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Are the Central European Stock Markets Still Different? A Cointegration Analysis[☆]

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Abstract

The Central European countries became members of the European Union (EU) in May 2004. Has their accession into the EU also resulted in a stronger financial integration with the global economy in general and with the "old" EU countries in particular? Based on a cointegration analysis applied to stock market movements, I detect for the period after the EU enlargement two new long-run equilibrium relations that indeed suggest a stronger inter-dependence of the markets, whereas no such relations can be observed before this date. In particular, one new relation links the Central European markets to the Western European market, reflecting tighter co-movements of the "new" and the "old" EU markets. The second relation points at the role of the US market for both the Central and the Western European markets.

Keywords: Transition Economies, Emerging stock markets, Central Europe, European integration, Cointegration, Long-run stock market linkages.

JEL classification: C5, F36, G11, G15.

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1. Introduction

The transition economies in Central Europe[1] offer a unique opportunity to study the effects of institutional changes on financial integration. First of all, many substantial legal and institutional reforms have been occurring in the Central European countries since the fall of communism in 1989. The countries quickly embarked on programs of liberalisation and privatisation, and their economies have undergone a relatively fast and successful transition process from centrally planned economies towards free markets. Second, an important role in the transition process has been played by large capital inflows to the region from developed countries. The Central European countries have attracted particularly significant amounts of foreign direct investment (FDI), originating mainly from Western Europe (Mora et al., 2002).

Furthermore, the institutional arrangements as well as fiscal and monetary policies have been strongly motivated by several criteria that set conditions for the European Union (EU) accession and have directed the adjustment of the Central European countries towards the EU standards since the mid 1990's.[2] The EU accession *per se* on May 1, 2004 was associated with the full removal of restrictions on movements of capital. Nevertheless, the restructuring process continues as policy makers in the new member countries attempt to join the Eurozone (McKinnon, 1999; Buiter and Grafe, 2002; Buiter, 2004)[3], and the institutional changes are accompanied by a convergence in macro-economic fundamentals of the recent EU members to the EU standards (Kocenda et al., 2006).

The ongoing institutional integration of the Central European countries with the global economy, the importance of foreign investment in these countries as well as their macro-economic developments suggest tightening of the stock market relations, as evidenced by extensive empirical literature on major developed stock 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). In particular, financial integration with the "old" EU[4] markets can be expected due to the EU enlargement as well as significant capital inflows from these countries. Surprisingly, the existing empirical literature on the Central European stock markets has delivered no (Gilmore and McManus, 2002, 2003) or only limited evidence (Egert and Kocenda, 2007; Syrioupoulus, 2006; Syllignakis and Kouretas, 2009) for such developments so far, when focusing on the long-run stock market linkages. In fact, this literature does not typically investigate the *developments* of the long-run relations; rather, it tends to fit only one model for the whole period of interest, tacitly assuming that the parameters of the model are constant over the whole time span. One exception is Voronkova (2004), who controls for structural breaks in the relations and indeed finds stronger evidence of long-run links than reported in the previous literature.

My approach follows Voronkova's work by assuming *a priori* that the characteristics of the linkages are likely to change. Therefore, I pay particular attention to the stability of the detected relations. I further extend the assumption of potential changes in the model and consider varying number of existing relations. In particular, I expect emergence of new long-run linkages related to the EU accession of the Central European countries in 2004. For this purpose, I compare the period before and after the EU enlargement[5] and find first strong evidence for increased stock market integration between the Central European markets and the developed markets associated with the EU accession. Moreover, the expected major role of the "old" EU in the integration process is confirmed by the results.

[1] The term "Central Europe" refers to the Visegrád Group of countries, namely the Czech Republic, Hungary, Poland and Slovakia.

[2] Hungary and Poland applied for the EU membership in 1994, followed by Slovakia in 1995 and the Czech Republic in 1996.

[3] Slovakia already adopted the Euro on January 1, 2009.

[4] The "old" EU refers to the EU-15 and comprises the following 15 countries: Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, Spain, Sweden, United Kingdom.

[5] Similar approach was used by Jochum et al. (1999) and Yang et al. (2006) to study the effects of the Russian financial crises.

More specifically, my analysis concentrates on the three largest Central European markets: the Czech, Hungarian and Polish markets.[6] The "old" EU is used as their mature counterpart. To capture other potentially 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, characterises the development in the largest emerging market in Europe with strong historical links to the Central European countries.

To model the long-term trends in the stock market co-movements, I use the Johansen cointegration technique in its multivariate setting. Although this technique enables the analysis of both the long-run and the short-run market structure, I focus on the long-run. The reasons are threefold. First, the potential long-run relations can be interpreted as equilibrium relations between asset prices and hence are a good measure of the degree of market integration. The asset prices may deviate from each other in the short run, but they will return to the equilibrium as a result of financial integration. Second, portfolios designed using only short-run correlations may not properly estimate the long-term gains. In fact, the standard risk-return analysis using the mean-variance approach ignores the long-term trends, since these are lost as the data are differenced. However, in case of tightening long-run linkages among the markets, potential benefits from international portfolio diversification can substantially decrease. Therefore, the long-run relations are of particular interest for stock market investors investing internationally. Finally, the estimates of the long-run relations within the Johansen cointegration framework are more reliable than the estimates of the short-run structure because the convergence to their true values is faster (see Juselius, 2007, p. 230).

