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**Nikola Gradojević**

Faculty of Business Administration,
Lakehead University,
Canada;

Faculty of Technical Sciences,
University of Novi Sad,
Serbia

✉ ngradoje@lakeheadu.ca

**Vladimir Djaković**

Faculty of Technical Sciences,
University of Novi Sad,
Serbia

✉ v_djakovic@yahoo.com

**Goran Andjelić**

Faculty of Technical Sciences,
University of Novi Sad,
Serbia

✉ gandjelic@uns.ac.rs

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# Random Walk Theory and Exchange Rate Dynamics in Transition Economies

**Summary**: This paper investigates the validity of the random walk theory in the Euro-Serbian dinar exchange rate market. We apply Andrew Lo and Archie MacKinlay's (1988) conventional variance ratio test and Jonathan Wright's (2000) non-parametric ranks and signs based variance ratio tests to the daily Euro/Serbian dinar exchange rate returns using the data from January 2005 - December 2008. Both types of variance ratio tests overwhelmingly reject the random walk hypothesis over the data span. To assess the robustness of our findings, we examine the forecasting performance of a non-linear, non-parametric model in the spirit of Francis Diebold and James Nason (1990) and find that it is able to significantly improve upon the random walk model, thus confirming the existence of foreign exchange market imperfections in a small transition economy such as Serbia. In the last part of the paper, we conduct a comparative study on how our results relate to those of other transition economies in the region.

**Key words:** Random walk, Forecasting, Exchange rates, Transition economies, Market efficiency, Artificial neural networks.

**JEL:** F31, G14, C53.

Since the seminal work of Richard Meese and Kenneth Rogoff (1983), scholars have paid considerable attention to exchange rate forecasting. This paper showed that a simple random walk model performed no worse than any of the standard macroeconomic exchange rate models. Even after including ex-post data on the fundamentals, out-of-sample forecasting performance at 1-, 6- and 12-month horizons was surprisingly low.

After more than 25 years of research, however, producing a short-run exchange rate forecasting model that would be more accurate than the random walk model has remained a major challenge to policy makers and practitioners. Initially the literature showed that the efforts were focused on linking the macroeconomic (fundamental) variables to exchange rates at medium to long forecast horizons. The findings by Meese and Rogoff (1983) were reinforced by a number of authors such as Meese and Andrew Rose (1991), Robert Flood and Rose (1995), Jeffrey Frankel and Rose (1995) and Min Qi and Yangru Wu (2003). Contrary to this literature, evidence that favors macroeconomic approach was found in Nelson Mark (1995), Menzie Chinn and Meese (1995), Mark and Donggyu Sul (2001), and Lutz Kilian and Mark Taylor (2003). While there has been a significant criticism related to the statistical robustness of these results (e.g., Kilian 1999), it has become apparent that some of the gains in the forecasting performance were due to accounting for non-linearities in the data (e.g., Kilian and Taylor 2003). Noteworthy, Meese and Rose (1991) and Qi and Wu (2003) did not find non-linearities and market fundamentals useful for lower frequency forecasting.

Recent literature has documented a significant short-run relationship between contemporaneous and lagged currency order flows and spot exchange rate movements (see e.g., Martin Evans and Richard Lyons 2005; Nikola Gradojević 2007). Moreover, order flows were found to be informative at medium (up to six months) and long run horizons (William Killeen, Lyons, and Michael Moore 2006). The new approach, often referred to as "New Micro Exchange Rate Economics" (Lyons 2001), calls attention to imperfections of financial markets: incomplete markets, sticky prices and various deviations from rational expectations due to over-reaction to news, noise or technical trading. This idea was pioneered by Albert Kyle (1985) and extended by many other authors.[1] In the foreign exchange market context, Flood and Rose (1995) suggested that more microeconomic detail should be taken into account. Similarly, Yin-Wong Cheung and Clement Yuk-Pang Wong (2000) conducted a survey of practitioners in the interbank foreign exchange markets and reported significant deviations from rational expectations: only 1% of the traders look at macroeconomic fundamentals to determine short-run exchange rate movements. Further, when Oliver Jeanne and Rose (2002) incorporated noise traders into the general equilibrium framework, they showed that for a fixed level of volatility fundamentals based on different levels of noise trading different levels of exchange rate volatility can occur. Within the partial equilibrium framework, Lyons and Evans (2002) included a variable reflecting the microeconomics of asset pricing to an exchange rate model. They introduced *order flow* as the proximate determinant of the exchange rate (using daily data over a four-month period) and were able to significantly improve on existing macroeconomic models. More precisely, they managed to capture about 60% of the daily exchange rate changes using a linear model. In a recent paper, using a linear order flow model on different data, Evans and Lyons (2005) managed to generate statistically significant forecasting improvements relative to the random walk model. However, certain concerns related primarily to order flow endogeneity (Martin Boyer and Simon van Norden 2006; Michael Sager and Taylor 2008) have been noted and cast doubt on the validity of their results.

