ARTICLES

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EXCHANGE RATE VOLATILITY AND UNCOVERED INTEREST RATE PARITY IN THE EUROPEAN EMERGING ECONOMIES

Dejan Živkov, Jovan Njegić, Mirela Momčilović, Ivan Milenković*

Abstract

This paper investigates whether UIRP principle holds and what is predominant driving force, which influences exchange rate movement – economic fundamentals or short-term speculative behaviour. Analysis covers seven East European transition countries and empirical data comprise weekly time series ranging from first week in January 2003 to last week in December 2013. The research method is Component-GARCH in Mean Model, which decomposes temporary and permanent element of volatility. The mean and variance equations have been adjusted for the structural breaks' presence in order to improve estimated parameters. The results suggested that UIRP principle does not hold in any country. After structural breaks inclusion, we have found that the permanent effect is significant in determination of exchange rate dynamics in five countries, but it does not apply for the transition effect. However, further outliers' purification revealed that only in Serbia short-term transition component plays an important role.

Keywords: East European countries, CGARCH-M, uncovered interest rate parity, structural breaks, carry-trade strategy

JEL Classification: C13, C58, F31, F32

1. Introduction

Most countries of the former planned economies in the Central and Eastern Europe have entered and finished the process of economic reform in the past 25 years and thus deepen the integration with the global financial market. As a result, linkages among national financial markets gradually strengthened, and an integrated regional capital market was starting to emerge. A theoretical approach suggests that investors should be able to allocate their capital freely in a one-world market, thereby reducing arbitrage opportunities across countries. Common knowledge is that most of these countries have had higher rates

* Dejan Živkov, Business School of Economics, University of Novi Sad, Novi Sad, Serbia (dejanzivkov@gmail.com);
Jovan Njegić, Business School of Economics, University of Novi Sad, Novi Sad, Serbia (jovan.nj@gmail.com);
Mirela Momčilović, Business School of Economics, University of Novi Sad, Novi Sad, Serbia (bizniscentar@gmail.com);
Ivan Milenković, University of Pristina, Faculty of Economics, Kosovska Mitrovica and University of Novi Sad, Faculty of Economics, Subotica, Serbia (imilenkovic@ef.uns.ac.rs).

of inflation compare to the European Monetary Union (EMU) – their main trading partner, either after entering the transition process as well as in later phases of preparation to join the European Union (EU). Central banks, which did not opt for fix exchange rate regime, carried out more restrictive monetary policy *via* relatively high referent interest rates. As a consequence, positive yield spread have been created between interest rates of these countries and referent interest rate of the European Central Bank (ECB).

The existence of interest rate differential (IRD) emerges like an opportunity for international investors to perform a risk-less arbitrage, *i.e.* a carry-trade portfolio strategy. It means lending in lower yield currency and investing in currency with the higher proceeds, concurrently taking an open position to nominal exchange rate risk, *i.e.* expectation of exchange rate appreciation. Particularly, Durčáková (2011) stated that increased market risk and higher values of the IRD cause the increase of the turnover in the Czech foreign exchange market. However, this approach does not stand in line with the efficient market theory. According to the Uncovered Interest Rate Parity (UIRP), high yielding currencies should depreciate in just exact amount to equalize expected returns on two different currency deposits. Numerous empirical researches performed both on developed and developing countries (Froot and Thaler, 1990; Chinn and Meredith, 2004; Sarno, 2005; Fisher, 2006; Burnside *et al.*, 2009) indicate that exchange rates fail to move one-for-one with IRD, leading to the conclusion that UIRP could not explain future exchange rate changes. Also, the study of Beker (2006) indicates that flexible exchange rates do not always reflect changes in fundamental macroeconomic variables. The presumption of exchange rate changes based upon IRD leads to conclusion that exchange rate changes should not be biased in the case of rational expectations.

Some studies suggest that UIRP does not work in the case of the appearance of non-linearities due to structural breaks and outliers, monetary policy regime changes, rational expectations failure or/and the factors which lie behind the Balassa-Samuelson effect. Further, the relative purchasing power parity hypothesis also has an important role in determination of future exchange rate, as argued by Žďárek (2012), which may contribute to the deviation from the UIRP theory. Yet, the factor that has been elaborated and often cited as a possible cause for UIRP failure is the risk premium existence, defined as *ex-ante* expected profit of the carry trade strategy (see *e.g.* Froot and Thaler, 1990; McCallum, 1994). If this assumption is taken *a priory*, it would mean that carry-trade portfolio speculations could consequently influence exchange rate movements in the short run. Some recent papers as Plantin and Shin (2008) argued that investors can engage in carry trade freely and flexibly if the speed of exchange rate reversal to its fundamental level is slow. Also, Bauer and Herz (2005) asserted that transition countries have strong incentives to stabilize their exchange rates with respect to the euro. In that sense, our research efforts seek the answer whether UIRP assumptions hold, and what is a predominant driving force for an exchange rate change in selected group of countries. A dilemma stands between macroeconomic fundamentals which influence long-range movements and market sentiments (*e.g.* carry-trade activities due to the difference in interest rates) which have a role in short-run determination.