The proper use of the cointegration technique relies on several assumptions, such as constancy of parameters or independence of residuals. Surprisingly, the literature applying this technique to the stock markets do not typically report *any* tests of these assumptions. I attempt to overcome this crucial shortcoming. As already addressed, I check the assumption of constant parameters by several recursive tests to detect possible structural changes and to avoid the distortion of the results by assumption violation. The stability of the relations is surely also a key issue for a plausible portfolio design. Furthermore, I carefully handle the assumption of independently and normally distributed residuals and, if necessary, I model too-large residuals caused by extraordinary shocks such as the terrorist attacks in September 11, 2001, by inclusion of proper dummy variables. In addition, I explicitly address the question of which markets are significantly involved in the long-run relations, which is also a commonly neglected, though very important, issue.[7]

In this way, I provide evidence for a similar degree of (co-)integration among the three Central European markets in both periods, before and after the EU enlargement. Nevertheless, no long-run linkages between the Central European markets and the two developed markets or the Russian market can be detected in the period before the enlargement. On the contrary, two new relations emerge after the EU enlargement and link the Central European markets to the other markets. In particular, one of these relations is identified as a "new EU relation", linking the movements of the Central European markets to the "old" EU. To check the degree to which generalising the results is possible, several robustness checks are conducted at the end of the analysis.

The remaining part of the paper is organised as follows. Section 2 describes the developments of the three

[6] The Slovakian stock market is not considered because of its minor size. For instance, its market capitalisation was approximately ten times smaller than that of the Czech stock market in 2007 (Standard & Poor's, 2008).

[7] A detection of a cointegration relation in a multivariate setting does not necessarily mean that a long-run equilibrium relation between the Central European and other markets exists. It might be the case that the relation involves only two markets - in an extreme case the two developed markets of Western Europe and the US. Hence, the study of Syllignakis and Kouretas (2009) involving seven Central and Eastern European markets and two developed markets does not deliver a clear picture of which of these markets are inter-linked.

largest Central European markets. Section 3 introduces the data and provides basic descriptive statistics. Section 4 explains the methodological approach. Two models for the pre- and post-accession period are estimated and compared in Section 5, and Section 6 concludes.

2. The Developments of the Central European Stock Markets

After the collapse of communism, the transition process of the Central European countries was accompanied by the establishment of stock markets. The first stock exchange in the region was reopened in Hungary in July 1990. In the next two years, the stock exchanges in Poland (1991) and in the Czech Republic (1992) started to operate as well. Consequently, all three markets underwent considerable growth in their size.

Market capitalisation grew relatively steadily in Hungary and Poland from around 5 per cent of GDP in 1995 to over 34 per cent of GDP in Hungary and nearly 50 per cent of GDP in Poland in 2007 (see Table 1). The EU accession in 2004 seems to have accelerated the growth of the markets since the market capitalisation nearly doubled in this year in both markets. The development in the Czech Republic reveals a somewhat different scenario. The high rates of market capitalisation and the large number of listed companies in the early stage of the transformation process reflect to a large extent the effects of a privatisation program that was carried out in the first half of the 1990s (see Hanousek et al., 2009). After the large waves of privatisation, the size of the stock market even decreased, but it started to considerably grow again in 2001. Following the scenario of the other two markets in the region, the Czech market experienced a steep jump in size in 2004. In 2007, the market capitalisation of the Prague stock exchange accounted for more than 40 per cent of GDP.

Compared to the developed markets, the rates of market capitalisation in per cent of GDP are still rather low, but they are significantly catching up.[8] The liquidity of the Central European stock markets appears to already be quite comparable to that of the developed markets. This is reflected by relatively high turnover ratios that even exceeded 100 per cent in 2007 in Hungary and in 2005 in the Czech Republic. The lower turnover ratios in Poland are similar, for instance, to turnover ratios in Austria, which have typically stayed below 50 per cent in recent years (Standard & Poor's, 2008).

The most interesting question from the point of financial integration is this: who invests in the markets, the domestic or the foreign investors? Are the stock markets of the Central European countries developed enough to attract significant foreign capital?