A number of other scholars have pursued exchange rate modeling and forecasting using various methodologies, but with mixed success. To model the observed conditional heteroskedasticity of exchange rates, ARCH (David Hsieh 1989) and GARCH (Tim Bollerslev 1990) models were employed, but the results were very discouraging. For instance, in Ramazan Gençay (1999), a GARCH model generated insignificant directional and mean-squared prediction error forecast improvements over a simple random walk. Paul Boothe and Deborah Glassman (1987), Hsieh (1988), Richard Baillie and Patrick McMahon (1989), and Diebold and Marc Nerlove (1989) reported that the exchange rate changes are leptocurtic and might be non-linearly dependent. Chung-Ming Kuan and Tung Liu (1995) used backpropagation

[1] See Maureen O' Hara (1995) for more information on this approach, generally known as market microstructure theory.

and recurrent artificial neural networks (ANNs) and detected non-linearities in the daily Japanese yen and British pound time series for the period 1980-1985. Some other studies involving ANNs, such as Gioqinang Zhang and Michael Hu (1998) and Hu et al. (1999), showed similar results: non-linear exchange rate forecasting based on its lagged values can be fruitful. Alternatively, technical trading signals can be constructed from the time series of spot exchange rates and used as forecasting variables in both linear and non-linear models. This intriguing possibility was researched and documented in Richard Levich and Lee Thomas (1993), Francesco Lisi and Alfredo Medio (1997), Gençay (1999), Lo, Harry Mamaysky, and Jiang Wang (2000). In all of these studies, the results contradicted the weak form of market efficiency.

This paper seeks to determine the appropriateness of the random walk hypothesis (RWH) for the Euro (EUR)-Serbian dinar (RSD) exchange rate market, i.e., we investigate whether successive RSD/EUR exchange rate changes are random and serially independent. In this setting, the rejection of the RWH would imply predictability of exchange rates based on an autoregressive structure. The contribution of our study is threefold. First, this is, to the authors' best knowledge, the first analysis concerning the RWH in the Serbian foreign exchange market for the transition period after 2000. Our work complements the study by Jesus Crespo-Cuaresma and Jaroslava Hlouskova (2005) that focuses on more developed transition economies, namely Hungary, Czech Republic, Poland, Slovakia and Slovenia; their findings reveal support for the RWH.[2] It would be important to confirm whether the RWH holds for a less mature and sluggish transition economy such as Serbia. Second, to investigate the RWH, as in Jeng-Hong Chen (2008), we apply a battery of tests that include Lo-MacKinlay's (1988) conventional variance ratio test and Wright's (2000) non-parametric ranks and signs based variance ratio tests. These methodologies are also supplemented by a robust non-parametric, non-linear artificial neural network (ANN) model that extends the approach by Diebold and Nason (1990). Finally, we utilize a very recent data set that covers the 2005-2008 period, making our results a useful input for current central bank policy making.

The rest of the paper is organized as follows: in Section 1, we briefly explain our data set and methodology. The results of our tests are reported in Section 2. Section 3 offers a detailed comparative study across transition economies in the region. Section 4 concludes.

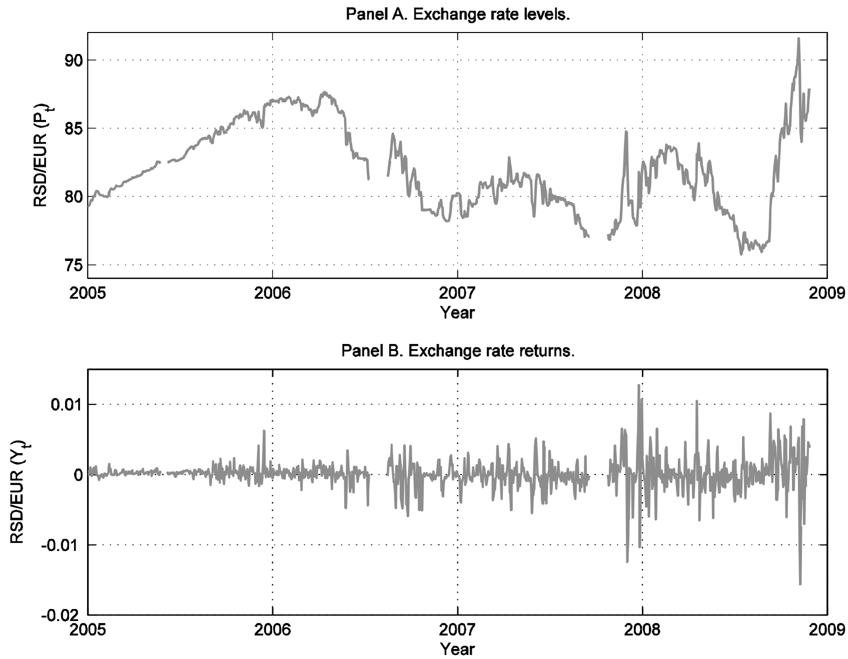
# 1. Data and Methodology Overview

## 1.1 Data

The data used in this research are from the National Bank of Serbia (NBS) and represent daily closing RSD/EUR exchange rates from January 4, 2005 to December 31, 2008 (Figure 1). The exchange rate returns ($Y_t$) are calculated in the standard fashion

[2] A few other contributions addressed the RWH in emerging and transition economies. Eui Jung Chang, Eduardo Lima and Benjamin Tabak (2004) and Suzanne Fifield, David Power, and Donald Sinclair (2005) uncovered evidence of market inefficiencies in emerging equity markets. In contrast, Claire Gilmore and Ginette McManus (2003) and, more recently, Nikolaos Giannellis and Athanasios Papadopoulos (2009) present mixed findings for equity and foreign exchange markets in Poland, Czech Republic, Slovakia and Hungary.

by taking the first differences of the natural logarithm of the exchange rates ($Y_t = \log P_t - \log P_{t-1}$).