Analysis has been focused on several ESEE transition countries, namely Croatia, the Czech Republic, Hungary, Poland, Romania, Russia and Serbia, within a time span of past eleven years. The Czech Republic, Poland, Romania and Serbia have led a *de facto* flexible exchange rate, Hungary focused its exchange rate policy on fix regime with wide bands, Croatia used fix regime with narrow bands and Russia used tight management till 2008 and looser exchange rate management afterwards. Russia and Croatia were included in the analysis since these countries pursued tightest exchange rate policies of all selected countries. Because of this fact, the intention was to test the hypothesis that neither permanent nor transitory factors influence Ruble or Kuna. The Croatian National Bank (CNB) conducted narrow band fluctuation of ±6% since the introduction of Kuna on 30 May 1994, followed by further reducing of the fluctuation span at later years. Since 2005 onwards, the CNB has been directing exchange rate policy within the narrow bandwidth around only ±2%. As for the Russia, the Central Bank of Russia (CBR) set up an explicit ceiling for real appreciation of the Ruble, given the strength of Russia's balance of payments. Therefore, Russia's exchange rate policy framework was often referred to as *de facto* very tightly managed nominal exchange rate regime. However, Russia's economy and financial markets were deeply affected after the financial crisis turmoil in 2008. The aftermath was that CBR entered into a transition path from an exchange rate-based monetary policy to the price stability.

Research has been conducted within the family of the generalized autoregressive conditional heteroscedasticity (GARCH) type models. Particularly, we used GARCH type specification known as Component-GARCH developed by Engle and Lee (1993) which decomposes volatility into two components – a stochastic long-run trend and short-run deviations from that trend. By separating permanent and transitory risk we managed to assess which factor primarily influences exchange rate volatility in these countries. Additionally, we utilized Bai-Perron test (Bai and Perron, 2003) for structural change detection in the mean of the series and Sanso test (Sanso *et al*., 2004) for break detection of the volatility process. The inclusion of the breaks in the models was done because our data encompasses world economic crisis during 2007–2009, and because of the fact that analysis deals with transition economies, which are less resilient to the structural changes than developed economies. Correcting these imperfections with suitable technique we obtained better results. To the best of our knowledge, this is first time that CGARCH-M methodology is used in this topic in this group of countries, also accounting structural breaks and outliers in the models.

Besides introduction, the remainder of the paper is organized as follows. Section 2 shortly describes theoretical background of the UIRP tenet. Section 3 explains CGARCH-M methodology and Section 4 briefly overview the dataset. Section 5 presents and discusses the results of the empirical analysis, taking into account structural breaks in the mean and the conditional variance, as well as outlier-corrected values. The last section concludes.

2. Theoretical Underpinning and UIRP Regression Equation

Usual way of UIRP biasedness testing in the literature is done by the regression of the exchange rate changes on the IRD (Froot and Thaler, 1990):

$$\Delta S_{t+k} = \log(S_{t+k} - S_t) = \alpha + \beta(i_{t,k} - i^*_{t,k}) + \varepsilon_{t+k} \tag{1}$$

where ΔS_{t+k} is the percentage depreciation or appreciation of the currency, or the change in the log of the spot exchange rate over k periods; $(i_{t+k} - i^*_{t+k})$ is the current k period home interest rate less the k period foreign interest rate (in our case ECB interest rate); and ε_{t+k} is a purely random error term, or the white noise. Relation (1) is viewed under the assumption of rational expectations where future exchange rate, S_{t+k}, is substituted by the actual exchange rate movements in k-th period. If rational expectation theory is valid, the null hypothesis for UIRP is that $\alpha = 0$, $\beta = 1$. However, numerous research has tested the unbiasedness hypothesis and found that coefficient β is not just less than one, but very frequently less than zero. In particular, if $0 < \beta < 1$ it means that 1 per cent increase in interest rate spread yield is associated with a less than 1 per cent expected drop in the home currency. More extremely, if $\beta < 0$, a unit increase in IRD would lead subsequently even to appreciation of the home currency and realization of the carry-trade strategy. Consequently, it means that the increase in the IRD combined with the strengthening path of the exchange rate create a risk premium and opportunities for risky investments.

So far, definite answer about UIRP credibility was not found. An extensive literature, including the initial studies by Bilson (1981), Longworth (1981), and Meese and Rogoff (1983) indicated that UIRP is at best useless principle for determination of future exchange rate change. Yet, Froot and Thaler (1990) presented comprehensive survey of 75 published estimates and reported few cases where the sign of the coefficient of the IRD in exchange rate prediction equations is consistent with the unbiasedness hypothesis under UIRP, and not a single case where it exceeds the theoretical value of unity. Chinn and Meredith (2004) also found strong evidence of the failure of UIRP if using short-horizon data in G-7 countries. On the contrary, some recent research suggested that uncovered interest parity tends to hold for financial instruments of long maturities, notably three years or more (Chinn, 2006; Alexius, 2001; Zhang, 2006; Chinn and Meredith, 2004). Particularly, Mc Callum (1994) claimed that unbiasedness of the UIRP regression could be explained with the policy response hypothesis. Monetary authorities could set up higher interest rate in order to attract more foreign capital and consequently resist to rapid change in exchange rate, which plays an important factor for inflation stability. However, this approach could also lead to currency overvaluation. Generally, ESEE countries usually set up a higher benchmark interest rate compared to the European central bank reference rate as a measure of restrictive monetary policies. This is true for almost all ESEE transitory economies with flexible exchange rate regime, since depreciation has the most immediate transmission channel on inflation and higher interest rate act as a preventive measure.

