# Serbian Foreign Exchange Market during 2004-2008

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**Abstract:** To ensure an efficient exchange rate intervention policy by a national bank the exchange market itself has to be efficient. Moreover, the functioning of exchange markets in transition countries is assumed to be highly relevant for the domestic economy. This investigation empirically points out the efficiency of the Serbian foreign exchange market and the intervention policy conducted by the National Bank of Serbia. The negative effect of exchange rate volatilities on the real economy is shown to be statistically valid for this potential EU candidate state.

**Key words:** Volatility, intervention efficiency, market efficiency, real effect

**JEL Code:** C22, E58

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# 1. Introduction

A fundamental role in macroeconomic stabilization in Serbia, as elsewhere in transition economies, has been played by the exchange rate. Due to the ongoing trade balance deficit - which is the "biggest problem and macroeconomic risk" according to the Government of the Republic of Serbia[1] - the relevance of imports and exports for the Serbian economy becomes obvious. They determine the supply and demand of the national/foreign currency and basically drive the exchange rate. Consequently, the determination of the exchange rate level is a matter of economic policy conducted by the Serbian Government.

Serbia is the only country in the Western Balkans with a truly flexible exchange rate, which enables swift and efficient adjustment to external and internal shocks. Since the introduction of a flexible exchange rate regime, the National Bank of Serbia (NBS) influences indirectly the foreign exchange market (forex market) primarily via its interest rates. By way of exception, the NBS intervenes in the forex market to limit daily exchange rate oscillations (explicitly not the exchange rate level) and thus ensures forex market stability. This issue becomes more important in times of global financial crises like the current US-subprime crisis, which substantially increases the exchange rate volatility. Consequently, the NBS equalizes short term exchange rate volatility with instability of the forex market. In principle, exchange rate volatility interpreted as an unstable forex market leads to uncertainty of market participants due to inferior economic planning. Moreover, transaction costs emerge from hedges against exchange rate fluctuations. Both arguments tend to result in investment restraints and therefore implicate negative effects on the economy. So, the NBS's duty is to ensure a stable and functioning forex market in which economic reforms implemented by the government can be efficient.

According to the Serbian Government a faster accession to the European Union (EU) has a strategic priority in 2009 and also leads to the desired exchange rate stability. Further integration into the Economic and Monetary Union requires the fulfillment of the Maasticht criteria, e.g. to ensure a stable exchange rate. Therefore, with or without the EU integration, the forex market stability is highly relevant for the Serbian economy.

Although there are several studies dealing with exchange rate volatilities in transition countries, see for example Kočenda (1998), Orlowski (2003), Kóbor and Székely (2004), Bulíř (2005), Kočenda and Valachy (2006) and Schnabl (2007), the Serbian perspective is empirically not considered, which this paper especially illuminates. One argument for neglecting the Serbian forex market is explicitly named by data availibility problems, see Schnabl (2007). To overcome this problem and to pick up the frequency of the NBS target indicator I use daily exchange rate data.

This empirical analysis raises the following questions: 1. Is the Serbian forex market efficient? 2. Is the intervention policy of the NBS efficient? 3. Is the Serbian real economy negatively affected by forex market volatilities?

[1] See page 4 of the public announcement "The Economic Crisis and its Impact on the Serbian Economy" from the office of the Serbian Prime Minister (December 2008).

To the first question. The EU enlargement process always requires stable economic circumstances. A prerequisite for stable economic circumstances is the existence of an efficient forex market. A market is efficient if historical return information does not improve forecasts of exchange rate returns, i.e. the exchange rate itself should be a random walk. The importance of the Euro is internationally accepted. A large number of national banks in the countries at the periphery of the European Monetary Union are using the Euro as an intervention and reserve currency. Therefore, if the Serbian forex market is efficient, the anchor currency Euro should not cause the Dinar exchange rate returns. This fact would imply a kind of "Euro efficiency". That means, that the market price considers all available information. To test this hypotheses with high frequency daily exchange rate data, a newly developed spectral Granger causality test of Lemmens, Croux and Dekimpe (2008) is used.

To the second question. Obviously the world is facing a heavy crisis, which began in the summer 2007 with the burst of the US-subprime bubble. After a period of more or less stable global demand, the first major bank ‚Bear & Stearns' collapsed at the beginning of 2008 and the crude oil price exploded. In September 2008, Lehman Brothers collapsed and the stock market indices shut down. These circumstances caused high forex market volatility worldwide. Due to almost zero interest rates, the Swiss National Bank conducts an exceptional currency devaluation intervention policy to push the performance of the national economy. To look at the Dinar/Euro exchange rate and judge that the intervention policy of the NBS is inefficient because of a certain volatility level, would not account for the global financial crisis. In order to analyse the comparative stability of daily exchange rates, the Dinar/US-Dollar and Euro/US-Dollar exchange rate will be used. An exchange rate is the price of a currency in another currency and therefore a relative price. Because of the same reference currency (US-Dollar), which is the global intervention and reserve currency so far, the Dinar and Euro exchange rates are comparable. If the NBS intervention policy is efficient, the above exchange rate volatilities (target indicators of the NBS), interpreted as a state of disequilibrium, should share a common trend (cointegration) during a choosen cadence or during periodes of excessive volatility. Finally, the current crisis offers a chance for the NBS to show the efficiency of the institution in terms of its exchange rate intervention policy.

To the third question. The inherent reason for the stabilization of large exchange rate volatilities by the NBS could be the assumption that large fluctuations in the financial sector cause negative effects on the real sector. Using aggregated volatility measures and the industrial production index this assumption can be tested in a time series perspective. If exchange rate volatilities have a statistically negative effect on the output, this fact would be an empirical justification of the NBS policy.