**Source**: Authors' calculations.

**Figure 1** RSD/EUR nominal exchange rate levels (Panel A) and returns (Panel B).

Table 1 shows the basic statistics for the research sample (daily return series of the RSD/EUR exchange rate). Both Dickey-Fuller and Phillips-Perron tests reject the null hypothesis of a unit root in $Y_t$ at the 1% significance level (p-value=0.000).

**Table 1** Summary Statistics for the RSD/EUR Exchange Rate Returns

| | |
|---|---:|
| Mean | 0,00011145 |
| Standard Deviation | 0,00484457 |
| Minimum | -0,03666535 |
| Maximum | 0,02915062 |
| 1st quartile | -0,00143438 |
| Median | 0,00004631 |
| 3rd quartile | 0,00150620 |
| Skewness | -0,39655689 |
| Kurtosis | 9,39609755 |

**Source:** Authors' calculations.

## 1.2 Variance Ratio Test by Lo and MacKinlay (1988)

The variance ratio test of Lo and MacKinlay (1988) is based on the property that the variance of increments of a random walk $X_t$ is linear in its data interval. That means, the variance of $(X_t - X_{t-q})$ is $q$ times the variance of $(X_t - X_{t-1})$. Therefore, the RWH can be checked by comparing $1/q$ times the variance of $(X_t - X_{t-q})$ to the variance of $(X_t - X_{t-1})$.

Suppose $P_t$ is the exchange rate at time $t$ and let a random walk series $X_t$ be the natural logarithm of $P_t$ [$X_t = \ln P_t$]. The variance ratio, $VR(q)$ is defined as:

$$VR(q) = \frac{\sigma^2(q)}{\sigma^2(1)} \tag{1}$$

where $\sigma^2(q)$ is $1/q$ times the variance of $(X_t - X_{t-q})$ and $\sigma^2(1)$ is the variance of $(X_t - X_{t-1})$. The null hypothesis is that $VR(q)$ is not statistically different from 1. The equations to calculate $\sigma^2(1)$ and $\sigma^2(q)$ are as follows:

$$\sigma^2(1) = \frac{1}{nq-1}\sum_{t=1}^{nq}(X_t - X_{t-1} - \hat{\mu})^2 \tag{2}$$

where

$$\hat{\mu} = \frac{1}{nq}\sum_{t=1}^{nq}(X_t - X_{t-1}) = \frac{1}{nq}(X_{nq} - X_0) \tag{3}$$

and

$$\sigma^2(q) = \frac{1}{m}\sum_{t=q}^{nq}(X_t - X_{t-q} - q\hat{\mu})^2 \tag{4}$$

where

$$m = q(nq - q + 1)(1 - \frac{q}{nq}) \tag{5}$$

$X_{nq}$ is the last observation of the data time series. The observation starts at $X_0$. There are nq+1 observations.

The asymptotically standard normal test statistic used to test the null hypothesis of random walk under the assumption of homoscedasticity is Z(q), calculated as:

$$Z(q) = \frac{VR(q) - 1}{\sqrt{\theta(q)}} \tag{6}$$

where

$$\theta(q) = \frac{2(2q-1)(q-1)}{3q(nq)} \tag{7}$$

The asymptotically standard normal test statistic that is heteroscedasticity-consistent, $Z^*$(q), is calculated as follows:

$$Z^*(q) = \frac{VR(q)-1}{\sqrt{\theta^*(q)}} \tag{8}$$

where

$$\theta^*(q) = \sum_{j=1}^{q-1}\left[\frac{2(q-j)}{q}\right]^2 \hat{\delta}(j) \tag{9}$$

and

$$\hat{\delta}(j) = \frac{\sum_{t=j+1}^{nq}(X_t - X_{t-1} - \hat{\mu})^2 (X_{t-j} - X_{t-j-1} - \hat{\mu})^2}{\left[\sum_{t=1}^{nq}(X_t - X_{t-1} - \hat{\mu})^2\right]^2} \tag{10}$$

### 1.3 Rank-Based Variance Ratio Tests by Wright (2000)

Wright (2000) indicates two potential advantages of ranks and signs based tests. First, it is relatively simple to calculate their exact distributions. Size distortions are not a concern due to no need to conform to any asymptotic approximation. Second, tests based on ranks and signs may be more powerful than other tests if the data are highly non-normal. Wright (2000) proposes the alternative non-parametric variance ratio tests using ranks and signs of return and demonstrates that they may have better power properties than other variance ratio tests.

Suppose that $Y_t$ is a time series of asset returns with a sample size of T. $Y_t = X_t - X_{t-1}$ . Let r ($Y_t$) be the rank of $Y_t$ among $Y_1$, $Y_2$,…, $Y_T$ . r ($Y_t$) is the number from 1 to T.