3. CGARCH Methodology

Component GARCH type model is very suitable for analysis of exchange rate variability, because it could split out the permanent and transitory component of the exchange rate change. The two components of volatility are driven by different factors: the long-run trend in volatility *e.g.* reflecting shocks to economic fundamentals, and transitory volatility, which is driven by market sentiment and short-term position-taking (Pramor and Tamirisa, 2006). Many authors verified its superior performances for modelling exchange rate volatility comparing to other GARCH type models (Black and McMilan, 2004; Pramor and Tamirisa, 2006; Li *et al.,* 2012). We have used CGARCH-M specification following the works of Pramor and Tamirisa (2006), Li *et al*. (2012) and Tai (1999). The GARCH-M [1] model was introduced as a generalized class of ARCH-in-Mean Models and it was designed to capture relationship between risk and return. Accordingly, in the mean equation we

1 All calculations in the paper were done *via* R-software.

added the conditional standard deviation (σ_{t+1}) as a time-varying risk premium. Therefore, the component GARCH – in Mean specification looks like:

$$\log(S_{t+1} - S_t) = \alpha + \beta(i_t - i_t^*) + \gamma\sigma_{t+1} + \varepsilon_{t+1} \tag{2}$$

$$q_{t+1} = \omega_1 + \omega_2(q_t - \omega_1) + \omega_3(\varepsilon_t^2 - \sigma_t^2) \tag{3}$$

$$\sigma_{t+1}^2 = q_{t+1} + \omega_4(\varepsilon_t^2 - q_t) + \omega_5(\sigma_t^2 - q_t) \tag{4}$$

$$\varepsilon_{t+1} \big| I_t \sim \text{i.i.d.}(0, \sigma_{t+1}^2) \tag{5}$$

Equation (2) differs from equation (1) by conditional component of the standard deviation of the error term. It is a mean equation with a conditional standard deviation (σ_{t+1}) as a measure of risk premium that directly influences exchange rate. If constant (α) and a time-varying components (γ) are both insignificantly different from zero, there is no risk premium. Otherwise, if $\alpha \neq 0$ but $\gamma = 0$, there is a constant risk premium. However, when $\gamma \neq 0$ the time-varying risk premium exists (Li, *et al.*, 2012). The conditional density function ($\varepsilon_t + 1 \| I_t$) is modelled as *i.i.d.*, and I_t is the information set available at time t. All models were estimated by *quasi* maximum likelihood (QMLE) technique which is robust to the non-normal distribution.

Equation (3) represents long-term relations, where q_{t+1} is the long-term component of the conditional variance, which reflects shocks of economic fundamentals and converges to the log-term, time invariable volatility level ω_1 with a magnitude of ω_2. The closer the estimated value of the ω_2 to one, the slower q_{t+1} approaches to ω_1 and *vice-versa*. Therefore, coefficient of ω_2 provides a measure of the long-term persistence. The $(\varepsilon_t^2 - \sigma_t^2)$ serves as driving force for the time dependent movement of the permanent component, thus ω_3 shows how shocks affect the permanent component of volatility. Equation (4) is a short-term relation and $(\sigma_t^2 - q_t)$ is a short-term or transitory component of the conditional variance and it indicates the degree of memory in transitory component. The $(\varepsilon_t^2 - q_t)$ measures the initial impact of a shock to the transitory component. As for parameters relations, an important requirement should be fulfilled to make model stable, *i.e.* permanent volatility ω_2 should exceed the sum of the coefficients ($\omega_4 + \omega_5$) in the transitory component, implying that short-run volatility converges faster than the long-run volatility. If that is not the case, than long-term volatility approaches to the mean faster, suggesting that market sentiment and market behaviour have greater impact on exchange rate changes, which makes model unstable.

4. Research Data and Test for Multiple Structural Breaks

Empirical data for selected countries (Croatia, the Czech Republic, Hungary, Poland, Romania, Russia and Serbia) comprise weekly series of exchange rates (Kuna, Koruna, Forint, Zloty, Leu, Ruble and Dinar expressed relative to Euro) and referent interest rates of these countries expressed in excess to ECB referent interest rate. The data was collected from the Global Financial Data. Time span ranges from the first week in January 2003 to the last week in December 2013. The weekly data were used in order to avoid unit root in the series of IRD, which is characteristic feature for the daily data. The Table 1 displays descriptive statistics of exchange rate changes (ΔS_{t+k}) expressed according to rational expectation theory, as well as Zivot-Andrews unit-root tests (Zivot and Andrews, 1992), robust to the existence of the structural breaks, for exchange rate changes and interest rate difference $(i_{t,k} - i^*_{t,k})$. It is apparent that all series exhibit non-normal behaviour with asymmetric distribution and leptocurticity. Excess kurtosis with fat tails in the unconditional distributions means that extreme changes occur more frequently. Thus, Jarque-Bera test rejects normality for all series. Accordingly, some improvements could be expected after GARCH conditional volatility adjustment. The LB(Q) tests found an autocorrelation presence in almost all exchange rate series. The LB(Q^2) statistics suggests the presence of time varying-variance in all series, presenting clear evidence of an ARCH pattern. Further, all series of exchange rate changes and all IRD are stationary at 1% level. Exchange rate changes for selected countries and autocorrelation functions of squared data are presented by Figure 1 and Figure 2, respectively, in Appendix of the paper.

Table 1 | Descriptive Statistics and Some Diagnostic Tests

	CRO	CZH	HUN	POL	ROM	RUS	SRB
Mean	0.001	−0.033	0.049	0.020	0.052	0.050	0.115
Standard deviation	0.374	0.914	1.449	1.471	1.045	1.204	1.029
Skewness	0.283	−0.177	0.279	0.810	0.338	0.742	0.448
Kurtosis	6.022	5.731	5.688	9.292	7.923	8.153	10.905
Jarque-Bera	214.2	171.9	170.9	957.1	559.8	651.9	1434.7
LB(Q)	0.004	0.003	0.000	0.000	0.175	0.085	0.092
LB(Q^2)	0.000	0.000	0.000	0.000	0.000	0.000	0.000
ZA test for ΔS_{t+k}	-14.643^*	-25.602^*	-27.272^*	-10.694^*	-25454^*	-11.434^*	-26.017^*
ZA test for $(i_{t+k} - i^*_{t+k})$	-6.955^*	-3.523^*	-3.141^*	-4.083^*	-4.362^*	-3.002^{**}	-3.651^*

Note: Labels CRO, CZH, HUN, POL, ROM, RUS and SRB are abbreviations for Croatia, the Czech Republic, Hungary, Poland, Romania, Russia and Serbia, respectively. LB-Q and LB-Q^2 tests denote p-values of Ljung-Box Q-statistics for level and squared residuals up to 20 lags, ZA-test stands for Zivot-Andrews unit-root test, *,** label one and five percent significance level for ZA-tests, respectively.