The remainder of the paper is organized as follows. Section 2 discusses modeling of forex market volatilities. Subsequently, Section 3 tests the Serbian "Euro efficiency" hypothesis. Section 4 investigates the efficiency of the NBS intervention policy and Section 5 considers the link between the Serbian exchange market volatilities and the real economy. The paper ends with concluding remarks.

## 2. Modeling of forex market volatilities

Asserting the efficient-market hypothesis (Fama (1970)) of daily exchange rate returns, it would not be possible to outperform the forex market and to achieve systematic returns at a specific date $t = 1,...,T$ . For the logarithmic Dinar/US-Dollar daily middle exchange rate $e_{d,t}$ and the logarithmic Euro/US-Dollar daily middle exchange rate $e_{e,t}$ , the correspondent returns $r_{d,t} = \Delta e_{d,t}$ and $r_{e,t} = \Delta e_{e,t}$ follow. The raw daily middle exchange rates are extracted from Thomson Datastream for the period 1 March 2004 to 31 December 2008 (NBS-Governor Jelašić). The unit root test procedure according to Dickey and Fuller (1979)[2] strongly rejects the null hypothesis of unit roots in the return series and characterize $e_{i,t}$ , $i = d, e$ , as random walks (test statistics 0.489 and 0.119, respectively). Therefore, the stationary error terms $u_{i,t} = \Delta e_{i,t}$ equal the return series $r_{i,t}$ and drive the stochastic trends of the exchange rates $e_{i,t}$ . As far as the test procedure characterizes the exchange rates as random walks, it is not possible to use past return information to outperform the forex market, i.e. the asserted efficient market hypothesis can not be rejected.

As the first step, the returns are modeled by two univariate regressions

$$r_{d,t} = \beta_1 + w_{d,t} \quad \text{and} \quad r_{e,t} = \beta_2 + w_{e,t} \,. \tag{1}$$

$w_d$ and $w_e$ are stationary stochastic processes. The OLS estimation results (Table 1) of these models show two important empirical facts. Firstly, the average Dinar return $\bar{r}_d = \hat{\beta}_1$ and the average Euro return $\bar{r}_e = \hat{\beta}_2$ are not significantly different from zero at any plausible level of significance, based on the standard errors of Newey and West (1987).

Table 1

**OLS estimation results of the univariate return models**

| Variable | Estimate | Standard error | p-value |
|---|---|---|---|
| $\beta_1$ | 0.00012 | 0.00023 | 0.6167 |
| $\beta_2$ | -0.00009 | 0.00018 | 0.6171 |

The column "p-value" indicates the Newey/West p-value of the hypotheses of zero coefficients. The p-values of the ARCH-LM test with 2 lags are for both equations smaller than $10^{-4}$. The asymptotic $x^2$ (2) distributed test statistics are for the first equation 161.24 and for the second equation 194.92.

This confirms the unit root test results. Secondly, the ARCH Lagrange multiplier (LM) test (Engle (1982)) for autoregressive conditional heteroskedasticity of the residuals $\hat{w}_d$ and $\hat{w}_e$ leads to the conclusion of heteroskedasticity. ARCH - which is quite common in financial time series - itself does not invalidate standard least-squares inference, but ignoring them may result in loss of efficiency.

[2] Kourogenis and Pittis (2008) show the validity of the traditional Dickey-Fuller test statistic with barely infinite variances. In terms of Monte Carlo studies the special case of IGARCH error processes confirm the general statement. As shown in the following, IGARCH errors have to be respected in this investigation. Then, the return series do not posses a finite variance, although they are still strictly stationary (Nelson (1990)).

Engle (1982) introduced autoregressive conditional heteroskedasticity models (ARCH) which are specifically designed to model and forecast conditional variances. These models were generalized as GARCH (generalized ARCH) by Bollerslev (1986). As shown above, the equations in (1) have to be modified due to insignificant constants and ARCH effects. These essential modifications and empirical necessities (Akaike information criterion) for the Dinar return series lead to the volatility equation

$$\sigma_{d,t}^2 = \alpha_1 + \alpha_2 r_{d,t-1}^2 + \alpha_3 \sigma_{d,t-1}^2 \tag{2}$$

and for the Euro return series to the volatility equation

$$\sigma_{e,t}^2 = \beta_1 + \beta_2 r_{e,t-1}^2 + \beta_3 r_{e,t-2}^2 + \beta_4 \sigma_{e,t-1}^2 \quad . \tag{3}$$

The mean equations $r_{i,t} = \sigma_{i,t} v_{i,t}$, $i = d,e$, contain innovations $v_{d,t}$ and $v_{e,t}$, which are independent and identically t-distributed processes with $\vartheta_d$ and $\vartheta_e$ degrees of freedom. This distributional assumption is in line with theoretical investigations[3] of financial markets, which postulate power-law distributions - like the t-distribution - of financial time series. The specific lag structures of the variance equations guarantee uncorrelated squared standardized residuals $\hat{v}_{i,t}^2 = (r_{i,t}/\hat{\sigma}_{i,t})^2$, $i = d,e$, according to Box and Ljung (1978) and no more ARCH effects. Therefore, the variance equations seem to be correctly specified. The estimated degrees of freedom of the t-distributions lead to the conclusion of fat-tailed exchange rate returns and is a very sensible result for financial time series. For the Dinar equation the condition $\alpha_2 + \alpha_3 < 1$ and for the Euro equation $\beta_2 + \beta_3 + \beta_4 < 1$ must be satisfied to ensure stationary volatilities. Based on the estimates of Table 2, $\hat{\alpha}_2 + \hat{\alpha}_3 \approx 1$ and $\hat{\beta}_2 + \hat{\beta}_3 + \hat{\beta}_4 \approx 1$ it follows, therefore, that the stationarity conditions are in fact not fulfilled. Hence, it can be supposed that the volatilities follow a stochastic trend and volatility shocks are persistent. Foreign exchange market interventions of central banks, which could be interpreted as shocks, will have persistent influence on the volatility evolution of exchange rates. This issue shows the importance of an appropriate intervention policy of central banks.

