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# Детерминанте номиналног динар-еуро девизног курса

Милан Недељковић Бранко Урошевић

# Determinants of the Dinar-Euro Nominal Exchange Rate

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**Детерминанте номиналног динар-еуро девизног курса**

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**Апстракт:** Овај рад истражује детерминанте дневних промена номиналног динар-еуро девизног курса у периоду од септембра 2006 до јуна 2010 године. Користећи иновативни семипараметрички приступ, различити типови нелинеарних веза су анализирани под нерестриктивим претпоставкама у вези неопаженог стохастичког процеса који генерише податке. У раду је идентификовано неколико фактора који утичу на дневне промене девизног курса и чији значај се мењао током времена. Информација у претходним промена девизног курса, промене у нивоу девизне штедње и промене у износу нето откупа девиза од стране банака су најзначајни фактори у периоду пре почетка светске економске кризе. Од септембра 2008 године други фактори везани за промену ризика земље и начин процесирања информација на девизном тржишту добијају на значају. Интервенције Народне Банке Србије имају ефекат на курс са периодом доцње.

**Кључне речи:** Девизно тржиште, парцијални линеарни модел, кернел оцењивање
**JEL Code: F31, C14, G18**

**Determinants of the Dinar-Euro Nominal Exchange Rate**

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**Abstract:** This paper studies drivers of daily dynamics of the nominal dinar-euro exchange rate from September 2006 to June 2010. Using a novel semiparametric approach we are able to incorporate the evidence of nonlinearities under very weak assumptions on the underlying data generating process. We identify several factors influencing daily exchange rate returns whose importance varies over time. In the period preceeding the financial crisis, information in past returns, changes in households' foreign currency savings and banks' net purchases of foreign currency are the most significant factors. From September 2008 onwards other factors related to changes in country's risk and the information processing in the market gain importance. NBS interventions are found to be effective with a time delay.

**Key words:** Foreign exchange market; Partially linear model; Kernel estimation
**JEL Code: F31, C14, G18**

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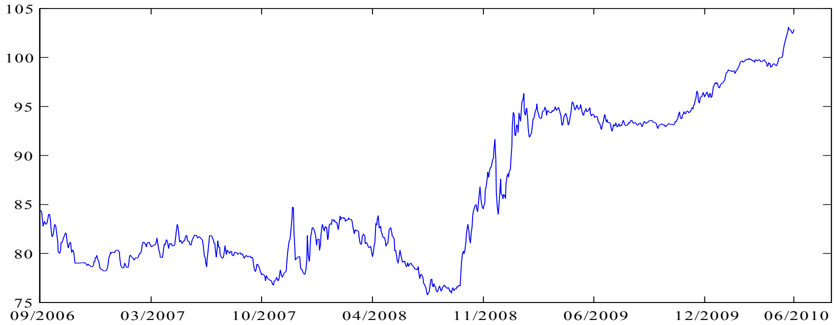
**Contents:**

## 1. Introduction

This paper present a first attempt to shed light on drivers of daily dynamics of the (nominal) Serbian dinar-euro exchange rate[1] from September 2006 to June 2010.[2] Figure 1 depicts evolution of the dinar-euro exchange rate over the analyzed period. Looking at the plot two periods are easily identified. The period before September 2008 is characterized by moderate volatility and slow appreciation of dinar vis-a-vis euro which reaches its maximum level at 75.754 on 07/08/2008. The reverse trend is observed following the beginning of the world-wide financial crisis, where after initial depreciation of 22% in three months from 6/10/2008 to 9/1/2009 and temporary stabilization, dinar is slowly, but continuously depreciating throughout the period.

**Figure 1**

**The Dinar-Euro Exchange Rate**




The observed behavior leads to two sets of interesting questions. The first set of questions that we aim to answer is related to the role of economic and financial fundamentals in explaining the observed behaviour of the nominal exchange rate.The literature on determinats of the nominal exchange rate in transition countries is relatively scarce. Crespo-Cuaresma, et al. (2005), Crespo-Cuaresma, et al. (2008) found support for the monetary model of the exchange determination as a long-run phenomen in a panel of Central European countries. The short-run dynamics of the exchange rate however are still largely unexplained. Ardic and Selcuk (2006), Egert and Komarek (2006) showed that changes in country risk and central bank

[1]Throughout the paper the exchange rate is defined as the number of units of domestic currency per 1 euro, hence increase in exchange rate implies dinars' depreciation.

