UDC: 339.743;339.738
JEL classification: C32, C82, F02, F31
Keywords: exchange rate volatility, TARCH, openness, news, regime, EMU integration

# Determinants of Exchange-Rate Volatility: The Case of the New EU Members

Juraj STANČÍK – CERGE-EI* Prague (juraj.stancik@cerge-ei.cz)**

*Abstract*

*Exchange-rate stability is not only a criterion for joining the Economic Monetary Union (EMU) but also a fundamental property of stable economic development. At present, new members of the European Union are trying to achieve this stability. However, there are several factors that could slow or interrupt these countries' EMU-integration process. For this reason, this paper analyzes key factors contributing to euro exchange-rate volatility in the new EU members: economic openness, the "news" factor, and the exchange-rate regime. A TARCH (threshold autoregressive conditional heteroskedasticity) model is employed to model the volatility of exchange rates. Although this paper focuses on each country separately, in general the results suggest that economic openness has a calming effect on exchange-rate volatility, news significantly affects volatility, and flexible regimes experience higher degrees of volatility. The extent of all these effects varies substantially across country, however.*

## 1. Introduction

One of the reasons for establishing the Economic and Monetary Union (EMU) was to promote exchange rate stability among member countries and to encourage trade inside the European Union (EU). Otherwise, exchange rate instability could have a negative impact on investment and trade. In the case of sudden movements of an exchange rate, domestic risk-averse companies could turn their focus on the domestic market rather than on the foreign one because the amount of their revenue would become unclear (Dell'Ariccia, 1999). In fact, this exactly opposes the aim of the EU.

As a result of EU enlargement, ten new countries joined the EU in May 2004. The process of their accession further continues as they prepare to join the EMU probably around 2009–2012. By that time, these countries will have to fulfill the Maastricht criteria. This research focuses on factors that can jeopardize the process of fulfilling the second of these criteria – the Exchange Rate Mechanism (ERM) criterion which defines the exchange rate of the participating currency against the euro. The currency can fluctuate around the central rate by ±15 %.

Thus, the goal of this paper is to analyze the sources of euro exchange rate volatility for five central and eastern European countries (CEEC-5) that acceded to

* a joint workplace of the Center for Economic Research and Graduate Education, Charles University, and the Economics Institute of the Academy of Sciences of the Czech Republic

** I would like to thank Evžen Kočenda for valuable consultations on this research, as well as Byeongju Jeong, Martin Čihák and two anonymous referees for useful comments. I am grateful also to Lawrence Smith and Eric Stanglin for editing the manuscript. This work was supported by the CERGE-EI/ /World Bank Fellowship and partial support from GAČR grant No. 402/06/1293 is gratefully acknowledged.

the EU in May 2004[1]. As possible sources, I am interested in the openness of an economy, the "news" factor, and the exchange rate regime due to their undisputed contribution to exchange rate movements. Since these countries are trying to fulfill exacting criteria imposed by the EU, including stable exchange rates, it is necessary and beneficial to know the source of their possible failure.[2]

Although there are already several studies dealing with the volatility of exchange rates in transition countries (Kočenda, 1998); (Orlowski, 2003); (Kóbor, Székely, 2004); (Bulíř, 2005); and (Kočenda, Valachy, 2006), the contribution of this study over the previous projects lies in investigating not only the volatility itself but also its determinants and their casual effects. Moreover, I employ the TARCH model for modeling the volatility of exchange rates because it allows for an asymmetric, i.e., more realistic, impact of news on exchange rate volatility.

In general, the results are consistent with natural expectations. They suggest that the openness has a lowering effect on exchange rate volatility in the case of Poland, Slovakia, and Slovenia. Furthermore, a less tight regime corresponds to higher volatility in the case of Hungary, Poland, and Slovenia. A significant effect of news on exchange rate volatility is found in all cases. However, the extent of all these effects varies substantially across countries.

The paper is structured as follows: The second section deals with previous studies relevant for this research. The methodology is explained in the third section. The fourth section comprises data description and the fifth one presents the empirical results. The last section concludes.

## 2. Literature Review

There is a variety of factors contributing to the fluctuation of an exchange rate, e.g., the openness of an economy, the domestic and foreign money supplies, the exchange rate regime, interest rates, central bank independence, levels of output, income, inflation, and unpredictable circumstances. The degree of the impact of each of these factors varies and depends on a particular country's economic condition. However, the countries that are in the process of transition (CEEC-5 group) are more vulnerable to being affected by these factors. Although this paper analyzes just a few of these factors, according to the empirical literature mentioned below they should be the ones with the biggest impact. In the following section, I explain my incentives for choosing particular factors as well as their validity.

### 2.1 Openness of an Economy

As was discussed above, one of the reasons for establishing the EMU was to promote the greater openness of economies and higher exchange rate stability among EU countries. However, you cannot achieve one without achieving the other. Thus, it is likely that there is a close link between these two factors.

[1] These are the Czech Republic, Hungary, Poland, Slovakia, and Slovenia. I do not include Malta and Cyprus here since these two countries are not in the process of transition and they are considered to be functioning market economies. Moreover, Estonian kroon, Latvian lats, and Lithuanian litas are firmly linked to the euro, and therefore, Estonia, Latvia, and Lithuania are not included in this research either.

[2] Although Slovenia has been a member of the EMU since January 1, 2007, it is included in this research for the sake of consistency.

One of the studies dealing with these factors was elaborated by Hau (2002). In particular, the author analyzes the openness of an economy and its impact on real exchange rate movements. He claims that trade integration and real exchange rate volatility are structurally linked and that there is a negative correlation between them. As support, he uses a small open economy model with a tradable and a non-tradable sector. The solution of this model indicates that economies which are more open have a more flexible aggregate price level. This flexibility reduces the effect of unanticipated money supply shocks. It further results in lower real exchange rate volatility for countries with greater openness of the economy. Hau further supports his claim with empirical research with a sample of 48 countries over a 19-year time period. As a proxy for openness he uses an import vs. GDP ratio. Real exchange rate volatility is measured as the standard deviation for the percentage changes of the effective real exchange rate over intervals of 36 months. The results confirm the impact of an economy's openness on exchange rate volatility when openness explains almost half of exchange rate variations. However, Hau's results do not say anything about particular countries because each country is represented only by arithmetic mean values over the whole period.

### 2.2 Unpredictable Circumstances

The next factor analyzed in this paper concerns unpredictable circumstances or news. This affects all real variables as well as asset yields. In stock markets simple information, often not even valid, might cause huge movements of stock prices. The behavior of exchange rates is very similar, and the consequences of events like government crises, market crises, industrial shocks and terrorist attacks are undisputed. The role of news as the predominant cause of exchange rate movements has already been emphasized in studies by Dornbusch (1978) and Frenkel (1981).