Table 2

**ML estimation results of the univariate GARCH return models**

| $\alpha_1$ | $\alpha_2$ | - | $\alpha_3$ | $\vartheta_d$ | $p_d$ |
|---|---|---|---|---|---|
| $7.20 \cdot 10^{-7}$ | $0.0416^{**}$ | - | $0.9470^{**}$ | $6.2537^{**}$ | 0.1775 |
| $(4.50 \cdot 10^{-7})$ | (0.0113) | - | (0.0167) | (1.1966) | - |
| $\beta_1$ | $\beta_2$ | $\beta_3$ | $\beta_4$ | $\vartheta_e$ | $p_e$ |
| $1.15 \cdot 10^{-7}$ | $-0.0531^{**}$ | $0.0945^{**}$ | $0.9557^{**}$ | $19.9961^{*}$ | 0.3856 |
| $(9.68 \cdot 10^{-8})$ | (0.0114) | (0.0139) | (0.0102) | (9.2330) | - |

Standard errors in parenthesis. ** and * indicate the rejection of the hypothesis of zero coefficients on the 99% and the 95% level. $P_d$ and $P_e$ symbolize the p-value of the ARCH-LM test with 2 lags.

[3] See, for example, chapter 9 and 10 of Aoki and Yoshikawa (2007).

To account for the insignificant constants in the volatility equations and the violation of the stationary conditions, integrated GARCH (IGARCH) models are specified according to Bollerslev and Engle (1986). The volatility equations (2) and (3) do not contain constants anymore and the restrictions $\alpha_2 + \alpha_3 = 1$ and $\beta_2 + \beta_3 + \beta_4 = 1$ will be considered. In comparison to Table 2, the estimation results of Table 3 underline the consistency of the ML estimator for GARCH models with nonstationary variances, see Jensen and Rahbek (2004).

Table 3

**ML estimation results of the univariate IGARCH return models**

| $\alpha_2$ | - | $\alpha_3$ | $\vartheta_d$ | $p_d$ |
|---|---|---|---|---|
| 0.0323** | - | 0.9677** | 6.9322** | 0.1946 |
| (0.0064) | - | (0.0064) | (1.1119) | - |
| $\beta_2$ | $\beta_3$ | $\beta_4$ | $\vartheta_e$ | $p_e$ |
| − 0.0525** | 0.0912** | 0.9614** | 10.4051** | 0.6630 |
| (0.0117) | (0.0131) | (0.0078) | (2.7687) | - |

Standard errors in parenthesis. ** indicate the rejection of the hypothesis of zero coefficients on the 99% level. $P_d$ and $P_e$ symbolize the p-value of the ARCH-LM test with 2 lags.

However, the previous models do not account for the anticipated high correlation between the two exchange rates. There should be a high contemporary correlation because the relative prices of the domestic currency are in expressed relation to the US-Dollar. To account for correlations between the exchange rates, a multivariate estimation framework has to be applied. Further extensions of univariate ARCH models to multivariate models, which are usually analagous to the extension from ARMA to vector ARMA models, come up in literature quite often. See, for example Bollerslev, Engle and Granger (1988), Engle and Rodrigues (1989), Kroner and Claessens (1991), among several others. Multivariate ARCH models allow the variances and covariances to depend on the information set in a vector ARMA manner. Therefore, multivariate models consider more information than univariate models, which abandon covariances. Hence, in spirit of a seemingly unrelated regression (Zellner (1962)), multivariate ARCH models could lead to a more efficient estimation. The preceding ideas can also be extended to a multivariate GARCH model, see Engle and Kroner (1995) or Bollerslev (1990).[4]

For the two exchange rate return models $[r_{d,t}, r_{e,t}]' = [u_{1,t}, u_{2,t}]'$ with the bivariate t-distributed error vector $u_t = [u_{1,t}, u_{2,t}]'$ conditioned on an available information set $\psi_{t-1}$, it is possible to specify the variance covariance matrix

$$\Sigma_t = \Omega + A \bullet u_{t-1}u_{t-1}' + B \bullet \Sigma_{t-1} \quad . \tag{4}$$

The coefficient matrices $A$, $B$ and $\Omega$ are $2 \times 2$ symmetric matrices, and the operator "•" is the element by element Hadamard product. The most general way is to allow the parameters in the matrices

[4] Although, the ML consistency for multivariate GARCH models with nonstationary variances is - as far as I know - not verified, the univariate consistency statement of Jensen and Rahbek (2004) is usually adopted in a multivariate framework, see for example Boswijk and Zu (2007).

to vary without any restrictions, i.e. parameterize them like here as indefinite matrices. In that case the model may be written in single equation form as:

$$\sigma_{ij,t} = \omega_{ij} + \alpha_{ij} u_{i,t-1} u_{j,t-1}' + \beta_{ij} \sigma_{ij,t-1} \tag{5}$$

where the index $i$ symbolizes the $i$-th row and $j$ the $j$-th column of the symmetric matrices mentioned above.

