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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

Milan Nedeljković   Branko Urošević

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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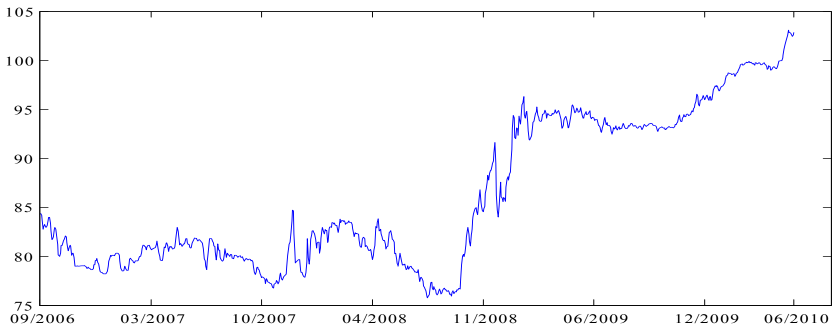
1. Introduction

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:

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

[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 Δe_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 Δe_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 ε_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 σ_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:

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

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: $\overline{g}_i(X_{t,i}) = \sum_{i=1}^{S_i} \overline{\phi}_{i,s} B_{i,s}(X_{i,t}) - n^{-1} \sum_{t=1}^{n} \sum_{i=1}^{S_i} \overline{\phi}_{i,s} B_{i,s}(X_{i,t})$.

The oracle (pseudo) responses are constructed as:

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

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 $\tau(\xi) = n^{-1}\sum_{t=1}^{n}\hat{\varepsilon}_t W(X_t, \xi)$, $W(\cdot)$ is a pre-specified function, $\hat{\varepsilon}_t$ are the residuals from estimation of model (1) and $d\chi(\xi)$ 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

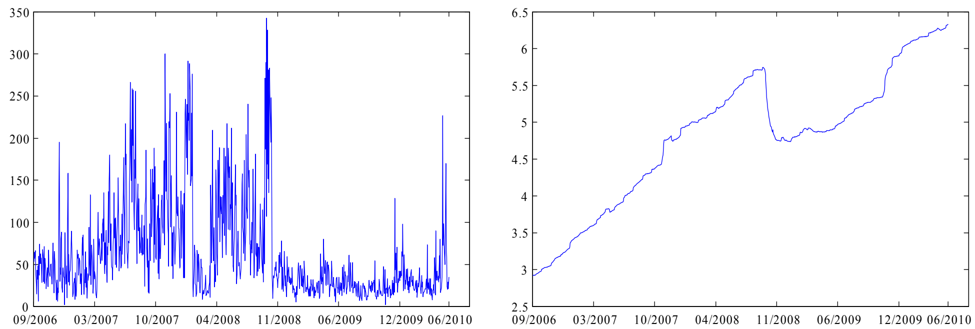
3.1 Data

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) and change in the index as a proxy for changes in perceived country risk.[7] Figures 2-4 present evolution of the selected variables over time.

Figure 3: Net purchases from FEO and household (EUR million); EMBI Index:

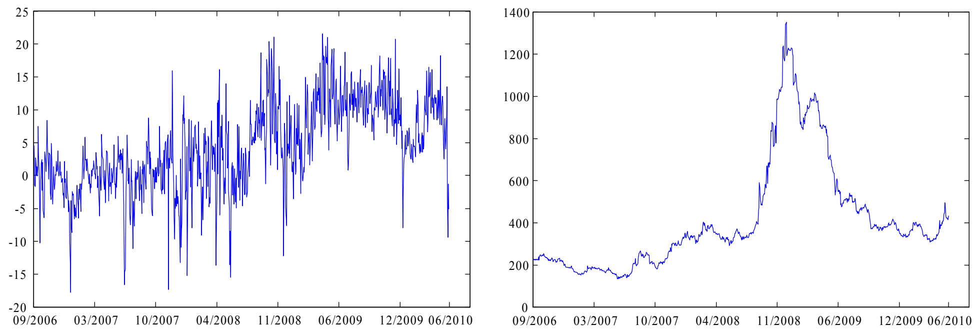
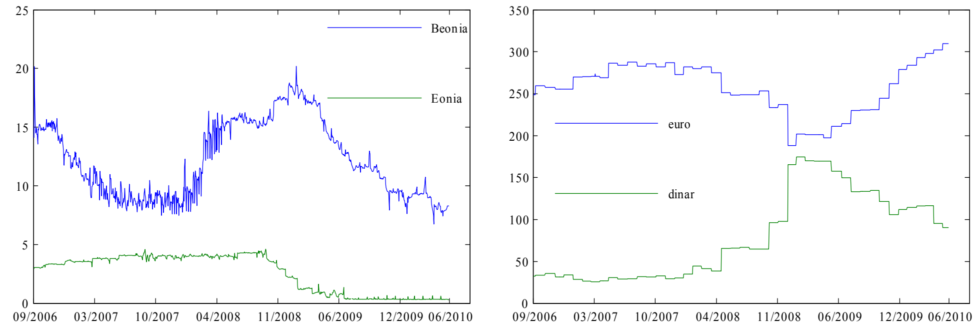