The latter one, by Frenkel, studies the volatility of the US exchange rates between GBP, FFR and DEM.[3] The first part of his paper concerns exchange rate movements and their predictability where he claims that the predicted changes in exchange rates capture only a small fraction of actual changes. Therefore, since most changes in exchange rates are unanticipated, most of them happen due to some new information. Frenkel supports this with an eight-year period of monthly data of the US/GBP, US/FFR, and US/DEM exchange rate movements. In the second part of his study, Frenkel, seeking a suitable instrumental variable for modeling news, discusses the relationship between exchange and interest rates. Although macroeconomic theory explains the negative impact of interest rates on exchange rates *via* capital/current accounts, he claims – based on empirical results from US data – that it does not hold in an inflationary environment, and the impact is actually positive.[4] Furthermore, in line with the rational expectations hypothesis indicating the predominant role of news in affecting real variables and Dornbusch (1978), who decomposes the news effects into "those which alter the expected future spot rate between the last period and the present, and those which lead to a reassessment of

[3] GBP = Great Britain Pound, FFR = French Franc, DEM = Deutsche Mark

[4] However, based on the empirical results from Frenkel's study, the macroeconomic theory holds in this case because the effect is negative for all three exchange rates, although it does not differ significantly from zero.

the one-period interest rate differential,"[5] Frenkel proposes a model for estimating the effect of news on exchange rate variability:

$$\ln S_t = a + b \ln F_{t-1} + \alpha\left[(i-i^*)_t - E_{t-1}(i-i^*)_t\right] + \omega_t$$

where $S_t$ is the spot rate, $F_{t-1}$ is the lagged forward exchange rate, $i$ is the interest rate in the domestic currency, $i^*$ is the interest rate in the foreign currency, and $E_{t-1}(.)$ represents the interest differential expected at time $t$ based on information available at time $t-1$. The first two components on the right-hand side represent the expected exchange rate and the term in brackets represents news. According to Frenkel, applying this regression on all three pairs of currencies (separately) indicates a positive correlation between news and exchange rates. However, there is a weakness in these results. Frenkel uses ln $S_t$ as a dependent variable but he does not consider that this time series is most likely non-stationary. He also does not take into account asymmetric effects of positive and negative news. Moreover, monthly data, used by Frenkel, cannot capture the moment of surprise caused by some new information. Therefore, I expect to obtain better and more significant results using higher frequency data.

The effect of news is discussed also in a study by Galati and Ho (2003) who investigate to what extent daily movements in the euro/dollar exchange rate are driven by news. Finding again a statistically significant correlation between them, good news results in the appreciation of currency, and vice versa. For modeling news they use a similar approach to Frenkel – the difference between the actual and forecasted values – although they measure it on various macroeconomic indicators.[6] The exchange rate movements are captured by the differences in values of the logarithm of the spot prices. Additionally, although Galati and Ho focus also on studying asymmetric behavior of an exchange rate with respect to good or bad news, they do not find any significant asymmetry. On the other hand, Engle and Ng (1993) claim that there is an asymmetric effect of news on volatility and suggest various modifications of the ARCH model [7] for emulating exchange rate volatility. For example, the EGARCH model allows different impacts of good and bad news, as well as major and minor news. In the spirit of this asymmetry, Sanchez-Fung (2003) studies daily returns, volatility, and news in the foreign exchange market of the Dominican Republic, concluding that impact on the volatility of exchange rate returns is higher for positive shocks (depreciations) than for negative ones (appreciations).

### 2.3 Exchange Rate Regime

The last but equally important factor is the exchange rate regime. It is a well-known fact that nominal exchange rate variability is lower in the case of fixed exchange rates than for floating ones. For my research, examples of countries that adopted a floating exchange rate are Slovakia, Poland, and the Czech Republic, while Slovenia and Hungary prefer variations of a pegged exchange rate.

[5] (Frenkel, 1981, p. 686)

[6] Change in non-farm payrolls, the unemployment rate, the employment cost index, durable goods orders, NAPM manufacturing, NAPM non-manufacturing, advance retail sales, industrial production, the consumer price index, and the producer price index.

[7] The autoregressive conditional heteroskedasticity model (ARCH) was introduced by Engle (1982). Later, this model was generalized (GARCH) by Bollerslev (1986).

## 2.4 Exchange Rate Volatility

There are further studies concerning exchange rate volatility, although mostly they investigate the impact of exchange rate volatility rather than sources of this volatility. However, among other things (such as the subject of study), they differ in the way of modeling exchange rate volatility. According to this modeling, they can be divided into two groups – the ones that use various modifications of standard deviations and the ones that use modifications of the ARCH approach.

Belke and Setzer (2003) belong to the first group. They study the impact of exchange rate volatility on the labor market. In their case, the exchange rate volatility is measured as the standard deviation of the 12 month-to-month changes in the logarithm of the spot rate. Dell'Ariccia (1999) studies the effect of exchange rate volatility on bilateral trade flows. He uses the standard deviation of the first difference of the logarithmic exchange rate as well, but he also employs two other measures – the sum of the squares of the forward errors and the percentage difference between the maximum and minimum nominal spot rate. Moreover, there are studies, such as (Kenen, Rodrik, 1986), (Koray, Lastrapes, 1989), (Chowdhury, 1993), (Kóbor, Székely, 2004), and (Bulíř, 2005) that model the exchange rate volatility as the moving sample standard deviation of the growth rate of the real exchange rate.

On the other hand, Baum et al. (2004), analyzing the impact of exchange rate volatility on the volume of bilateral exports, and Choudhry (2005), investigating the influence of exchange rate volatility on real exports, apply the GARCH model for measuring volatility. Further modification of the ARCH approach can be found in Orlowski (2003, 2004). Here, for modeling exchange rate volatility the TARCH model is employed. Kočenda and Valachy (2006), analyzing recent developments in exchange rate volatility in the Visegrad Group countries,[8] suggest usage of the leverage GARCH model.[9] Moreover, exchange rate volatility and the TARCH model are analyzed also in studies by Kočenda (1998) and McKenzie (2002). Although the TARCH approach is mostly employed in papers analyzing stock price movements, Kočenda (1998) claims that with regard to risk there is almost no difference between holding foreign exchange and equity. For this reason, he stresses the justification of using the TARCH approach also for modeling exchange rate volatility.

Thus, this latter approach – ARCH – is plausible also for this research because its modification allows for an asymmetric, i.e., more realistic, impact of news on exchange rate volatility.

## 2.5 Previous Literature about Exchange Rate Volatility

The issue of the stability of the exchange rate in new EU member countries preparing for EMU accession is well researched. The following paragraphs describe some of the most important papers focusing on this topic.

Kočenda (1998) studies the exchange rate of the Czech Koruna against six major currencies. He finds the somewhat surprising result that the Czech exchange rate is less volatile with a wider fluctuation band. Orlowski (2003) examines the impact of monetary policies on exchange rate risk premiums and inflation in the Czech

[8] the Czech Republic, Hungary, Poland, and Slovakia.

[9] The leverage GARCH model is in fact the TARCH model.

Republic, Poland, and Hungary. He concludes that the governments of these countries succeeded mainly in lowering inflation rather than exchange rate volatility. Orlowski (2004) then continues in his research by analyzing the effect of interest rates and inflation on exchange rate movement, which is defined as the differential of the log of the spot exchange rate. He claims that the Czech exchange rate is more affected by inflation expectations, while the opposite is true for Poland, where the impact of the interest rate differential is more pronounced. Regarding Hungary, its currency is affected by both of those factors. A further analysis of exchange rate volatility reveals that in the Czech Republic and Poland it is driven mainly by the persistency effect. In Hungary it is dominated by asymmetric shocks. All three of these papers allow for asymmetric effects of shocks on volatility but only within the TARCH model through a leverage term.