Source: Author's calculation

The existence of structural breaks is a common problem in macroeconomic series, which is especially true for less resistant transition economies. Furthermore, all selected economies suffered tremendous shock when world crisis spilled over to their markets (*i.e.* significant exchange rate depreciation took place). Thus, we expected that our models could suffer from structural breaks presence in the mean as well as in the variance equation. Leaving this problem disregarded might cause significant estimation problems in form of spurious regression and parameter bias. Consequently, it would spell misspecification and erroneous conclusions. In order to estimate multiple break dates in the mean of the exchange rate series without prior knowledge when those breaks happened, we employed statistical technique of Bai and Perron (2003). For the variance process we utilized test proposed by Sanso *et al*. (2004). Using these methods we searched for breaks in the mean and variance at 5% significance level. Table 2 reports dates when breaks occurred for every country.

Table 2 | Break Dates in Mean and Variance Equation

CRO	CZH	HUN	POL	ROM	RUS	SRB
Panel A. Structural break test in mean						
–	31/12/2008 12/01/2011	–	–	07/03/2007 15/10/2008	–	30/08/2006
Panel B. Structural break test in variance						
30/03/2005 16/01/2008 23/09/2009	25/06/2008 13/01/2010	28/07/2004 15/03/2006 24/09/2008 04/08/2010	20/08/2008 21/07/2010	04/02/2009	01/09/2004 19/07/2006 24/09/2008 14/04/2010	12/07/2006 29/04/2009

Source: Author's calculation

According to the break test results, the Czech Republic and Romania have two and Serbia has one structural break in the mean of the series, but all countries have at least one break in the variance. In order to avoid poor specification in CGARCH-M models, we used dummy variables to accommodate the extraordinary changes, similarly as Fang and Miller (2009), Cunado *et al.* (2006), Ewing and Malik (2013) tackled their problems with breaks in GARCH models. For every break in the mean and variance equation a dummy variable was created and inserted into the models. Dummy variable equals unity from the break date forward and zero otherwise. CGARCH-M specification with included dummy variables looks like:

$$\log(S_{t+1} - S_t) = \alpha + \beta(i_t - i_t^*) + \gamma\sigma_{t+1} + \sum_{j=0}^{2} d_j D_j + \varepsilon_{t+1} \tag{6}$$

$$q_{t+1} = \omega_1 + \omega_2(q_t - \omega_1) + \omega_3(\varepsilon_t^2 - \sigma_t^2) \tag{7}$$

$$\sigma_{t+1}^2 = q_{t+1} + \omega_4(\varepsilon_t^2 - q_t) + \omega_5(\sigma_t^2 - q_t) + \sum_{i=1}^{4} \varphi_i D_i \tag{8}$$

$$\varepsilon_{t+1} \middle| I_t \sim \text{i.i.d.}(0, \sigma_{t+1}^2) \tag{9}$$

where j and i represent number of breaks in the mean and variance equation. If there are any improvements it would be read off in the conditional distribution characteristics. According to the distributional assumption of the GARCH specification, implementation of GARCH models should capture unconditional skewness and excess kurtosis to some extent and thus ameliorate normal distribution in the models. The model adequacy is recognized *via* its standardized residuals, which should have mean zero and unit standard deviation. If non-normality stays persistent after GARCH employment, the likely reason could be the presence of structural changes that may distract parameter estimation in the GARCH models. Mikosch and Starica (2004) argued that structural changes may confound persistence estimation in GARCH models. Hillebrand (2005) provided the evidence that discrete shift in the unconditional variance could lead to the bias GARCH estimates. Thus, the introduction of structural breaks *via* dummy variables could significantly improve model adequacy. After structural break exclusion any non-normality left-over could indicate to the outliers. As Franses and Ghijsels (1999) argued, neglecting outliers also leads to biased parameter estimates in conditional mean equations. In the following chapter our intention is to estimate CGARCH-M models with and without dummies and to make mutual comparison.

5. Research Results

5.1 Component GARCH-M models estimates

This chapter provides the results of UIRP sustainability *via* Component GARCH-M methodology, and answers the question what is the main driving force of the exchange rate – underlying fundamentals or transitory shifts mirrored in market sentiment. Considering selected countries and different exchange rate regimes in these countries, *a priori* assumption could be that neither permanent nor transitory effects will be significant in the case of Croatia and Russia due to the tight exchange rate policy, *i.e*. a strictly managed exchange rate within the narrow bands. Table 3 gives the results of CGARCH-M model neglecting the structural breaks along with diagnostic tests which indicate the model adequacy. In order to gain confidence about the assessed parameters, the models accurateness has been examined firstly.