In general, the Central European countries have been receivers of large capital inflows from developed countries. The capital flows to these countries constituted around 5 to 6 per cent of their GDPs in the period between 1993–1999. The main source of foreign financing was direct investment, followed by portfolio investment (see Mora et al., 2002; Koeke and Schroeder, 2003). In 1995, for instance, Hungary attracted the largest amount of FDI per capita of any country outside the developed market economies, and as a result, the share of FDI in its GDP exceeded 10 per cent (Lankes and Stern, 1999). Most of the capital flows to Central Europe originated in Western European countries such as Germany or Austria (Marin, 2004). Direct investment from Germany accounted for about one third of cumulative FDI to the broader eastern European region by 1996, and its majority indeed flew to Central Europe.

Nevertheless, foreign investment figures in Table 1 indicate that foreigners indeed also held significant amounts of stock market assets (other than FDI-related holdings).[9] In 2007, the value of foreign investment

[8] The market capitalisation in 2006 represented 57, 59, 148 and 160 per cent of GDP for Germany, Austria, the US and the UK, respectively (Standard & Poor's, 2008).

[9] I follow the work of Koeke and Schroeder (2003) and use the international investment position in equity securities from the International Financial Statistics (IMF, 2008b, 2009, line 79 ldd) for measuring foreign investment in the stock markets. This approach

Table 1: The developments of the three Central European stock markets

	1995	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005	2006	2007
Prague stock exchange (Czech Republic)													
Number of listed companies	1635	1588	276	261	164	131	94	78	63	54	36	29	32
Market capitalization (billions of US $)	15.7	18.1	12.8	12.0	11.8	11.0	9.3	15.9	17.7	30.9	38.3	48.6	73.4
Market capitalization (% of GDP)	30.1	31.2	24.1	21.4	21.6	21.7	15.3	21.5	19.5	28.8	30.8	34.0	42.2
Turnover ratio	32.9	50.3	47.9	38	36.7	60.3	34.1	48.7	52.5	78.5	120.7	75.1	71.7
Foreign investment (billions of US $)	2.6	3.4	3.0	3.8	2.7	3.1	3.6	4.3	5.5	9.3	9.0	11.6	14.5
Foreign investment (in % of market cap.)	16.9	18.8	23.7	31.5	23.1	27.8	38.1	26.7	31.1	30.3	23.4	23.8	19.8
Foreign investment (in % of GDP)	5.1	5.9	5.7	6.7	5.0	6.0	5.8	5.8	6.1	8.7	7.2	8.1	8.3
Budapest stock exchange (Hungary)													
Number of listed companies	42	45	49	55	66	60	57	48	49	47	44	41	41
Market capitalization (billions of US $)	2.4	5.3	15.0	14.0	16.3	12.0	10.4	13.1	16.8	28.7	32.6	42.0	47.7
Market capitalization (% of GDP)	5.4	11.7	32.8	29.3	34.0	26.3	20.0	20.0	20.1	28.5	29.8	37.1	34.4
Turnover ratio	17.3	41.6	73.4	113.9	95.8	90.7	44.4	52.2	57.6	59.9	79.2	86.8	107
Foreign investment (billions of US $)	.	.	2.5	2.3	4.4	3.0	2.9	3.8	5.6	11.4	13.2	18.4	15.3
Foreign investment (in % of market cap.)	.	.	16.7	16.7	26.7	24.8	28.0	28.9	33.5	39.6	40.4	43.8	32.0
Foreign investment (in % of GDP)	.	.	5.5	4.9	9.1	6.5	5.6	5.8	6.7	11.3	12.0	16.3	11.0
Warsaw stock exchange (Poland)													
Number of listed companies	65	83	143	198	221	225	230	216	203	225	248	267	328
Market capitalization (billions of US $)	4.6	8.4	12.1	20.5	29.6	31.3	26.0	28.8	37.2	71.1	93.9	149.1	207.3
Market capitalization (% of GDP)	3.6	5.9	8.5	12.9	19.1	19.8	14.0	15.0	17.7	29.3	31.0	44.0	48.9
Turnover ratio	71.5	84.8	78.4	54.4	45.8	49.9	26.1	22.4	26.6	33.1	37.3	45.7	49.2
Foreign investment (billions of US $)	.	2.3	2.7	5.0	5.0	5.4	4.3	4.4	6.7	13.7	18.8	22.6	33.2
Foreign investment (in % of market cap.)	.	27.2	22.0	24.3	16.8	17.1	16.5	15.3	18.0	19.3	20.0	15.1	16.0
Foreign investment (in % of GDP)	.	1.6	1.9	3.1	3.2	3.4	2.3	2.3	3.2	5.7	6.2	6.7	7.8

Sources: Author's calculation based on the Emerging and Global Stock Market Factbooks (Standard & Poor's, 2002, 2008) and International Financial Statistics (IMF, 2008b, 2009).
Notes: Foreign investment refers to the international investment position, liabilities, and equity securities in the International Financial Statistics (IMF, 2008b, 2009, line 79 ldd).

in stock market assets represented around 8 per cent of GDP in the Czech Republic and Poland and even 11 per cent of GDP in Hungary. Especially in Hungary, foreign investment has constituted around 40 per cent of market capitalisation in recent years, which even exceeds foreign ownership holdings in some of the developed markets.[10]

Foreign investment followed the scenario of market capitalisation and experienced a sharp one-time increase in 2004. This is very likely a result of increased confidence and willingness of foreign investors to participate in the Central European markets associated with the accession of the countries to the EU in 2004. From a legal point of view, the stock markets had already been open to foreign investors prior to the EU accession. The restrictions on foreign investment were gradually lifted between 1994 to 1999 (see Dvorak and Podpiera, 2006; Syllignakis and Kouretas, 2009). However, the increased interest of foreigners after the EU accession indicates that some foreign investors may have refrained from the markets before the EU enlargement in particular because of institutional or political risk.