The following three papers are similar in focusing on the same objects – the Visegrad Group countries. Kóbor and Székely (2004) study volatility using a Markov regime-switching model which allows them to identify periods of highly and lowly volatile exchange rates. Not surprisingly, their results say that volatility between these periods changes and is lower in lowly volatile periods. They also claim that there are substantial differences in volatility among the four countries. Bulíř (2005) looks at the relationship between exchange rate volatility and financial market liberalization and concludes that liberalization significantly contributes to the stability of the exchange rates in all four countries. Finally, Kočenda and Valachy (2006) compare exchange rate volatility between fixed and floating regimes. Their findings confirm natural expectations that volatility increases under a less tight, i.e., floating, regime. Moreover, they augment the TARCH model by inclusion of an interest rate differential and its intertemporal change in order to account for their impact on volatility. They claim asymmetric decreasing effects of news on exchange rate volatility, as well as contemporaneous impact of the interest differential.

The previous literature dealing with exchange rate volatility is quite comprehensive. Nevertheless, the previous studies concentrate mainly on volatility itself leaving the question of its determinants unresolved. Although there are some attempts to solve this problem, they are mostly implicit. Literature focusing primarily and explicitly on determinants of exchange rate volatility is still missing.

## 3. Methodology

The paper analyzes the volatility of the exchange rate between the euro and the domestic currency for five different countries. The crux of this project lies in properly choosing the way to model the analyzed factors, especially the openness of an economy and news and, more importantly, to approximate an otherwise unobservable volatility.

### 3.1 Factors

Starting with independent variables, for modeling the openness of an economy, I follow Hau (2002) and use a proxy defined as the ratio of quarterly imports and quarterly gross domestic product. However, in order to observe the effects of openness on exchange rate volatility caused only by structural changes in openness and not by business cycles noise, the Hodrick-Prescott filter[10] is applied to quarterly openness time series. Then, since I need a daily frequency time series, the resulting time series is extended so that it comprises only four different values for every year, and the same smooth ratio of quarterly imports and quarterly GDP is assigned to each day in a particular quarter.

Since it is difficult to observe and quantify unpredictable circumstances or news, I build on the specification proposed by Frenkel (1981), who, knowing the fact that asset markets clear fast and react immediately to news, creates a new variable

$$NEWS_t = (i - i^*)_t - E_{t-1}(i - i^*)_t \tag{1}$$

where $i$ is the interest rate in the home currency and $i^*$ is the interest rate in the foreign currency; the first term in this difference denotes the innovation in the interest differential and the second one denotes the interest differential which was expected to prevail in period $t$ based on the information available at $t$–1. Partially following Frenkel, the latter term is estimated from a regression of the interest differential on the constant, two-lagged values of the differential and the logarithm of the lagged spot exchange rate. Frenkel is followed only *partially* because originally he suggests using the forward exchange rate instead of the spot exchange rate. However, since forward exchange rate markets are either not developed sufficiently or do not have a long history in the CEEC-5 group, the spot rate is used instead. I justify this modification using Frenkel's own argumentation when he claims the correlation between the forward and spot exchange rate to be more than 0.99 in the case of his data. This is also the case for CEEC-5 where the correlation for available periods is more than 0.98 for every country.

For modeling different exchange rate regimes, I create a set of dummy variables for different regimes.

### 3.2 Measure of Volatility

Regarding the dependent variable, i.e., the volatility of exchange rates, I employ the threshold autoregressive conditional heteroskedasticity (TARCH) model. This model comprises a leverage term that allows for the asymmetric effects of good and bad news. The general *TARCH(p, q)* model is specified as:

$$r_t = a_0 + \sum_{i=1}^{P} a_i r_{t-i} + \sum_{i=0}^{Q} b_i \varepsilon_{t-i};\ \varepsilon_t \sim N\left(0, \sigma_t^2\right)$$

$$\sigma_t^2 = \omega + \sum_{i=1}^{p} \alpha_i \varepsilon_{t-i}^2 + \sum_{i=1}^{q} \beta_i \sigma_{t-i}^2 + \xi d_{t-1} \varepsilon_{t-1}^2$$

where variable $r_t$ is the exchange rate change over two consecutive trading days, and $\sigma_t^2$ is the conditional variance that is a function of not only the previous realizations of $\varepsilon_t$, but also the previous conditional variances and the leverage term. The core of this leverage term is the dummy variable $d_{t-1}$ that equals 1 in the case of a negative shock ($\varepsilon_{t-1} \leq 0$) and 0 in the case of a positive shock ($\varepsilon_{t-1} > 0$). Thus, the positive value of the coefficient $\xi$ indicates an increased conditional variance by $\varepsilon_{t-1}^2$ in the case of

[10] This is a smoothing method used by macroeconomists to obtain a smooth estimate of the long-term trend component of a series, first used by Hodrick and Prescott (1997). The penalty parameter is set to 1600 since the filter is applied to quarterly data.

negative shocks or news that occur at time $t$–1, while the negative value of coefficient $\xi$ indicates a decreased conditional variance. The additional restriction $\sum_{i=1}^{p} \alpha_i + \sum_{i=1}^{q} \beta_i < 1$ is a sufficient and necessary condition for stability of the conditional variance.

In order to get a properly specified model and correctly conditioned volatility, the most appropriate *ARMA(P,Q)* model of the exchange rate return is estimated using the *Box-Jenkins* methodology[11]. Then the *Ljung-Box Q-test*[12] is applied to test squared residuals of the *ARMA(P,Q)* model for the presence of conditional heteroskedasticity. The next step is to identify the orders of the *TARCH(p,q)* process by experimenting with different orders $p$ and $q$; estimating the whole *ARMA(P,Q)*-*TARCH(p,q)* model; checking the significance of the estimated coefficients; and then diagnosing the standardized residuals. Once the presence of conditional heteroskedasticity is detected and the orders $p$ and $q$ of the TARCH process are chosen, the whole *ARMA(P,Q)-TARCH(p,q)* model is estimated using the maximum likelihood estimation where the log-likelihood function has the form

$$L = -\frac{1}{2T}\sum_{t=1}^{T} \log \sigma_t^2 + \varepsilon_t^2/\sigma_t^2$$

Finally, the standardized residuals are diagnosed by applying the *Ljung-Box Q-test* and the *LM test for the presence of an ARCH process.*[13] If the estimated model is a correct one, then these residuals should be white noise[14] and no further GARCH process should be present.

### 3.3 Model for Estimation of the Effects

Having estimated all the necessary variables, I perform the actual analysis of the impact of various factors by estimating the following model using OLS:

$$ERV_t = \alpha + \beta\left(\frac{Im_t}{GDP_t}\right) + \gamma_G G_t * NEWS_t + \gamma_B B_t * NEWS_t + \delta REGIME_t + \varepsilon_t \qquad (2)$$

where $ERV_t$ denotes exchange rate volatility estimated in the previous TARCH model, $G_t$ is a dummy variable that equals 1 in case of good news ($NEWS_t < 0$), $B_t$ is a dummy variable that equals 1 in case of bad news ($NEWS_t > 0$) and $REGIME_t$ denotes a set of dummy variables for exchange rate regimes. This process is repeated for each of the five countries in order to determine the different impacts on each particular currency. The interpretation of coefficient $\beta$ is straightforward, a positive value of $\beta$ results in increased volatility in the case of increased openness. Similarly, a positive

[11] (Box, Jenkins, 1976)
[12] (Ljung, Box ,1978)
[13] (Engle, 1982)
[14] The latest literature suggests an even stronger condition. The standardized residuals should be tested for being *iid* because there might be hidden nonlinear patterns that are not detected if a white noise test is applied. For this purpose, one can apply the BDS test developed by Brock et al. (1987) or, as an alternative, Kočenda's test, which was devised by Kočenda (2001). However, these two tests cannot be applied here, since they require the standardized residuals to have a normal distribution and the residuals estimated in this study do not have such a property.