All estimated models do not have problems with serial correlation, which is depicted by Ljung-Box Q-statistics at level residuals. LB(Q^2) confirms the absence of time varying volatility. Additionally, we have employed a portmanteau test for time-based dependence in a series known as the BDS test (see Brock *et al*., 1996) which has its null hypothesis that the data are pure white noise (completely random). It is argued that it has power to detect a variety of departures from randomness – linear or non-linear stochastic processes, deterministic chaos, *etc*. Plotted BDS tests suggest that estimated residuals are independent and identically distributed (*iid*) at very high probability rate for all countries except for Serbia. The standardized residuals are moving around mean zero and unity for standard deviation except for Croatia whose standard deviation is over one. Likewise, the GARCH model implementation reduced kurtosis coefficient significantly in all cases except for Croatia which report only modest improvement. Nevertheless, Croatia, Romania and Serbia still report non-normal distribution depicted in relatively high Jarque-Bera coefficient. Due to imperfections in Croatian, Romanian and Serbian models, in the following we will not interpret its parameters presented in Table 3 due to possible coefficient bias. However, other CGARCH-M models make great improvements in fitting empirical data. They will be addressed below.

According to assessed parameters in Table 3, Czech and Poland models have negative coefficient but insignificant. Negative β coefficient would mean that increase in the IRD will lead to an appreciation in the expected exchange rate. On the other hand, positive β coefficient between zero and one in case of Hungary and Russia imply currency depreciation but less than one-for-one as UIRP theory proposes, whereby only β coefficient in Hungarian model is significant. Our results assert that UIRP theory does not hold in these four countries. Also, our findings are in line with the conclusion of Li *et al.* (2012) who investigated ten develop and emerging markets and found similar results. They contended that none of the emerging market have negative coefficient in front of IRD but quite opposite. Three out of five emerging markets have positive and statistically significant β coefficient. It means that central banks of emerging markets do not have power to influence the exchange rate movement *via* higher interest rate, but probably other factors have more decisive role, *i.e.* inflation rate, export intensity, capital inflows *etc*. On the contrary, four out of five developed countries have negative and significant β coefficient. The rationale is that these countries have much more stable macroeconomic environment, where higher interest rate lure capital and thus contribute to exchange rate appreciation. Analysing risk premium in selected countries, only Serbia has negative and statistically significant γ coefficient as presented in Table 3, which is in favour of UIRP theory but due to probable misspecification we could not take it as reliable indicator. However, it could mean that an increase in risk perception leads to the decrease of the depreciation of the home currency resulting in increased returns from holding this domestic currency. This could be further corroborated after the introduction of structural breaks in the models and their general quality improvement.

Table 3 | CGARCH-M Model Estimation Equation (2–4)

	CRO	CZH	HUN	POL	ROM	RUS	SRB
Panel A. Parameter estimates							
α	0.0061	−0.1397	0.1840	−0.2450	−0.1220	0.0913	0.2489
β	0.0755	−0.4990	0.3746*	−0.3434	0.1536*	0.2071	−0.0482
γ	−0.0549	0.0988	−0.1425	0.1700	0.1915	−0.0783	−0.1399***
ω_1	0.0880	0.7293*	2.0760*	1.7262*	2.8182	6.9888	3.9383**
ω_2	0.9983*	0.9752*	0.9706*	0.9488*	0.9669*	0.9996*	0.9996*
ω_3	−0.0159*	0.0799*	0.1161*	0.1279*	0.2992*	0.0640*	0.0638*
ω_4	0.1693*	0.0453	0.0060***	0.0261	0.2386*	0.1097*	0.1116*
ω_5	0.7035*	−0.8647*	−0.9912*	−0.9148*	0.0780	0.7413*	0.8187*
Panel B. Diagnostic tests							
LB-Q(20)	0.152	0.867	0.293	0.925	0.617	0.505	0.369
LB-Q²(20)	0.722	0.912	0.674	0.985	0.914	0.997	0.654
BDS	0.8547	0.9730	0.7606	0.2777	0.5368	0.3977	0.0025
Av.st.res.	0.0195	0.0237	0.0476	0.0321	−0.0092	0.0275	−0.0411
St. dev.	1.2987	0.99789	0.9978	1.0010	1.0037	1.0162	1.0579
Skew.	−0.0819	−0.2560	0.4144	0.4891	0.2255	0.1699	0.1404
Kurt.	5.8997	3.4567	4.1661	4.1319	6.0802	4.4131	6.1642
JB	191.191	10.693	46.478	50.825	220.07	47.969	228.305

Notes: LB-Q and LB-Q² stand for Ljung-Box Q statistics testing for autocorrelation up to 20 legs. All LB values in the table are probabilities. ARCH (LM) test is calculated up to 20 lags and table contains p-values. BDS test are done for standardized residuals and represents p-values. Av.st.res. stands for average standardized residuals. JB values are coefficients of Jarque-Bera test for normality. *,**,*** denote 1%, 5% and 10% significance level, respectively.

Source: Author's calculation

As for GARCH parameters, it could be seen that ω_2 coefficient of the lagged permanent volatility is large and highly significant at 1% level for the Czech Republic, Hungary, Poland and Russia, which verifies the presence of long-run volatility persistence. Particularly, all coefficients are very close to one indicating that permanent volatility converges to its mean very slowly, *i.e*. permanent conditional volatility exhibits long memory. The transitory component, ω_4, is significant only for Hungary and Russia. It measures the initial impact of a shock to the CGARCH transitory component. The parameter ω_5 indicates the degree of memory in the transitory component and it is significant for all four countries. These results could imply that short-term factors, *i.e*. market sentiment, also affect exchange rate movement. However, since the value of the permanent component (ω_2) is greater than the sum of the transitory components ($\omega_4 + \omega_5$), it indicates that mean reversion is slower in the long run, which makes our models stable. Actually, Grilli and Kaminsky (1991), and Glen (1992) contended that purchasing power parity might support the slow mean reversion of the exchange rate in the long run. Nevertheless, further conclusions could be made only after the introduction of dummy variables in the models. These results are presented in the Table 4 in the next chapter.