The developments of the Central European markets appear to be significantly influenced by their accession to the EU in May 2004. Around this date, a sharp increase in the size as well as in the attractiveness of the markets to foreign investors can be observed. I turn to investigate in the following whether this date also meant stronger integration of the markets into the global economy.

3. Data

For capturing the stock market movements, I collected data of weekly closing price indices for the three Central European markets, the Western European market, and the US and Russian markets. All the data have been obtained from the Thompson Financial Datastream database. To avoid the distorting effects of using different types of local stock market indices for the emerging markets, the standardised IFC Investable (IFCI) indices are used for representing the Czech, Hungarian, Polish and Russian markets. Moreover, these indices are designed to feature the type of assets that are legally and practically available to a foreign portfolio investor. Since stronger linkages between the recent and the "old" EU members are expected, the Western European countries are of particular interest for analysis. The limitations regarding a reasonable number of markets in the cointegrated VAR model suggest including only a single representative of the Western European market; thus, the DJ Stoxx 600 is used. The S&P 500 is chosen, as usual, to represent the US market.

Following Jochum et al. (1999) and Voronkova (2004), all indices are measured in local currency.[11] The data are converted to the natural logarithms, and denoted by LCZ, LHN, LPO, LWE, LUS, and LRU for the Czech, Hungarian, Polish, Western European, US and Russian markets, respectively.

The data were collected for the time period between October 30, 1998, and May 4, 2007. The end of the period is limited by the data availability when writing the paper, but the sample period still covers three years after the EU enlargement on May 1, 2004. The choice of the start of the 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 intention 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 Central European countries had already been opened, on March 31, 1998, a significant turning point in the negotiations appeared to be November 2001. In this month, the European Commission announced the EU enlargement in its Annual Progress Reports on Enlargement and provided

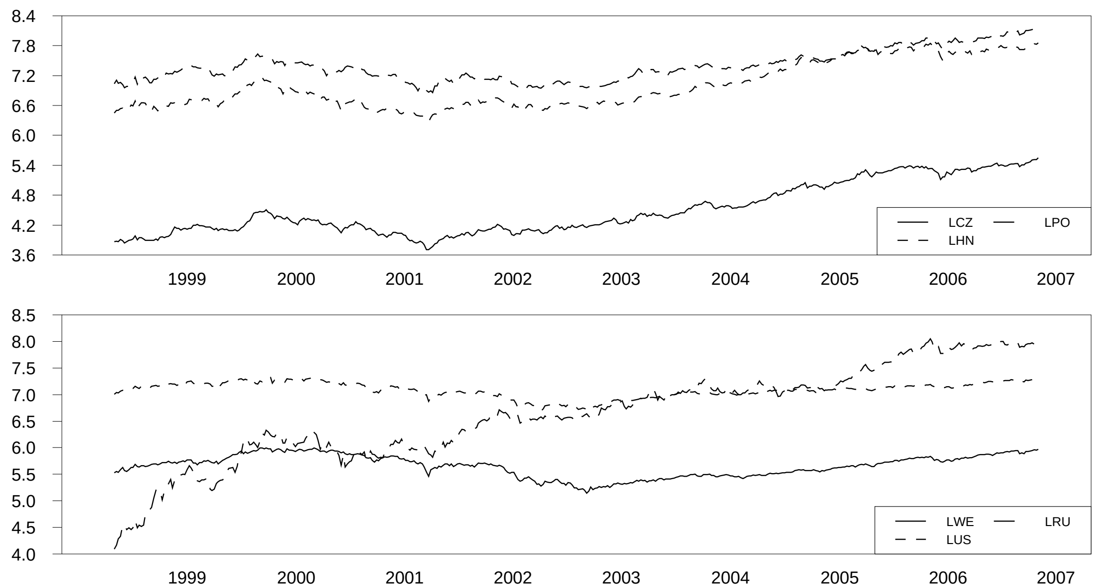
might underestimate the true holdings by foreigners, as some equity holdings are part of FDI.

[10] The corresponding figures for Austria, Germany and the US in 2006 were 45, 37 and 14 per cent, respectively.