Figure 4: Beonia and Eonia Index; Banks' dinar and euro required reserves (RSD billions):



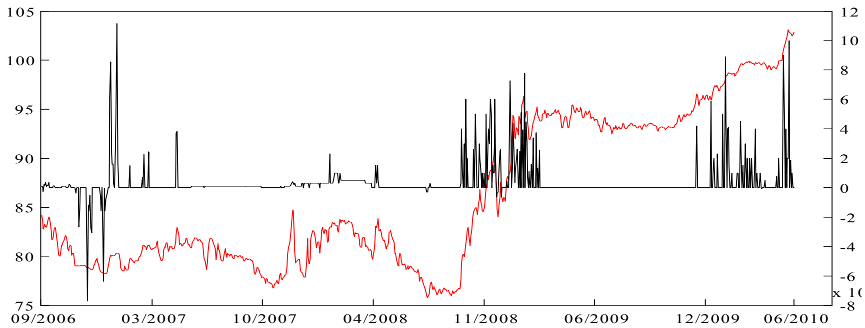
The second set of variables that we consider are the instruments used by the NBS: volume of daily interventions in the foreign exchange market and changes in banks' euro-denominated required reserves.8 A sterilized central bank' intervention can influence the daily exchange rate movements through a variety of channels (see Sarno and Taylor (2003), Chapter 7). In this paper, we are interested in the aggregate effect that interventions have on the exchange rates. In the sample, NBS intervened on 357 days in total (327 days were net purchases of dinar and 30 days were net sales) with the average amount of 11.35 (14.73, respectively) million EUR. Figure 5 shows the interventions and the corresponding exchange rates on the day following the

[7]Note that since EMBI Serbia is based on the yield on the long-run goverment bond with embedded put option, the level of EMBI index might overestimate the true risk premium for Serbia. Changes in EMBI index however should be less affected by the measurement error.

[8]Since the movements in banks' dinar denominated reserves follow closely the movements in euro denominated reserves with an opposite sign we do not include both variables to avoid potential multicolinearity.

intervention. This is in line with a method that is used for determining the official dinar-euro exchange rate. The exchange rate et on day t is defined as the average rate from transactions (including the NBS intervention, if present) on the day t-1. Figure 5 shows that most of the net sales of dinar occurred in the pre-crises period, while the post crisis period is characterized by a larger number of net purchase interventions. Second plot in Figure 4 illustrates changes in the required reserve policy (both variables are dinar-denominated).

Figure 5: The RSD/EUR exchange rate and Intervention (10 million EUR):



3.2 Formulation of the Model

Given a potential structural break in the series at the beginning of the recent crisis which would invalidate assumption of strict stationarity we split the sample and estimate the model separately over the two periods. The first period is from 1/9/2006 to 12/9/2008 and the second from 15/9/2008 to 4/6/2010. The break point corresponds to the collapse of Lehman Brothers. We start by testing for stationarity of variables of interest. As expected, unit root tests[9] imply that nominal exchange rates, EMBI index, overnight interest rate spread, and household savings are first difference stationary, while other variables are found to be weakly stationary. Based on the above discussion, we start with the following model:

