Impact of the European Union Enlargement on the Cointegration of the Central European Stock Markets

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ABSTRACT

This study investigates empirically the long-run price relationships among the three largest Central European (CE) stock markets (Czech, Hungarian, Polish), the Russian and two major world markets (US and West European) using weekly prices of stock market indexes from November 1998 to May 2007 and applying the Johansen method of cointegration. Particular attention is paid to the impact of the European Union (EU) enlargement process before May 2004. Results show that the number of cointegration relations among the three CE markets did not change after the enlargement, though the characteristics of the linkages were modified. Regarding the long-run equilibrium relations among all investigated markets, two new relations emerged after the EU enlargement, meaning that potential long-run portfolio diversification benefits have been reduced. In particular, one new cointegration relation after the EU enlargement could be identified as a new EU relation linking the West European and the three CE markets.

KEYWORDS: emerging stock markets, international, market linkages, cointegration, Central Europe

JEL CLASSIFICATION: G10, G11, G15

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1 Introduction

The issue of international portfolio diversification is of particular importance for stock market investors. In case of strong relationships among the markets, the possibility of portfolio diversification can be very limited. A lot of empirical literature has already addressed the linkages among the major developed markets. The results show that the relationships have been tightening due to the deregulation and liberalisation in capital and money markets (see Eun and Shim (1989), Koch and Koch (1991), Taylor and Tonks (1989), Kasa (1992), Masih and Masih (1992), Longin and Solnik (1995), Bessler and Yang (2003)). The lower diversification benefits among developed markets have redirected the attention to emerging markets. This paper follows the recent work in this field (e.g. Arshanapalli et al. (1995), Choudhry (1997), Tuluca and Zwick (2001), Manning (2002) Chen et al. (2002)) by investigating linkages between emerging and developed stock markets.

Since the number of empirical studies involving the emerging markets in Europe is still rather limited compared to the emerging markets in Asia or Latin America, this study focuses on Central European (CE) markets. Moreover, concentrating on the CE markets is an interesting issue nowadays, because many substantial institutional changes have been occurring in the CE countries in recent years, mainly in connection with the accession to the European Union (EU) on May 1, 2004. The institutional arrangements as well as fiscal and monetary policies have been strongly motivated by several criteria that set conditions for the EU accession and directed the adjustment of the CE countries towards the EU standards. The EU accession on itself was then associated with the removal of all restrictions on movement of capital and the CE markets have therefore become more opened and liberalised. Nevertheless, the restructuring process is going on as the policy makers in the new member countries attempt to join the Eurozone soon (McKinnon (1999), Buiter and Grafe (2002), Buiter (2004)). The institutional changes are accompanied by the convergence in macro-economic fundamentals of the recent EU members to the EU standards. For instance, E. Kocenda and Yigit (2006) report significant strong nominal convergence towards the EU standards, though the convergence is found to be slower regarding the per-capita real income.

Therefore, it is of particular interest, whether stronger linkages between the CE and the West European countries can be found regarding the financial markets as well. Dvorak (2007) investigates this question focusing on financial integration and finds evidence for significant financial integration of the new EU members that joint EU in 2004 towards the old EU countries. The aim of this paper is to approach the issue from a different point of view by concentrating on the long-run stock market linkages using the Johansen cointegration technique (see Johansen (1991)). We focus on the three largest CE markets, the Czech, Hungarian and Polish market and the West European markets as counterpart. To capture potentially other influential stock market movements, the US and the Russian market are added to the analysis as well. The US market obviously represents the largest developed stock market in the world. The inclusion of the Russian market on the contrary captures the development in the largest emerging market in Europe with strong historical link to the CE countries. Since changes in the long-run equilibrium relations due to the EU accession are expected, we focus on the comparison of the period before and after the EU enlargement. In particular, we attempt to answer two key questions: (1) What were the characteristics of the long-run market linkages before the EU accession? (2) How did they change after the EU enlargement? For that purpose, two models for the different periods (pre- and post- accession) are set up, interpreted and compared in this study. Although this analysis is by far not the first study of the long-run market linkages involving the CE countries, it delivers the first evidence about the long-run equilibrium changes in the investigated stock markets associated with the EU enlargement.

Most of the previous studies concentrated on the Russian Financial crises in 1998 and its impact. For instance, the first study dealing with the long-run co-movements among Central and Eastern European (CEE) markets and the US market (see Jochum et al. (1999)) has found out that a long-run relationship within the Russian, Hungarian, Polish, Czech and the US market is present for the period 1995–1997, but the relationship disappears during the crises of emerging markets in 1997–1998. According to Gilmore and McManus (2002) no long-run equilibrium relation can be found among the US, Czech, Polish and Hungarian markets for the period 1995–2001. Yuca and Simga-Mugan (2000) have reported no evidence of cointegration among the CEE countries and a limited evidence of linkage with developed stock markets. Similarly, Syrioupoulus (2006) has found out that the long-run linkages of the CEE stock markets with developed markets are stronger than the linkages among the CEE markets. Allowing for instability in the cointegration relations, Voronkova (2004) has reported a stronger link between the CE markets and between the CE markets and their mature counterparts than in the previous literature. The most recent paper Yang et al. (2006) has concentrated on the same markets as Jochum et al. (1999) plus the German market and the period has been extended to June 2002. The results show that the long-run price relationship among the markets was strengthened after the Russian financial crisis.

The existing studies differ in the investigated time-spans as well as the data frequency chosen or the stock market indexes used. These differences might explain some of the conflicting results or even lead to non-comparable results at all. Nevertheless, an additional source of the mixed evidence seems to be caused by the instability of the equilibrium relations. Since shifts and changes in the equilibrium relations are highly likely to occur in emerging markets, it is challenging to follow the studies of Jochum et al. (1999), Yang et al. (2006) and Voronkova (2004) and allow for the changes in the equilibrium relations as this study does. Moreover, the proper use of the cointegration technique relies on several assumptions, such as the constancy of the parameters or independence of the residuals. Surprisingly, only a few studies report tests of these assumptions. In case that the assumptions are not fulfilled, the results concluded by these studies might be unreliable and this can further contribute to the mixed evidence.

To fill this gap in the literature, this paper pays particular attention to the assumptions for the use of the Johanson cointegration technique. Especially, the assumption of constant parameters is checked by several recursive tests in order to avoid the distortion of the results by its violation. Other assumptions regarding independently and normally distributed residuals are handled carefully as well and, if necessary, too large residuals caused by extraordinary shocks (e.g. 11th September) are modelled by inclusion of proper dummy variables. Although the Johansen cointegration technique enables to analyse both the long-run and the short-run market structure, we focus on the long-run equilibrium relations. The reasons are twofold. First, the long-run relations are relevant for the long-run gains from international portfolio diversification and allow to investigate the potential integration of the markets. Second, the estimates for the long-run relations are more reliable than the estimates of the short-run structure because the convergence to their true values is faster (Juselius (2007)). It will be shown in our models that the coefficients of the short-run structure will indeed be more unstable than the estimates of the long-run, since the investigated time-spans are rather short. In order to check the degree to which the generalisation of the obtained results is possible, several robustness checks are conducted at the end of the analysis.

The rest of the paper is organised as follow: Section 2 describes the statistical model used and the notation throughout the paper. Section 3 introduces the choice of the information set and delivers basic summary statistics of the data. Two models for the pre- and post-accession period are estimated and compared in Section 4. Finally, Section 5 concludes.

2 Methodology

For modelling the long-run relationships among the stock markets, the cointegrated VAR model (see Johansen (1991), Juselius (2007)) is used. Generally, let $\boldsymbol{X}_t$ denote a vector of p time series integrated of order one (I(1) hereafter) or stationary time series. Furthermore, assume that $\boldsymbol{X}_t$ can be modelled by the VAR(k) model, i.e.

$$\boldsymbol{X}_t = \boldsymbol{\Pi}_1 \boldsymbol{X}_{t-1} + \cdots + \boldsymbol{\Pi}_k \boldsymbol{X}_{t-k} + \varepsilon_t, \qquad \text{where} \qquad \varepsilon_t \sim \boldsymbol{IN}_p(0, \boldsymbol{\Omega}) \qquad \text{and} \qquad t = 1, \ldots, T. \tag{1}$$

More specifically, the time series used in our setting are the logarithms of the stock market indexes that are almost always found to be I(1) processes and thus appropriate for the chosen method. The error terms $\varepsilon_t, t = 1, \ldots, T$ are assumed to be independently normally distributed with constant variance-covariance matrix.

More convenient way of working with the VAR(k) model in the cointegration framework is to rewrite the model in the vector equilibrium correction model (hereafter VECM) form:

$$\begin{aligned} \Delta \boldsymbol{X}_t &= \boldsymbol{\Pi} \boldsymbol{X}_{t-1} + \boldsymbol{\Gamma}_1 \Delta \boldsymbol{X}_{t-1} + \cdots + \boldsymbol{\Gamma}_{k-1} \Delta \boldsymbol{X}_{t-k+1} + \varepsilon_t, \qquad (2) \\ \text{where} \quad \Delta \boldsymbol{X}_t &= \boldsymbol{X}_t - \boldsymbol{X}_{t-1}, \\ \boldsymbol{\Pi} &= -(\boldsymbol{I} - \sum_{j=1}^{k} \boldsymbol{\Pi}_j), \\ \boldsymbol{\Gamma}_i &= -\sum_{j=i+1}^{k} \boldsymbol{\Pi}_j \end{aligned}$$

and $\boldsymbol{I}$ denotes the $p \times p$ identity matrix. Apart from the fact that the typically strong multicollinearity effect present in the VAR form is reduced in the VECM form, the VECM representation allows to directly deal with the nonstationary pattern in the data. Note that the $\boldsymbol{\Pi}$ matrix captures now all information about the long-run effects, because the VECM form contains only one term $\boldsymbol{\Pi}\boldsymbol{X}_{t-1}$ in the levels $\boldsymbol{X}_{t-1}$. Since a stationary process $(\Delta\boldsymbol{X}_t)$ cannot be equal to a nonstationary process, the rank of $\boldsymbol{\Pi}$ cannot be full provided at least one of the time-series is nonstationary. In such a case, $\boldsymbol{\Pi}$ can be partitioned as

$$\boldsymbol{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}',$$

where $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are $p \times r$ matrices and $r < p$. Then, $\boldsymbol{\beta}'\boldsymbol{X}_{t-1}$ represents the r stationary cointegration relations. The coefficients of $\boldsymbol{\alpha}$ capture the adjustment of variables to the cointegration relations and are called loadings. The $\boldsymbol{\Gamma}_i$ matrices contain information about the short-run linkages.