TABLE 1 Openness after Smoothing – Data Summary

| Country | Min | Max | Mean | St. dev. |
|---|---|---|---|---|
| Czech Republic | 0.5706 | 0.6345 | 0.6153 | 0.0193 |
| Hungary | 0.5687 | 0.6241 | 0.6059 | 0.0155 |
| Poland | 0.2882 | 0.3538 | 0.3112 | 0.0207 |
| Slovakia | 0.6093 | 0.7380 | 0.6903 | 0.0402 |
| Slovenia | 0.4886 | 0.5229 | 0.5060 | 0.0098 |

value of coefficient $\delta$ results in increased volatility in the presence of a particular regime with respect to a base regime. However, the manner of constructing the $NEWS_t$ variable requires an opposite interpretation – if it is assumed that the interest rate in the foreign currency does not change due to news in the home country and good news in the home country leads to a decrease in the home currency interest rate, then good news results in a negative value of the $NEWS_t$ variable. Thus, a negative value of coefficient $\gamma$ results in increased volatility in the case of good news.

## 4. Data

All the data used in this project were collected with daily frequency during the period of January 1, 1999–December 31, 2004 from several sources – IMF-IFS, Eurostat, national banks, and central statistical offices. Although there is another strand of literature that studies intraday volatility[15] using high frequency data with a several-minute-long time interval (Černý, Koblas, 2005), for the purpose of this project it is sufficient to use daily data. Using lower frequency data, e.g., monthly or quarterly, might result in the failure of this study because it would often not be able to capture any effects of news on exchange rate movements.

The extent of openness differs substantially in the CEEC-5 group. While the value of Polish imports corresponds on average to 31 % of GDP, in the case of Slovenia it is about 50 %. The economies of the Czech Republic and Hungary are on average even more open (61 %), but the biggest share of imports over GDP can be found in Slovakia (almost 70 %). *Figure 1* displays the openness path in each country before (dotted line) and after (solid line) smoothing. All patterns exhibit an increasing trend, except Hungary with its decreasing trend since the fourth quarter of 2001. Basic characteristics are summarized in *Table 1*.

The factor of news is modeled from a particular country's interest rates (IBORs) with maturity of three months,[16] the Central European Bank's interest rates (EURIBOR) with the same maturity, and the spot exchange rates against the euro.

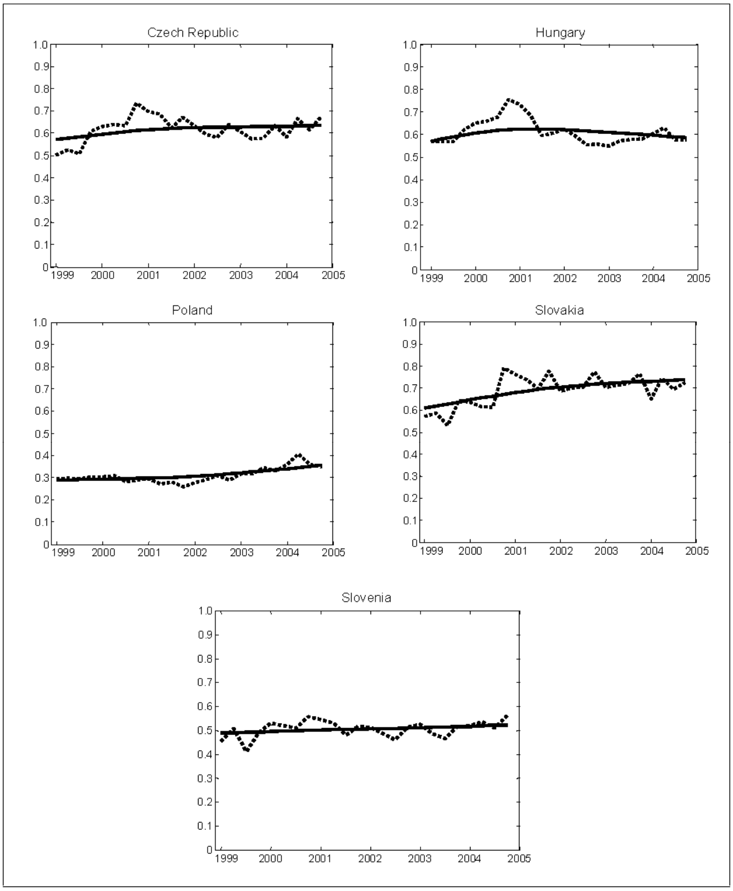
An overview of adopted official exchange rate regimes in each country can be found in *Table 2*. Since Slovakia and the Czech Republic did not change their exchange rate regime during the whole time span, the regime factor is not analyzed as a source of exchange rate volatility in their case.[17]

Finally, the basic characteristics of the nominal exchange rates of each CEEC-5 country's currency vis-à-vis the euro are summarized in *Table 3*. Corresponding ex-

[15] Typically concerning stock market indices.

[16] PRIBOR for the Czech Republic, BUBOR for Hungary, WIBOR for Poland, and BRIBOR for Slovakia. In the case of Slovenia, interbank money market rates for deposits up to 30 days are used instead due to the lack of SITIBOR data.

FIGURE 1 Openness in the CEEC-5 Group During 1999–2004

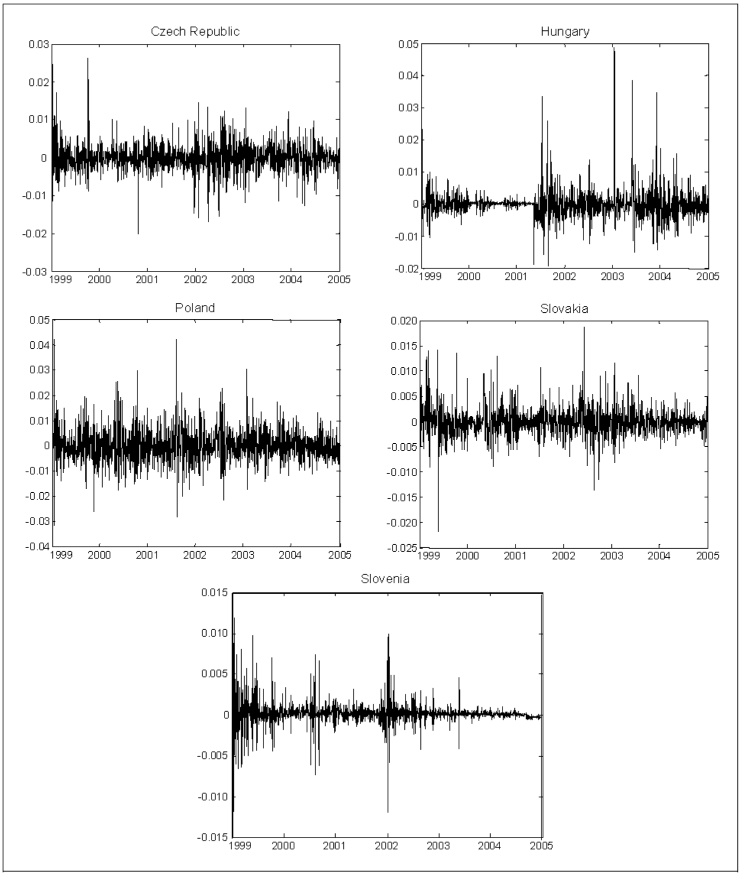



*Source*: IMF-IFS; Eurostat; author's calculations

change rate changes are displayed in *Figure 2*. Looking at this figure, there is a general trend of decreasing volatility at the end of the time span. Regarding the connection between volatility and real-life events, there is a tendency for increased volatility prior to presidential or parliamentary elections, although this is not always the case. In the case of Hungary, there is a visible change in the regime in May 2001 with much higher volatility afterwards. On the other hand, Slovenia has extremely low volatility, which is the result of a tight exchange rate regime during the whole time span.