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

a timetable for the enlargement (see Dvorak and Podpiera, 2006). Since then, foreign investors anticipated that the three Central European countries would enter the EU in 2004. Figure 1) shows that a dramatic rise in the Central European market prices can be observed after this date, whereas the developed markets did not follow this pattern.

Figure 1: Logarithms of the stock market indices



Notes: LCZ, LHN, LPO, LWE, LUS and LRU stand for the logarithms of the Czech, Hungarian, Polish, Western European, US and Russian indices, respectively.

The timing of the different influential events led me to the decision to split the data into three periods:

- the pre-accession period, from October 30, 1998, to November 2, 2001 (three years),
- the accession period, from November 3, 2001, to April 30, 2004 (two and a half years),
- the post-accession period, from May 1, 2004, to May 4, 2007 (three years).

Since the accession to the EU is to a large extent a gradual process, 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, I investigate this period only marginally and focus rather on the comparison of the pre- and post-accession periods. Using weekly data, both the pre- and the post accession periods cover exactly 158 observations and thus are well comparable.[12]

The logarithmic transformation enables the interpretation of the first differences as continuous stock market returns. Table 2 provides means and standard deviations of the return series in the pre- and post-accession periods. Except for Russia, all means were higher in the post-accession period, though the increase

[12]The time span of three years is similar to the lengths of periods used in other studies (e.g., Jochum et al., 1999; Yang et al., 2006) that also analyse long-run stock market equilibrium relations.

was never statistically significant. Regarding the standard deviations, the volatility of all the markets in the post-accession period was significantly lower and indicated a more stable situation in all the markets.

Table 2: Descriptive statistics of the stock market returns

	Pre-accession period October 30, 1998 - November 2, 2001		Post-accession period May 1, 2004 - May 4, 2007		T-test for difference p-values	
	mean	S.D.	mean	S.D.	mean	S.D.
Czech Republic	0.00074	0.034	0.0057	0.031	0.18	0.09
Hungary	0.00094	0.046	0.0051	0.035	0.37	0.00
Poland	0.00045	0.042	0.0048	0.029	0.30	0.00
Russia	0.01139	0.082	0.0048	0.040	0.37	0.00
West Europe	0.00092	0.028	0.0030	0.015	0.43	0.00
US	0.00010	0.030	0.0018	0.014	0.53	0.00

Notes: The table presents means and standard deviations of the continuous (log) return series in the pre- and post- accession periods. In the last two columns, p-values of the t-test for difference in means and standard deviations of the two periods are reported.

This period is also characterised by higher correlations among the series, as evidenced in Table 3. This suggests that the short-run linkages among the markets were stronger after the EU enlargement as the markets became more synchronised. Surprisingly, a significant increase occurred, especially in the correlations between the Central European markets and the Russian market, but not between the Central and Western European markets. However, the correlations with the Russian market started at very low levels in the pre-accession period, which might still reflect a rather non-standard behaviour of the Russian market after the financial crises in 1998. Since the short-run correlations provide only a limited tool for measuring stock market integration, I study the long-run linkages in the following.

Table 3: Correlations of the returns

	Pre-accession period October 30, 1998 - November 2, 2001						Post-accession period May 1, 2004 - May 4, 2007					
	CZE	HUN	POL	RUS	WE	US	CZE	HUN	POL	RUS	WE	US
CZE	1						1					
HUN	0.54	1					0.59	1				
POL	0.49	**0.52**	1				0.58	**0.72**	1			
RUS	**0.23**	**0.31**	**0.27**	1			**0.52**	**0.55**	**0.48**	1		
WE	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

The table shows correlations of the continuous (log) return series in the pre- and post- accession periods. CZE, HUN, POL, RUS, WE and US stands for Czech, Hungarian, Polish, Russian, Western European and US markets, respectively. Significantly different correlations (at a 10 per cent level) in the two periods are indicated by boldface.

4. Methodology

For modelling the long-run relationships among the stock markets, I applied the cointegrated VAR model introduced by Johansen (1991). This method is very applicable for my type of data because it is specifically designed for non-stationary stochastic processes, and stock market prices are indeed usually integrated of order one (I(1) hereafter). Our stock market indices follow this pattern, as indicated by Figure 1 and by the results of the augmented Dickey Fuller tests.[13]

The cointegration method assumes that the time series can be modelled by a VAR(k) model. Denoting the vector of the six stock market indices in logarithms in period t by $\boldsymbol{X}_t$, this means that

$$\boldsymbol{X}_t = \boldsymbol{\Pi}_1 \boldsymbol{X}_{t-1} + \cdots + \boldsymbol{\Pi}_k \boldsymbol{X}_{t-k} + \boldsymbol{\varepsilon}_t, \qquad \text{where} \qquad \boldsymbol{\varepsilon}_t \sim \boldsymbol{IN}_6(0, \boldsymbol{\Omega}) \qquad \text{and} \qquad t = 1, \ldots, T. \tag{1}$$

Hence, the error terms $\boldsymbol{\varepsilon}_t$ are assumed to be independently normally distributed with a constant variance-covariance matrix. To meet these assumptions, I use weekly frequency data rather than daily data. The reason is that the lower-frequency data suffer less from the "stylised facts" of the financial time series such as heavy-tailed distributions or ARCH effects. Moreover, the information loss due to the lower frequency is not very important in the cointegration framework, since the length of the period, and not the frequency, is important for the detection of the long-run relations. Furthermore, the disturbing effects of different market closing times (European vs. Russian vs. US market) are eliminated.