$$\Delta e_t = \alpha + \sum_{i=1}^{p_1} g_{1,i}\left(\Delta e_{t-i}\right) + \sum_{i=1}^{p_2} g_{2,i}\left(\Delta i_{t-i}\right) + \sum_{i=1}^{p_3} g_{3,i}\left(vol_{t-i}\right) + \sum_{i=1}^{p_4} g_{4,i}\left(INT_{t-i}\right) + \sum_{i=1}^{p_5} g_{5,i}\left(\Delta EMBI_{t-i}\right)$$
$$+ \sum_{i=1}^{p_6} g_{6,i}\left(\Delta S_{t-i}\right) + \sum_{i=1}^{p_7} g_{7,i}\left(RR_{t-i}\right) + g_8\left(NP_{t-1}\right) + g_9\left(Temp_{t-1}\right) + \sum_{i=1}^{4} \delta_i D_{t,i} + \sigma_t \varepsilon_t$$

where Δi_{t-1} is the change in overnight interest rate spread, volt-1 is the volume of interbank trades, INT_{t-1} is the NBS intervention, $\Delta EMBI_{t-1}$ is the log-difference in perceived risk, ΔS_{t-1} is the log-change in foreign savings, NP_{t-1} is net bank purchases of foreign currency, RR_{t-1} is the required euro reserves and Dt is the day of the week dummy. We also include variable $TEMP_{t-1}$ which is the average temperature during

[9]The results are availble on request.

the day in Belgrade collected from the Republic Hydrometeorological Service of Serbia in order to control for possible seasonal effects in the exchange rate. Note that given the method in which the officially reported nominal exchange rate is determined, the first lag of explanatory variables has a contemporaneous effect on the exchange rate.

The presence of contemporaneous relationships, however, introduces a potential endogeneity bias in some of the variables as the amount of intervention at time *t-1* may depend on the evolution of the exchange rate over the same period whose average trading value is reported as the official rate for time *t*. Likewise, the volume of trades at time *t-1* may be influenced by the movements in the exchange rate over the same period. The behaviour of other explanatory variables is less likely to be influenced by the current movements in the exchange rate. In order to control for endogeneity of intervention we follow the instrumental variable two stage approach (Humpage, 1999, Galati, et al., 2005, Fatum and Pedersen, 2009) and estimate a central bank reaction function in order to capture the expected component of the intervention variable and use predicted value of intervention from this estimation in place of the original intervention values. The truncated characteristic of the intervention variable as the dependent variable introduces additional problem when estimating the parameters of the reaction function, leading to potential sample selection bias in estimation. We follow Humpage (1999) and estimate the parameters through the two step procedure. In the first step the probability of intervention is specified as a function of the deviation of the today's exchange rate from the trend and the exchange rate volatility using the probit model:

$$P\left(C_t = 1 \mid e_t - \breve{e}_t, \sigma_t\right) = \Phi\left(\gamma_1 + \gamma_2\left(e_t - \breve{e}_t\right) + \gamma_3 \sigma_t\right) \tag{7}$$

where C_t is the zero-one indicator variable for the decision to intervene, $\breve{e}_t$ is the trend variable computed as a 10-day moving average of the exchange rate and volatility is approximated using the rolling 10-day standard deviation of the exchange rate. We have also experimented with using 20-day trend and volatility estimates, but the obtained results are qualitatively similar. The variables determining the decision to intervene are those commonly found in empirical studies of the central bank's reaction function and are in line with an intention of the NBS to intervene in order to prevent high daily oscilations in dinar-euro rate.[10] In the second step a simple OLS regression is performed:

$$INT_t = \beta_1 + \sum_{j=1}^{p} \beta_{2,j} INT_{t-j} + \sum_{j=0}^{p-1} \beta_{3,j} \Delta e_{t-j} + \beta_4 M_t \tag{8}$$

[10] Another reason to pursue intervention may be to increase liquidity in the foreign exchange market with a small depth. However, without utilizing high frequency data with the exact timing of the intervention it is not possible to capture the importance of this channel. The values of the previous day volumes are not found to be statistically significant in the probit equation.

where M_t is the Mills ratio from the probit estimation in order to control for the sample selection bias. The equation (8) models the amount of intervention as a function of the past interventions and the previous changes in exchange rate. The fitted values from the equation (8) are then used in place of the observed interventions. Similarly, we control for the endogeneity of the volume series using a two step IV procedure where the lagged volume and the exchange rate changes are used as instruments for the observed volume. Note that since we use a linear specification to control for the endogeneity the first step estimation error is of $n^{-1/2}$ order and, therefore, following Li and Wooldridge (2002), it can be neglected in the second step semiparametric estimation. Table 1 presents the results from two-step estimation of the NBS reaction function using heteroscedasticity robust standard errors. Both variables in probit equation are found to be statistically significant. The volume of intervention, in turn, depends on the past intervention values (the so-called intervention clustering phenomenon) and on the past changes in exchange rate. The R^2 from the OLS regression in both samples is relatively high (for time series estimation) and significantly higher than the values reported in Humpage (1999) and Galati, et al. (2005). As a result, we can be relatively confident that the weak instrument problem is not present.