Table 4

**ML estimation results of the bivariate GARCH return model**

| Diagonal elements of $\Omega$, $A$ and $B$ | | | | | |
|---|---|---|---|---|---|
| $\omega_{11}$ | $\omega_{22}$ | $\alpha_{11}$ | $\alpha_{22}$ | $\beta_{11}$ | $\beta_{22}$ |
| $2.24 \cdot 10^{-6**}$ | $1.20 \cdot 10^{-6**}$ | $0.1159^{**}$ | $0.0886^{**}$ | $0.8675^{**}$ | $0.9028^{**}$ |
| $(5.12 \cdot 10^{-7})$ | $(3.72 \cdot 10^{-7})$ | (0.0171) | (0.0163) | (0.0142) | (0.0162) |
| Remaining coefficients | | | | | |
| $\omega_{12}$ | $\alpha_{12}$ | $\beta_{12}$ | $\vartheta$ | | |
| $1.41 \cdot 10^{-6**}$ | $0.0955^{**}$ | $0.8921^{**}$ | $4.2384^{**}$ | | |
| $(3.40 \cdot 10^{-7})$ | (0.0145) | (0.0131) | (0.4602) | | |

Standard errors in parenthesis. **indicate the rejection of the hypothesis of zero coefficients on the 99% level. $\vartheta$ symbolizes the degree of freedom of the t-distribution.

Table 4 shows the estimation results of the multivariate GARCH(1;1) model, which exhibits the best empirical fit. The system portmanteau tests for autocorrelations of the standardized residuals show at any lag autocorrelations using square root of covariance (Urzua (1997)).

Like in the previous univariate case, the bivariate system could also contain nonstationary variances. The eigenvalues of $A \otimes A + B \otimes B$ have to be less than one in absolute value to ensure covariance stationarity of the system ("$\otimes$" indicates the Kronecker product). Based on the coefficient estimates the eigenvector $[-0.0115, -0.0111, 0.0001, 3.1980]'$ gives advice for nonstationary variances, again. Therefore, shocks of the exchange rates seem to be persistent. Although, the calculated Dinar volatility $\hat{\sigma}_{d,t}^2$ and the calculated Euro volatility $\hat{\sigma}_{e,t}^2$ from the bivariate GARCH model are not integrated of the order 1 ($I(1)$) in a conventional sense, these volatilities are nonstationary in the GARCH framework.

## 3. Forex market efficiency

Section 2 shows empirically that historical information about individual exchange rate returns does not lead to better forecasts of exchange rate returns. In that sense, the markets are efficient. In contrast, "Euro efficiency" is defined as follows:

**Definition 1**

*The Serbian forex market is "Euro efficient' if the European exchange rate does not affect the Serbian exchange rate neither in a long-run nor in a short-run perspective.*

Investigating causality is a topic of great interest in scientific research, especially in economics. Granger (1969) introduced an econometric concept, which was modified in several ways. In this context, I use a Granger causality test over the spectrum (Lemmens et al. (2008)). In contradiction to the traditional methodology of a one shot Granger causality, the new developed procedure allows for a causality test between two processes at different frequencies. Concerning exchange rates it is possible to test a short-, medium- or long-run causality of the Euro/US-Dollar exchange rate for the Dinar/US-Dollar exchange rate. If the European rate leads the Serbian rate, the Serbian forex market would be de facto inefficient, because the market would not assimilate all available information.

Let $r_{d,t}$ be the stationary Serbian exchange growth rate series and $r_{e,t}$ be the stationary European exchange growth rate series of section 2, $t = 1,...,T$ . $v_t$ and $u_t$ are the univariate innovations series derived from $r_{d,t}$ and $r_{e,t}$ . These innovations have to be white-noise processes and are possibly correlated with each other at different leads and lags. In order to obtain the innovations, usually ARMA models are used as a filter. Concidering the specific properties of daily financial data, the univariate IGARCH models of Table 3 are used as a filter. Therefore, $u_t$ and $v_t$ represent the standardized innovations of the univariate return series. The spectra of $u_t$ and $v_t$ at Fourier frequencies $\lambda_j = j/T$ , $0 \leq j \leq T/2$ , are defined by

$$S_u(\lambda_j) = \sum_{-\infty}^{\infty} \gamma_u(k) e^{-i\lambda_j k} \quad \text{and} \quad S_v(\lambda_j) = \sum_{-\infty}^{\infty} \gamma_v(k) e^{-i\lambda_j k} \tag{6}$$

$\gamma_u(k) = Cov(u_t, u_{t-k})$ and $\gamma_v(k) = Cov(v_t, v_{t-k})$ represent the autocovariances of $u_t$ und $v_t$ at lag $k$ . The correspondent cross-spectrum $S_{uv}(\lambda_j)$ between $u_t$ and $v_t$ at lag $k$ is

$$S_{uv}(\lambda_j) = \sum_{-\infty}^{\infty} \gamma_{uv}(k) e^{-i\lambda_j k} \tag{7}$$

with the cross-covariances $\gamma_{uv}(k) = Cov(u_t, v_{t-k})$ . Based on the cross-spectrum and the single spectra it is possible to define the coefficient of coherence

$$h_{uv}(\lambda_j) = \left| S_{uv}(\lambda_j) \right| \Big/ \sqrt{S_u(\lambda_j) S_v(\lambda_j)} \tag{8}$$

which gives a measure of strength of linear association between $r_{e,t}$ and $r_{d,t}$ , frequency by frequency ( $h_{uv} \in [0;1]$ ). Unfortunately, it does not provide any information on the direction of the relationship between the two processes. In order to calulate estimations of the above theoretical concepts, I use for the autocovariances, $k = 0,1,2,...$ ,

$$\hat{\gamma}_u(k) = \sum_{t=1}^{T-k} (\hat{u}_t - \bar{\hat{u}})(\hat{u}_{t+k} - \bar{\hat{u}}) \Big/ T \quad , \quad \hat{\gamma}_v(k) = \sum_{t=1}^{T-k} (\hat{v}_t - \bar{\hat{v}})(\hat{v}_{t+k} - \bar{\hat{v}}) \Big/ T, \tag{9}$$

for the lags, $k = 0,-1,-2,...$ , and for the leads, $k = 0,1,2,...$ ,

$$\hat{\gamma}_{uv}(k) = \sum_{t=1}^{T+k} (\hat{v}_t - \bar{\hat{v}})(\hat{u}_{t-k} - \bar{\hat{u}}) \Big/ T \quad , \quad \hat{\gamma}_{uv}(k) = \sum_{t=1}^{T-k} (\hat{v}_{t+k} - \bar{\hat{v}})(\hat{u}_{t-k} - \bar{\hat{u}}) \Big/ T, \tag{10}$$