If necessary, the model can be further extended by inclusion of deterministic components, such as constant, deterministic trend or several dummy variables. Let $\boldsymbol{D}_t$ denote the matrix of the deterministic components and $\boldsymbol{d}_t$ the vector of dummy variables. Then

$$\boldsymbol{D}_t = \begin{bmatrix} 1 \\ t \\ \boldsymbol{d}_t \end{bmatrix}$$

and the model from Equation 2 is extended as follows:

$$\Delta\boldsymbol{X}_t = \boldsymbol{\alpha}\boldsymbol{\beta}'\boldsymbol{X}_{t-1} + \boldsymbol{\Gamma}_1\Delta\boldsymbol{X}_{t-1} + \cdots + \boldsymbol{\Gamma}_{k-1}\Delta\boldsymbol{X}_{t-k+1} + \boldsymbol{\Phi}\boldsymbol{D}_t + \boldsymbol{\varepsilon}_t. \tag{3}$$

The deterministic components can be further partitioned into those entering the cointegration relations, i.e. restricted to appear only in the cointegration relations, and the unrestricted ones (see Juselius (2007) for more details). Formally,

$$\boldsymbol{\Phi}\boldsymbol{D}_t = (\boldsymbol{\alpha}\tilde{\boldsymbol{\beta}}' + \tilde{\boldsymbol{\gamma}})\boldsymbol{D}_t = \left(\boldsymbol{\alpha}[\tilde{\boldsymbol{\beta}}_0, \tilde{\boldsymbol{\beta}}_1, \tilde{\boldsymbol{\beta}}_2{}'] + [\tilde{\gamma}_0, \tilde{\gamma}_1, \tilde{\gamma}_2]\right) \begin{bmatrix} 1 \\ t \\ \boldsymbol{d}_t \end{bmatrix}.$$

Our modelling strategy is to allow at the beginning of the analysis for a relatively rich structure concerning the cointegration relationships. Hence, an unrestricted constant $(\boldsymbol{\alpha}\tilde{\boldsymbol{\beta}}_0 + \tilde{\gamma}_0)$ is included in all introduced models and the eventual dummy variables are generally regarded as unrestricted $((\boldsymbol{\alpha}\tilde{\boldsymbol{\beta}}_2{}' + \tilde{\gamma}_2)\boldsymbol{d}_t)$ as well. Nevertheless, the time-trend is restricted to enter only the cointegration relations, i.e. $\tilde{\gamma}_1 = 0$ and the specification becomes $(\boldsymbol{\alpha}\tilde{\boldsymbol{\beta}}_1 + \boldsymbol{0})t$. The reason for restriction of the time trend is to avoid a quadratic trend in the level data $\boldsymbol{X}_t$. Although a quadratic trend could improve the fit within the sample, it would lead to an implausible economic result that the stock markets follow quadratic trends. The resulting model from Equation 3 then becomes

$$\begin{aligned} \Delta\boldsymbol{X}_t &= \boldsymbol{\alpha}(\boldsymbol{\beta}'\boldsymbol{X}_{t-1} + \tilde{\boldsymbol{\beta}}_1 t) + \boldsymbol{\Gamma}_1\Delta\boldsymbol{X}_{t-1} + \cdots + \boldsymbol{\Gamma}_{k-1}\Delta\boldsymbol{X}_{t-k+1} + (\boldsymbol{\alpha}\tilde{\boldsymbol{\beta}}_0 + \tilde{\gamma}_0) + (\boldsymbol{\alpha}\tilde{\boldsymbol{\beta}}_2{}' + \tilde{\gamma}_2)\boldsymbol{d}_t + \boldsymbol{\varepsilon}_t \\ &= \boldsymbol{\alpha}\left[\boldsymbol{\beta}', \tilde{\boldsymbol{\beta}}_1\right]\boldsymbol{Z}_{t-1} + \boldsymbol{\Gamma}_1\Delta\boldsymbol{X}_{t-1} + \cdots + \boldsymbol{\Gamma}_{k-1}\Delta\boldsymbol{X}_{t-k+1} + (\boldsymbol{\alpha}\tilde{\boldsymbol{\beta}}_0 + \tilde{\gamma}_0) + (\boldsymbol{\alpha}\tilde{\boldsymbol{\beta}}_2{}' + \tilde{\gamma}_2)\boldsymbol{d}_t + \boldsymbol{\varepsilon}_t, \end{aligned}$$

where $\boldsymbol{Z}_{t-1} = \left[\boldsymbol{X}'_{t-1}, t\right]'$. Later on, the rich models can be reduced in more parsimonious ones by examining the significance of the corresponding coefficients.

Using directly the VECM in Equation 3 leads to $\boldsymbol{\alpha}\boldsymbol{\beta}'$ estimates that are distorted by the simultaneous estimation of the other coefficients $\boldsymbol{\Gamma}_1, \ldots, \boldsymbol{\Gamma}_{k-1}$ and $\boldsymbol{\Phi}$. The idea of the Frisch-Waugh Theorem (Frisch and Waugh (1933)) can be used in order to get "cleaner" estimates of the long-run structure captured by $\boldsymbol{\alpha}\boldsymbol{\beta}'$, where the estimation of the short-run effects and the deterministic components has been concentrated out. More specifically, the estimates of $\boldsymbol{\alpha}\boldsymbol{\beta}'$ can be obtained by using the concentrated model

$$\boldsymbol{R}_{0t} = \boldsymbol{\alpha}\boldsymbol{\beta}'\boldsymbol{R}_{1t} + \boldsymbol{\varepsilon}_t, \quad t = 1, \ldots, T, \tag{4}$$

where $\boldsymbol{R}_{0t}$ and $\boldsymbol{R}_{1t}$ are residuals from regressing $\Delta\boldsymbol{X}_t$ and $\boldsymbol{X}_{t-1}$, respectively, on $\boldsymbol{X}_{t-1}, \ldots, \boldsymbol{X}_{t-k+1}, \boldsymbol{D}_t$ (for more details see Juselius (2007)). The estimates of $\boldsymbol{\alpha}\boldsymbol{\beta}'$ based on the VECM ("X-form") and on the concentrated model ("R-form") will be distinguished later on in the tests of constant parameters.

Aside from the VECM form, important insights can be gained by inverting the VAR(k) model into its moving average (hereafter MA) representation[1], where $\boldsymbol{X}_t$ can be expressed as

$$\boldsymbol{X}_t = \boldsymbol{C}\left(\sum_{s=1}^{t}\boldsymbol{\varepsilon}_s\right) + \boldsymbol{C}\boldsymbol{\Phi}\boldsymbol{D}_t t + \boldsymbol{C}^*(L)\boldsymbol{\varepsilon}_t + \boldsymbol{X}_0 + \boldsymbol{C}^*(L)\boldsymbol{\varepsilon}_0. \tag{5}$$

Then, $\boldsymbol{C}(\sum_{s=1}^{t}\boldsymbol{\varepsilon}_s)$ refers to nonstationary stochastic trends, $\boldsymbol{C}\boldsymbol{\Phi}\boldsymbol{D}_t t$ to deterministic components, $\boldsymbol{C}^*(L)\boldsymbol{\varepsilon}_t$ to stationary stochastic components and $\boldsymbol{X}_0 + \boldsymbol{C}^*(L)\boldsymbol{\varepsilon}_0$ contains initial values. Similarly to the partition of the $\boldsymbol{\Pi}$ matrix, the $\boldsymbol{C}$ matrix can be further decomposed (see (Johansen, 1996, Chapter 4)) to

$$\boldsymbol{C} = \boldsymbol{\beta}_\perp(\boldsymbol{\alpha}'_\perp\boldsymbol{\Gamma}\boldsymbol{\beta}_\perp)^{-1}\boldsymbol{\alpha}'_\perp = \bar{\boldsymbol{\beta}}_\perp\boldsymbol{\alpha}'_\perp, \quad \text{where} \quad \bar{\boldsymbol{\beta}}_\perp = \boldsymbol{\beta}_\perp(\boldsymbol{\alpha}'_\perp\boldsymbol{\Gamma}\boldsymbol{\beta}_\perp)^{-1} \quad \text{and} \quad \boldsymbol{\Gamma} = -(\boldsymbol{I} - \sum_{j=1}^{k-1}\boldsymbol{\Gamma}_j).$$

Now, $\alpha'_\perp\sum_{s=1}^{t}\boldsymbol{\varepsilon}_s$ determines $p-r$ common stochastic (driving) trends and captures the driving forces of the system. $\bar{\boldsymbol{\beta}}_\perp$ denote the loadings of these trends that refer to the size of the market movements related to each stochastic trend.

3 Data

The data consists of weekly closing price indexes for the Czech, Hungarian, Polish, Russian, US and West European market. All the data used have been obtained from the Thompson Financial Datastream database. In order to avoid the distorting effects of using different types of local stock market indexes for the emerging markets, the standardised IFC Investable (IFCI) indexes are used for representing the Czech, Hungarian, Polish and Russian markets. Since the focus of this study is on the EU enlargement, the West European countries, most of them being the old members of the EU, are of a particular interest for the analysis. The limitations regarding the reasonable number of markets in the VAR model suggest to use only a single representative of the West European markets and, thus, the DJ Stoxx 600 is used. Following most of the related studies, the S&P 500 is chosen as a representative of the US market. The indices are measured in local-currency terms[2].

The data was collected for the time period between October 30, 1998, and May 4, 2007. The end of the time period is limited mainly by the data availability when writing the paper. The sample period covers three years after the EU enlargement on May 1, 2004, when the three involved CE countries entered the EU. The beginning of the sample period is motivated, first, by the attempt to avoid the distorting impacts of the emerging market crises in 1997/1998 and the Russian financial crises in August/September 1998 and, second, by the request to obtain a comparable period of three years before the EU enlargement when the accession date was still unclear. Although the accession negotiation for all of the three investigated CE countries were opened already on March 31, 1998, a significant turning point in the negotiations appeared to be November 2001, when the European Commission announced the EU enlargement in its Annual Progress Reports on Enlargement (see Dvorak and Podpiera (2006) for more details). Hence, the availability of the data and the timing of the different influential events suggest to split the data into the three following periods:

- pre-accession period: 10/30/1998 - 11/2/2001, covering 3 years,
- accession period: 11/2/2001 - 4/30/2004, covering 2.5 years,
- post-accession period: 4/30/2004 - 5/4/2007, covering 3 years.

Since the accession to the EU is a gradual process to a large extend, the accession period is viewed as a transition period, in which the long-run equilibrium relations are likely to be unstable, modified or changed. Therefore, it is investigated only marginally and the focus is on the comparison of the pre- and post-accession periods. Their time span of 3 years is comparable to other studies (e.g. Jochum et al. (1999), Yang et al. (2006)) analysing long-run market equilibrium relations. Using the weekly frequency of the data, both periods cover exactly 158 observations and, thus, are well-comparable.

[1]The derivation of the MA representation formula is left out here, but can be found in (Juselius, 2007, Chapter 5).

[2]In case of the DJ Stoxx 600, the used currency is Euro.

The reason for using weekly frequency data rather than daily data is the expectation that weekly data suffer less from the "stylised facts" of the financial time series such as heavy-tailed distributions or ARCH effects than the daily data. Moreover, the disturbing effects of different market closing-times (European, Russian, US market) are eliminated.