[2] September 2006 marks the beginning of the free float regime of the dinar-euro exchange rate. Before that and since the beginning of transition in October 2000, the exchange rate had fixed parity before January 2003 and from January 2003 to September 2006 the exchange rate was defined through crawling peg regime.

interventions are the main drivers of the nominal exchange rate returns in Turkey and Czech Republic. The fundamentals considered in this paper include the main determinants of the supply and demand on the foreign exchange market as well as the measures of risk. We do not take into consideration traditional macro fundamentals since the small length of the sample and the fact that macro fundamentals are measured (at least) at monthly frequency preclude us from using them without relying on (subjective) interpolation methods. This is also in line with vast evidence in empirical literature that macro fundamentals perform poorly in explaining the short-run movements in floating exchange rates, although their importance increases in the long-run, as documented in the aforementioned studies.[3] The second set of questions is related to the operations of the National Bank of Serbia. In particular, we investigate the effectiveness of the daily sterilized interventions on the foreign exchange market and the role of changes in the bank reserve policy. The contribution of this paper is therefore twofold. First, to the best of our knowledge, this is the first empirical analysis of the determinants of the daily movements in dinar-euro exchange rate. Given the importance that the nominal exchange rate movements have on the real economy in small transition country like Serbia[4], identification of its determinants could provide useful information for both policy makers and market participants. Second, we use semiparametric techniques to estimate the relation of interest and highlight the benefits of such approach in comparison to standard linear models in the presence of different types of nonlinearities.

The rest of this paper is organized as follows: Section 2 presents econometric methodology. Section 3 describes the data and empirical results and Section 4 concludes the paper. All empirical results are presented in Appendix.

## 2. Methodology

We use partially linear additive model (PLAM) with (conditional) heteroscedasticity to model the determinants of the nominal exchange rate:

$$\Delta e_t = \alpha + D_t^{'}\delta + \sum_{j=1}^{p} g_j\left(X_{t,j}\right) + \sigma_t \varepsilon_t,\ t = 1..n \tag{1}$$

[3] For recent surveys on the short and long-run performance of purchasing power parity models see Taylor and Taylor (2004), and for monetary models of the exchange rate determination see Cheung, et al. (2005).

[4] See, for example, Hsing and Hsieh (2010) who include the exchange rate as on of the determinants of changes in the real output in Serbia. The importance of exchange rate as a determinant of the inflation rate through the exchange rate pass-through has received a considerable attention in the literature, for recent applications in transition countries see for example Cozmanca and Manea (2010) and references therein.

where $\Delta e_t$ is the percentage change in nominal exchange rate (log-return), $X_t \in R^p$ is a p-dimensional vector of independent variables which have a nonlinear effect on the nominal exchange rate and may also include the lagged values of $\Delta e_t$, $D_t \in R^d$ is a d-dimensional vector of independent variables that have linear influence on the exchange rate and random term $\varepsilon_t$ is assumed to be *i.i.d.* $g_i(\cdot)$ are unspecified univariate function of each variable $X_{t,j}$. Since we are primarily interested in the conditional mean movements we leave conditional variance $\sigma_t^2$ unspecified, but take into account conditional heteroscedasticity in estimation of the model parameters.

Specification given in equation (1) encompasses standard linear model that can be obtained by assuming that functions $g_i(\cdot)$ are linear. The assumption that the true data generating process (dgp) in the conditional mean is linear is relatively strong. Recent empirical literature on exchange rate modelling (see, for example, survey in Sarno and Taylor, 2003) documents that nonlinear effects are important in explaining the behaviour of nominal and real exchange rates in both developed and developing countries. Heterogeneity of agents in the foreign exchange market, the presence of (asymmetric) transaction costs, institutional rigidities may all lead to such situations. The common approach in all the non-linear studies is to modify linear specification and estimate a particular non-linear model for the relationship of interest, often from the class of smooth transition models. This, however, again puts relatively strong assumption that the researcher knows the type of nonlinear underlying dgp.