With respect to the spectra the estimates are calculated by

$$\hat{S}_u(\lambda_j) = \sum_{k=-M}^{M} w_k \hat{\gamma}_u(k) \cos 2\pi\lambda_j k \quad , \quad \hat{S}_v(\lambda_j) = \sum_{k=-M}^{M} w_k \hat{\gamma}_v(k) \cos 2\pi\lambda_j k \tag{11}$$

and

$$\hat{S}_{uv}(\lambda_j) = \sum_{k=-M}^{M} w_k \hat{\gamma}_{uv}(k) \cos 2\pi\lambda_j k \quad , \tag{12}$$

with $M = \sqrt{T}$ and the Parzen weights (truncation point = 36)

$$w_k = \begin{cases} 1-6\,|k/T|^2 + 6\,|k/T|^3 & , \quad |k| \le T/2 \\ 2(1-|k/T|)^3 & , \quad T/2 < |k| \le T \\ 0 & , \quad else. \end{cases} \tag{13}$$

The null hypothesis $h_{uv}(\lambda_j) = 0$ against $h_{uv}(\lambda_j) > 0$ is rejected if

$$\hat{h}_{uv}(\lambda_j) > \sqrt{\chi^2_{2,1-\alpha} \Big/ 2(n-1)} \quad , \; n = T \Big/ \left( \sum_{-M}^{M} w_k^2 \right), \tag{14}$$

and $\chi^2_{2,1-\alpha}$ being the $1-\alpha$ fractile of the chi-squared distribution with 2 degrees of freedom. If one is interested in the direction of the relationship between $r_{e,t}$ and $r_{d,t}$, a simple modification of the coefficient of coherence is necessary. Instead of $\hat{S}_{uv}(\lambda_j)$ the causality can be captured by

$$\hat{S}_{u \Rightarrow v}(\lambda_j) = \sum_{k=-M}^{-1} w_k \hat{\gamma}_{uv}(k) \cos 2\pi\lambda_j k \tag{15}$$

and the estimated Granger coefficient of coherence is given by

$$\hat{h}_{u \Rightarrow v}(\lambda_j) = |\,\hat{S}_{u \Rightarrow v}(\lambda_j)\,| \Big/ \sqrt{\hat{S}_u(\lambda_j)\hat{S}_v(\lambda_j)}. \tag{16}$$

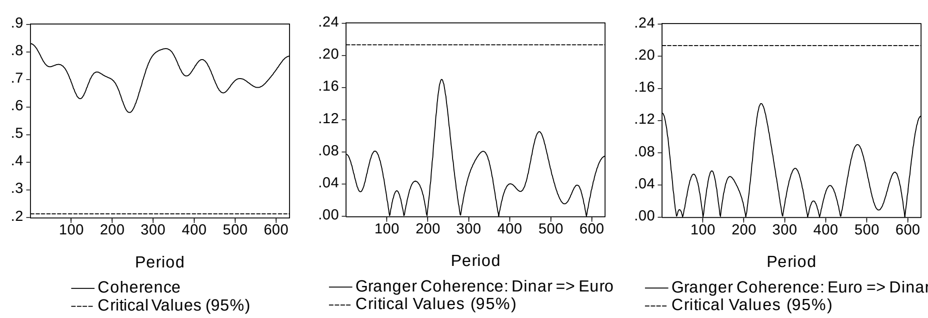
The null hypothesis of no Granger causality – which would implicate market efficiency at frequency $\lambda_j$ - versus $\hat{h}_{u \Rightarrow v}(\lambda_j) > 0$ is rejected if

$$\hat{h}_{uv}(\lambda_j) > \sqrt{\chi^2_{2,1-\alpha} \Big/ 2(n'-1)} \quad , \; n' = T \Big/ \left( \sum_{-M}^{-1} w_k^2 \right). \tag{17}$$

The empirical application of the concepts (programmed in EViews 6) are visualized in Chart 1. The test concerning the coefficient of coherence rejects the hypothesis of no linear association between the two exchange rates at any frequence. This supports the results of Section 4 and leads to the conclusion of a high linear association between the short-, medium- and long-run exchange rates at a specific frequency. Additionally, it is not possible to reject the hypothesis of an "Euro efficient" Serbian forex market due to Granger coefficients of coherence which lie below the critical bound.

Chart 1

**Estimated coefficient of coherence and estimated Granger coefficients of coherence**



## 4. Intervention efficiency of the NBS

Each year, the NBS publishes a monetary program in advance to inform the general public. For the year 2005 this program contains the following paragraph:

*"The National Bank of Serbia shall further liberalize the foreign exchange market through deregulation of the said market, with the aim of increasing its competitiveness, so that the flow of foreign exchange is directed towards commercial channels and direct interventions of the National Bank of Serbia in this market are scaled down."*

and for the year 2007 the paragraph:

*"The exchange rate of the Dinar will be formed freely, with reference to supply and demand in the foreign exchange market. (...) National Bank of Serbia may intervene in the foreign exchange market: (1) to limit daily oscillations (...)"*