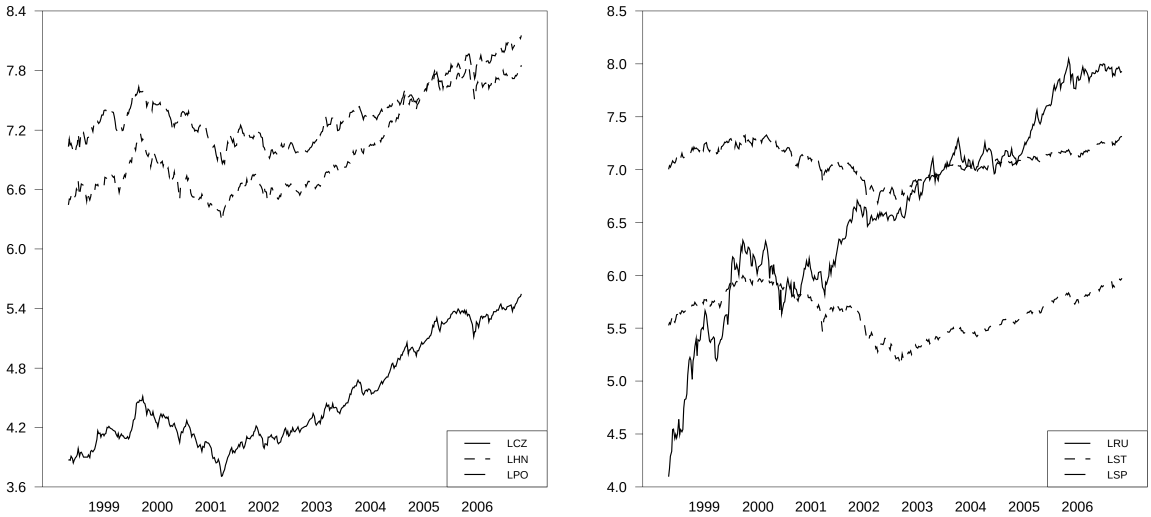

Figure 1: Logarithms of the stock market indexes

All the data have been converted to the natural logarithms as is usual in the literature. The notation used is lcz, lhn, lpo, lru, lst, lsp for the logarithms of the Czech, Hungarian, Polish, Russian, West European and US index, respectively. The development of the logarithms over the whole period is captured in Figure 1 and indicate a clear non-stationary behaviour. This allows us to use the logarithms of the indexes in the later analysis as the variables for the cointegration analysis. Note that a dramatic rise in the CE market prices can be observed after November 2001 that coincides with the EU enlargement announcement.

Descriptives of the returns	Pre-accession period 10/30/1998 - 11/2/2001		Post-accession period 4/30/2004 - 5/4/2007		T-test for differences (p-values)	
market	mean	S.D.	mean	S.D.	mean	S.D.
Czech Rep.	0.00074	**0.03490**	0.00567	**0.03051**	0.183	0.094
Hungary	0.00094	**0.04645**	0.00514	**0.03489**	0.365	0.000
Poland	0.00045	**0.04237**	0.00475	**0.02930**	0.296	0.000
Russia	0.01139	**0.08194**	0.00482	**0.03967**	0.366	0.000
West Europe	0.00092	**0.02824**	0.00296	**0.01512**	0.426	0.000
US	0.00010	**0.02982**	0.00176	**0.0144**0	0.530	0.000

Table 1: Mean and standard deviation of the continuous (log) return series in the pre- and post- accession periods, t-test for difference in means and standard deviations (p-values reported). Significant differences at 10% in boldface.

The logarithmic transformation enables to interpret the first differences as continuous stock market returns. Table 1 provides the means and standard deviations of the return series in the pre- and post- accession periods. Except for Russia, all means of the returns are higher in the post-accession period. Nevertheless, none of the changes is statistically significant. Regarding the standard deviations, the volatility of all the markets is significantly lower at 10 % in the post-accession period that indicates more stable situation on all markets in this period.

Table 2 displays correlations between the returns again in both periods. Nearly all correlations are higher in the post-accession period, meaning that the short-run linkages among the markets are stronger after the EU enlargement. Surprisingly, a significant increase occurred especially in the correlations between the CE markets and the Russian market and not between the CE and the West European as expected. Nevertheless, note that the correlations between the Russian and CE markets were on very low levels in the pre-accession period and that these correlations in the post-accession period got to a very similar levels as the correlations between the CE and the West European markets.

Correlation of returns	Pre-accession period 10/30/1998 - 11/2/2001						Post-accession period 4/30/2004 - 5/4/2007					
market	Czech	Hun.	Pol.	Rus.	West E.	US	Czech	Hun.	Pol.	Rus.	West E.	US
Czech R.	1						1					
Hungary	0.54	1					0.59	1				
Poland	0.49	**0.52**	1				0.58	**0.72**	1			
Russia	**0.23**	**0.31**	**0.27**	1			**0.52**	**0.55**	**0.48**	1		
West Europe	0.40	0.54	0.47	0.34	1		0.50	0.51	0.57	0.44	1	
US	**0.23**	0.39	0.38	0.31	0.74	1	**0.43**	0.49	0.49	0.36	0.76	1

Table 2: Correlations of the continuous (log) return series in the pre- and post- accession periods. Significantly different correlations at 10 % in boldface.

The level time-series data $\boldsymbol{X_t}$ in our setting are required to be at most I(1) which is equivalent to stationarity requirement of the returns (data in first differences $\Delta\boldsymbol{X_t}$). Hence, several specifications of an augmented Dickey Fuller test[3] were used on the returns data in both periods. Since the null hypothesis of a unit root was always rejected at any usual significance level, the data were found to be suitable, i.e. at most I(1).

4 Results

In this section, two models are introduced in order to compare the long-run relations in the pre- and post-accession periods. The analysis proceeds in several steps for both models. First, the unrestricted VAR model is presented and further adjusted in order to get a satisfactory specification of the model. Second, the cointegration rank is chosen. Third, several hypothesis on the parameters are tested. In addition, the MA representation and the common driving trends are analysed for the model in the pre-accession period.

4.1 Model 1: The pre-accession Period

Unrestricted VAR model

First of all, a basic VAR(3) model as described in Section 2 with an unrestricted constant, a restricted trend and no dummy variable is estimated for the logarithms of all 6 ($= p$) investigated stock market indexes using data for the pre-accession period. The theoretical specification in Equation 1 requests two assumptions concerning the residuals ε_t: independence and normality. A necessary condition for the independence assumption is no autocorrelation in any moment which can be tested by Lagrange Multiplier tests (Anderson (2003) or Rao (1973)). The results in Table 3 (Model 1 basic) indicate that the assumption of no autocorrelation is violated for lag 2[4]. The multivariate test for the normality of residuals proposed in Hansen and Doornik (1994) is rejected, too. Similarly, the tests for ARCH effects reject the null hypothesis of no autocorrelation in second moments and detect heteroscedasticity in the residuals. Moreover, several large standardised residuals (over 3.5) could be detected. Since both residual assumptions are violated, the specification of our model is not satisfactory at this stage and the situation also does not improve by inclusion of additional lags.

Working with financial data, we cannot expect to entirely get rid of the heavy tailed (non-normal) distribution as well as the strong ARCH effects. This is not a crucial obstruction, since the estimates of the VAR model are generally robust to the deviations from normality (see Juselius (2007)) and presence of the ARCH effects as long as the higher order moments (2nd and 4th moments) exists (see Gonzalo (1994), Lee and Tse (1996)). However,

[3] ADF test with constant, with constant and trend, both for lag 1 to 3

[4] The usual 5% significance level is used for all conducted tests, when not stated differently.

Test	DF	Model 1 basic	Model 1	Model 2
Autocorrelation:				
LM(1):	36	44.784 [0.150]	34.139 [0.557]	35.639 [0.486]
LM(2):	36	56.338 [0.017]	34.255 [0.552]	35.988 [0.469]
LM(3):	36	41.130 [0.256]	33.702 [0.578]	34.002 [0.564]
Normality:	12	62.343 [0.000]	17.154 [0.144]	13.960 [0.303]
ARCH effects:				
LM(1):	441	545.652 [0.000]	516.100 [0.008]	364.795 [0.997]
LM(2):	882	1038.643 [0.000]	943.420 [0.074]	905.753 [0.282]
LM(3):	1323	1519.055 [0.000]	1479.320 [0.002]	1319.228 [0.524]

Table 3: Results of the misspecification tests for autocorrelation (1st to 3rd lag), multivariate normality and ARCH effects (1st to 3rd lag) in the residuals. All the test statistics are χ^2 distributed. DF denotes degrees of freedom, LM means Lagrange Multiplier. P-values reported in parenthesis.

one can still try to account for large residuals and improve the skewness and kurtosis of the residual distribution by adding appropriate dummy variables. The following dummy variables appeared to be economically relevant and statistically significant in our VECM model:

- January 8 and 15, 1999 (especially significant for the West European and CE markets) - an unrestricted transitory shock dummy[5] related to a stock market overreaction after the introduction of the Economic and Monetary Union (EMU) on January 1, 1999 in most of the EU countries
- April 14, 2000 (negatively significant for the US and the West European markets) - an unrestricted blip (impulse) dummy[6] capturing a temporal drop of the US market on April 14, 2000[7]
- September 14, 21, 28, 2001 (especially significant for the US and West European markets) - three unrestricted blip dummy variables accounting for the market instability after the September 11, 2001 terrorist attacks

The lag length of the model is generally set to be 3, because no autocorrelation in residuals is rejected for a VAR(2) model, but not for a VAR(3) model. However, by examining the coefficients of the Γ_1 and Γ_2 matrixes in the VECM specification[8], it has turned out that some columns contained only insignificant coefficients. Using the notation $\Gamma_i = [\Gamma_{i,lcz}; \Gamma_{i,lhn}; \ldots; \Gamma_{i,lsp}]$, the deletion of the columns $\Gamma_{1,lru}, \Gamma_{2,lst}, \Gamma_{2,lsp}$ was not rejected while conducting consecutive likelihood ratio tests for the suggested submodels. Also the joint hypothesis

$$H_0 : \Gamma_{1,lru} = \Gamma_{2,lst} = \Gamma_{2,lsp} = 0$$

with the likelihood ratio test statistic $LR = 2(3333.778 - 3322.686) = 22.184 < \chi^2_{0.95}(18) = 28.879$ was not rejected.

The misspecification tests for the resulting model (i.e. model with the listed dummies and the specific lag length structure) reported again in Table 3 (Model 1) show that both no autocorrelation and normality of the residuals have improved substantially after the inclusion of the dummy variables, since they are not rejected this time. Hence, the extended model is preferred to the basic one and considered in the following analysis.