Instead of imposing a particular type of non-linear relationship between the variables of interest we consider a semiparametric additive partially linear model given in (1). Specification (1) thus provides a flexible generalization of the linear model and allows one to investigate non-linear effects without imposing any structure on the type of nonlinearity. At the same time, the specification is also robust to the curse of dimensionality that arises in case of fully nonparametric estimation[5].

The model (1) is estimated using recently proposed (Ma and Yang, 2010) spline backfitted kernel smoothing (SBKS) estimator. The estimator combines the benefits of the previously proposed spline and kernel estimators of the PLAM. Li (2000) showed that spline estimators of the PLAM, introduced by Hastie and Tibshirani (1990), are computationally efficient and consistent. However, their distributional properties remain unknown. Kernel estimators, on the other hand, have a well-defined limiting distribution, but require an additional step to control for the additive structure

[5]The curse of dimensionality arises if the conditional mean is specified as $\Delta e_t = \alpha + D_t\delta + G(X_t) + \sigma_t\varepsilon_t$, where function $G(\cdot)$ is *p*-dimensional. Since fully non-parametric estimation of function $G(\cdot)$ is performed over the $R^p$ space, the rate of convergence of the estimator thus depends on the dimension *p* and is much slower than $n^{1/2}$ for moderate levels of *p*, leading to potential inaccuracy of the small sample estimates, see more in Fan and Yao (2003).

of the data generating process. A common approach is to use a two-step kernel marginal integration (Fan, et al., 1998, Fan and Li, 2003). Since kernel marginal integration requires estimation of a high dimensional nonparametric function E(Δet|Xt) (without imposing additive structure) in the first step, this might introduce finite sample imprecision given the number of variables in the model, which may not be fully recovered by the second step estimation.

SBKS estimator combines the benefits of the two approaches. In the first step (undersmoothed) spline estimators $\overline{g}_j(\cdot)$ of the unknown functions and parameters {α, β} are obtained thus avoiding the need for kernel estimation of high dimensional nonparametric function. Next, these estimators are used to construct "oracle" responses $\overline{\Delta e_{t,j}}$ (as if the influence of variables other then Xj were known). The good distributional properties of the kernel estimators are then exploited using a local linear regression estimation on pairs {$\overline{\Delta e_{t,j}}$, $X_{t,j}$} to obtain the final estimate $\hat{g}_j(x_j)$.

In particular, the first step series estimation of the model in (1) is based on idea that each smooth function $g_j(x_j)$ can locally be well approximated (in the mean square sense) by a linear combination of the base functions $B_{j,s}$:

$$g_j(x_j) \approx \sum_{s=1}^{S_j} \phi_{j,s} B_{j,s}(x_j) \tag{2}$$

where the base functions Bj,s are assumed to be sufficiently smooth and satisfy the condition $B_{j,s}(x_j=0)=0$ in order for the functions $g_j(x_j)$ to be identified. We employ spline approximation where $B_{j,s}(\cdot)$ are linear B-spline basis (see de Boor, 2001) and Sj denotes the number of knots. The estimates $\left\{\overline{\alpha}, \overline{\delta}, \overline{\phi}_{j,s}, j=1..p;\ s=1..S_j\right\}$ are obtained by minimizing the least squares criterion:

$$\sum_{t=1}^{n}\left(\Delta e_t - \alpha - D_t^{'}\delta - \sum_{s=1}^{S_j}\phi_{j,s}B_{j,s}\left(X_{j,t}\right)\right)^2 \tag{3}$$

Pilot (first-step) estimators of parameters {α,β} are the least squares estimates $\{\overline{\alpha}, \overline{\delta}\}$. Pilot estimators $\overline{g}_j(X_j)$ are obtained after additional recentering as: $\overline{g}_i\left(X_{t,i}\right) = \sum_{i=1}^{S_i}\overline{\phi}_{i,s}B_{i,s}\left(X_{i,t}\right) - n^{-1}\sum_{t=1}^{n}\sum_{i=1}^{S_i}\overline{\phi}_{i,s}B_{i,s}\left(X_{i,t}\right)$.