It follows from the above two paragraphs that the NBS changed its exchange rate policy from a rather managed to a free floating regime sometime in 2006. Therefore, the Serbian exchange market should be stable in comparison to the most important local European exchange markets until this policy breakpoint, i.e. the computed nonstationary volatilities $\hat{\sigma}_{d,t}^2$ and $\hat{\sigma}_{e,t}^2$ should follow a common trend, so that $\sigma_{d,t}^2$ and $\sigma_{e,t}^2$ are cointegrated in a GARCH sense. Consequently, the following definition holds:

**Definition 2**

*The Serbian forex market is stable in comparison to the European exchange market, if the nonstationary Serbian forex market volatility* $\sigma_{d,t}^2$ *and the nonstationary European exchange market volatility* $\sigma_{e,t}^2$ *share a common trend.*

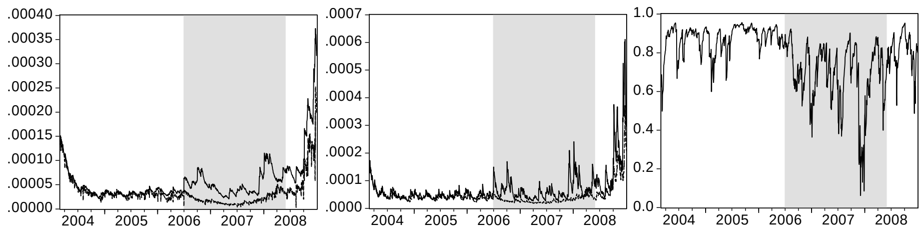
After the policy breakpoint in 2006, the existence of a cointegration relation between the volatility series is not expected, due to the freely formed Dinar exchange rate. The more interesting question is how the NBS exchange rate policy works in times of trouble, such as the current financial market crisis. This crisis bares the chance for the NBS to show how stable this institution is. According to its monetary program and the supposed importance of the exchange market for the domestic economy, the NBS should intervene actively to limit daily exchange rate oscillations and return to a stable common volatility trend with the European exchange rate, which is supposed to be the local anchor currency. This would show that the NBS is able to drive the exchange market efficiently even in hard times. As mentioned in the introduction, this return to a cointegration relation should take place in 2008.

Unfortunately, the contemporary state of the art econometrics offers not a volatility cointegration test in a GARCH sense. Prevailing literature broach the issue of cointegration with nonstationary volatility, but not a test procedure for common volatility trends, see for example Boswijk and Zu (2007) or Cavaliere, Rahbek and Taylor (2008). Due to the methodological deficiencies - which shows potential for further econometric research - the conventional trace and maximum eigenvalue statistics of Johansen (1988, 1991) are the best available test procedures and provide an informative basis for the current problem.

Nevertheless, inspecting the line graphs of the computed volatilities in Chart 2, the impression of a common trend during two periods prevails.

Chart 2

**Development of forex market volatilities and correlations**




Calculated exchange market volatilities $\hat{\sigma}^2_{d,t}$ (more volatile line, left) and $\hat{\sigma}^2_{e,t}$ for the IGARCH model and the bivariate GARCH model (middle), calculated exchange market correlations of the bivariate GARCH model.

From March 2004 to June 2006 ( $T_1 = 589$ ) the volatility series seem to share a common stochastic trend without a deterministic trend in the data. The shaded period from July 2006 to May 2008 ( $T_2 = 500$ ) seems to be conspicuous in terms of a more asynchronous evolution of the volatilities. Additionally, the conditional correlations between the exchange rate returns (Chart 2) decrease. During the remaining period of this sample (June 2008 to December 2008, $T_3 = 153$ ), both volatility series seem to contain a linear deterministic trend and a common stochastic trend. To apply the Johansen test, it is necessary to consider a VAR(p) model $x_t = \Phi_1 x_{t-1} + ... + \Phi_p x_{t-p} + \varepsilon_t$, which can be reformulated as a VEC model

$$\Delta x_t = \Pi x_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta x_{t-i} + \varepsilon_t \quad . \tag{18}$$

Here $x_t = [\hat{\sigma}^2_{d,t}, \hat{\sigma}^2_{e,t}]'$, $\varepsilon_t = [\varepsilon_{d,t}, \varepsilon_{e,t}]'$, $\Gamma_i = -(\Phi_{i+1} + ... + \Phi_p)$ with $i = 1,..., p-1$, $\Pi = (\Phi_i + ... + \Phi_p) - I_2$ and the $2 \times 2$ coefficient matrices $\Phi_i$, $i = 1,..., p$ hold. As long as $\varepsilon_t$ is not white noise, the Johansen procedure claims admittance of lagged $x_t$. In the conditional variance framework, the additional lags will not cure the heteroskedasticity issue. While the data generating process of the conditional volatilities satisfies a specific AR representation, system autocorrelation tests are affected by the autoregressive description of heteroskedasticity. Hence, the determination of the maximal lag p based on serial correlation system tests of the residuals would be quite problematic. If instead we use the correlogram of the computed volatilities, the AIC of the VEC model and consider one month observations, the maximal lag p will be defined by 20 (days).

Table 5 documents the test results of the trace and eigenvalue procedure for the computed IGARCH and bivariate GARCH volatilities and for the different time periods mentioned above. Barring from period $T_3$ the test procedures assume neither a deterministic trend in the data nor a constant in the cointegrating relation. In case of $T_3$ the deterministic trend assumption of the data is linear and the cointegrating relation of the IGARCH volatilities contains a constant. For the bivariate GARCH volatilities, the cointegrating relation contains no constant.