5.2 Component GARCH-M models regarding structural breaks and outliers

Following findings of Bai-Perron and Sanso tests we embedded multiple dummy variables in our CGARCH-M models. The improvements have been recognized in the models. Similarly to the previously assessed CGARCH models in Table 3, the models with included dummy variables do not have problems with serial correlation, which is confirmed by high rate of LB(Q) test. The LB(Q^2) test confirms the absence of heteroscedasticity and BDS test suggests that residuals are *iid,* which now applies for Serbia as well. Average standardized residuals are close to zero and standard deviations move around unity for all countries. Likewise, the improvements are obvious at all kurtosis coefficients, which are reflected by lower Jarque-Bera coefficients. However, even after introduction of structural breaks, Romania, Serbia and to some extent Croatia still have relatively large, but significantly lower JB indicators. As previously stated, the likely reason of higher JB in those countries could be the presence of outliers in the series. These disadvantages, as well as structural breaks, could lead to distorted outcome and unreliable inference.

As for the mean equation in the Table 4, the results have not changed much comparing to Table 3. Hungary and Romania still have statistically significant β coefficient ranging between zero and one, which confirms UIRP failure. Important change could be noticed in Serbian risk premium where γ parameter is no longer significant. However, more significant changes are present in GARCH equations. For instance, after the introduction of dummy variables, neither permanent nor transitory influences are statistically significant in case of Croatia, as being suggested by Table 3. The same applies for Russia. This is profoundly different result comparing to basic models in Table 3, and it stands in line with our initial hypothesis. It can be concluded that GARCH results from Table 3 are erroneous in case of Croatia and Russia due to structural breaks, and that some other factor plays more important role than fundamental or transitory shifts. The rationale could lie behind officially pursued exchange rate policy of these countries. As explained in the introduction, these countries led tight exchange rate policies. This empirical fact confirms the findings from our CGARCH-M models, since economic fundaments and/or market sentiment did not have enough space to influence exchange rate fluctuation and thus they proved insignificant.

Table 4 | CGARCH-M Model Estimation with Structural Breaks Implementation Equation (6–8)

	CRO	CZH	HUN	POL	ROM	RUS	SRB
Panel A. Parameter estimates							
α	0.0105	−0.2887**	0.0950	−0.2388	−0.1394	0.0506	0.2293*
β	0.0732	−0.4280	0.3391*	−0.2949	0.1643**	0.3031	−0.0137
γ	−0.0526	0.2798	−0.0678	0.1664	0.2306	−0.0571	−0.0551
d_1	–	−0.1045	–	–	−0.1853***	–	−0.1311
d_2	–	0.2266**	–	–	0.1877***	–	–
ω_1	0.3507*	0.4796*	0.9552*	1.0295*	5.0302**	2.5988*	0.1750*
ω_2	0.6611	0.8162*	0.7411*	0.9068*	0.9679*	0.3008	0.9325*
ω_3	0.8760	0.1321	0.0808	0.2201	0.3530*	0.7935	0.1046
ω_4	−0.7078	−0.0644	0.0618	−0.2103	0.2536*	−0.7132	0.0867
ω_5	1.3252	0.5718	−0.3465	0.9683*	0.1114	0.7312	0.7865*
φ_1	−0.1024	0.9241	−0.5915	1.2060**	−1.7275***	−1.6009**	0.4666**
φ_2	0.0230	−0.8322	1.5876**	−1.0262**	–	−0.7276*	−0.3666**
φ_3	−0.0269	–	3.7813**	–	–	1.9912*	–
φ_4	–	–	−3.6572**	–	–	−0.5348	–
Panel B. Diagnostic tests							
LB-Q(20)	0.125	0.815	0.205	0.846	0.652	0.591	0.567
LB-Q^2(20)	0.981	0.832	0.496	0.981	0.971	0.985	0.494
BDS	0.6815	0.9078	0.8169	0.6590	0.2528	0.8154	0.0544
Av.st.res.	0.0202	0.0131	0.0340	0.0204	−0.0004	0.0481	−0.0039
Std. Dev.	1.0028	0.9999	1.0037	0.9906	0.9772	1.0029	0.9968
Skew.	0.2906	−0.2601	0.2747	0.3038	0.2231	0.2099	0.0777
Kurt.	4.1346	3.2988	3.5448	3.4889	5.8521	3.5653	5.4252
JB	36.839	8.174	13.592	13.812	189.243	11.257	133.623

Note: See Table 3

Source: Author's calculation

Unlike relatively fixed exchange rate regime as in the case of Croatia and Russia, other selected countries have conducted *de facto* flexible regimes (the Czech Republic, Poland, Romania and Serbia), Hungary applied target zone with wide band ±15% till February 2008 and free float afterwards. CGARCH-M models indicate that exchange rates are affected by permanent volatility in all these countries. This is in line with expectations, because these exchange rates were not influenced by discretionary monetary policy, but predominantly by macroeconomic factors. For the Czech Republic and Hungary neither of the transitory components is significant. Only ω_5 for Poland and Serbia and ω_4 for Romania are significant. In general, overall findings prove our doubts that structural break presence would distort results and yield unreliable outcomes. In order to further improve models for Croatia, Romania and Serbia we utilized the method for detection and correction of outliers in the GARCH models developed by Chen and Liu[2] (1993). Results for outliers corrected Component GARCH-M models are presented in Table 5.