[17] According to Reinhart and Rogoff (2004), it is necessary to be careful while modeling different exchange rate regimes and not to blindly follow official classification of these regimes. For this reason, they study dual and parallel exchange rate markets on a sample of 153 countries over a 55-year time period. Based on this, they claim that a majority of official pegs are actually floats, and vice versa. As a result, they provide a new system for classifying exchange rate regimes with the accent on real and proclaimed regimes.
Fortunately, this is not the case in the CEEC-5 group. These countries either comply with their proclaimed regimes or the changes are only superficial. Only in the case of Poland is there a pre-announced crawling band of ±12.5 %, later changed to ±15 %, while according to Reinhart and Rogoff (2004) it is de facto ±5 % in both cases. Therefore, I employ official exchange rate regimes for the purpose of this study.

FIGURE 2 Exchange Rate Changes in the CEEC-5 Group




## 5. Empirical Results

The first stage of analyzing the effects of determinants of exchange rate volatility consists of estimating the corresponding TARCH model for each currency. The results of these estimations are summarized in *Table 4*. With the exception of Slovenia, the results suggest that in the CEEC-5 group the exchange rate volatility is statistically significantly asymmetrically affected by unpredictable circumstances. The sign of leverage term implies lower conditional variance in the case of a negative shock for Hungary, Poland, and Slovakia, while for the Czech Republic it implies higher conditional variance. In the case of Slovenia, a simple ARCH model with variance as an ARCH-M term is estimated instead. For each country the sum of $\alpha$'s and $\beta$'s in the variance equation satisfies the stability condition, which implies that exchange rate changes converge to the steady-state level, although this convergence is slow in the Czech Republic with sums close to one. Additionally, the value of coefficient $\beta$ in the GARCH term close to one suggests that there is a high persistence of conditional variance in these two countries. Regarding asymmetric effects and variance persistency, these results are in accordance with previous studies by Orlowski

TABLE 2 Exchange Rate Regimes – An Overview

| Country | Variable | Period | Official regime |
|---|---|---|---|
| Czech Rep. | – | 27. 5. 1997– . . . | managed floating |
| Hungary | REG1 | 1. 1. 1999–31. 12. 1999 | crawling band around basket (±2.25 %) (basket=USD 30 %, EUR 70 %) |
| | REG2 | 1. 1. 2000–3. 5. 2001 | crawling band around EUR (±2.25 %) |
| | REG3 | 4. 5. 2001–3. 6. 2003 | crawling band around EUR (±15 %) (central rate 276.10 HUF/EUR) |
| | base | 4. 6. 2003– . . . | crawling band around EUR (±15 %) (central rate 282.36 HUF/EUR) |
| Poland | REG1 | 1. 1. 1999–24. 3. 1999 | crawling band around basket (±12.5 %) (basket=USD 45 %, EUR 55 %) |
| | REG2 | 25. 3. 1999–11. 4. 2000 | crawling band around EUR (±15 %) |
| | base | 12. 4. 2000– . . . | managed floating |
| Slovakia | – | 1. 10. 1998– . . . | managed floating |
| Slovenia | REG1 | 1. 1. 1999–27. 6. 2004 | crawling band around EUR (±2 %) |
| | base | 28. 6. 2004– . . . | ERM II |

*Source*: (Reinhart, Rogoff, 2004); (Kočenda, Valachy, 2006); national banks

TABLE 3 Nominal Exchange Rates vis-à-vis the EUR – Data Summary

| Country | Min | Max | Mean | St. dev. |
|---|---|---|---|---|
| Czech Republic | 28.9590 | 38.5830 | 33.5196 | 2.2916 |
| Hungary | 234.7200 | 273.9200 | 252.9331 | 7.7495 |
| Poland | 3.3433 | 4.9346 | 4.1169 | 0.3439 |
| Slovakia | 38.5450 | 47.4840 | 42.3693 | 1.5187 |
| Slovenia | 187.1333 | 240.0300 | 219.6494 | 15.6962 |

*Source*: author's calculations

(2003) and Kočenda and Valachy (2006), although these studies do not deal with Slovenia.

Once the correct ARCH model is specified, conditional variance from this model is calculated. Moreover, the *NEWS* variable is estimated from equation (1). Basic characteristics about variance and news are presented in *Tables 5* and *6,* respectively. Both of these variables are further depicted in *Figures 3* and *4,* respectively.

The final stage is to run a regression (2). This regression is run without the *REGIME* variable for the Czech Republic and Slovakia, since these countries did not change their exchange rate regime during the whole time span. I also allow for lags of the *NEWS* variable in this regression in order to capture delayed effects of news. The results of these estimations are summarized in *Table 7*.

The negative sign of coefficient $\beta$ corresponds to the previously mentioned theory that countries with more open economies tend to have lower exchange rate volatility. The only exception here is Hungary with a positive, but statistically insignificant, value of $\beta$. The effect of openness is insignificant also in the Czech Republic. On the other hand, openness has statistically significant effects on exchange rate volatility in the other three countries. The smallest effect is in Slovakia, where a 1 % increase in the ratio of import over GDP decreases variance by 3.1 % of its mean value.

TABLE 4 TARCH Model Estimations

| | Czech Rep. | Hungary | Poland | Slovakia | Slovenia |
|---|---|---|---|---|---|
| $\sigma^2$ | – | – | – | – | 55.8525$^{a}$ |
| | – | – | – | – | (9.4673) |
| $a_1$ | – | 0.0857$^{b}$ | – | 0.0855$^{a}$ | –0.2439$^{a}$ |
| | – | (0.0424) | – | (0.0316) | (0.0604) |
| $a_2$ | – | – | –0.0781$^{a}$ | – | – |
| | – | – | (0.0291) | – | – |
| $a_5$ | –0.4526$^{a}$ | –0.0766 | – | – | – |
| | (0.1446) | (0.0472) | – | – | – |
| $a_6$ | – | –0.0779$^{a}$ | – | – | – |
| | – | (0.0287) | – | – | – |
| $a_{10}$ | – | – | 0.0846$^{a}$ | – | – |
| | – | – | (0.0262) | – | – |
| $b_5$ | 0.4768$^{a}$ | – | – | – | – |
| | (0.1424) | – | – | – | – |
| $\omega$ | $2\cdot10^{-7}$ | $9\cdot10^{-6a}$ | $2\cdot10^{-5a}$ | $5\cdot10^{-6a}$ | $5\cdot10^{-7}$ |
| | $(1\cdot10^{-7})$ | $(3\cdot10^{-6})$ | $(2\cdot10^{-6})$ | $(7\cdot10^{-7})$ | $(8\cdot10^{-6})$ |
| $\alpha_1$ | 0.1203$^{b}$ | 0.4318$^{a}$ | 0.2000$^{a}$ | 0.2201$^{a}$ | 0.6542$^{a}$ |
| | (0.0570) | (0.1667) | (0.0502) | (0.0648) | (0.1544) |
| $\alpha_2$ | –0.0849 | 0.3263$^{b}$ | 0.1276$^{a}$ | 0.1208$^{b}$ | 0.2167$^{b}$ |
| | (0.0520) | (0.1505) | (0.0467) | (0.0580) | (0.1053) |
| $\alpha_3$ | – | – | 0.1753$^{a}$ | 0.0495 | – |
| | – | – | (0.0491) | (0.0338) | – |
| $\alpha_4$ | – | – | 0.0809$^{b}$ | – | – |
| | – | – | (0.0373) | – | – |
| $\beta_1$ | 0.9114$^{a}$ | – | – | – | – |
| | (0.0216) | – | – | – | – |
| $\xi$ | 0.0779$^{b}$ | –0.3332$^{b}$ | –0.1995$^{a}$ | –0.1347$^{c}$ | – |
| | (0.0352) | (0.1550) | (0.0628) | (0.0822) | – |
| # of obs. | 1 497 | 1 507 | 1 438 | 1 469 | 1 497 |
| adj. $R^2$ | 0.0125 | 0.0084 | 0.0070 | 0.0069 | –0.0379 |
| AIC | –8.5043 | –8.3460 | –7.3190 | –8.9829 | –11.0550 |
| SIC | –8.4795 | –8.3213 | –7.2896 | –8.9613 | –11.0372 |