Define

$$r_{1t} = \frac{\left(r(Y_t) - \frac{T+1}{2}\right)}{\sqrt{\frac{(T-1)(T+1)}{12}}} \tag{11}$$

$$r_{2t} = \Phi^{-1}\left(\frac{r(Y_t)}{T+1}\right) \tag{12}$$

where $\Phi$ is the standard normal cumulative distribution function ($\Phi^{-1}$ is the inverse of the standard normal cumulative distribution function).

The series $r_{1t}$ is a simple linear transformation of the ranks, standardized to have sample mean 0 and sample variance 1. The series $r_{2t}$, known as the inverse normal or van der Waerden scores, has sample mean 0 and sample variance approximately equal to 1. Wright substitutes $r_{1t}$ and $r_{2t}$ in place of the return ($X_t - X_{t-q}$) in the definition of Lo-MacKinlay's variance ratio test statistic (assuming homoscedasticity), $Z(q)$ in equation (6). The rank-based variance ratio test statistics $R_1$ and $R_2$ are defined as:

$$R_1 = \left( \frac{\frac{1}{Tk}\sum_{t=k}^{T}\left(r_{1t} + r_{1t-1} \ldots + r_{1t-k+1}\right)^2}{\frac{1}{T}\sum_{t=1}^{T} r_{1t}^2} - 1 \right) \times \left( \frac{2(2k-1)(k-1)}{3kT} \right)^{-1/2} \tag{13}$$

$$R_2 = \left( \frac{\frac{1}{Tk}\sum_{t=k}^{T}\left(r_{2t} + r_{2t-1} \ldots + r_{2t-k+1}\right)^2}{\frac{1}{T}\sum_{t=1}^{T} r_{2t}^2} - 1 \right) \times \left( \frac{2(2k-1)(k-1)}{3kT} \right)^{-1/2} \tag{14}$$

Note that $\frac{1}{T}\sum_{t=1}^{T} r_{1t}^2 = 1$ so that this term may be omitted from the definition of $R_1$ in equation (13), whereas $\frac{1}{T}\sum_{t=1}^{T} r_{2t}^2 \approx 1$. The exact sampling distributions for $R_1$ and $R_2$ should be simulated and their critical values are listed in Wright (2000).

### 1.4 Sign-Based Variance Ratio Tests by Wright (2000*)*

For any series $Y_t$, let $u(Y_t, q) = 1(Y_t > q) - 0.5$. So, $u(Y_t, 0)$ is ½ if $Y_t$ is positive and −½ otherwise. Let $s_t=2u(Y_t, 0) = 2u(\varepsilon_t, 0)$ . Clearly, $s_t$ is an independently and identically distributed (iid) series with mean 0 and variance 1. Each $s_t$ is equal to 1 with probability ½ and is equal to −1 otherwise. The signed-based variance ratio test statistic $S_1$ is defined as

$$S_1 = \left( \frac{\frac{1}{T_k}\sum_{t=k}^{T}\left(s_t + s_{t-1} \ldots s_{t-k+1}\right)^2}{\frac{1}{T}\sum_{t=1}^{T} s_t^2} - 1 \right) \times \left( \frac{2(2k-1)(k-1)}{3kT} \right)^{-1/2} \tag{15}$$

Similar to the above, the sampling distribution for $S_1$ should be simulated and the critical values can be found in Wright (2000).

## 1.5 Backpropagation Artificial Neural Networks (ANNs)

ANNs represent a general class of non-parametric, non-linear models that had been originally conceptualized for pattern recognition and system control purposes, but subsequently found applications in finance and economics.[3]

Suppose that a single hidden layer ANN is composed of $s$ input and $q$ hidden nodes whereas the $i^{th}$ independent variable is denoted by $x_{it}$ ($i$=1,...,$s$). The hidden and the output layers are characterized by two arbitrary types of non-linearities: $\psi$ and $\sigma$, respectively. Backpropagation learning algorithm requires continuous differentiable non-linearities and the most commonly used type is the sigmoid logistic (or logsig) function:

$$f(w) = \frac{1}{1+e^{-w}} \tag{16}$$

The dependent variable ($y_t$) is written as

$$y_t = \sigma\left( \beta_0 + \sum_{j=1}^{q} \beta_j \psi\left( \alpha_{j0} + \sum_{i=1}^{s} \alpha_{ij} x_{it} \right)\right) + \varepsilon_t \tag{17}$$

where $\alpha_{ij}$ and $\beta_j$ denote appropriate connection weights between the adjacent layers. Subscripts 0 for $\alpha$ and $\beta$ stand for ANN biases. Other types of transfer functions used in this paper are hyperbolic sigmoid tangent and linear.

Studies by George Cybenko (1989) and Ken-Ichi Funahashi (1989) show that the non-linear representation given by equation (17), with $\psi$ given by equation (16) can approximate a large number of mappings between $x_t$'s and $y_t$ reasonably well.

In the first step of the implementation of the ANN model, the data is divided into training, validation, and testing parts, roughly in the ratio 6:3:1.[4] The selection of the number of hidden layers and nodes in them is guided by the ANN's performance on the validation data with respect to the MSPE. Cross validation revealed that the optimal number of hidden layers is {2} and the number of hidden nodes in the two layers {3 and 5}. The parameters are estimated using the standard Levenberg-Marquardt algorithm. Overfitting is prevented by early stopping, i.e., stopping the training process when the validation set error starts to increase. To control for data snooping biases, as in Rene Garcia and Gençay (2000), the robustness of the ANN model is explored from the aspect of repeating the parameter estimation from five different sets of starting values.