Cointegration rank

The rank of Π matrix (i.e. the cointegration rank, denoted by r hereafter) in Equation 2 indicates the number of cointegration relations that can be interpreted as the long-run equilibrium relations. It can be determined by the

[5] $\boldsymbol{d}_{tr} = (\ldots, 0, 1, -1, 0, \ldots)$, 1 and -1 corresponds to January 8 and 15, respectively. For more details on the dummy variables see Juselius (2007)

[6] $\boldsymbol{d}_p = (\ldots, 0, 1, 0, \ldots)$, 1 corresponds to April 14

[7] Nasdaq and Dow Jones indexes logged their biggest single-day losses ever. Possible causes could have been too fast growth of economy and stock markets in the prior months, the announcement of the higher core rate of inflation and heavy sales of Microsoft shares after April 3, 2000 when Judge Jackson ruled that Microsoft violated the Sherman Antitrust Act.

[8] Note that the lag length of 3 in the VAR form corresponds to the lag length of 2 in the VECM form.

trace test introduced in Johansen (1991). Since the model contains a trend in the cointegration relation and several dummy variables, an asymptotic trace test distribution was simulated by the program developed in Nielsen (2004) using length of random walks 158 (the same as the length of the sample) and 5000 replications. The results reported in Table 4 suggest rank 1, because $H_0 : r = 0$ is rejected, but $H_0 : r = 1$ cannot be rejected. The graphical analysis

p-r	r	Eig.Value	Trace	Trace*	Frac95	P-Value	P-Value*
6	0	0.273	121.742	111.210	108.699	0.005	0.032
5	1	0.153	72.219	65.454	82.325	0.229	0.469
4	2	0.121	46.465	38.739	59.546	0.401	0.760

Table 4: Model 1: Asymptotic trace test and the eigenvalue roots. Trace* and P-Value* denotes the results of small sample Bartlett correction introduced in Johansen (2002). Frac95 is the 95% quantile from the simulated trace test distribution.

of the first cointegration relation in Figure 2 proposes stationarity as well, contrary to the other relations. Other indicators for the right choice of cointegration rank (e.g. the number of unit roots in the companion matrix) are not further presented because the evidence seems to be sufficient for the conclusion that just one long-run equilibrium relation is found among the indexes. Hence, the cointegration rank is set to 1 in the following.

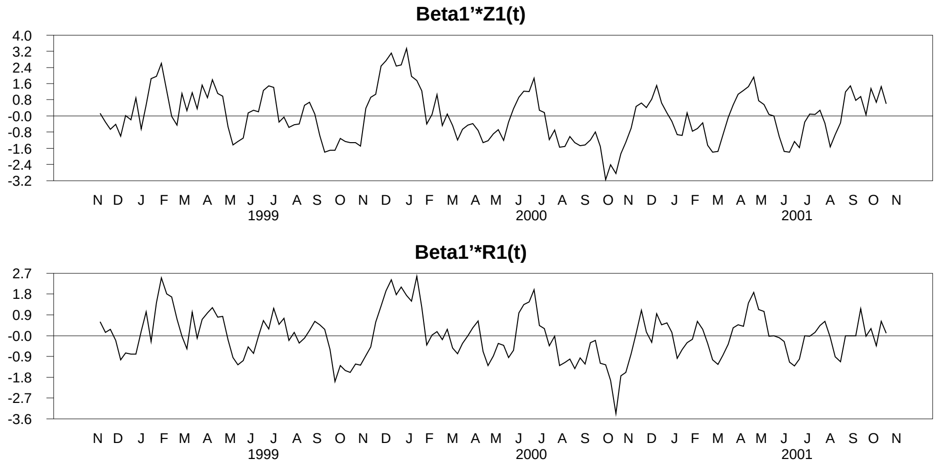

Figure 2: Model 1: First cointegration relation in "X-" and "R-form".

After the rank determination, crucial model assumptions, namely constancy of parameters can be examined by several recursive tests (see Juselius (2007) for the overview of conducted tests). Figure 5 in Appendix displays a stable behaviour of the first eigenvalue corresponding to the stationary relation. The constancy of the eigenvalue seems to be fulfilled as it approximately lies within the narrowest confidence bounds. The constancy of parameters is further supported by the test of β constancy in Figure 6 in Appendix, since the constancy could not be rejected for any subsample. Similarly, constancy of the loglikelihood function captured in Figure 7 in Appendix is not rejected for most of the subsamples. The same recursive tests conducted backwards also did not indicated serious violation of the constant parameters assumption. Since the results of all these tests suggest good stability of the cointegration relation in the examined period, further adjustment of the model (e.g. inclusion of structural breaks in the deterministic trend or constant) is not relevant.

Hypothesis testing

Since the rank was determined to be 1, the long-run structure is already identified. Nevertheless, it is still possible to impose further restrictions on the α and β coefficients and test them. This allows us to investigate several hypotheses, e.g. which markets can be excluded from the cointegration relation or which markets are the adjusting ones. The restrictions should fulfil again the criterion of stability over time, because it can happen that some hypothesis is not rejected for the whole sample though it is rejected for smaller subsamples. The stability of restrictions can be again tested by recursive tests.

First, restrictions on the β coefficients are investigated. We start with several tests of the exclusion from the long-run relation. This helps us to find out whether the cointegration relation involves all the markets simultaneously or only some of them. In particular, we are interested whether the cointegration relation links the CE markets to the other markets such as the West European or not. Test of exclusion in Table 5 clearly indicates that lru, lst, lsp and the time trend can be individually excluded from the cointegration relation. Also the joint hypothesis of exclusion of lru, lst and lsp ($\mathcal{H}_1$)[9] is not rejected. The estimated coefficients are reported in Table 6. Since the coefficient of the time trend is significant in this relation, it is not further excluded. Hence, the results suggest that the cointegration relation involves only the three CE markets while the other world markets can be excluded.

Next, we look more in detail at the characteristics of the relation. The estimated magnitude of the remaining β coefficients suggests also a test for the homogeneity of the Czech, Polish and Hungarian markets $\mathcal{H}_2$[10] that is again not rejected in Table 6 and this result is robust also in shorter subsamples. The found homogeneity of the markets means that the three markets share common driving trends that forces them to move in the same direction and by similar amount[11]. This consequently limits the portfolio diversification. Since the coefficient for the Hungarian market is relatively low, we also test for the stationarity of the spread between the Czech and Polish market[12] and do not reject the test of $\mathcal{H}_3$ in Table 6. In this case $(lcz - lpo) \sim I(0)$ and there are no portfolio diversification benefits in the long-run between the two markets.[13] The reason is that the MA representation shows that the two markets are driven by exactly one common driving trend in the same direction and by the same amount[14]. Nevertheless, considering $\mathcal{H}_3$ implies the exclusion of the Hungarian market from the cointegration relation that is not consistent with the result of the test for individual exclusion of the Hungarian market. Moreover, the p-value of the test for $\mathcal{H}_3$ is quite low compared to the test of $\mathcal{H}_2$. Therefore, the homogeneity restrictions in $\mathcal{H}_2$ are preferred to the spread restrictions in the later analysis.

In addition, we test for the stationarity of individual variables[15] and report the results in Table 5. As expected already from Figure 1 that displays clear non-stationary behaviour of the indices, none of the variable is found to be stationary. Having at least two non-stationary variables in the model justifies relevance of the cointegration method application.

Test	DF	5% C.V.	lcz	lhn	lpo	lru	lst	lsp	trend
exclusion	1	3.841	23.135	4.864	19.674	0.280	0.078	0.021	0.692
			[0.000]	[0.027]	[0.000]	[0.597]	[0.780]	[0.885]	[0.406]
stationarity	5	11.070	38.114	40.850	39.302	39.489	38.710	36.317	
			[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	
weak	1	3.841	21.560	2.263	0.006	5.893	0.773	0.745	
exogeneity			[0.000]	[0.133]	[0.939]	[0.015]	[0.379]	[0.388]	
unit vector	5	11.070	7.816	42.179	27.463	34.407	35.365	35.650	
in α			[0.167]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	

Table 5: Model 1: Likelihood ratio tests of different restrictions on β and α under the rank 1 assumption. DF denotes the degree of freedoms, p-values in the parenthesis.

Second, the restrictions on α coefficients are analysed. Tests of the exogeneity of the variables help us to find out, which of the markets are adjusting to the long run relation and which follows only their stochastic trends.

[9] $\mathcal{H}_1 : \beta_{1,lru} = \beta_{1,lst} = \beta_{1,lsp} = 0$

[10] $\mathcal{H}_2 : \beta_{1,lcz} + \beta_{1,lhn} + \beta_{1,lpo} = 0$ & $\beta_{1,lru} = \beta_{1,lst} = \beta_{1,lsp} = 0$

[11] This is related to the fact that a vector of $(1, 1, 1, 0, 0, 0, 0)$ is orthogonal to the found "homogeneous" β_1 vector under $\mathcal{H}_2$. However, the loadings $\tilde{\boldsymbol{\beta}}_\perp$ in the MA representation consist of two parts, not only of $\boldsymbol{\beta}_\perp$, but also of $(\boldsymbol{\alpha}'_\perp \boldsymbol{\Gamma} \boldsymbol{\beta}_\perp)^{-1}$ that makes the interpretation of the loadings and their exact size difficult.

[12] $\mathcal{H}_3 : \beta_{1,lcz} = \beta_{1,lpo}$ & $\beta_{1,lhn} = \beta_{1,lru} = \beta_{1,lst} = \beta_{1,lsp} = 0$

[13] We tested also stationarity of other spreads, but none of them was found to be stationary.

[14] Note that a vector of $(1, 0, 1, 0, 0, 0, 0)$ is orthogonal to the β_1 under $\mathcal{H}_3$.

[15] $H_0 : \beta_{1,i} = 0 \ \forall i$ except one

Test	lcz	lhn	lpo	lru	lst	lsp	trend	$\chi^2(\nu)$	p-value
$\mathcal{H}_1$	1	-0.2320	-0.7443	0	0	0	-0.0003	0.9490 (3)	0.8136
$\mathcal{H}_2$	1	-0.2247	-0.7753	0	0	0	-0.0003	0.8859 (4)	0.9266
$\mathcal{H}_3$	1	0	-1	0	0	0	0	7.202 (6)	0.303

Table 6: Model 1: Testing stationarity of restrictions on β under the rank 1 assumption. The estimated coefficients and results of likelihood ratio tests.

We expect that the CE and the Russian market are more likely to be the adjusting ones due to their rather small size compared to the US and the West European market. Test of individual weak exogeneity[16] detects only two possible adjusting markets to the long-run relation, namely, the Czech and the Russian market (see again Table 5). All the other markets are found to be weakly exogeneous, meaning that they follow only their own stochastic trends. Furthermore, the tests for a unit vector in α[17] suggest the presence of one unit vector related to a single adjusting market, the Czech market. This finding is not in line with the significance of $\alpha_{lru,1}$ from the individual weak exogeneity test. However, the significance is rather borderline (at 5 %) in smaller subsamples. Considering also the good stability of the unit vector restriction in Czech market in smaller subsamples, it seems reasonable to regard the Czech market to be the exclusively adjusting one. This means that shocks to the Czech market have no permanent effect on any market in the system, event not on itself, and the "random walk" movements of the Czech market are driven by permanent shocks to other markets.