A 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 (VECM) form:

$$\Delta \boldsymbol{X}_t = \boldsymbol{\Pi} \boldsymbol{X}_{t-1} + \boldsymbol{\Gamma}_1 \Delta \boldsymbol{X}_{t-1} + \cdots + \boldsymbol{\Gamma}_{k-1} \Delta \boldsymbol{X}_{t-k+1} + \boldsymbol{\varepsilon}_t, \tag{2}$$

$$\text{where} \qquad \Delta \boldsymbol{X}_t = \boldsymbol{X}_t - \boldsymbol{X}_{t-1}, \qquad \boldsymbol{\Pi} = -(\boldsymbol{I} - \sum_{j=1}^{k} \boldsymbol{\Pi}_j), \qquad \boldsymbol{\Gamma}_i = -\sum_{j=i+1}^{k} \boldsymbol{\Pi}_j$$

and $\boldsymbol{I}$ denotes the identity matrix. This representation allows one to directly deal with the non-stationary pattern in the data that is now concentrated exclusively in $\boldsymbol{\Pi X}_{t-1}$, as it is the only term in levels in Equation 2. So the $\boldsymbol{\Pi}$ matrix captures all information about the long-run effects, and its rank r cannot be full. Supposing that the rank were full, a stationary process $\Delta \boldsymbol{X}_t$ would then be equal to a non-stationary term $\boldsymbol{\Pi X}_{t-1}$ (plus several stationary terms), which leads to a contradiction. Hence, $\boldsymbol{\Pi}$ can be partitioned as

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

where $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are $6 \times r$ matrices and $r < 6$. The rank r (or the cointegration rank) can be determined using the trace test, also called the Johansen test. This procedure discriminates between (r) significant and $(6 - r)$ insignificant eigenvalues λ_i, $i = 1, \ldots, 6$, on which the maximum likelihood estimation of the model is based (see Johansen, 1996, Chapter 6). Consequently, each significant eigenvalue is related to one stationary cointegration relation, which can be viewed as a long-run equilibrium relation among the markets. The r stationary relations are represented by $\boldsymbol{\beta}' \boldsymbol{X}_{t-1}$. The coefficients of $\boldsymbol{\alpha}$ capture the adjustment of markets to

[13] I conducted the univariate augmented Dickey Fuller tests with a constant and with a constant and a time trend, both for lag 1 to 3. In particular, the tests applied on the return series revealed that the data in levels are at most I(1) and thus suitable for cointegration analysis in the I(1) framework. This pattern was later confirmed when analysing the multivariate models.

the cointegration relations and are called loadings. The $\mathbf{\Gamma}_i$ matrices contain information about the short-run linkages.

The model can be extended by the inclusion of deterministic components (a constant, a deterministic trend or different dummy variables) that are partitioned into those restricted to enter only the cointegration relations and the unrestricted ones (see Juselius, 2007, Chapter 6). My modelling strategy is to allow for a relatively rich structure at the beginning of the analysis. Hence, I include an unrestricted constant and, eventually, unrestricted dummy variables. The time trend is restricted to enter only the cointegration relations to avoid a quadratic trend in the level data.[14] Later on, I try to reduce the rich model into a more parsimonious one by examining the significance of the corresponding coefficients.

In general, the stability of the model can be investigated by several recursive tests. The idea is to choose a baseline period (e.g., the first year), on which the first model is estimated, and then recursively test whether additional observations follow the same model. In this way, I study the constancy of $\boldsymbol{\beta}$ and significant λ_i estimates as well as the stability of the full model using the log-likelihood function. In addition, I check the validity of imposed restrictions throughout the different time periods.

A useful tool for the recursive tests of constancy is to distinguish between two types of $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ estimates, the first based on the VECM ("X-form") and the others based on the concentrated model ("R-form"). The latter model is motivated by the idea of the Frisch-Waugh Theorem (Frisch and Waugh, 1933) and enables the obtaining of "cleaner" estimates of the long-run structure after the short-run dynamics and the deterministic components have been concentrated out (see Juselius, 2007, Chapter 7). In particular, if the constancy of the "X-form" coefficients is rejected as opposed to the "R-form" coefficients, the non-stability is likely to come from the short-run structure.