Table 1: Estimates of the NBS reaction function

Variable	Subsample 1	Subsample 2
Probit equation		
[illegible]	-0.318***	-1.422***
e_t-$\breve{e}_t$	11.781 ***	21.634 ***
δ_t	51.184 ***	81.021 ***
Intervention equation		
β_1	0.138	1.804***
INT_{t-1}	0.525***	0.062
INT_{t-2}	-0.121	0.098*
INT_{t-4}	-0.029	0.137**
INT_{t-5}	0.085*	0.012
Δe_{t-1}	13.749***	72.005***
Δe_{t-2}	6.545	-28.818**
Δe_{t-4}	5.602*	9.819
M_t	-0.035	-0.945***
R^2	0.324	0.328

The specification (1) potentially includes a large number of variables as all lags of the variables are treated as separate variables in semiparametric estimation. In order to obtain a parsimonious representation of the model we employ Huang and Yang (2004) *BIC* criterion to select variables for semiparametric estimation. We follow the (adjusted) three step procedure from Jansen, et al. (2008) where we assume that all continuous variables from specification (1) are present in the semiparametric model for Δe_t:

Step (i): start with one predictor variable and determine the lag order for that variable. Specifically, choose p_{11} to minimize *BIC* for $1\leq p_{11}\leq P$ where P is the pre-specified maximum lag length. Next, choose p_{11} and p_{12} to minimize *BIC* for $1\leq p_{11}$, $p_{12}\leq P$ (here p_{11} may or may not equal the value obtained in the previous search, depending on the dimensionality of the problem). Conduct a sequence of such searches until the addition of one additional lag no longer improves the model fit.

Step (ii): Given the lag order structure for the first variable, determine the lag structure for the second predictor variable using the same procedure as in step (i). Given the lag structures for the first two variables, specify the lag order for the third variable. Finally, proceed with inclusion of other (non-lagged) variables.

Step (iii): Repeat steps (i)-(ii) for different orderings of the predictor variables.

3.3 Results

We start from the specification in (5) and use *BIC* to select the significant variables where we set the maximum lag length $P=5$. We perform the selection procedure separately over the two subsamples. The number of interior knots in spline estimation is fixed to five. The bandwidth parameter in local linear estimation is set as $b_j=c_j std(X_j)n^{-1/5}$ where $std(X_j)$ is the sample standard deviation of the variable X_j and constant c_j is selected via generalised cross-validation with 50 grid points. We used standard quartic kernel in estimation. Since the asymptotic variance of $\hat{g}_j(x_j)$ depends on the unknown quantities which need to be separately (nonparametrically) estimated, the confidence intervals are obtained through boostrap. We use fixed regressor wild boostrap with 1000 bootstrap repetitions in order to control for conditional heteroscedasticity in the data (Yang, 2008). The confidence intervals are obtained as the appropriate percentiles of the boostrap distribution at each evaluation point x_j.

The results from estimation together with the 5% confidence bands are presented in Figures 6-7 and Table 2. On each plot the values of the independent variable are on the *x*-axis and the corresponding value of the effect on exchange rate returns is given on *y*-axis. Table 2 presents estimates of the parameters from the linear part. We only present the results for the variables selected by *BIC*.

In the first subsample characterized by moderate volatility and small medium run appreciation, we find evidence of significant nonlinearity in the autoregressive part of the model. In addition, we find that changes in savings in foreign currency (which almost tripled over the period) and banks' net purchases of foreign currency had significant effect on the exchange rate. The other variables are not found to improve the model fit via the *BIC* criterion (the final value of *BIC* is -10.72). The finding is related to the stability of the exchange rate over the first subsample and characteristics of the Serbian economy. In particular, the significance of the lagged exchange rate returns implies a certain degree of inefficiency of the foreign exchange market (see, e.g., Gradojevic, et al., 2010), using a different statistical approach). The shape of the estimated coefficient function implies a degree of mean-reverting behaviour of the exchange rate where smaller positive changes in the one period lagged exchange rate (depreciation) are associated with the further depreciation of the exchange rate, but the large past depreciations have appreciation effect.