The general properties of outliers-corrected models have been improved greatly. It could be seen that all kurtosis coefficients are lower in panel B of Table 5 comparing to its counterparts in Table 4. Similarly, all JB values are much lower, which leads to the conclusion that all three models have *i.i.d*. distribution. Since all misspecifications have been amended, we can believe much more firmly in parameter correctness. As for Croatia there is not much difference, however in case of Romania and Serbia changes are noticeable. In the Romanian model, coefficient in front of IRD is no longer significant in the mean equation. GARCH parameters retained their significance but with slightly lower values. As for Serbia, the mean equation remains unchanged, but all GARCH parameters are statistically significant after outlier's elimination. It seems that outlier's presence warped the results in the foregoing model in Table 4 and the new outcome suggests both permanent and transitory influence on Serbia's exchange rate. Coefficient that depicts long-run time-varying trend, ω_2, is higher than sum ω_4 + ω_5, implying slow mean reversion in the long-run, and making Serbia's model stable. Therefore, the long–run volatility is more durable than the short-run, indicating that exchange rate movement is mainly driven by shocks from economic fundamentals, *e.g*. PPP. However, transitory component is also significant and these newly results can be interpreted in the sense of proclaimed monetary policy of the National Bank of Serbia (NBS). NBS introduced the new monetary instrument 2W-REPO rate on 1st September 2006 which serves as measure for excess money sterilization and exchange rate stabilization. Many domestic as well as foreign investors invest in newly created REPO securities, which contribute to short-term position taking and transitory effect on exchange rate. From this point of view, we could clearly say that adopting conclusions based on the underlying CGARCH-M models in Table 3 would be wrong, *i.e.* without recognizing structural breaks and outlier's presence. By correcting these deficiencies, we showed that models with structural breaks introduction and outlier's purification yield better and more trustworthy results.

2 This procedure consists of three stages. Firstly, all potential outliers are identified based on preliminary model's point estimates. Secondly, all point estimates of the model along with outlier effects are acquired using the accumulated outlier information of the first step. In final stage, the outliers' effect is estimated again taking into account less-contaminated estimates of model parameters obtained in the second step.

Table 5 | Outliers-Corrected CGARCH-M Model Estimation

	CRO	ROM	SRB
Panel A. Parameter estimates			
α	−0.0162	−0.1856*	0.2333*
β	0.0700	0.3212	−0.0548
γ	0.0555	0.3151	−0.1422
ω_1	0.2381*	0.6558*	0.1839*
ω_2	0.7000	0.9526*	0.9321*
ω_3	0.5528	0.0834**	0.1321*
ω_4	−0.4461	0.1780**	0.1587**
ω_5	1.0501	−0.0607	−0.4077***
Panel B. Diagnostic tests			
LB-Q(20)	0.441	0.297	0.011
LB-Q²(20)	0.316	0.822	0.727
BDS	0.5874	0.8711	0.8460
Av.st.res.	0.0096	−0.0103	−0.0042
St. dev.	1.0008	0.9989	1.0001
Skew.	0.2282	0.0287	−0.0221
Kurt.	3.5675	3.5645	3.7674
JB	11.826	6.950	12.605

Note: See Table 3. Using the outlier-corrected series, Bai-Perron and Sanso tests were conducted. Structural breaks detection in the mean and variance did not much change, comparing to the findings in the Table 2.
Source: Author's calculation

