries and in interest differential corrected for country risk is 4 for the Czech Republic, 1 for Poland, and 0 for Hungary and Romania. It is determined by the SC criterion. Constant and trend are used as deterministic components, except that for Poland only constant is included. The 5% critical values are -3.44 for the model with constant and trend and -2.88 for the model with constant only. The appropriate ARDL model is selected based on AIC. Several dummy variables are included in the modelling, defined in the following way. For the Czech Republic three dummies are used (one transitory to take -1 and 1 for 2008M7 and 2008M8 respectively and 0 otherwise, and two impulse dummies to take only non-zero value 1 for 2009M1 and 2013M11). For Poland three impulse dummies are employed having the non-zero value of 1 only for 2008M10, 2008M12, and 2009M2. The model for Hungary has two impulse dummies: one takes only the non-zero value 1 for 2008M10 and the other one two non-zero values of 1 for 2009 M1 and M2. Romania has one impulse dummy designed to have only a non-zero value of 1 for 2009M1. The symbol ** denotes significance at the 0.05 level. The Emerging Market Bond Index (EMBI) is used as a proxy for country risk for Poland and Hungary and credit default swap (CDS) for the Czech Republic and Romania.

It is evident that FX changes are a stationary time series, whereas the interest rate differential corrected for country risk has one unit root in all countries except the Czech Republic, where it is stationary. Hence, the necessity of applying the ARDL procedure arises. As all calculated F-statistics are above the upper bound at the 5% significance level (Table 2), the level relationship is confirmed for all countries. The last columns of Table 11 give estimated coefficients. In three economies the slope estimate β is negative, and it is only significant in the case of Romania. For the Czech Republic the slope is estimated to be positive, but appears to be insignificant. The constant term is insignificant in all countries.

Recall that results for Serbia are quite different from those for most of the other considered CEE countries. The Serbian results are in line with the findings reported for Romania, although the slope coefficient is estimated to be twice as low as in Romania (-1.35 in Serbia and -0.77 in Romania). These are the only two countries that have significantly negative coefficients of the interest rate differential corrected for country risk. This is not surprising, given the tremendously high interest rate differentials in both countries even when corrected for country risk. Romania is also characterized by weak macroeconomic performance and a volatile economy that results in higher interest rates, which might be a potential target for carry-trades in the short run. On the other hand, Poland and the Czech Republic are the most developed countries in the CEE region, with stable economic conditions and relatively good performance compared with the rest of the region. This explains why the interest differential is lower in these countries.

We confirm the existence of the forward premium puzzle in most of the considered countries, with significant slopes in only Serbia and Romania. These results correspond to other evidence in the literature showing that the interest rate puzzle appears when testing the UIP condition in a short-term horizon (one week to a quarter).

7. CONCLUSION

This paper examines an ex post uncovered interest rate parity condition in Serbia over a one-month horizon. The money market interest rates BELIBOR and EURIBOR and EUR/RSD exchange rates are used for the period September 2005 to December 2016. Country risk is incorporated a priori in the model by taking into account the Emerging Market Bond Index (EMBI). First,

econometric results are derived from long-run parameter estimation based on the autoregressive distributed lag model. In Serbia a positive interest differential corrected for country risk is found to cause the dinar to appreciate, instead of depreciating as UIP predicts. Hence, the interest parity puzzle is confirmed, as previously detected by the OLS approach (Božović and Talijan 2015). The results are in line with the growing body of literature, mainly derived for developed markets.

Unexpected shocks to interest rate differentials corrected for country risk lead to dinar appreciation with one lag, which continues in the following four months. On the other hand, the effect of unanticipated shocks in FX changes on interest rate differential responses seems to be negligible. This result is derived from the impulse response function computed from the bivariate vector autoregressive model.

The two-regime Markov switching parameter model is also estimated. It suggests that regime switches in the relationship between FX changes and interest rate differentials are significant in determining FX changes. Stronger dinar appreciation occurred as a result of higher interest rates during lower-variability sub-periods of the dinar. Higher-variability sub-periods are characterized by much lower appreciation.

Our results underline that there was a huge incentive for short-term carry trades in Serbia, regardless of the substantial risk. This finding might be useful when evaluating existing capital controls and tailoring new ones, so that policymakers pay particular attention to short-term lengths.

Finally, preliminary econometric analysis conducted for other four European emerging economies only documents similar results for Romania.

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