The oracle (pseudo) responses are constructed as:

$$\overline{\Delta e_{t,j}} = \Delta e_t - \overline{\alpha} - D_t^{'}\overline{\delta} - \sum_{i=1, i\neq j}^{p}\overline{g}_i\left(X_{t,i}\right)$$

Final estimates $\hat{g}_j(x_j)$ are obtained using a kernel regression estimator on pairs {$\overline{\Delta e_{t,j}}$, $X_{t,j}$ }. It is well known that the standard kernel regression (Nadaraya-Watson) method suffers from the poor boundary behavior and larger asymptotic bias in comparison to local linear smoothers (Fan and Gijbels, 1996). Given the importance of the large movements in explanatory variables in the second part of the sample which are by construction located close to the boundary of the support, we use the local linear estimator to obtain $\hat{g}_j(x_j)$.

$$\hat{g}_j\left(x_j\right) = e\begin{pmatrix} S_0\left(x_j\right) & S_1\left(x_j\right) \\ S_1\left(x_j\right) & S_2\left(x_j\right) \end{pmatrix}^{-1} \begin{pmatrix} Z_0\left(x_j\right) \\ Z_1\left(x_j\right) \end{pmatrix} = S\left(x_j\right)^{-1} Z\left(x_j\right). \tag{4}$$

Here, $e$ is a selection vector $e = [0\ 1]$. For the index k=0,1,2:

$$S_k\left(x_j\right) = n^{-1}\sum_{t=1}^{n}\left(X_{t,j} - x_j\right)^k K_b\left(X_{t,j} - x_j\right)$$

$$Z_k\left(x_j\right) = n^{-1}\sum_{t=1}^{n}\left(X_{t,j} - x_j\right)^k K_b\left(X_{t,j} - x_j\right)\overline{\Delta e_{t,j}}$$

and $K_b\left(v\right) = b^{-1}K\left(\frac{v}{b}\right)$ is kernel function with bandwidth $b$.

Using Theorem 1 and 2 in Ma and Yang (2010) the confidence intervals (sets) for the estimated functions (parameters) can be obtained. The constructed confidence intervals are then employed for testing linearity of each explanatory variable. Under the null hypothesis of linearity the fitted linear function is covered by the semiparametric confidence interval. Any deviation of the linear fit outside the confidence interval thus provides evidence against the linearity of the effect of a particular variable.

Even though additive partially linear model is more flexible than the (parametric) linear model it is still possible that the model is misspecified. To guard against this possibility, we test the adequacy of the additive model which is equivalent to testing the null hypothesis:

$$H_o \;:\; E\left(\Delta e_t \mid X_t, D_t\right) = \alpha + D_t'\delta + \sum_{j=1}^{p} g_j\left(X_{t,j}\right) \; a.s \tag{5}$$

for some $\delta \in R^d$ and some smooth class of functions $g_j()$. The alternative hypothesis is negation of the null. To test the null hypothesis we employ the Cramer-von Mises (CM) type of statistic proposed by Li, et al. (2003)[6] and consider the following test statistic:

$$CM = \int_{R^p} \left|n^{1/2}\tau\left(\xi\right)\right|^2 d\chi\left(\xi\right) \tag{6}$$

[6] Li, et al. (2003) assume that the data is i.i.d. The proof of the limiting distribution with the dependent data can be established using central limit theorem for β-mixing Hilbert-valued variables, similar to Chen and Fan (1999).

where $\tau\left(\xi\right)=n^{-1}\sum_{t=1}^{n}\hat{\varepsilon}_t W\left(X_t,\xi\right)$, $W\left(\cdot\right)$ is a pre-specified function, $\hat{\varepsilon}_t$ are the residuals from estimation of model (1) and $d\chi\left(\xi\right)$ is a probability measure on Rp. Since the limiting distribution of the test statistic is unknown and conditional heteroscedasticity is present in the data we employ the fixed regressor wild bootstrap (Chen and Fan, 1999) to obtain the p-values of the test statistic.