The explanatory variables in the model are lagged dependent variables and, thus, the forecasting model becomes a non-linear autoregressive one:

[3] More extensive review of various applications of ANNs in finance and economics can be found in Qi (1996).

[4] During the robustness analysis, due to a smaller sample size, this ratio was maintained at 5:3:2 for individual years.

$$r_t = \phi(r_{t-1},...,r_{t-p}) + \varepsilon_t,\ p \in \{3,5\} \tag{18}$$

# 2. Results

## 2.1 Conventional Variance Ratio Test by Lo and MacKinlay (1988)

Tables 2 and 3 list the test statistics of the RWH for the entire data span and individual years based on the methodology of conventional variance ratio test by Lo and MacKinlay (1988). The results in Table 2 indicate support for rejecting the null hypothesis that variance ratio is not statistically different from one, except for q=16. Therefore, the RWH for the RSD/EUR exchange rate returns is rejected.

**Table 2** Lo-MacKinlay's Test Statistics of the RWH for the Entire Data Period (4/01/2005 - 31/12/2008)

| q | 2 | 4 | 8 | 16 |
|---|---|---|---|---|
| Z(q) | 11.2235555* | 7.6373054* | 2.4904837** | 0.3503222 |
| Z*(q) | 4.5303668* | 3.3784079* | 1.2205851 | 0.1928132 |

**Notes:** q denotes q-day returns, Z(q) is variance ratio test statistics assuming homoskedasticity, Z*(q) is variance ratio test statistics assuming heteroskedasticity. Under the random walk null hypothesis, the test statistic is asymptotically distributed as standard normal. (*), (**) and (***) indicates the test statistic is significant at 1%, 5% and 10% significance level, respectively.

**Source:** Authors' calculations.

Table 3 reports different results for individual years: for the most of the years and values of q, the RWH holds. Nevertheless, it appears that in 2006 the evidence of random walk starts disappearing. This is followed by a stronger rejection of the RWH in 2007 and 2008, when q=2 and q=4. We conclude that the findings in Table 2 are in general driven by violations of the random walk model in more recent years.

Given that data are characterized by changing volatility, more weight should be given to the variance ratio test corrected for heteroskedasticity.[5] Therefore, based on Z*(q), we conclude that the results for all years are not as strong for longer horizons (q=8 and q=16 days), but still reject the RWH for q=2 and q=4. In the same vein, the RWH is rejected only for short horizons in 2007 and 2008, while the rejections of the RWH in 2006 can be ignored due to the observed insignificant Z*(q) test statistics for that year.

**Table 3** Lo-MacKinlay's Test Statistics of the RWH for Individual Years

| | | | | |
|---|---|---|---|---|
| 2005 | | | | |
| q | 2 | 4 | 8 | 16 |
| Z(q) | 0.5679029 | 0.1471480 | -1.5199489 | -1.6619191*** |
| Z*(q) | 0.32715856 | 0.08883021 | -0.95604774 | -1.12076179 |
| 2006 | | | | |
| q | 2 | 4 | 8 | 16 |
| Z(q) | 6.454338* | 6.503556* | 5.428026* | 4.175687* |
| Z*(q) | 1.220191 | 1.368004 | 1.363370 | 1.322564 |

[5] We thank the anonymous referee for this and other useful suggestions.

| | | | | |
|---|---|---|---|---|
| 2007 | | | | |
| q | 2 | 4 | 8 | 16 |
| Z(q) | 5.8890030* | 3.8108268* | 0.9940270 | -0.7012695 |
| Z*(q) | 2.4384726** | 1.7726138*** | 0.5477701 | -0.4553079 |
| 2008 | | | | |
| q | 2 | 4 | 8 | 16 |
| Z(q) | 4.8176981* | 3.1437711* | 0.6681710 | 0.1326565 |
| Z*(q) | 2.64473243* | 1.88701517*** | 0.44647631 | 0.09889118 |

**Notes**: q denotes q-day returns, Z(q) is variance ratio test statistics assuming homoskedasticity, Z*(q) is variance ratio test statistics assuming heteroskedasticity. Under the random walk null hypothesis, the test statistic is asymptotically distributed as standard normal. (*), (**) and (***) indicates the test statistic is significant at 1%, 5% and 10% significance level, respectively.

**Source**: Authors' calculations.

## 2.2 Signs and Ranks Test by Wright (2000)

The results of the Wright's sign-based variance ratio test are presented in Tables 4 and 5. As in subsection 3.2, we first show the test statistics for the full sample (Table 4) and then concentrate on individual years (Table 5). Table 4 suggests that the RWH is strongly rejected for all q values at the 5% significance level.

**Table 4** Wright's Sign-Based Test Statistics of the RWH for the Entire Data Period (4/01/2005 - 31/12/2008)

| Q | 2 | 4 | 8 | 16 |
|---|---|---|---|---|
| $S_1$(q) | 8.027764** | 8.009905** | 6.811306** | 7. 345206** |

**Notes:** q denotes q-day returns, $S_1$(q) is variance ratio sign-based test statistics. Under the random walk null hypothesis, the critical values are based on Table 1 from Wright (2000). (**) indicates the test statistic is significant at the 5% significance level.