*Note*: Standard errors are in parentheses; significance at the 1%, 5%, and 10% levels is denoted by a, b, and c superscript, respectively.

Bigger effects are seen in Poland (8.8 %) and in Slovenia with its huge 98 % decrease.[18] Although the impact on Slovenia seems to be too dramatic, one has to keep in mind that openness in Slovenia is the most stable among these five countries and a 1% increase in the overall trend is quite unlikely. Moreover, these numbers are provided here only for comparison in order to see the differences between particular countries – Slovenia is much more affected by its foreign trade than are the other four countries.

The results for regimes reflect natural expectations – a less tight regime corresponds to higher volatility. In May 2001 Hungary changed its exchange rate regime from a ±2.25 % crawling band to ±15 % and the results suggest that the former regi-

[18] These results are obtained by comparing the estimated coefficient from *Table 7* with the corresponding mean value of estimated conditional variance from *Table 5*.

FIGURE 3 Estimated Conditional Variance in the CEEC-5 Group

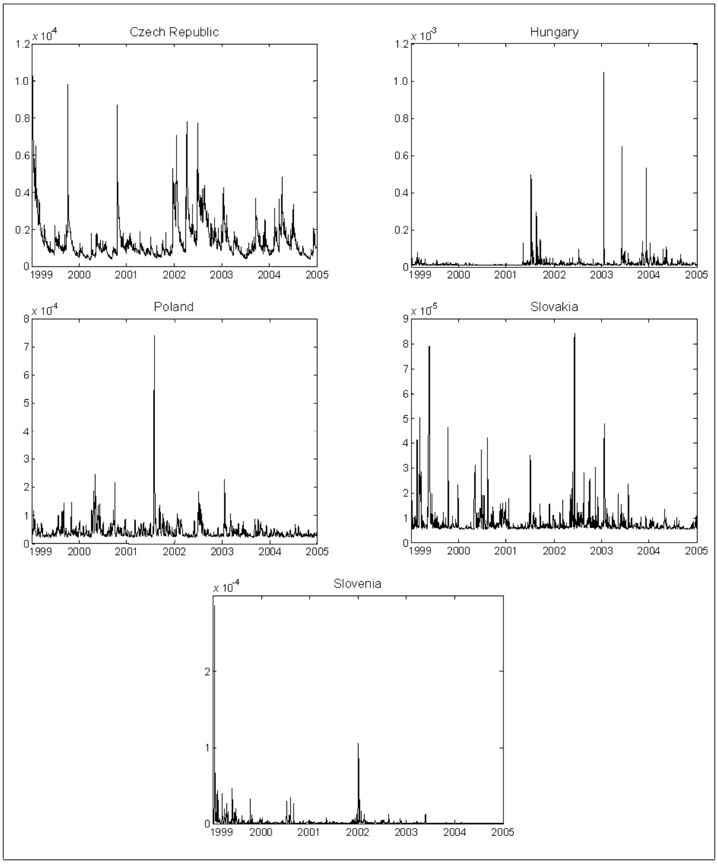



TABLE 5 Estimated Conditional Variance – Data Summary

| Country | Min | Max | Mean | St. dev. |
|---|---|---|---|---|
| Czech Republic | $3.68 \cdot 10^{-6}$ | 0.000117 | $1.42 \cdot 10^{-5}$ | $1.15 \cdot 10^{-5}$ |
| Hungary | $9.12 \cdot 10^{-6}$ | 0.001042 | $2.01 \cdot 10^{-5}$ | $4.92 \cdot 10^{-5}$ |
| Poland | $2.24 \cdot 10^{-5}$ | 0.000740 | $4.42 \cdot 10^{-5}$ | $3.98 \cdot 10^{-5}$ |
| Slovakia | $5.37 \cdot 10^{-6}$ | 0.000084 | $8.10 \cdot 10^{-6}$ | $6.15 \cdot 10^{-5}$ |
| Slovenia | $4.96 \cdot 10^{-7}$ | 0.000286 | $2.28 \cdot 10^{-6}$ | $1.03 \cdot 10^{-5}$ |

me significantly lowers conditional variance by its mean value compared to the latter regime. In the case of Poland, a change of the exchange rate regime from a ±12.5 % crawling band to ±15 % has no significant effects on conditional variance with respect to a base floating exchange rate regime. Similarly, in Slovenia a change from a ±2% crawling band to Exchange Rate Mechanism II (ERM II) has no significant effects on volatility either. This implies that Slovenia does not use the whole ±15% band that is allowed by ERM II.

With regard to the news effects,[19] the results are mostly consistent with the results concerning the leverage effect from the TARCH model. The exceptions are Po-