The influence of changes in foreign currency-denominated savings and net bank purchases likely stems from the high level of euroization of the economy and large amount of foreign currency in non-bank channels and grey economy. Increase in foreign currency savings results in increased demand for foreign currency leading to the depreciation pressures in the presence of low market liquidity. Conversely, negative changes in foreign currency savings lead to appreciation pressures, although the trend is reversed for large negative changes which may signal a loss of confidence in the banking system and less demand for home currency.

Figure 6: Plots of the SBKS estimator (solid lines), bootstrap confidence intervals (dashed lines), linear regression estimator (dotted lines):

Figure 6.1: One Period Lagged Change In Exchange rate:

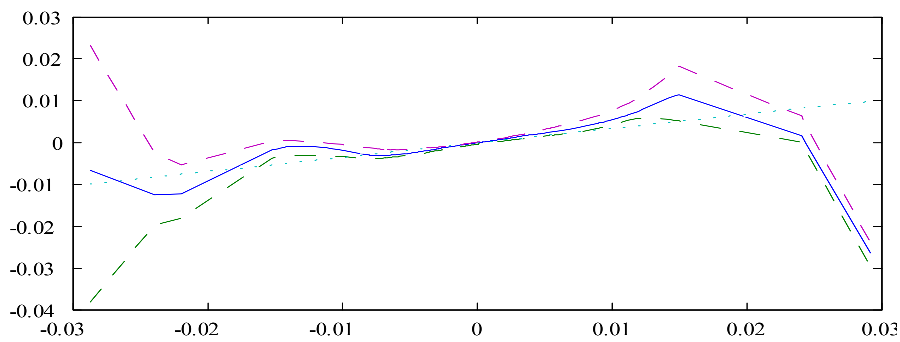


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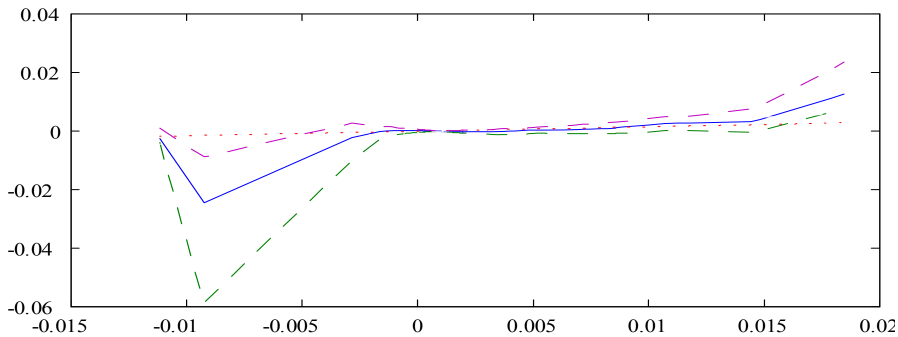
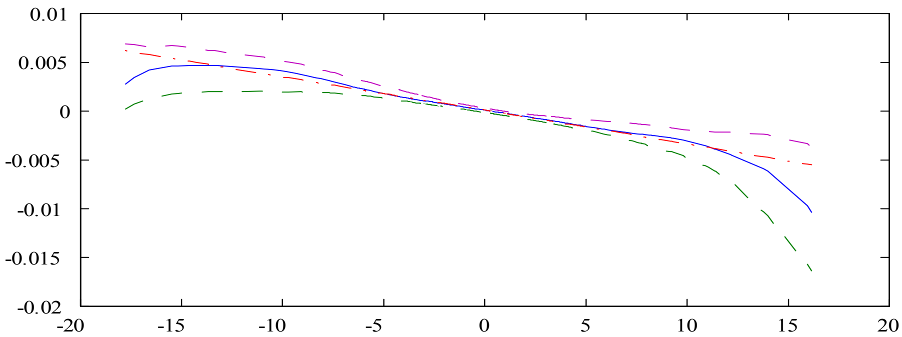
Figure 6.2: Savings in foreign currency:



Analogously, changes in bank's net purchases signal the importance of the unofficial channels in the economy and its effects on the foreign exchange demand. As Figure 6.3 shows the null that the net bank purchases have a linear effect is not rejected and the final specification incorporates this variable in the linear part of the model together with day of the week effects, while semiparametric estimates of the nonlinear part are those presented in Figure 6.1-6.2. The presence of semi-parametric component does not allow computation of R^2 as the standard measure of the fit and to gauge the relative success of the model we compute the correlation between the fitted values and observed values of the exchange rate returns.