**Source:** Authors' calculations.

The sign-based test results from Table 5 give strong rejections of the null hypothesis in each year. These findings therefore reinforce somewhat weak evidence against the RWH found by applying the Lo-MacKinlay's test on individual years.

**Table 5** Wright's Sign-Based Test of the RWH For Individual Years

| q | 2 | 4 | 8 | 16 |
|---|---|---|---|---|
| 2005 | | | | |
| $S_1$(q) | 4.860197** | 5.938022** | 6.017391** | 7.664609** |
| 2006 | | | | |
| $S_1$(q) | 15.65299** | 24.89738** | 36.14193** | 50.31838** |
| 2007 | | | | |
| $S_1$(q) | 5.343904** | 4.671115** | 3.028660** | 2.553083** |
| 2008 | | | | |
| $S_1$(q) | 2.451909** | 2.217939** | 1.222091 | 1.131926 |

**Notes:** q denotes q-day returns, $S_1$(q) is variance ratio sign-based test statistics. Under the random walk null hypothesis, the critical values are based on Table 1 from Wright (2000). (**) indicates the test statistic is significant at the 5% significance level.

**Source:** Authors' calculations.

The rank-based tests statistics for all four years are shown in Table 6. Clearly, the RWH is overwhelmingly rejected for q's ranging from 2 to 16, which is in line with the conclusions based on Tables 2 and 4.

**Table 6** Wright's rank-based test statistics of the RWH for the entire data period (4/01/2005 - 31/12/2008)

| Q | 2 | 4 | 8 | 16 |
|---|---|---|---|---|
| $R_1(q)$ | 10.199353** | 8.216612** | 4.219845** | 2.963116** |
| $R_2(q)$ | 11.009022** | 8.406201** | 3.761043** | 2.192301** |

**Notes:** q denotes q-day returns, $R_1(q)$ and $R_2(q)$ are variance ratio rank-based test statistics. Under the random walk null hypothesis, the critical values are based on Table 1 from Wright (2000). (**) indicates the test statistic is significant at the 5% significance level.

**Source:** Authors' calculations.

Almost all of the test statistics in Table 7 are statistically significant at 5%. These results are stronger than the ones obtained by the Lo-MacKinlay's test (Table 3). In all, Tables 2-7 uncover compelling evidence of the violations of the RWH in the Serbian foreign exchange market.

**Table 7** Wright's Rank-Based Test of the RWH for Individual Years

| | | | | |
|---|---|---|---|---|
| 2005 | | | | |
| q | 2 | 4 | 8 | 16 |
| $R_1(q)$ | 2.9606816** | 2.9275567** | 0.9813035 | -0.2990660 |
| $R_2(q)$ | 2.2856833** | 2.1864336** | 0.2989632 | -0.8443084 |
| 2006 | | | | |
| q | 2 | 4 | 8 | 16 |
| $R_1(q)$ | 15.40102** | 24.09902** | 33.84267** | 44.01952** |
| $R_2(q)$ | 14.84317** | 22.60606** | 30.44595** | 37.29629** |
| 2007 | | | | |
| q | 2 | 4 | 8 | 16 |
| $R_1(q)$ | 6.250553** | 4.716401** | 1.893261** | 1.033080 |
| $R_2(q)$ | 6.5084786** | 4.8290712** | 1.8392112 | 0.5506108 |
| 2008 | | | | |
| q | 2 | 4 | 8 | 16 |
| $R_1(q)$ | 4.884942** | 3.239261** | 1.329731 | 1.215336 |
| $R_2(q)$ | 4.7080317** | 3.0297958** | 0.8430305 | 0.6060859 |

**Notes:** q denotes q-day returns, $R_1(q)$ and $R_2(q)$ are variance ratio rank-based test statistics. Under the random walk null hypothesis, the critical values are based on Table 1 from Wright (2000). (**) indicates the test statistic is significant at the 5% significance level.

**Source**: Authors' calculations.

## 2.3 ANN model

We run equation (18) to forecast out-of-sample in each of the individual years as well as for the whole sample (all years). The results for years 2005, 2006, 2007 and 2008 are listed in Table 8 (panels A, B, C and D). From the MSPE ratio[6] column, it is apparent that forecasting improvements are present in each year and for all experiments. However, although frequently sizable (15%-20%), due to a smaller sample size (roughly 40 out-of-sample observations), the forecasting improvements are not always statistically significant across years according to the Diebold and Roberto Mariano (1995) (DM) statistics.

When all 2005-2008 data are processed together, the results are consistently statistically significant and reveal substantial evidence against the RWH in the EUR-RSD market. Clearly, for the deviations from the RWH to become evident, the ANN model requires more estimation (training) data. Considering the results for individual years, this indicates that the RWH is not strongly rejected.