5. Results

5.1. The Pre-accession Period

I start with the estimation of the VAR(k) model using data for the pre-accession period. The basic model includes an unrestricted constant, a restricted trend, but no dummy variable. As a starting point, I use a lag length of 3. The theoretical specification in Equation 1 requests two residual assumptions, namely, normality and independence. The results of misspecification tests indicate that both assumptions are clearly violated for this basic model. To improve the model specification, I include several dummy variables, which turned out to be economically relevant as well as statistically significant.

- An unrestricted transitory shock dummy for the first two weeks in January 1999 (January 8 and 15)[15] is related to a stock market over-reaction after the introduction of the Economic and Monetary Union (EMU) on January 1, 1999 in most of the "old" EU countries. It corrects particularly for the volatile behaviour of the Western European markets in the first two weeks in January.
- An unrestricted blip (impulse) dummy on April 14, 2000[16] captures a temporal drop of the US market following the burst of the dot-com bubble. Therefore, the corresponding coefficient is highly negatively significant, especially for the US market.

[14] Although a quadratic trend could improve the fit within the sample, it would lead to the implausible economic result that the stock markets follow quadratic trends.

[15] A transitory dummy is modelled by the inclusion of $\boldsymbol{d}_{tr} = (\ldots, 0, 1, -1, 0, \ldots)$ to the explanatory variables in the VECM form. 1 and -1 correspond to January 8 and 15, respectively. For more details on the dummy variables, see Juselius (2007, Chapter 6).

[16] An impulse dummy is modelled by the inclusion of $\boldsymbol{d}_p = (\ldots, 0, 1, 0, \ldots)$, where 1 corresponds to April 14.

- Three unrestricted blip dummy variables for three weeks after the September 11 terrorist attacks (September 14, 21, 28, 2001) account for the substantial market instability, mainly in the US market, followed by the Western European markets.

Furthermore, I exclude insignificant short-run coefficients and thus adjust the lag length of the model. See Appendix A.1 for a more detailed discussion on the specification of the model. The resulting model meets the required assumptions and is used in the following.

As an indicator for the number of cointegration relations, I use the trace test (Table 4). Since the model contains a trend in the cointegration relation and several dummy variables, I do not report the results of the standard test; instead, I simulate an asymptotic trace test distribution by the program developed in Nielsen (2004). The results suggest rank 1, because $H_0 : r = 0$ is rejected, but $H_0 : r = 1$ cannot be rejected. In addition, the graphical analysis of the first cointegration relation in Figure 3 in Appendix B proposes stationarity. Hence, the evidence suggests the existence of one long-run equilibrium relation among the indices in the pre-accession period.

Table 4: Trace test for the pre-accession period

H_0 :	Eigenvalue	Trace	Trace*	Frac95	P-Value	P-Value*
r=0	0.27	121.7	111.2	108.7	0.01	0.03
r=1	0.15	72.2	65.4	82.3	0.23	0.47
r=2	0.12	46.5	38.7	59.5	0.40	0.76

Notes: The results of the asymptotic trace test and corresponding eigenvalues are reported. A length of 158 random walks (the same as the length of the sample) and 5000 replications were used for the simulation. Frac95 denotes the 95% quantile from the simulated trace test distribution. Trace* and P-Value* refers the results of a small sample Bartlett correction introduced in Johansen (2002).

After the rank determination, I examine the stability of the model. All tests of constancy in Figures 5, 6 and 7 shown in Appendix C suggest a good stability of the model. In particular, Figure 7 confirms constancy of the long-run equilibrium relation. Due to the satisfactory results of these tests, further adjustment of the model such as inclusion of structural breaks is not necessary.

I turn now to the question of whether the cointegration relation involves all the markets simultaneously or only some of them. In particular, I am interested whether the cointegration relation links the Central European markets to the other markets such as the Western European market, thereby suggesting the integration of these markets. A test of exclusion clearly indicates that this is not the case, as LRU, LWE, LUS can be individually (Table 5) as well as jointly ($\mathcal{H}_1$ in Table 6)[17] excluded from the cointegration relation, contrary to the Central European markets. This means that the cointegration relation involves only the three Central European markets and that there are no long-run linkages to their mature counterparts or the Russian market in the pre-accession period.

The magnitude of the coefficients under $\mathcal{H}_1$ in Table 6 lead me to further test for the price homogeneity of the Czech, Polish and Hungarian markets, denoted by $\mathcal{H}_2$.[18] The accepted homogeneity means that a common stochastic trend exists that drives the markets to move in the same direction by a similar amount.[19] This is particularly inconvenient for the stock market investors, since the portfolio diversification among the

[17] $\mathcal{H}_1 : \beta_{1,LRU} = \beta_{1,LWE} = \beta_{1,LUS} = 0$

[18] $\mathcal{H}_2 : \beta_{1,LCZ} + \beta_{1,LHN} + \beta_{1,LPO} = 0$ & $\beta_{1,LRU} = \beta_{1,LWE} = \beta_{1,LUS} = 0$. The time trend, though very small in magnitude, is significant, and its exclusion is rejected.