The joint restrictions on α and β coefficients (homogeneity restrictions $\mathcal{H}_3$ on β and a unit vector in α for the Czech market)[18] are not rejected with a high p-value of 0.52. The estimated coefficients after the restrictions are summarised in Table 7. Moreover, Figure 8 in the Appendix shows that the likelihood ratio test of restrictions based on the "R-form" estimates is also not rejected for all of the shorter subsamples. Regarding the "X-form" estimates, the rejection of the restrictions for the shortest subsamples can be caused by the instability of the short-run coefficients (e.g. coefficients of Γ_1 and Γ_2) in small samples and, thus, does not appear to be a serious distortion regarding the results relevant in the long-run.

TEST OF RESTRICTED MODEL:				$\chi^2(9) = 8.143$ [0.52]			
	lcz	lhn	lpo	lru	lst	lsp	TREND
β_1'	**1.000**	**-0.267**	**-0.733**	0	0	0	**-0.0004**
	(.NA)	(-2.975)	(-8.155)	(.NA)	(.NA)	(.NA)	(-2.566)
	Δ lcz	Δ lhn	Δ lpo	Δ lru	Δ lst	Δ lsp	
α_1'	**-0.283**	0	0	0	0	0	
	(-5.746)	0	0	0	0	0	

Table 7: Model 1: Coefficients of the restricted model (t-statistics in the parenthesis)

Having a reasonable model with one cointegration relation for the pre-accession period, the applicability of the model can be tested also for other periods, e.g. for the accession period. A recursive test of $H_0 : \beta =$ "known" β, where "known" β means the estimated β based on the pre-accession period, provides an evidence whether the cointegration relation found in the pre-accession period remained similar also in the following accession period. Figure 3 shows that the cointegration relation persisted for approximately one year in the accession period and later on, around November 2002, the relation changed permanently. This finding is consistent for both the "R-form" and the "X-form", since both test statistics crosses the 5% rejection line around November 2002 and stay above the line permanently. [19]. Hence, no clear support for a dramatic change in the cointegration relation due to the EU enlargement announcement in November 2001 is delivered by this test. Nevertheless, a permanent change occurred later in the accession period and the timing of the change corresponds roughly to the end of the admission negotiations on December 13, 2002.

[16] $H_0 : \alpha_{i,1} = 0$ for some i

[17] $H_0 : \alpha_{i,1} = 0 \ \forall i$ except one

[18] $H_0 : \beta_{1,lcz} = \beta_{1,lhn} + \beta_{1,lpo} \ \ \& \ \ \beta_{1,lru} = \beta_{1,lst} = \beta_{1,lsp} = 0 \ \ \& \ \ \alpha_{i,1} = 0 \ \forall i, i \neq lcz$

[19] The problematic rejection at the beginning of the displayed sample regarding the "X-form" can be again explained by the short-run

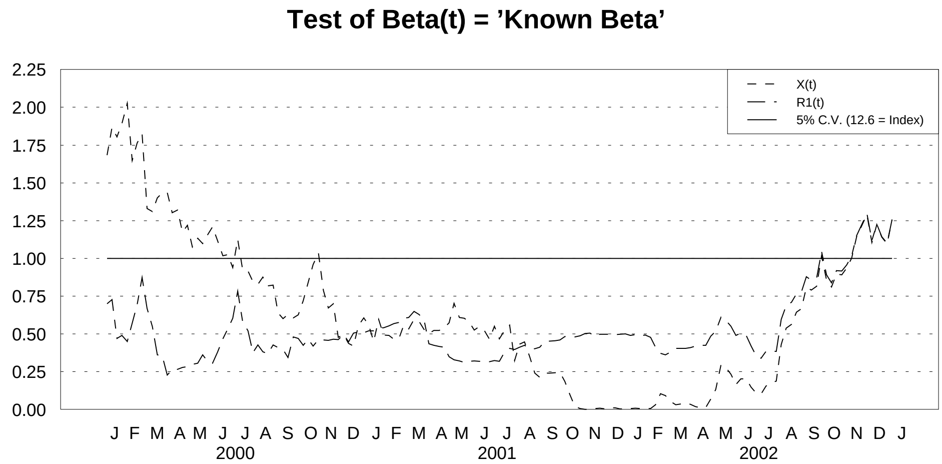

Figure 3: Model 1: Test of β = "known" β, where the "known" β estimates are based on the pre-accession period. The scaling of the test is consistent with 1 being the 5 % rejection line.

Common driving trends

The VAR model was further inverted into its MA representation in Equation 5 keeping the unit vector restrictions on α[20]. In general, every weakly exogeneous variable from the VAR representation translates to one common driving (stochastic) trend in the MA representation. Since the model allows for the same number of weak exogeneous variables as common driving trends (i.e. $p - r = 5$), every common driving trends can be identified as consisting only of the shocks to the corresponding weak exogeneous variable (see $\alpha_\perp$ in Table 8) and the resulting trends can be labelled by the individual markets. The estimates of the loadings $\bar{\beta}_\perp$ to the common trends suggest that the development of the Czech market is driven mainly by the Polish and less by the Hungarian trend, both trends having positive impacts on the Czech market. All the other markets should follow only their own trends, since they were found to be weakly exogeneous. This seems to be a clear case for most of these markets, i.e. for the Polish, Russian, West European and US market. Nevertheless, the significance (t-value over 2) of the $\bar{\beta}_{lhn,ct5\perp}$ leads to the suspicion that the Hungarian market could be partly adjusting to the US common driving trend. Similar situation concerns the Czech market. These results are slightly not in line with the previously imposed restrictions on β. More specifically, given the exclusion of the US market from the β vector, $\bar{\beta}_{ct5\perp}$ should contain a unit vector in the US market, i.e. just one significant coefficient $\bar{\beta}_{lsp,ct5\perp}$. However, the significance of $\bar{\beta}_{lhn,ct5\perp}$ and $\bar{\beta}_{lcz,ct5\perp}$ are rather borderline, especially compared to the high significance of $\bar{\beta}_{lsp,ct5\perp}$. Hence, we could expect that a joint hypothesis of a unit vector in the US market would not be rejected and the results of the VECM and MA representation would be consistent.

A more serious problem arising with the MA representation in our model is that the shocks are not contemporaneously uncorrelated and, therefore, cannot be interpreted as unanticipated, unique and invariant. The attempt to correct for this by the structural MA model estimation where shocks are orthogonalised was not very helpful for further interpretation of the system structure.

Finally, the impulse response analysis was conducted and indicated convergence to the long-run equilibrium after 19 periods that corresponds to 4 - 5 months.

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parameters instability, since the test for "R-form" does not reject the null hypothesis.

[20]The restrictions on β were released for computational reasons

$\alpha'_{\perp}$	lcz	lhn	lpo	lru	lst	lsp
CT 1	0	**1**	0	0	0	0
CT 2	0	0	**1**	0	0	0
CT 3	0	0	0	**1**	0	0
CT 4	0	0	0	0	**1**	0
CT 5	0	0	0	0	0	**1**

$\boldsymbol{\beta}_{\perp}$	CT 1	CT 2	CT 3	CT 4	CT 5
lcz	**0.299**	**1.005**	-0.040	-0.285	**0.222**
	(2.621)	(3.776)	(-0.186)	(-1.803)	(2.102)
lhn	**0.797**	0.445	-0.061	-0.201	**0.264**
	(7.258)	(1.737)	(-0.291)	(-1.320)	(2.595)
lpo	0.118	**1.209**	-0.033	-0.316	0.207
	(0.956)	(4.221)	(-0.141)	(-1.855)	(1.819)
lru	0.041	-0.039	**0.962**	0.274	0.216
	(0.191)	(-0.078)	(2.338)	(0.915)	(1.078)
lst	0.015	0.013	-0.037	**0.855**	0.055
	(0.270)	(0.098)	(-0.353)	(11.249)	(1.079)
lsp	0.016	-0.131	-0.016	0.126	**0.780**
	(0.292)	(-1.030)	(-0.150)	(1.665)	(15.381)

Table 8: Model 1: Common driving trends $\alpha'_{\perp}$ (CT) and their loadings $\bar{\beta}_{\perp}$ (t-statistics in parenthesis)

4.2 Model 2: The post-accession Period

Unrestricted VAR model

Similarly to Model 1, the residuals in the basic model for the post-accession period suffered from autocorrelation, non-normality and heteroscedasticity. Most of the difficulties were solved by the inclusion of the following unrestricted blip dummy variables that accounted for the largest residuals in the model:

- March 18, October 14, November 11, 2005 - correcting for temporal instability in Czech and Hungarian markets
- May 19, June 9, 16, 30, 2006 - correcting for the instability mostly in the CE and Russian markets[21]
- March 2, 2007 - a global downturn of the markets following a sharp fall of China's domestic stock markets

Having included the above dummies in the model, the lag length of the VAR model was set to 1. Note that this implies zero lag length in the VECM form and, consequently, the $\boldsymbol{\alpha\beta}'$ estimates in the "X-" and "R-form" are the same. The reasons for the choice of the lag length 1 were twofold. First, the misspecification tests in Table 3 do not detect any residual autocorrelation, non-normality or even any presence of ARCH effects. Second, the likelihood ratio tests for reducing lag length from 3 to 1 ($\chi^2(72) = 82.947$ with p-value 0.178) and from 2 to 1 ($\chi^2(36) = 39.977$ with p-value 0.298) did not reject the hypothesis of the submodel with lag length 1. Since the specification of the model appears to be very good, it is considered in the following.

Cointegration rank

Again we investigate several indicators available for the right choice of the cointegration rank r. First, the simulated values for the trace test in Table 9[22] indicate rank 2 at 5 % and rank 3 at 10 % confidence level. Moreover, considering a very small difference between the magnitude of the second (0.170) and the third eigenvalue root (0.157), it seems reasonable to prefer rank 3 to rank 2. Second, Figure 4 suggests stationarity of the third long-run

[21] In May and June 2006, we observe high volatility of the markets, especially emerging. This might be due to a healthy market correction after almost 3 years of big gains. The IMF Financial market update, June 2006, writes: "Since the release of the April 2006 Global Financial Stability Report, global financial markets have experienced increased volatility and a sharp correction in the price of riskier assets."

[22] Simulation was the same as for Model 1. Hence, the simulated values differ only negligibly.

p-r	r	Eig.Value	Trace	Trace*	Frac95	P-Value	P-Value*
6	0	0.216	125.934	123.508	108.136	0.002	0.003
5	1	0.170	87.700	86.324	81.365	0.016	0.021
4	2	0.157	58.468	57.756	59.579	0.058	0.067
3	3	0.101	31.737	31.461	40.450	0.272	0.285

Table 9: Model 2: Asymptotic trace test and the eigenvalue roots. Trace* and P-Value* denotes the results of small sample Bartlett correction introduced in Johansen (2002). Frac95 is the 95% quantile from the simulated trace test distribution.

equilibrium relation. Similarly, the magnitude (0.762) of the fourth largest root in the companion matrix under rank 3 assumption does not indicate a close unit root and, thus, support stationarity of the third cointegration relation. Hence, rank 3 appears to be the right choice for the following analysis.