## 3. Empirical Results

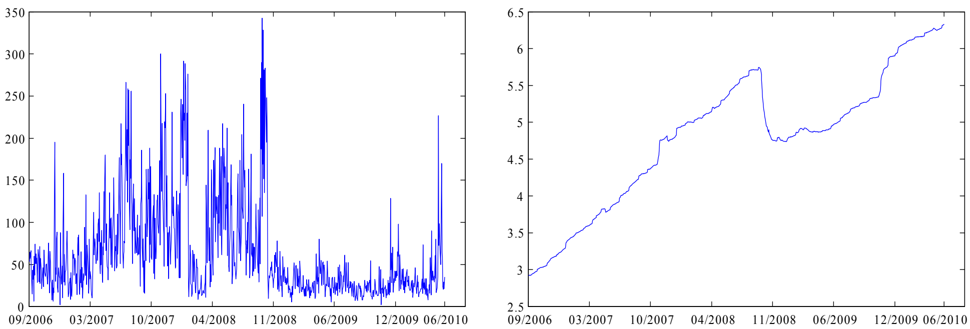
### 3.1 Data

Daily data from 1/9/2006 to 4/6/2010 is obtained from the National Bank of Serbia (NBS) database. We consider the following variables as determinants of the nominal exchange rate: the volume of the interbank currency trades (excluding NBS intervention), the interest rate differential, total household savings in foreign currency and net banks' purchases of foreign currency from the foreign exchange offices (FEO) and households. The choice of variables is based on data availability and the expected role.

There are several theoretical explanations for the impact of the volume of interbank currency trades on exchange rates. The first is the mixture of distribution hypothesis (Clark, 1973) which is based on the assumption that asset prices and volumes are jointly driven by an unobserved stochastic process. If more information arrives to the market in a given time interval, the prices would respond more strongly. That implies the existence of a contemporenous relationship between the exchange rate returns and trading volumes. The second theory is the sequential information arrival model (Copeland, 1976) where new information becomes available to one investor at a time. Due to sequential information flow, current and lagged values of trading volume can convey useful information about future exchange rates. The third theory focuses on the presence of technical (noise) traders in the market (see De Long et al., 1990). Their trading strategies may impact the exchange rates, implying a positive relationship between the volume and the exchange rates. Related to the first two arguments is the concept of order flow and macro news dissemination. Lyons (1995, 2001) and Evans and Lyons (2008) study order flow and discuss how it can be related to information processing in the foreign exchange market. In particular, they find that a significant part of surprises about fundamentals is transmitted to exchange rates via order flow and that this channel is more important in explaining short-run movements in exchange rates than direct modelling of news. Since order flows and market expectation data are unavailable for the period of study, using the volume of the interbank currency trades can also be viewed as a (crude) proxy for information processing in the dinar-euro market.

The importance of other variables follows from the characteristics of the Serbian economy and financial markets. Changes in the interest rate differential can influence the nominal exchange rate primarily via the standard uncovered interest rate parity channel. Given the low level of development of the bond market in Serbia, alternative measures of interest rate differential are required. We use overnight interest rate spread calculated as the difference between Beonia (Serbian equivalent of Eonia) and Eonia as a proxy for the interest rate differential. We have also experimented with the spread between 2-week dinar repo rate and the respective Euribor rate as well as with the spread between the reference rates, but the variables are not found to be statistically significant. Due to the high level of euroization of the economy, a significant part of household savings is euro-denominated and hence may have short-run effects on the foreign exchange market in the presence of low liquidity. McKinnon (1982) showed that higher degree of dolarization may lead to more instability in the nominal exchange rate and stronger depreciation expectations. This in turn may impact depreciation of the actual exchange rate depending on the expectation formation mechansim and the level of liquidity in the market (see for example, Akcay, et al., 1997, in case of Turkey). The net banks' purchases of foreign currency from the FEOs and households should capture the additional pressures on the currency from the non-bank savings and grey economy.

Figure 2: **Volume of interbank trades (EUR million); Household savings (EUR billion):**




In addition to "direct" determinants of supply and demand on the foreign exchange market we also consider whether changes in the risk perception of international investors with respect to Serbian economy have influence on the dinar-euro exchange rates. Since credit default swap (CDS) spreads for Serbia are not available for the entire sample period we use Emerging Market Bond Index (EMBI)