**Table 8** Forecasting Performance of the ANN ModelPanel A. Year: 2005

| RSD/EUR | | MSPE ratio | DM |
|---|---|---|---|
| *p*=3 | | | |
| | Experiment (1) | 0.9275 | -1.29*** |
| | Experiment (2) | 0.9049 | -1.71** |
| | Experiment (3) | 0.9553 | -1.04 |
| | Experiment (4) | 0.9080 | -1.41*** |
| | Experiment (5) | 0.9732 | -0.38 |
| *p*=5 | | | |
| | Experiment (1) | 0.9431 | -1.04 |
| | Experiment (2) | 0.9568 | -0.52 |
| | Experiment (3) | 0.9301 | -1.67** |
| | Experiment (4) | 0.8952 | -1.31*** |
| | Experiment (5) | 0.9071 | -1.44*** |

**Panel B. Year: 2006**

| RSD/EUR | | MSPE ratio | DM |
|---|---|---|---|
| *p*=3 | | | |
| | Experiment (1) | 0.8975 | -1.03 |
| | Experiment (2) | 0.8321 | -1.38*** |
| | Experiment (3) | 0.8764 | -1.65** |
| | Experiment (4) | 0.8627 | -1.19 |
| | Experiment (5) | 0.9621 | -1.02 |
| *p*=5 | | | |
| | Experiment (1) | 0.7199 | -1.37*** |
| | Experiment (2) | 0.8869 | -0.75 |
| | Experiment (3) | 0.9331 | -1.03 |
| | Experiment (4) | 0.7344 | -1.69** |
| | Experiment (5) | 0.8588 | -1.45*** |

**Panel C. Year: 2007**

| RSD/EUR | | MSPE ratio | DM |
|---|---|---|---|
| *p*=3 | | | |
| | Experiment (1) | 0.7682 | -1.72** |
| | Experiment (2) | 0.8951 | -1.35*** |
| | Experiment (3) | 0.9609 | -0.63 |

[6] The ratio of the ANN model's mean-squared prediction error to that of the random walk (no change) model over the last 20% of the observations. When all data are used, the last 10% of the observations are kept out-of-sample (about 100 observations).

| | | | |
|---|---|---|---|
| | Experiment (4) | 0.9273 | -0.98 |
| | Experiment (5) | 0.8990 | -1.29*** |
| *p*=5 | | | |
| | Experiment (1) | 0.8450 | -1.66** |
| | Experiment (2) | 0.9122 | -1.33*** |
| | Experiment (3) | 0.9321 | -0.87 |
| | Experiment (4) | 0.8181 | -2.21** |
| | Experiment (5) | 0.9359 | -1.02 |

**Panel D. Year: 2008**

| RSD/EUR | | MSPE ratio | DM |
|---|---|---|---|
| *p*=3 | | | |
| | Experiment (1) | 0.9740 | -0.99 |
| | Experiment (2) | 0.9075 | -3.77* |
| | Experiment (3) | 0.9663 | -0.61 |
| | Experiment (4) | 0.8599 | -1.99** |
| | Experiment (5) | 0.8758 | -0.62 |
| *p*=5 | | | |
| | Experiment (1) | 0.8808 | -0.60 |
| | Experiment (2) | 0.9583 | -0.81 |
| | Experiment (3) | 0.9147 | -0.50 |
| | Experiment (4) | 0.9638 | -0.23 |
| | Experiment (5) | 0.9311 | -0.39 |

**Panel E. Year: 2005-2008 (all years)**

| RSD/EUR | | MSPE ratio | DM |
|---|---|---|---|
| *p*=3 | | | |
| | Experiment (1) | 0.9248 | -2.83* |
| | Experiment (2) | 0.9372 | -2.37* |
| | Experiment (3) | 0.9680 | -1.90** |
| | Experiment (4) | 0.9256 | -1.55*** |
| | Experiment (5) | 0.8227 | -2.25** |
| *p*=5 | | | |
| | Experiment (1) | 0.9231 | -1.29*** |
| | Experiment (2) | 0.8660 | -2.06** |
| | Experiment (3) | 0.8964 | -1.66** |
| | Experiment (4) | 0.8817 | -1.69** |
| | Experiment (5) | 0.9646 | -2.51* |

**Notes:** Experiment (1)-(5) denotes that the out-of-sample MSPEs are obtained from ANNs for which the parameters were estimated from five random seeds. The MSPE ratios column are the ratios of the ANN model's MSPE to that of the random walk model. DM denotes the Diebold and Mariano (1995) test statistic. Its one-sided critical values are -2.33, -1.645 and -1.282 for confidence levels of 99%, 95% and 90%, respectively. (*), (**) and (***) indicates the DM statistic is significant at 1%, 5% and 10% significance level, respectively.

**Source:** Authors' calculations.

## 3. Comparative Study: Serbian Dinar (RSD), Hungarian Forint (HUF), Czech Koruna (CZK), Slovak Koruna (SKK), Slovenian Tolar (SIT) and Polish Zloty (PLN)

Hungary, the Czech Republic, Poland, Slovakia, and Slovenia became transition economies long before Serbia (1989-1991) and completed the transition process by joining the European Union in 2004. A sensible comparative study should therefore utilize foreign exchange market data from the late 1990s for those countries. The contributions by Cuaresma and Hlouskova (2005) and Claire Gilmore and Ginnette McManus (2003) are from the period 1995-2000 which makes them ideal candidates for such a study. Giannellis and Papadopoulos (2009) focus on the subsequent period (1999-2006) and will also be used in our discussions.