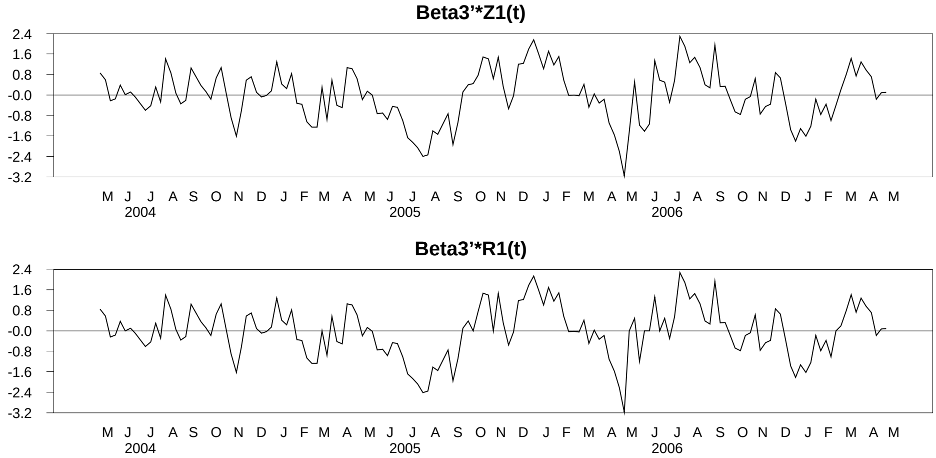

Figure 4: Model 2: Third cointegration relation

Several recursive tests are again conducted in order to check the assumption of constant parameters. The tests for the constancy of the loglikelihood function and of the β parameters (see Figures 11 and 10 in Appendix) do not indicate violation of the constancy assumption. However, the eigenvalue fluctuation tests in Figure 9 in Appendix clearly detect nonconstancy in the two largest eigenvalues corresponding to the first two stationary relations. Since the eigenvalues are linear functions of the corresponding α and β parameters and the constancy of β parameters does not seem to be violated, the rejection occurs highly likely due to the nonconstant α parameters. Consequently, the following analysis concentrates on the examination of the β coefficients, since the estimates of the α's unlike the estimated β-parameters cannot be considered as reliable.

Hypothesis testing and identification of the long-run structure

Compared to the pre-accession period, the higher number of stationary relations suggests stronger linkages between the six markets in general, though it is not clear at this stage whether this happens due to tighter relation between the CE and West European markets. In order to learn more about this, the long-run structure need to be identified. The identification can be achieved by a pure rotation of vectors in $\boldsymbol{\beta}$. However, such identification is relevant only for its statistical purpose and does not provide further economical insights about the long-run relation structure. Therefore, several hypotheses on the β coefficients are tested that allow us to distinguish the three cointegration relations and label them. We are interested in particular, whether new relations between the CE and the West European markets emerged, whether the linkages among the CE markets strengthened or not and which role is played by other markets, the US and the Russian markets.

We start again with the test for exclusion of an individual β coefficient. The results in Table 10 show that no variable apart from the trend can be excluded from all three cointegration relations simultaneously. In addition, we check whether some of the variable is stationary on itself. The test for stationarity of a single variable in Table 10 show that it is not the case, since stationarity is always rejected.

Test	DF	5% C.V.	lcz	lhn	lpo	lru	lst	lsp	trend
exclusion	3	7.815	10.807	9.326	17.186	14.440	14.931	16.675	4.908
			[0.013]	[0.025]	[0.001]	[0.002]	[0.002]	[0.001]	[0.179]
stationarity	3	7.815	18.653	17.083	11.602	18.604	8.612	10.298	
			[0.000]	[0.001]	[0.009]	[0.000]	[0.035]	[0.016]	

Table 10: Model 2: Likelihood ratio tests of different restrictions on β under the rank 3 assumption. DF denotes the degree of freedoms, p-values in the parenthesis.

A further interesting hypothesis arises in relation to the already obtained results in the pre-accession period, namely, whether the same or a similar cointegration relation can be found among the CE markets also in the post-accession period. Therefore, the test for exclusion of all the world markets from a single relation $\mathcal{H}_1$[23] is conducted and not rejected as shown in Table 11. The homogeneity of the CE markets $\mathcal{H}_2$[24] is not rejected as well. This result indicates that the three CE markets still share a common driving trends that limit the diversification possibilities. However, note that the estimated coefficients in $\mathcal{H}_2$ have substantially changed from the previous estimates in $\mathcal{H}_1$ and the p-value decreased and, thus, the homogeneity restrictions appear to be too restrictive. The change in the magnitude of the coefficients was already indicated by Figure 3, when the estimated β from the pre-accession period did not remain constant in the accession period. Consequently, the magnitudes of the estimated coefficients in the post-accession period are very different from those estimated in the pre-accession period (compare results of $\mathcal{H}_2$ in Table 11 and in Table 6). Particularly, the magnitude of $\beta_{1,lhn}$ has increased at the expense of $\beta_{1,lpo}$.

Consequently, the low estimate of $\beta_{1,lpo}$ further suggests to test also for the stationarity of the spread of the Czech and Hungarian markets. When a time trend is included, the spread stationarity is not rejected (see $\mathcal{H}_3$ in Table 11), though the p-value of the test further decreased. This result means, on the one side, that the long-run diversification benefits between the Czech and Hungarian markets are very limited in the post-accession period, since $(lcz - lhn - 0.001t) \sim I(0)$ and the markets share a common driving trend. On the other side, the diversification possibilities between the Czech and Polish market has increased compared to the pre-accession period, because their spread is not found to be stationary anymore. Regarding other markets, stationarity of their spreads is always rejected meaning that no other markets are so strongly linked as the Czech and Hungarian market in the post-accession period.

Furthermore, we investigate whether the initial cointegration relation among the CE markets ($\mathcal{H}_1$) could be a linear combination of two "smaller" separate stationary relations containing less than the three CE markets. However, this is not the case, since no other linear combination involving only the pair of the Czech and Polish or the pair of Polish and Hungarian market is found to be stationary. This means that the cointegration relation under $\mathcal{H}_1$ is irreducible and that both of the two remaining cointegration relations involve also (but not necessarily only) the remaining markets. We prefer the initial cointegration relation under $\mathcal{H}_1$ to the homogeneity ($\mathcal{H}_2$) and spread ($\mathcal{H}_3$) restricted relations due to the highest p-value in the following. We regard this relation as one identified cointegration relation in our model and label it as "CE markets relation". To sum up the findings regarding the CE markets, the number of cointegration relations has remained the same after the EU enlargement. In particular, no new relation could be indicated. Nevertheless, the characteristics of the relation have changed. More specifically, the evidence suggests a shift from the strong linkage between the Czech and Polish market towards a stronger tight between the Czech and Hungarian market.

Now we turn to analyse the remaining two cointegration relations. First, we test whether a stationary relation can be detected among the remaining (i.e. non-CE) markets. The result of testing $\mathcal{H}_4$[25] in Table 11 indicates no cointegration related only to the three non-CE markets. Therefore, the two remaining cointegration relations bridge the two groups of markets, the non-CE and the CE markets, which is a novelty arising in the post-accession period.

[23] $\mathcal{H}_1 : \beta_{1,lru} = \beta_{1,lst} = \beta_{1,lsp} = 0$. Note that the result of the test is invariant for the relation chosen.

[24] $\mathcal{H}_2 : \beta_{1,lcz} + \beta_{1,lhn} + \beta_{1,lcz} = 0 \;\; \& \;\; \beta_{1,lru} = \beta_{1,lst} = \beta_{1,lsp} = 0$

[25] $\mathcal{H}_4 : \beta_{1,lcz} = \beta_{1,lhn} = \beta_{1,lpo} = 0$

A plausible economic prior for further identification can be the expectation that the CE markets are more strongly linked to the West European markets after the EU accession. Therefore, we look for a stationary relation capturing the relationship between these in the following.

Test	lcz	lhn	lpo	lru	lst	lsp	trend	$\chi^2(\nu)$	p-value
$\mathcal{H}_1$	1	-0.7800	-0.5745	0	0	0	0.0007	0.5906 (1)	0.442
$\mathcal{H}_2$	1	-0.8144	-0.1856	0	0	0	-0.0012	2.254 (2)	0.324
$\mathcal{H}_3$	1	-1	0	0	0	0	-0.0014	5.020(3)	0.170
$\mathcal{H}_4$	0	0	0	-0.0839	1.0000	-0.8167	-0.0014	5.0127 (1)	0.0252
$\mathcal{H}_5$	-1.0525	0.8654	0	0	1	0	-0.0012	0.0102(1)	0.9197
$\mathcal{H}_6$	-1	0.8206	0	0	1	0	-0.0013	0.0102(2)	0.9949
$\mathcal{H}_7$	-1	1	0	0	1	0	-0.0022	3.352(3)	0.3402

Table 11: Model 2: Testing restrictions on β under the rank 3 assumption. The estimated coefficients and results of likelihood ratio tests. ν denotes degree of freedoms.

One convenient candidate for such a relation turned out to consist of the Czech, Hungarian and West European market ($\mathcal{H}_5$) because of the very high p-value.[26] Moreover, the p-value increased by additional restriction $\beta_{2,lcz} = -\beta_{2,lst}$ ($\mathcal{H}_6$). Since the coefficient for $\beta_{2,lhn}$ is also close to unity, we test for the homogeneity in the Czech and Hungarian markets in $\mathcal{H}_7$[27]. Although $\mathcal{H}_7$ is not rejected, the p-value decreased substantially by the additional restriction and, thus, we prefer the relation in $\mathcal{H}_6$. Since the joint hypothesis for $\mathcal{H}_6$ and the "CE markets" relation ($\mathcal{H}_1$) is also not rejected ($\chi^2(3) = 2.256$ with p-value 0.521), the relation under $\mathcal{H}_6$ is included in the cointegration space and labelled as the "new EU relation". It captures a new linkage between the CE and the West European markets that emerged in the post-accession peroad and could not be detected in the pre-accession period.

Because of no economic prior for the third relation, it is, first, just identified only by the exclusion of the Czech and the Hungarian market and links the Polish market with the remaining markets. The resulting estimate of the whole $\boldsymbol{\beta}$ matrix is reported in Table 12. Although the third cointegration relation contains also the Russian market and the market cannot be excluded according to the t-test, the estimated coefficient of $\beta_{3,lru}$ is relatively small. On the contrary, the coefficients for the other markets, the Polish, West European and the US market are substantially larger and indicate that the third relation is mainly capturing the linkages between these markets. Note that just-identification of the third cointegration relation could be achieved also by, for instance, the exclusion of the Czech and Polish or Polish and Hungarian market. Consequently, the estimated coefficients of the relation would be changed. However, the relatively small size of the Russian market is found also under the other identification schemes. This suggests a rather minor importance of the Russian market in the whole long-run scheme. Even in the case that the Russian market represented an important driving force in the model, the deviations from the long-run equilibrium relations due to the Russian market would not be very large.