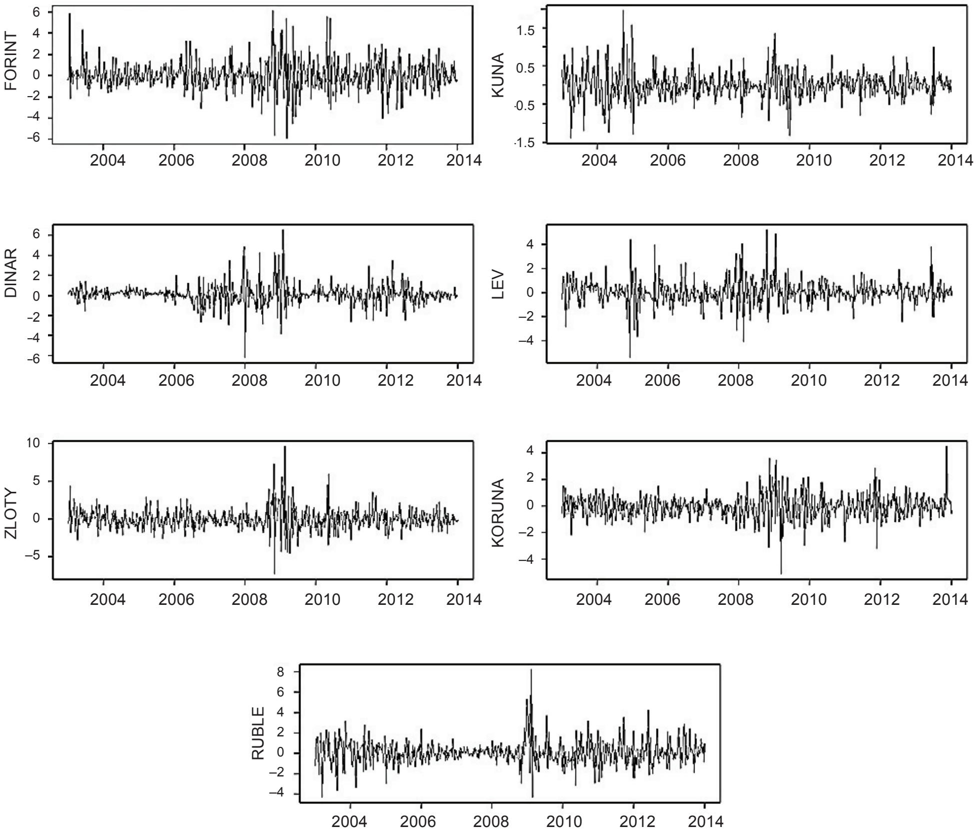
5. Conclusion

This paper investigates whether UIRP principle holds and what is predominant driving force that influences its movement – economic fundamentals or short-term speculative behaviour. Analysis covers seven ESEE transition countries which exchange rates are not fixed but range from *de facto* flexible regime (the Czech Republic, Poland, Romania and Serbia), pegged regime with wide bands (Hungary) to pegged regime with narrow bands (Croatia, Russia). We applied CGARCH-M methodology, since it gives us the ability to determine the main causes of exchange rate volatility. Preliminary findings suggest that all unconditional distributions of exchange rate series are prone to asymmetric distribution and leptocurticity. The implementation of GARCH models reduces non-normal properties to some extent, but not completely. The likely causes are structural breaks and outliers, due to world crisis 2007–2009 and the fact that we investigated the transition economies which are less resilient to external shocks. In order to clear that doubt, we utilized Bai-Perron and Sanso tests, which confirmed that all models contain several structural breaks both in the mean and the variance series. We have circumvented this problem in later models by using dummy variables.

Structural breaks and outliers refinements have been done in order to gain confidence in estimated parameters' values. It turned out that it was justifiable, because findings from purified models significantly differentiate from basic models' findings. However, both basic CGARCH-M models as well as break-corrected models suggested that UIRP principle did not hold in any country. CGARCH-M findings also confirmed that despite the fact that there was an interest rate span, carry-trade opportunity did not influence the dynamics of the exchange rates in the short run. It turns out that only Hungarian Forint has significant coefficient in front of IRD, but with much lower value than ideal value of one, which UIRP propose. These findings stand in line with other studies which consider UIRP in emerging and transition countries. As for GARCH parameters, we found conflicting results offered by basic and break-refined models. Basic models suggested that exchange rate volatility is driven by permanent component in all analysed countries and by transition component in four out of seven countries. However, after the inclusion of dummy variables, we found that permanent effect is significant in five countries, while the transition affect is not present in any country whatsoever. Additionally, we have discovered that short-term transition component played important role after further outliers' purification only in Serbia, while the possible cause could be actively conducted monetary policy of NBS, which included money sterilization and exchange rate stabilization.

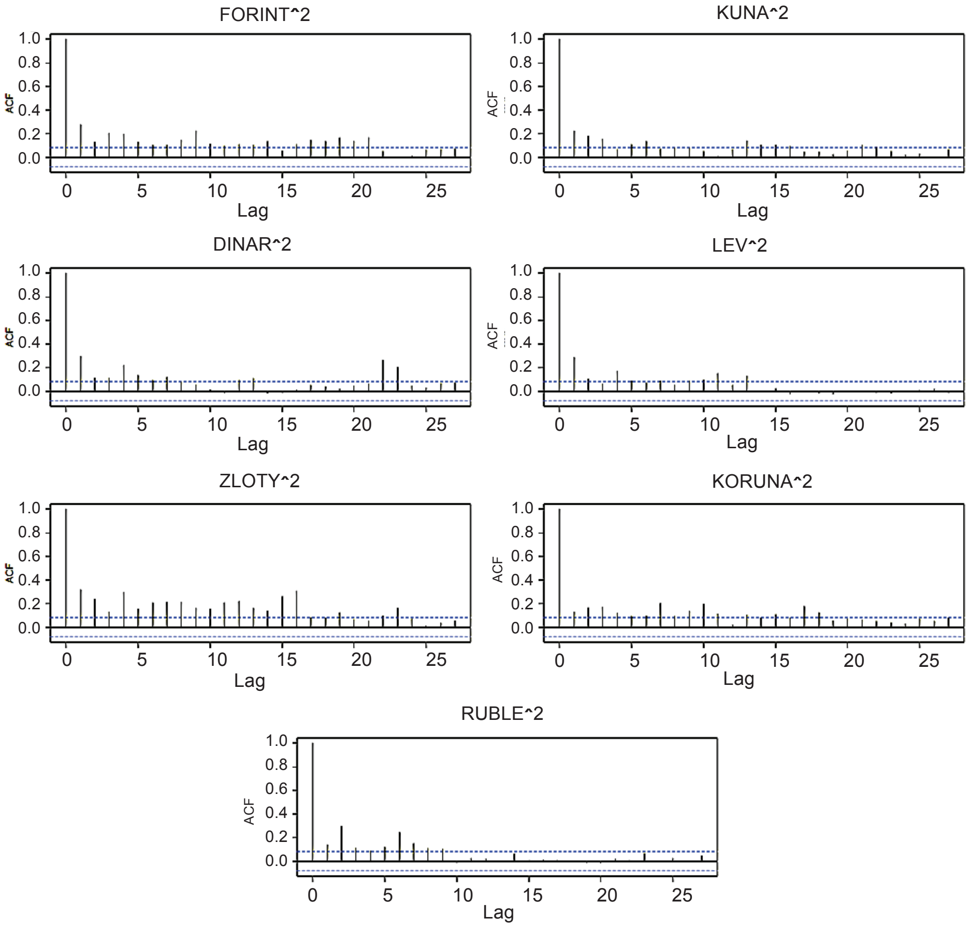
Appendix

Figure 1 | Exchange Rate Changes for Selected Countries



Source: www.globalfinancialdata.com

Figure 2 | Autocorrelation Functions of Squared Data



Source: Author's calculation

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