Figure 6.3: Net foreign currency purchases by banks:



The correlation between the fitted and observed values when the net bank purchases enter the model linearly is 0.669, implying a relatively good fit of the model.We then test whether additive partially linear model adequately captures nonlinearities in the data by computing the Li, et al. (2003) test. We choose $W(X_t, \xi) = \exp(iX_t^{'}\xi)$ and $d\chi(\xi)$ to be standard normal p-variate vector which yields a simple closed form expression for the underlying integral (Escanciano, 2007). The statistic has value 0.738 and running the 1000 bootstrap replications yields critical value 1.412 and corresponding p-value 0.72. Therefore, we fail to reject the additive partially linear model for the first subsample.

Table 2:
Estimates of the Linear part:

Variable	Subsample 1	Subsample 2
A	-0.0013**	-0.0261**
D_1	0.0018**	0.0007
D_2	0.0020**	0.0004
D_3	0.0014 **	0.0001
D_4	0.0010**	-0.0005
NP_{t-1}	-0.0004**	-0.0002**
$EMBI_{t-1}$	-	0.0179**

Notes: *,**,*** imply significance at the 10%, 5% and 1% level.

The set of influential variables is larger in the second subsample. Results are reported in Figure 7 where we find that in addition to autoregression effects, country risk, banks' net purchases, volume of interbank trades, interest rate differential and NBS intervention have effect on the dinar-euro rate. Conversely, the changes in foreign currency savings and the required reserve are not found to improve the fit (the final value of *BIC* is -10.63). Also, we find no support for the seasonal effects using temperature as a proxy (the length of the subsamples may be too small to capture the seasonal effects). The effect of banks' net purchases is presented in Figure 7.1 and is of smaller magnitude relative to the first subsample although the null of linearity is again not rejected.

Figure 7: Plots of the SBKS estimator (solid lines), bootstrap confidence intervals (dashed lines), linear regression estimator (dotted lines):

Figure 7.1: Net foreign currency purchases by banks:

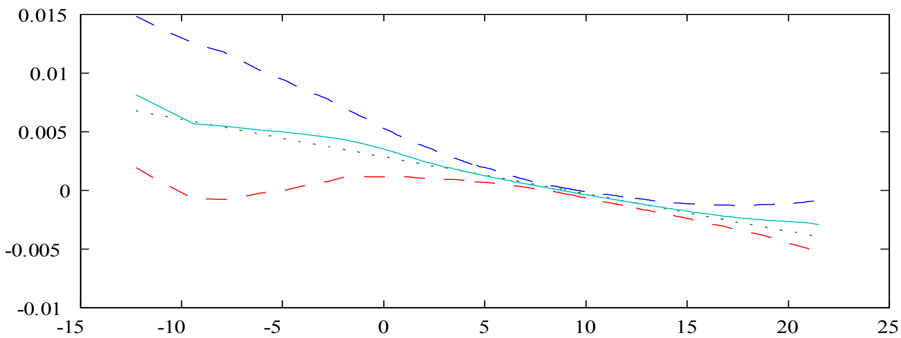


Figure 7.2: EMBI Index:

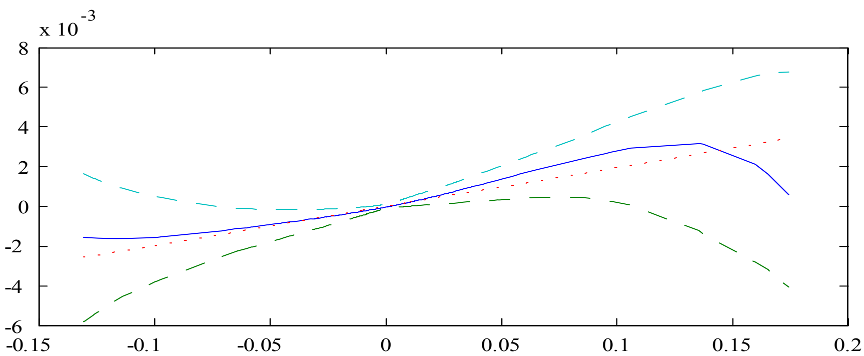


Figure 7.2 plots the estimated coefficient function with respect to changes in the country risk. We can see a small degree of nonlinearity in response of the exchange rate to changes in the EMBI index. The positively slopped linear fit cannot be rejected at the 5% significance level, implying that larger changes in the EMBI index have proportionally larger effect towards depreciation. Given the evidence of linearity in the first two variables the model is re-estimated with inclusion of these variables in