TEST OF RESTRICTED MODEL: $\chi^2(3) = 2.2721$ [0.5179]

	lcz	lhn	lpo	lru	lst	lsp	TREND
β_1'	**1**	**-0.771**	**-0.628**	0	0	0	**0.001**
	(.NA)	(-13.496)	(-7.900)	(.NA)	(.NA)	(.NA)	(1.999)
β_2'	**-1**	**0.820**	0	0	**1**	0	**-0.001**
	(.NA)	(13.943)	(.NA)	(.NA)	(.NA)	(.NA)	(-4.035)
β_3'	0	0	**-0.364**	**-0.049**	**1**	**-0.349**	**-0.001**
	(.NA)	(.NA)	(-5.608)	(-5.116)	(.NA)	(-7.899)	(-1.541)

Table 12: Model 2: Coefficients of the restricted model (t-statistics in the parenthesis)

We can again test the stability of the imposed over-identifying restrictions recursively by likelihood ratio test. The joint restrictions as imposed in Table 12 are not rejected for most of the subsamples as can be seen in Figure

[26] Note that the exclusion of lru and lsp from one relation leads to a just identified relation in our model achieved by rotation, but not to testable over-identifying restrictions.

[27] $\mathcal{H}_7 : \beta_{2,lcz} = -\beta_{2,lhn} = -\beta_{2,lst}$ & $\beta_{2,lpo} = \beta_{2,lru} = \beta_{2,lsp} = 0$

12 in Appendix. The only problematic period appears in August 2005, when the restrictions are rejected, though only borderline and temporarily. Since the low t-value of the trend in the third cointegration relation suggests an insignificant coefficient, a model without this coefficient was also estimated. However, the situation regarding the recursive tests of the restrictions worsened.

The estimated α coefficients are not reported and further interpreted because of the serious non-constancy detected. For the same reason, also the tests on the α coefficients and the MA representation are left out. Nevertheless, we report at least the results of the impulse response analysis. It suggests convergence to the long-run equilibrium in 23 steps corresponding to five months. The adjustment speed to the long-run equilibrium is found to be a little slower in the post-accession period than in the pre-accession period (19 weeks). However, the post-accession period contains two more equilibrium relations and therefore it is not surprising that the time to achieve the equilibrium is longer.

4.3 Robustness Check

Several different robustness checks were conducted in order to show to what degree the generalisation of the results is possible. First, different model specifications are examined. Second robustness check covers than the issue of the data frequency chosen. Finally, the use of local currency compared to the dollar terms is discussed.

Model specification

The models for the pre- and post-accession period were adjusted by the inclusion of several dummies in order to better satisfy the assumptions for the Johansen cointegration method. Similarly, an appropriate lag length was chosen. Nevertheless, in this part, different model specifications constructed without the focus on the requested assumptions are investigated to show the robustness of the results.

First, look at the alternative models for the pre-accession period. Considering Model 1 with "full" lag length of 3 in the VAR form (e.g. with $\Gamma_{1,lru}, \Gamma_{2,lst}, \Gamma_{2,lsp}$ included), the obtained estimates vary only negligibly and all conducted tests of coefficient restrictions deliver the same results. Models with lag lengths of 2 and 4 detect also rank 1. The estimation of Model 1 without any dummy variable suggests again rank 1, though the support for preferring rank 1 to rank 0 is rather weak based on the trace test results. The situation improves substantially by the inclusion of the three dummies related to the 11th September. Under the rank 1 assumption, the tests of restrictions on α and β differs a little for the latter models[28] by suggesting a stronger adjustment of the Russian market (i.e. no weak exogeneity) and possible exclusion of the Hungarian market from the cointegration relation. However, these possibilities were already indicated and discussed also in relation to the original model.

Second, alternative model specifications are examined for the post-accession period. Again more lags are added or dummies are eliminated. Here we find a stronger support for preferring rank 2 to rank 3 according to trace test. Both relations bridge the CE and the non-CE markets, but the "pure CE relation" is not found to be stationary anymore. Hence, the main result that a new linkages between the CE and the other markets emerge in the post-accession period stays unchanged. On the contrary, the degree of cointegration between the CE markets seems to decrease under some of the alternative model specifications. However, most of these specifications violate the residual or constancy parameter assumptions and, hence, deliver further evidence about the importance of the assumptions for statistical inference.

Data frequency

Throughout the analysis, the data frequency chosen is weekly though the daily data is easily available. The reason for this choice is the fact that the lower frequency stock market data suffer less from the stylized facts of financial time-series (heavy tailed and skewed distribution, ARCH effects). The elimination of the stylized facts is crucial for an appropriate application of the Johansen cointegration method since the method is based on the assumption of independent and normally distributed residuals. In addition, using weekly frequency eliminates the impact of different closing times of the markets on the results. Nevertheless, the interesting question is whether a lot of important information is lost due to the lower frequency and what the impact on the estimates is. Therefore, an alternative weekly data is considered, namely, the Tuesday closing prices instead of the Friday closing prices. As

[28] Models with lag 2, lag 4, no dummies, dummies for 11th September

expected, models based on the alternative data need to be adjusted especially for the appropriate dummy variables inclusion accounting for the short-run extraordinary deviations. The most obvious example is the 11th September 2001. Since the 11th September 2001 was Tuesday, the dummy used in Friday data is shifted 3 days ahead and the magnitude of the corresponding dummy coefficient for the US market turn out to be much higher.

The results show that the cointegration rank and the identified long-run relations are robust to the day of the week used in both pre- and post-accession period, since the same restrictions on the β coefficients are not rejected and the magnitude of the estimated coefficients remains similar. Nevertheless, certain differences in the results appear regarding the estimated α coefficients. Considering the pre-accession period, unit vector in lcz is still not rejected at 5% level, but the tests for individual weak exogeneity suggest adjustment of the Hungarian market to the long-run relation in addition to the Russian and Czech market. The α estimates in the post-accession period appear to be similarly unstable as for the Friday data and, hence, are regarded also as unreliable.

Currency used

The data in US dollar are used for the last robustness check. Our analysis confirms the finding in Yang et al. (2006), Koch and Koch (1991) and Bessler and Yang (2003) that the results do not depend substantially on using the local currency or dollar terms.

5 Conclusions

In order to investigate the impact of the EU enlargement process before May 2004 on the long-run stock market linkages, this study develops for each, the pre- and the post-accession period, one cointegrated VAR model involving three largest CE (Czech, Hungarian, Polish), the Russian and major world (West European and US) stock markets for two different periods. The main result is that the stock market linkages between the CE markets and the other markets were strengthened after the EU enlargement, though the linkages among the three CE markets remained on similar level. In particular, a new "EU relation" between the CE and West European markets could be identified in the post-accession period. This leads to the conclusion that the potential long-run portfolio diversification benefits across the CE and the other markets were reduced after the EU enlargement.

Since the EU accession is not a sudden event but a gradual process, the first investigated period (pre-accession period) covers 3 years before the EU enlargement announcement by European Commission in November 2001. The second (post-accession) period includes again 3 years, starting on 30th April 2004, 1 day before the official EU accession of the Czech Republic, Hungary and Poland on 1st May 2004. Due to the same time span, the two presented models assigned as Model 1 and Model 2 are well-comparable.

Concerning the long-run equilibrium relations, Model 1 detected only 1 cointegration relation. Moreover, this relation turned out to involve only the three CE markets. All major world markets were jointly excludable from this relation. On the other hand, three equilibrium relations were found in Model 2. One of the relation could have been identified as the pure "CE market" relation and the degree of cointegration among the CE markets appeared to be very similar as in Model 1. The two remaining relations were both bridging the two groups of markets, the CE and the other markets, and thus delivering the evidence of newly formed relations linking the CE markets to the West European, US and Russian market. In particular, the West European stock market index was strongly involved in both relations, showing that the linkages between the old and the new EU members had significantly increased after the EU enlargement. Moreover, one of the relations could be interpreted as the new "EU" relation. The link to the US market appeared to be less pronounced and the Russian market was involved at least.

Comparing the extend to which the assumptions of the two models were satisfied, both models after the inclusion of appropriate dummy variables passed most of the misspecification tests for the residual assumptions. Furthermore, the estimated coefficients in Model 1 and the β coefficients in Model 2 featured good constancy. However, severe problems in parameter nonconstancy were detected in Model 2 regarding the α coefficients and alternative specifications of the model did not help to solve the problem. One possible conclusion is that the driving forces of the markets in the post-accession period are still quite unstable compared with the pre-accession period. In addition, the available time-span of 3 years is rather short for reliable α estimates that converge to their true values slower that the β coefficients. Therefore, the interpretation and tests regarding the unreliable α estimates were left out in Model 2 and the characteristics of the adjustment to the long-run relations cannot be compared for the two periods.

The results of the impulse response analysis for the two models indicated similar adjustment speed to the equilibrium relations. The adjustment in Model 1 was a little faster, since 19 weeks (4 to 5 months) were needed for the convergence to the long-run equilibrium relation. However, 23 weeks (5 to 6 months) in Model 2 is a very similar result, considering the fact that the convergence in this case is achieved for three instead of only one equilibrium relations. On the other side, an interesting difference in the two models can be found in the used lag length. In model 1, the necessary lag length for a good model specification is 3, at least for all CE markets. However, model 2 requires only lag length 1. This indicates less persistence and faster adjustment in the short-run in the post-accession period for the CE markets.

To sum up, this study contributes to the existing literature on long-run stock market linkages with further evidence on structural changes. The focus is on the changes associated with the EU accession of the three CE countries in May 2004. It is shown that the long-run linkages between the CE markets and the West European markets as well as the US and Russian market were strengthened after the EU accession and, thus, the long-run portfolio diversification benefits across the markets were reduced.

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A Appendix

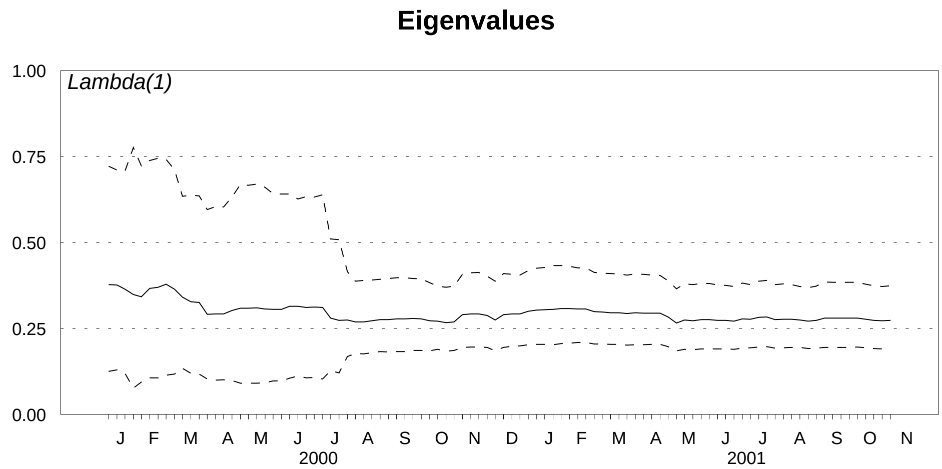

Figure 5: Model 1: Development of the first eigenvalue. Dashed line displays 5 % confidence bounds.