FIGURE 4  Estimated NEWS Variable in the CEEC-5 Group

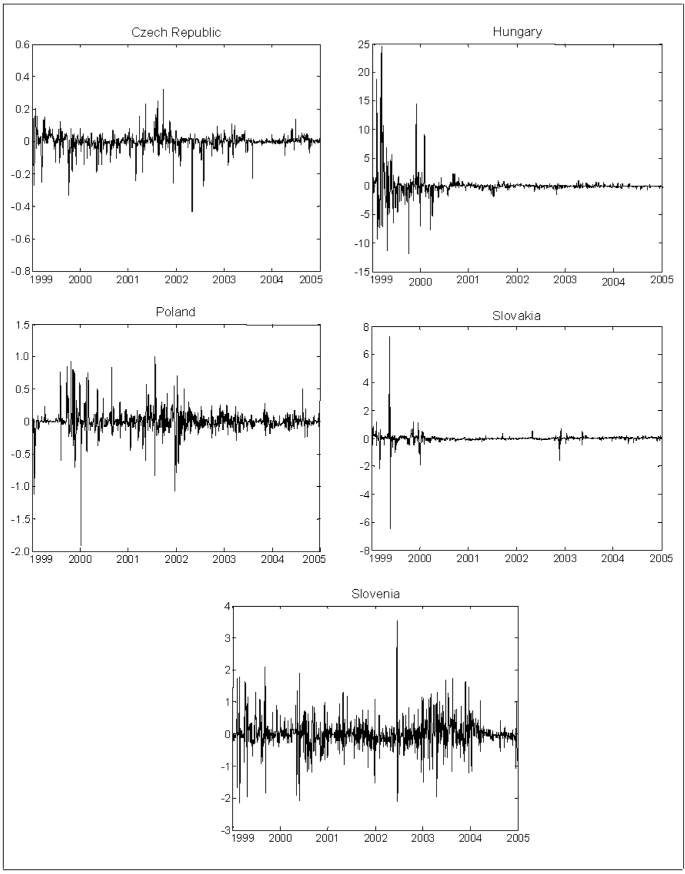



TABLE 6  Variable NEWS – Data Summary

| Country | Min | Max | Mean | St. dev. | Median |
|---|---|---|---|---|---|
| Czech Republic | −0.7872 | 0.3244 | $-5.85 \cdot 10^{-16}$ | 0.0564 | 0.0033 |
| Hungary | −12.3328 | 24.5462 | $1.03 \cdot 10^{-14}$ | 1.9964 | 0.0222 |
| Poland | −1.9082 | 1.0049 | $1.15 \cdot 10^{-15}$ | 0.1789 | −0.0024 |
| Slovakia | −6.4948 | 7.2563 | $-4.19 \cdot 10^{-15}$ | 0.3832 | 0.0021 |
| Slovenia | −2.1471 | 3.5324 | $6.41 \cdot 10^{-16}$ | 0.4378 | −0.0303 |

*Note*: A negative value indicates good news, a positive values indicates bad news.

land and Hungary with lagged effects of news. In both cases, the effects of good news are opposite those estimated by the leverage term from the TARCH model. The reason for this discrepancy may be that the TARCH model uses the residuals only from exchange rate changes, while the approach in equation (1) accounts for changes in expectations about exchange rates as well as interest rates. However, the com-

[19] Recall that a negative sign of coefficient $\gamma$ in the case of good news is interpreted so that it increases exchange rate volatility.

TABLE 7 Sources of Exchange Rate Volatility – Results

$$ERV_t = \alpha + \beta\left(\frac{Im_t}{GDP_t}\right) + \gamma_G G_t * NEWS_t + \gamma_B B_t * NEWS_t + \delta REGIME_t + \varepsilon_t$$

| | Czech Rep. | Hungary | Poland | Slovakia | Slovenia |
|---|---|---|---|---|---|
| $\alpha$ | $2 \cdot 10^{-5}$<br>$(3 \cdot 10^{-5})$ | $-1 \cdot 10^{-4}$<br>$(2 \cdot 10^{-4})$ | $2 \cdot 10^{-4a}$<br>$(5 \cdot 10^{-5})$ | $3 \cdot 10^{-5a}$<br>$(8 \cdot 10^{-6})$ | $1 \cdot 10^{-4c}$<br>$(6 \cdot 10^{-5})$ |
| $\beta$ | $-1 \cdot 10^{-5}$<br>$(6 \cdot 10^{-5})$ | $3 \cdot 10^{-4}$<br>$(4 \cdot 10^{-4})$ | $-4 \cdot 10^{-4a}$<br>$(2 \cdot 10^{-4})$ | $-3 \cdot 10^{-5b}$<br>$(1 \cdot 10^{-5})$ | $-2 \cdot 10^{-4c}$<br>$(1 \cdot 10^{-4})$ |
| $\gamma_G$ | $-7 \cdot 10^{-6}$<br>$(6 \cdot 10^{-6})$ | –<br>– | –<br>– | –<br>– | –<br>– |
| $\gamma_{G\ lag\#1}$ | –<br>– | –<br>– | $-7 \cdot 10^{-6b}$<br>$(4 \cdot 10^{-6})$ | –<br>– | –<br>– |
| $\gamma_{G\ lag\#3}$ | –<br>– | –<br>– | –<br>– | $1 \cdot 10^{-6b}$<br>$(4 \cdot 10^{-7})$ | $-2 \cdot 10^{-6c}$<br>$(1 \cdot 10^{-6})$ |
| $\gamma_{G\ lag\#5}$ | –<br>– | $-4 \cdot 10^{-6c}$<br>$(2 \cdot 10^{-6})$ | –<br>– | –<br>– | –<br>– |
| $\gamma_{B\ lag\#3}$ | –<br>– | –<br>– | –<br>– | –<br>– | $7 \cdot 10^{-7}$<br>$(5 \cdot 10^{-7})$ |
| $\gamma_{B\ lag\#4}$ | –<br>– | –<br>– | –<br>– | $2 \cdot 10^{-6d}$<br>$(1 \cdot 10^{-6})$ | –<br>– |
| $\gamma_{B\ lag\#5}$ | $-1 \cdot 10^{-5c}$<br>$(7 \cdot 10^{-6})$ | $4 \cdot 10^{-7d}$<br>$(3 \cdot 10^{-7})$ | –<br>– | –<br>– | –<br>– |
| $\gamma_{B\ lag\#7}$ | –<br>– | –<br>– | $-3 \cdot 10^{-5c}$<br>$(1 \cdot 10^{-5})$ | –<br>– | –<br>– |
| $\delta_{REG1}$ | –<br>– | $-1 \cdot 10^{-5a}$<br>$(4 \cdot 10^{-6})$ | $-2 \cdot 10^{-5}$<br>$(2 \cdot 10^{-5})$ | –<br>– | $-2 \cdot 10^{-6}$<br>$(1 \cdot 10^{-6})$ |
| $\delta_{REG2}$ | –<br>– | $-2 \cdot 10^{-5b}$<br>$(1 \cdot 10^{-5})$ | $-7 \cdot 10^{-6}$<br>$(8 \cdot 10^{-6})$ | –<br>– | –<br>– |
| $\delta_{REG3}$ | –<br>– | $-8 \cdot 10^{-6}$<br>$(1 \cdot 10^{-5})$ | –<br>– | –<br>– | –<br>– |
| $\rho$ | $0.8511^a$<br>(0.0310) | $0.5228^a$<br>(0.1158) | $0.8170^a$<br>(0.1075) | $0.6993^a$<br>(0.0550) | $0.6630^a$<br>(0.2272) |
| # of obs.<br>adj. $R^2$ | 1 495<br>0.7405 | 1 506<br>0.2901 | 1 437<br>0.6753 | 1 464<br>0.5303 | 1 493<br>0.4594 |

*Note*: Standard errors are in parentheses; significance at 1%, 5%, 10%, and 15% level is denoted by a, b, c, and d superscript respectively. Parameter $\rho$ is included to account for serial correlation in residuals.

plexity of this latter approach guarantees more accurate measures of the news effects. In all five countries, the results suggest that news statistically significantly impacts exchange rate volatility; there is no statistically significant difference between good and bad news at the level of the effect, and there are huge differences in vulnerability across the CEEC-5 group when taking into account the extreme values of estimated news.[20] Good (bad) news increases (decreases) exchange rate volatility in the Czech Republic (about 30 % of its mean) and Poland (about half the mean). The situation is the opposite in Slovakia, where good news decreases and bad news increases exchange rate volatility by more than the mean of its exchange rate volatility. Hungary and Slovenia

[20] The following results are obtained by multiplying the estimated coefficients from Table 7 with the corresponding extreme values from *Table 6* (min. for good news and max. for bad news) and then comparing them with the mean value of estimated conditional variance from *Table 5*.

are the only countries where any news, good or bad, increases exchange rates volatility, again by more than the mean of their exchange rate volatility.