the linear part of the model and the results are reported in Table 2. We do not find statistically significant day of the week effects in the second subsample. As expected, the estimated parameters for country risk and net banks' purchases are significant at the 5% level (boostrapped confidence interval). Estimates of the nonlinear part are given in Figures 7.3-7.6.

Results in Figure 7.3 imply that the information from the contemporaneous interbank trade volumes have effect on the exchange rate movements. Smaller trade volumes have effect towards appreciation, while larger volumes lead to proportionally larger dinar depreciation. Note that foreign exchange market in Serbia is mostly demand driven (supply of foreign currency is relatively stable). Thus, smaller demand for euro results in smaller trade volumes and leads to appreciation of dinar while the opposite holds in case of large demand for euros.

Figure 7.3: Volume in the foreign exchange market:

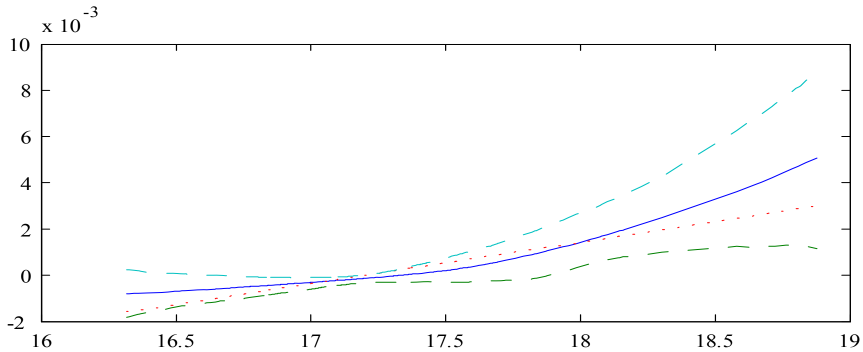


Figure 7.4: One period lagged change in the exchange rate:

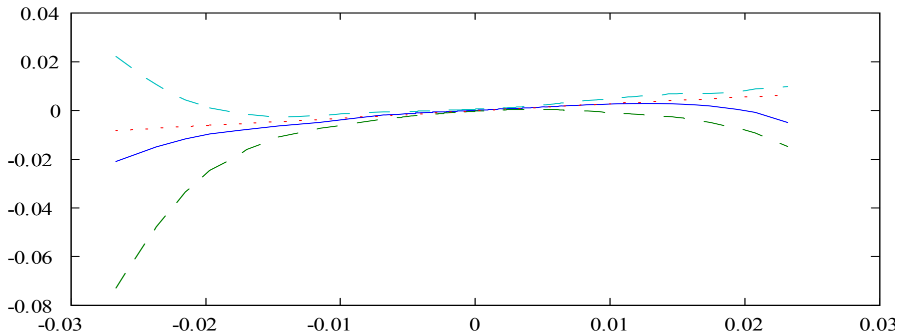


Figure 7.4 plots the estimated autoregression coefficient function. As in the first subsample, it implies the presence of inefficiencies in the foreign exchange market, while the magnitude of the estimated coefficients is smaller relative to the first subsample signalling the importance of other factors in the second period. Figures 7.5 shows that the NBS interventions were effective over this period. The effect is nonlinear and depends on the size of the intervention, where larger sales of euros have stronger appreciation effect on dinar, which is in line with a signalling effect that large sales have over the most recent period. The impact of interventions on the

exchange rate movements comes with a two day lag, while the inclusion of the contemporaneous value of intervention does not improve the fit of the model. This is in line with the evidence on (lagged) effectiveness of interventions in other transition countries (Egert and Komarek, 2006, Disyatat and Galati, 2007).