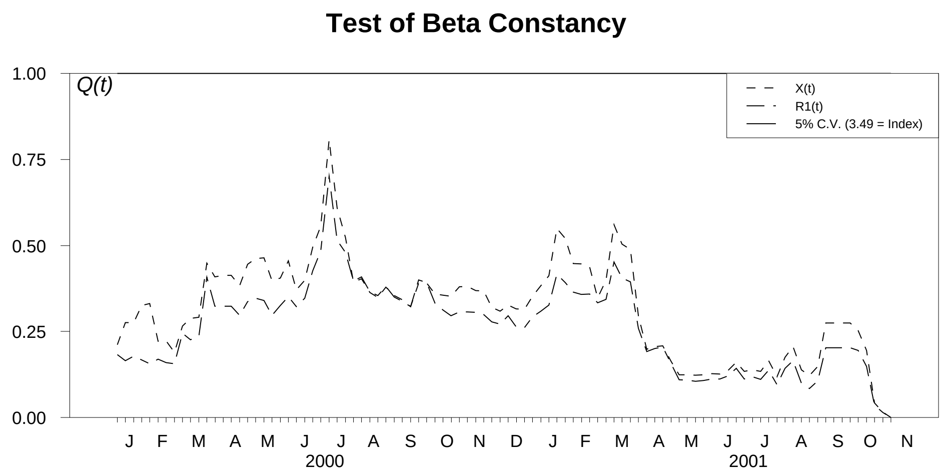

Figure 6: Model 1: Test of β constancy. The scaling of the test is consistent with 1 being the 5 % rejection line

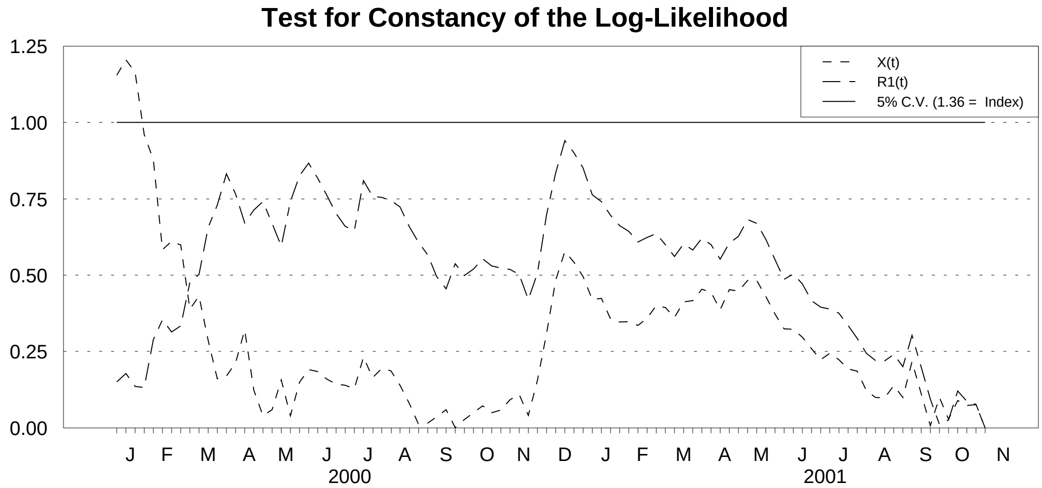

Figure 7: Model 1: Test for constancy of loglikelihood function. The scaling of the test is consistent with 1 being the 5 % rejection line.

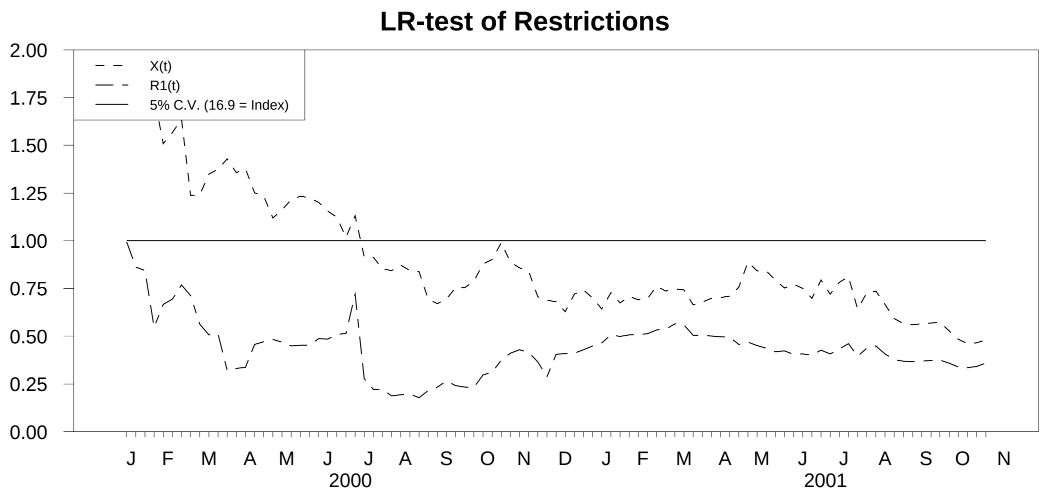

Figure 8: Model 1: Likelihood ratio test of restrictions. The scaling of the test is consistent with 1 being the 5 % rejection line.

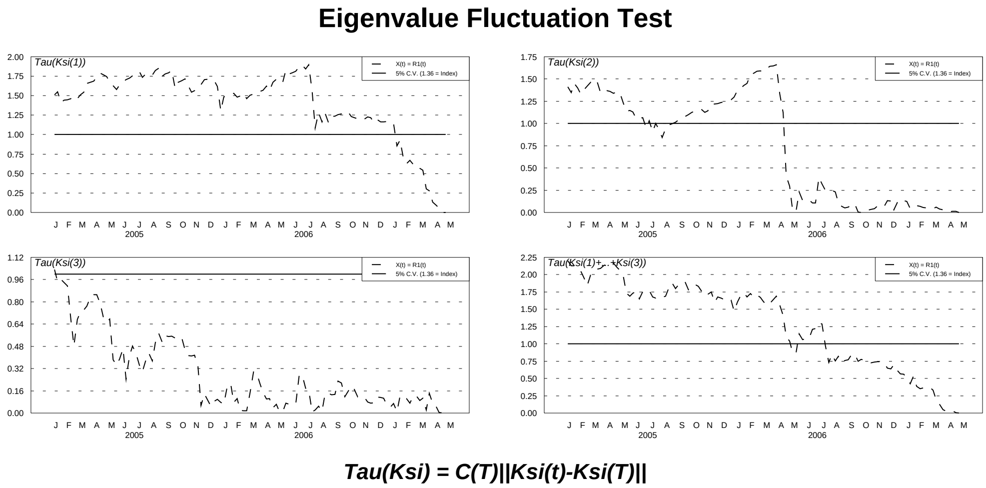

Figure 9: Model 2: Fluctuation test for the three largest eigenvalues. The scaling of the test is consistent with 1 being the 5 % rejection line.

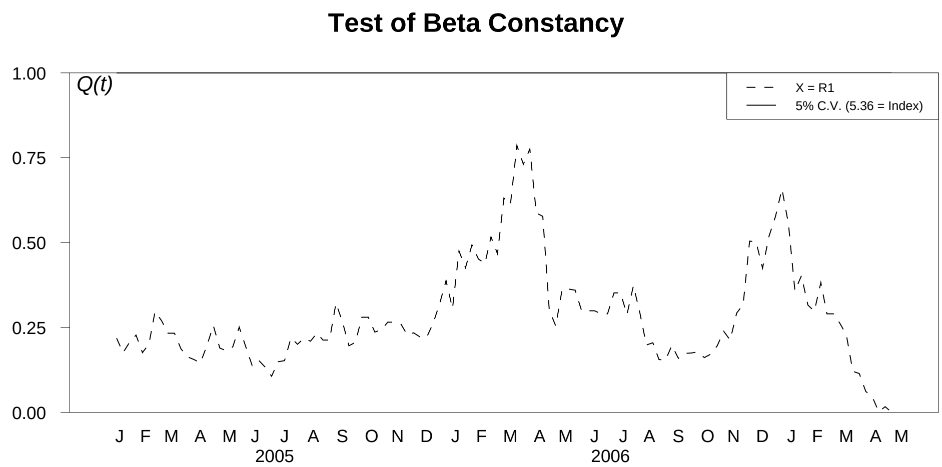

Figure 10: Model 2: Test of β constancy. The scaling of the test is consistent with 1 being the 5 % rejection line

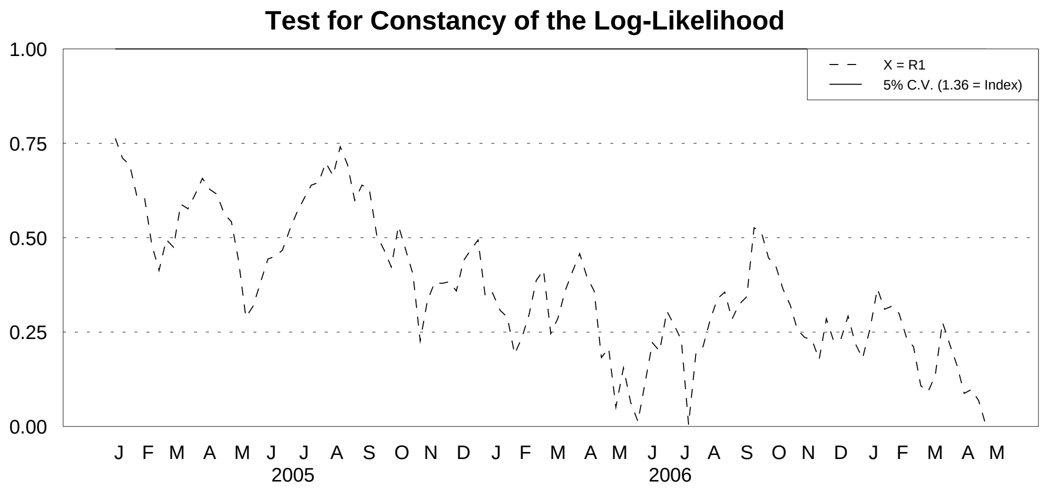

Figure 11: Model 2: Test for constancy of loglikelihood function. The scaling of the test is consistent with 1 being the 5 % rejection line.

Figure 12: Model 2: Likelihood ratio test of restrictions. The scaling of the test is consistent with 1 being the 5 % rejection line.