Table 5

**Trace and maximum eigenvalue cointegration tests**

| | Hypothesized Cointegration Rank | Eigenvalue | Trace Statistic | Maximum Eigenvalue Statistic |
|---|---|---|---|---|
| **IGARCH** | | | | |
| $T_1$ | 0 | 0.017902 | 14.80146 | 10.63988 |
| | | | (0.0188) | (0.0632) |
| | 1 | 0.007041 | 4.161578 | 4.161578 |
| | | | (0.0491) | (0.0491) |
| $T_2$ | 0 | 0.015716 | 8.061715 | 7.920551 |
| | | | (0.2319) | (0.1795) |
| | 1 | 0.000282 | 0.141163 | 0.141163 |
| | | | (0.7565) | (0.7565) |
| $T_3$ | 0 | 0.081668 | 13.20984 | 13.03501 |
| | | | (0.1073) | (0.0775) |
| | 1 | 0.001142 | 0.174825 | 0.174825 |
| | | | (0.6759) | (0.6759) |
| **Bivariate GARCH** | | | | |
| $T_1$ | 0 | 0.023899 | 16.38752 | 14.24715 |
| | | | (0.0099) | (0.0143) |
| | 1 | 0.003627 | 2.140371 | 2.140371 |
| | | | (0.1692) | (0.1692) |
| $T_2$ | 0 | 0.016273 | 8.265634 | 8.203564 |
| | | | (0.2168) | (0.1618) |
| | 1 | 0.000124 | 0.062070 | 0.062070 |
| | | | (0.8381) | (0.8381) |
| $T_3$ | 0 | 0.123086 | 20.14488 | 20.09594 |
| | | | (0.0092) | (0.0054) |
| | 1 | 0.000320 | 0.048941 | 0.048941 |
| | | | (0.8249) | (0.8249) |

p-values according to MacKinnon, Haug and Michaelis (1999) in parenthesis.

The visual inspection of the volatility evolutions and the associated impression of a cointegrating relation between the calculated volatilities is basically confirmed by the trace and eigenvalue tests. The whole sample can be subdevided into three periods and lead to the conclusion of volatility breakpoints. It is noteworthy that the Serbian exchange rate volatility returns to a common trend with the European exchange rate volatility during a difficult time of the global US-subprime crisis. Against the background of frequent foreign exchange market interventions of the National Bank of Serbia in $T_3$, the monetary policy goal of stabilization of daily exchange rate volatilities is achieved from a comparative point of view. Hence, the hypothesis of an efficient exchange rate policy of the NBS can not be rejected.

In general, it is interesting to note that the NBS reacts - in terms of market interventions - to uncertainty inside an important domestic market, i.e. the forex market. On the other hand, the Federal Reserve Bank (FED) reacts - in terms of interest rate adjustments - to uncertainty inside another domestic market, i.e. the stock market (see Jovanović and Zimmermann (2008)). So, regardless of the degree of development a country has reached, it seems to be an empirical fact that national banks try to moderate uncertainty inside their important markets.

## 5. Real effect of forex market volatilnost

Fundamentally, prices of currencies are driven by macroeconomic events, but developments in the Serbian forex market can also have profound effects on the aggregate Serbian economy. A very good example for an important macroeconomic event is the current global financial crisis started in the US, which affects the evolution of the Serbian forex market volatility investigated in the previous section. Due to the supposed fundamental impact of the exchange rate channel for the aggregate Serbian economy, the NBS monetary policy pursues the limitation of extreme exchange rate volatilities. Hence, the NBS objective is not the exchange rate level but the shocks of the exchange rate volatility, which are probably expected to influence the aggregate economy in a negative sense. Although this is a quite plausible logic in terms of economic theory, the question of a statistical significance arises.

To handle the abstract assumed relationship that shocks of the Serbian forex market influence the Serbian real economy negatively, one requires a shock and an aggregate output definition. Concerning the first definition, stationary forex market volatility shocks are computed by $\Delta\hat{\sigma}_t^2 := \hat{\sigma}_t^2 - \hat{\sigma}_{t-1}^2$ with a monthly time index t. In order to calculate monthly exchange rate volatilities, the computed nonstationary daily Dinar exchange rate volatilities $\hat{\sigma}_{d,t}^2$ of the IGARCH model mentioned above are aggregated in terms of arithmetic means. These monthly volatilities are consequently nonstationary and symbolized by $\hat{\sigma}_t^2$.[5]

Using the monthly industrial production index (2007 = 100) for Serbia from January 2005 to December 2008 (Source: Statistical Office of the Republic of Serbia) as a proxy for aggregate output, the logarithm of the seasonal adjusted time series is symbolized by $y_t$. Comparing $y_t$ with the logarithm of the seasonal adjusted monthly industrial production index (2002 = 100) for the US (Source: Database Feri) in Chart 3, a spillover of the US financial crisis in terms of a decrease in the Serbian industrial production time series during 2008 becomes obvious.

[5] The following results are also robust, if one uses the volatilities of the bivariate GARCH model or first calculate $\Delta\hat{\sigma}_t^2$ with a daily time index t and computes in a second step monthly volatility shocks.

Chart 3

**Monthly seasonal adjusted Serbian and US industrial production index (logarithm)**


Source: Statistical Office of the Republic of Serbia, Database Feri

The utilization of logarithmic monthly industrial production indices to measure aggregate output is quite common in the New Keynesian literature, see for example Clarida, Galí and Gertler (1998). Usually the deterministic part of the stochastic process $Y_t$ of the industrial production is characterized as trendstationary around a linear or quadratic trend in the long run. But facing the data availibility problem on account of the young history of the Republic of Serbia and the global financial crisis which caused the current decline in the aggregate output, it seems more adequate to specify a reciprocal time trend to describe $y_t$ . This reciprocal trend copes with the present saturation chracteristic of the Serbian (and US) industrial production index. Assuming this deterministic trend structure, the forex market shocks $\Delta\hat{\sigma}_t^2$ should have a statistically significant influence on the Serbian industrial production $y_t$ . Therefore, the following regression models are specified:

$$y_t = \alpha_1 + \alpha_2 \cdot t + u_{1,t} \tag{19}$$

$$y_t = \beta_1 + \beta_2 \cdot t^{-1} + u_{2,t} \tag{20}$$

$$y_t = \gamma_1 + \gamma_2 \cdot t^{-1} + \gamma_3 \Delta\hat{\sigma}_t^2 + u_{3,t} \tag{21}$$