Figure 7.5: Two period lagged intervention:

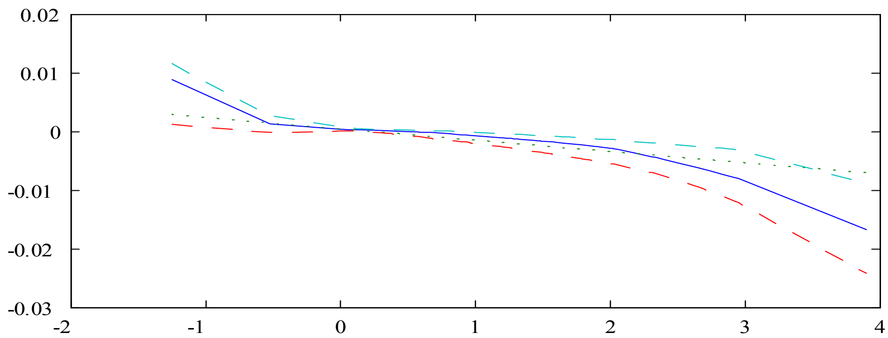
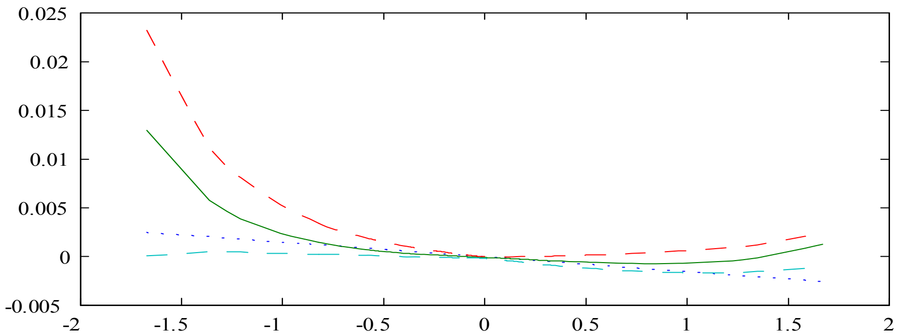


Figure 7.6: Overnight interest rate spread:



Finally, in Figure 7.6 we plot the effect of changes in the overnight interest rate spread. The effect is positive and larger for negative changes in the spread. Negative changes in overnight spread lead to the standard carry-trade type effect on the exchange rate. Depreciation effects from increases in the spread over the period can be associated with the increase in the liquidity risk premium and consequently smaller capital inflows.

To assess the adequacy of the model in the second subsample we again perform Li, et al. (2003) test using 1000 bootstrap replications and obtain *p*-value 0.61. The correlation between the fitted and observed values is 0.586.

4. Conclusions and future research

In this paper we have studied drivers of daily dynamics of the nominal Serbian dinar-euro exchange rate from September 2006 to June 2010 using linear and nonlinear statistical models. We identified several factors influencing daily exchange rate changes. In the period before the financial crisis, in addition to information in past returns, changes in banks' net purchases of foreign currency and household savings have statistically significant explanatory power. From September 2008 onwards other factors gain importance. In particular, changes in country's risk and volume of interbank trades are associated with depreciation of dinar. NBS interventions are found to be effective with a time delay.

There is a lot more work to be done in the present framework. First, due do data limitation, we have not included any macro variables or macro (news) surprises in the model. As discussed in the main text, the former should not influence the results judging from the experience from other markets. The latter, however, might have an effect, given that significant relationship between the high-frequency exchange rate returns and news was found in developed (Andersen and Bollerslev, 1998, Andersen, et al., 2003 among others) and transition countries (Disyatat and Galati, 2007, Frömmel et al., 2011). Since we work with daily data, the effect of news should not be as large as in the high-frequency framework, but this need to be empirically validated. Second, we have only included a proxy for Serbia's country risk. Exchange rates of some of the neighboring countries with free float have experienced a similar behavior as Serbian dinar. For example, Romanian lei depreciated from 3.673 lei per euro on 12/09/2008 to 4.225 on 04/06/2010 with a similar time pattern. Hungarian forint also depreciated in the same period with a rapid 31% fall between September 2008 and March 2009. This implies that broader set of regional risks might have influence on the exchange rate, however we abstract from the inclusion of EMBI indices for other countries given its high correlation with EMBI index for Serbia (correlations between EMBI index for Serbia and those for Bulgaria, Ukraine and Turkey is above 0.9 in both subsamples). Further research is required to capture additional measures of the regional risk. Third, we used a simple linear approach to control for potential endogeneity of intervention and volume variable. A semi-parametric approach to control for endogeneity is an extension that can be of interest beyond the present paper.

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