The results of short-horizon exchange rate forecasting tests for Hungary for the period 1993-2000 are quite disappointing (Cuaresma and Hlouskova 2005). Some violations of the RWH are evident only at longer horizons (9-12 months), when the HUF/USD exchange rate was used. In contrast, Hungarian equity markets within the period of 1995-2000, show a different picture: daily data for the Budapest Stock Exchange reject the random walk model (Gilmore and McManus 2003). This indicates that transition reforms in Hungary were first reflected in the foreign exchange market, while the development of their equity market was lagging. It is worth noting that these results are in agreement with those observed for Slovakia, while forecasting exercises for the Czech Republic show no predictability of the CZK/EUR and CZK/USD exchange rates across all horizons.

In transitional Poland, an improved predictability in the foreign exchange market (the PLN/EUR and PLN/USD exchange rates) can be observed, but the sizable improvements apply only to longer horizons. Similar to the findings for Hungary and the Czech Republic, the evidence for the Warsaw Stock Exchange was in accord with the RWH. Surprisingly, there appears to be some Granger causality running from the Czech and Hungarian equity markets to the Warsaw Stock Exchange which may be explained by the higher levels of foreign direct investment in Hungary and the Czech Republic, relative to Poland (Gilmore and McManus 2003). An alternative explanation for this phenomenon could be that the former two economies were more successful in their transition efforts.

Of all the considered transition economies in the region, only Slovenia demonstrates considerable non-random walk effects for the SIT/EUR and SIT/USD exchange rates (Cuaresma and Hlouskova 2005), but not for very short horizons such as one to three months. This may be surprising given that Slovenia was one of the best performing transition economies, but can be explained by substantial central bank intervention and sluggish privatization during 1995-2000.

In the subsequent years, until 2006, the evidence suggests that the PLN/EUR market achieved efficiency, while certain market imperfections were still evident for the CZK/EUR (inefficient) and SKK/EUR (quasi-efficient) exchange rates. In the case of CZK/EUR exchange rate, one can argue that the observed deviations from market efficiency were caused by tight monetary policy (Giannellis and Papadopoulos 2009).

In comparison to the above, the results for Serbia are rather unique. At the very short (daily and weekly) and short horizons (1-3 months), when all of the other foreign exchange markets in the region were consistent with the RWH, the RSD/EUR exchange rates did not follow the random walk. Nevertheless, the results for the longer horizons are comparable and indicate significant departures from the RWH. In particular, this is the case with respect to the SIT/EUR exchange rate that showed predictability at 6-, 9-, and 12-month horizons. One natural explanation for such persistent non-random walk behaviour in the EUR-RSD market could be found in the efforts of the NBS to correct misalignments from equilibrium exchange rates in recent years. Any central bank intervention results in exchange rate fluctuations that are not entirely market-driven and may be inconsistent with efficiency. Furthermore, government interventions can be perceived by speculators as evidence of inefficiency. This in turn may initiate speculative attacks and induce excessive violations of the RWH. It is important to note, however, that due to political turbulence and war devastation, Serbia started with transition from a much weaker economy than other countries in the region. Thus, Serbian financial markets and institutions could be at the natural evolutionary stage of their development that is characterized by limited support for the RWH.[7]

## 4. Concluding Remarks

This paper has tested the RWH for the Serbian foreign exchange market. Specifically, by employing an array of methodologies, we focus on the predictability of the RSD/EUR exchange rate daily returns over the 2005-2008 period. First, we applied the standard Lo-MacKinlay's (1988) conventional variance ratio test and Wright's (2000) non-parametric ranks and signs based variance ratio tests. Although not as strong and consistent for individual years, the results suggested that RSD/EUR exchange rate movements can be predicted by using historical information. This evidence is particularly strong across methodologies when the whole data set is used. Thus, taken together, the findings in general reject the RWH.

Next, we explored the predictability of the RSD/EUR exchange rate by extending the autoregressive nearest-neighbors model by Diebold and Nason (1990) to a non-linear ANN specification. The results of this forecasting exercise are similar to the previous tests: statistically significant forecasting improvements are not robust for individual years, but non-random walk effects are dominant when the ANN model is estimated using the data for all years. We conclude that the observed potential market inefficiencies are present in the Serbian foreign exchange market only to a certain extent. We are also unable to detect any patterns in the degree of market efficiency across years. The absence of year-to-year regularities is inconsistent with the adaptive market hypothesis (Lo 2004) that one may expect to hold in a developing transition economy.

Further, our results generally contradict Cuaresma and Hlouskova (2005) in that they do not reject the RWH for the Czech Koruna, Hungarian Forint, Slovak Koruna, Slovenian Tolar and Polish Zloty. From the market efficiency perspective, this implies that the Serbian foreign exchange market is underdeveloped relative to other foreign exchange markets in the region, which is expected, since Serbia was one of the last former centrally planned economies to become a transition economy. It should also be noted that during the data span the NBS has extensively committed to maintaining currency stability and to taking preliminary steps toward full-fledged inflation targeting. Hence, our evidence of deviations from the RWH could also be an artifact of the economic interventionism and proactive policies of the NBS.

[7] Our unreported preliminary results also showed persistent violations of the RWH in the Serbian equity market – Belgrade Stock Exchange (BELEX15 and BELEXline indices).

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