The previous results reveal an interesting fact. While good news contributes to increased volatility in every country except Slovakia, the impact of bad news is negative or close to zero when positive values are either not significant or significant only at the 15% level in the case of Hungary, Slovakia, and Slovenia. Nevertheless, although this fact seemingly contradicts natural expectations, it has a reasonable explanation. Regarding good news – the exchange rate changes also in this case and even if it declines – this appreciation contributes to increased volatility. On the other hand, the impact of bad news is suppressed because bad news is usually accompanied by expectations of active policy responses that hamper any potential movements in the exchange rate (Orlowski, 2003).

Several further tests are performed to ensure that the previous results are robust for changes in the methodology of either the dependent or independent variable. At first, openness in regression (2) is replaced with its original form, i.e., without the Hodrick-Prescott filter. The results for news and regimes remain unchanged – the sign as well as significance. Regarding openness, the results are in general lower in magnitudes (2–10 times), with the sign and significance unchanged. This is probably caused by higher variation in unfiltered openness when the overall impact is diminished. This only supports the application of the Hodrick-Prescott filter. The second variation in regression (2) is the change in the dependent variable. Now exchange rate volatility is modeled using a simpler GARCH model so the asymmetric effects are allowed only by the *NEWS* variable.[21] The results of all variables are consistent with the baseline specification also in this case. There are only minor differences in magnitudes of news.

## 6. Conclusion

In this paper I have analyzed the sources of euro exchange rate volatility separately for every country in the CEEC-5 group. As possible sources, I am interested in the openness of an economy, the news factor, and the exchange rate regime. Exchange rate volatility is estimated from a TARCH model with emphasis on the asymmetric effects of news. However, these asymmetric effects are confirmed only in the sense of their sign, not their value. This study further confirms the assumption that more open economies tend to have lower exchange rate volatility when this result holds in most countries.

Looking at the results for particular countries, news has a large effect on exchange rate volatility in Hungary, Slovakia, and Slovenia. But Slovenia has huge potential in its openness, which has a substantial decreasing impact on its exchange rate volatility. On the other hand, Hungary and Slovakia cannot rely on such a tool because openness has almost no effect on their exchange rate volatility. The other two countries, Poland and the Czech Republic, cannot rely on openness in decreasing their exchange rate volatility either. However, these countries' exchange rate volatility is affected by news only slightly. Regarding regimes, only key changes in exchange rate regimes have significant effects on exchange rate volatility, while minor and superficial changes are not reflected in volatility at all.

[21] This robustness test is not performed for Slovenia since the TARCH model was not employed in this case.

The main contribution of this study is that it sheds some light on a few potential pitfalls that may occur during the CEEC-5 group's EMU integration process. The fact that the vulnerability of these countries varies may be explained by the different strength of each country's currency or by different policies adopted by each country's central bank. Either way, further research is needed in order to distinguish between these two cases and to see their impact on other Maastricht criteria.

REFERENCES


Baum CF, Caglayan M, Ozkan N (2004): Nonlinear Effects of Exchange Rate Volatility on the Volume of Bilateral Exports. *Journal of Applied Econometrics*, 19(1):1–23.

Belke A, Setzer R (2003): Exchange Rate Variability and Labor Market Performance in the Visegrad Countries. *Economics of Planning*, 36(2):153–175.

Bollerslev T (1986): Generalized Autoregressive Conditional Heteroskedasticity. *Journal of Econometrics*, 31:307–327.

Box G, Jenkins G (1976): *Time Series Analysis, Forecasting, and Control*. San Francisco (California), Holden day.

Brock WA, Dechert WD, Scheinkman JA (1987): A Test for Independence Based on the Correlation Dimension. *University of Wisconsin at Madison*, Department of Economics *Working Paper.*

Bulíř A. (2005): Liberalized Markets Have More Stable Exchange Rates: Short-Run Evidence from Four Transition Countries. *Finance a úvěr-Czech Journal of Economics and Finance*, 55(5-6): 206–231.

Černý A, Koblas M (2005): Stock Market Integration and the Speed of Information Transmission: the Role of Data Frequency in Cointegration and Granger Causality Tests. *Journal of International Business and Economics*, 1:110–120.

Choudhry T (2005): Exchange Rate Volatility and the United States Exports: Evidence from Canada and Japan. *Journal of the Japanese and International Economies*, 19(1):51–71.

Chowdhury AR (1993): Does Exchange Rate Volatility Depress Trade Flows? Evidence from Error-Correction Models. The *Review of Economics and Statistics*, 75(4):700–706.

Dell'Ariccia G (1999): Exchange Rate Fluctuations and Trade Flows: Evidence from the European Union. *IMF Stuff Paper*, 46(3):315–334.

Dornbusch R (1978): Monetary Policy under Exchange Rate Flexibility. In: Managed Exchange-rate Flexibility: The Recent Experience. *Federal Reserve Bank of Boston Conference Series*, 20.

Engle RF, Ng VK (1993): Measuring and Testing the Impact of News on Volatility. The J*ournal of Finance*, 48(5):1749–1778.

Engle RF (1982): Autoregressive Conditional Heteroskedasticity with Estimates of the Variance of U.K. Inflation. *Econometrica*, 50:987–1008.

Frenkel JA (1981): Flexible Exchange Rates, Prices, and the Role of "News": Lessons from the 1970s. *Journal of Political Economy*, 89(4):665–705.

Galati G, Ho C (2003): Macroeconomic News and the Euro/Dollar Exchange Rate. *Economic Notes*, 32(3):371–398.

Hau H (2002): Real Exchange Rate Volatility and Economic Openness: Theory and Evidence. *Journal of Money, Credit and Banking*, 34(3):611–630.

Hodrick RJ, Prescott EC (1997): Postwar U.S. Business Cycles: An Empirical Investigation. *Journal of Money, Credit and Banking*, 29(1):1–16.

Kenen PB, Rodrik D (1986): Measuring and Analyzing the Effects of Short-Term Volatility in Real Exchange Rates. The *Review of Economics and Statistics*, 68(2):311–315.

Kočenda E (1998): Altered Band and Exchange Volatility. *Economics of Transition*, 6(1):173–181.

Kočenda E (2001): An Alternative to the BDS Test: Integration across the Correlation Integral. *Econometric Reviews*, 20(3):337–351.

Kočenda E, Valachy J (2006): Exchange Rate Volatility and Regimes Change: A Visegrad Comparison. *Journal of Comparative Economics*, 34:727–753.

Kóbor Á, Székely IP (2004): Foreign Exchange Market Volatility in EU Accession Countries in the Run-Up to Euro Adoption: Weathering Uncharted Waters. *IMF Working Paper*, 16.

Koray F, Lastrapes WD (1989): Real Exchange Rate Volatility and U.S. Bilateral Trade: A Var Approach. The *Review of Economics and Statistics*, 71(4):708–712.

Ljung G, Box G (1978): On a Measure of Lack of Fit in Time Series Models. *Biometrica*, 65: 297–303.

McKenzie M (2002): The Economics of Exchange Rate Volatility Asymmetry. *International Journal of Finance & Economics*, 7(3):247–260.

Orlowski LT (2003): Monetary Convergence and Risk Premiums in the EU Accession Countries. *Open Economies Review*, 14:251–267.

Orlowski LT (2004): Exchange Rate Risk and Convergence to the Euro. *ZEI Working Paper*, B25.

Reinhart CM, Rogoff KS (2004): Modern History of Exchange Rate Arrangements: A Reinterpretation. *Quarterly Journal of Economics*, 119(1):1–48.

Sanchez-Fung JR (2003): Non-linear Modeling of Daily Exchange Rate Returns, Volatility, and News in a Small Developing Economy. *Applied Economics Letters*, vol. 10(4):247–250.