$u_{i,t}$ , $i = 1,2,3$ , stand for the specific error terms. The estimation results in Table 6 and the illustration in Chart 4 lead to the conclusion of an adequate specification of $y_t$ in terms of equation (21) during the period observed. Several statements can be derived from the description of $y_t$ . Firstly, exchange rate shocks have a statistically significant negative effect on the industrial production. This statistical fact can be viewed as an empirical justification of the NBS monetary policy concerning the stabilization of exchange rate volatilities. Secondly, an exchange rate shock in period t influences the aggregate output only in t, because of the absence of autoregressive elements. Thus, a specific volatility shock has no persistent negative effect on the real economy. Thirdly, on a saturation level, the industrial production can be systematically described by

$$y_t \approx \gamma_1 + \gamma_3 \Delta\hat{\sigma}_t^2 \tag{22}$$

for large t. Beside the saturation level $\gamma_1$ the exchange rate shocks $\Delta\hat{\sigma}_t^2$ determine the level of industrial production. Therefore, a stabilization of these shocks by the NBS would support economic stability in terms of the variability of $y_t$ .

Table 6

**OLS estimation results of the industrial production models**

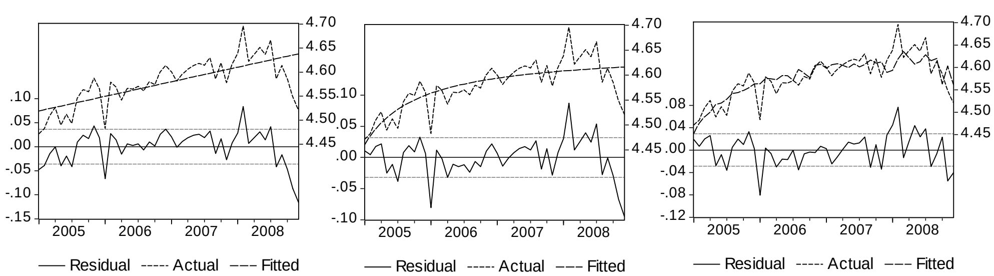
| $\alpha_1$ | $\alpha_2$ | $\beta_1$ | $\beta_2$ | $\gamma_1$ | $\gamma_2$ | $\gamma_3$ |
|---|---|---|---|---|---|---|
| $4.494^{**}$ | $0.0025^{**}$ | $4.649^{**}$ | $-1.884^{**}$ | $4.658^{**}$ | $-2.026^{**}$ | $-637.850^{**}$ |
| (0.021) | (0.0007) | (0.015) | (0.259) | (0.010) | (0.208) | (182.724) |

$T = 48$ observations (January 2005 - December 2008). Standard errors according to Newey/ West in parenthesis. $^{**}$ indicates the rejection of the hypothesis of zero coefficients on the 99% level. $\overline{R}_1^2 = 0.485$ of equation (19) and $\overline{R}_2^2 = 0.598$ of equation (20). Concerning equation (21): $\overline{R}_3^2 = 0.669$ , p-value of the Breusch/Godfrey test (lag 2) = $0.4499$ , p-value of the White test = $0.8685$ , p-value of the Ramsey RESET test (1 fitted term) = $0.1047$ , p-value of the Jarque/Bera test = $0.5123$ .

In times of economic crisis with excessive market uncertainty an all-too liberal exchange rate floating hardly affects the real sector. Therefrom, the policy implication is obvious: The NBS should maintain its exchange rate policy, which proved to be efficient in the cointegration analysis.

Chart 4

**Fit of the linear, reciprocal and reciprocal plus volatility regression model**



## 5. Conclusion

This paper empirically investigates the Serbian forex market and its impact on the real economy. The Dinar/US-Dollar and the Euro/US-Dollar exchange rates are modeled by univariate and bivariate volatility models. Based on this specifications, daily exchange rate volatilities are calculated.

A cointegration relation between the exchange rate volatilities can not be rejected for the period March 2004 to June 2006 and June 2008 to December 2008. Until June 2006, the NBS had a rather free floating exchange rate policy with frequent interventions in the forex market. It reached this self-imposed goal in terms of a common trend of the Dinar in comparison to the Euro. Since June 2008, in the current global turbulence, the NBS has stabilized the Serbian daily exchange rate via forex market interventions and has implemented an efficient exchange rate policy even in hard times. The period in-between is characterized by the self-imposed goal of a free floating exchange rate without frequent interventions. Hence, the exchange rate volatilities do not share a common trend during this period.

Based on the exchange rate modeling and a Granger causality test over the spectrum, the hypothesis of an efficient Serbian forex market can not be rejected. Neither short- nor long-run causal effects of the Euro for the Dinar exchange rate are significant. All available information about the Euro exchange rates are shown vis-a-vis the Dinar.

The statement that exchange rate stability has a positive effect on the real economy, in particular for the emerging European economies, i.e., the central, eastern and south-eastern European countries, holds true especially in case of Serbia. Due to the specification of the industrial production index as a reciprocal trendstationary process, the negative effect of exchange rate volatilities on the real economy is statistically significant.

Summing up the findings and the raised research questions, it may be concluded that the NBS exchange rate intervention policy is empirically justified due to the negative effect of exchange rate fluctuations on the real economy. Furthermore, the NBS intervention policy is efficient, even in times of global crisis and leads us to conclude that the NBS is a stable institution. Its policy can be effective because of an efficient forex market in Serbia.

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