[19] This interpretation can be derived from the MA representation of the model (see also Juselius, 2007, Chapter 14).

Table 5: Tests of restrictions on β and α in the pre-accession period

Test	DF	CV	LCZ	LHN	LPO	LRU	LWE	LUS
Exclusion	1	3.84	23.13	4.86	19.67	0.28	0.08	0.02
			[0.00]	[0.03]	[0.00]	[0.60]	[0.78]	[0.89]
Unit vector	5	11.07	7.82	42.18	27.46	34.41	35.37	35.65
in α			[0.17]	[0.00]	[0.00]	[0.00]	[0.00]	[0.00]

The table reports likelihood ratio test statistics of exclusion restrictions on β (H_0 : $\beta_{1,i} = 0$ for a particular market i) and of unit vector in α ($H_0 : \alpha_{i,1} = 0$ for all i except one) under the rank 1 assumption. LCZ, LHN, LPO, LWE, LUS and LRU stands for the logarithms of the Czech, Hungarian, Polish, Western European, US and the Russian indices, respectively. All the test statistics are χ^2 distributed, DF denotes the degree of freedoms and CV the corresponding 5% critical value. P-values are reported in the parentheses.

three Central European markets is strongly limited. I also examine the stationarity of the spread between the Czech and Polish market ($\mathcal{H}_3$).[20] The result suggests that $(LCZ - LPO)$ can be regarded as stationary, which implies a strong integration of the two markets and even no benefits of portfolio diversification between them in the long run.[21] The reason is that under this scenario, the two markets are driven by exactly one common stochastic trend, which pushes them to move in the same direction and by the same amount, meaning that they follow on average the same random walk. Nevertheless, $\mathcal{H}_3$ implies the exclusion of the Hungarian market, which is in conflict with 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 for $\mathcal{H}_2$. Therefore, I consider the homogeneity restrictions in $\mathcal{H}_2$ to be more plausible than the spread restrictions, and I use them in the following.

Table 6: Estimates of $\boldsymbol{\beta}$ under different restrictions in the pre-accession period

Test	LCZ	LHN	LPO	LRU	LWE	LUS	trend	DF	χ^2	p-value
$\mathcal{H}_1$	1	-0.23	-0.74	0	0	0	-0.00	3	0.95	0.81
$\mathcal{H}_2$	1	-0.22	-0.78	0	0	0	-0.00	4	0.88	0.93
$\mathcal{H}_3$	1	0	-1	0	0	0	0	6	7.20	0.30

Notes: The table presents the estimated coefficients under different restrictions on $\boldsymbol{\beta}$ and the results of likelihood ratio test statistics of these restrictions. LCZ, LHN, LPO, LWE, LUS and LRU stands for the logarithms of the Czech, Hungarian, Polish, Western European, US and the Russian indices, respectively. All the test statistics are χ^2 distributed. DF denotes the degree of freedoms.

The analysis of the long-run relation captured by $\boldsymbol{\beta}$ does not indicate which of the markets are adjusting to the cointegration relation and which are following only their own stochastic trends. For this, the $\boldsymbol{\alpha}$ coefficients are investigated. The tests for a unit vector in $\boldsymbol{\alpha}$ suggest that the Czech market can be considered the only adjusting one. This means that shocks to the Czech market have no permanent effect on any market in the system, not even on itself, and that the "random walk" movements of the Czech market are driven by permanent shocks to the other two markets involved in the equilibrium relation.

This finding is not surprising, when considering that the Czech market in the pre-accession period is the smallest of the three Central European markets and also the most open to foreign capital (see Table 1). The differences are especially pronounced when comparing the Czech and the Polish markets. In 2000, the

[20] $\mathcal{H}_3 : \beta_{1,LCZ} - \beta_{1,LPO} = 0 \;\; \& \;\; \beta_{1,LHN} = \beta_{1,LRU} = \beta_{1,LWE} = \beta_{1,LUS} = \beta_{1,trend} = 0$

[21] I also tested the stationarity of other market spreads, but none of them was found to be stationary.

capitalisation of the Polish stock market exceeded nearly three times the Czech market capitalisation, but only 17 per cent of investment going to the Polish market originated abroad, in contrast to 28 per cent of foreign investment going to the Czech market. Hence, the Polish market is much more likely to reflect only the local developments, and the Czech market is adjusting accordingly. The importance of the Polish market to the Czech market adjustment is also manifested by the large coefficient of $\beta_{1,LPO}$ as compared to $\beta_{1,LHN}$ and the borderline stationarity of the Czech-Polish spread.

The joint restrictions on $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ coefficients (i.e., the unit vector in $\boldsymbol{\alpha}$ for the Czech market and the homogeneity restrictions $\mathcal{H}_2$ on $\boldsymbol{\beta}$) are also not rejected, and the recursive likelihood ratio test based on the "R-form" estimates indeed suggests a validity of the imposed restrictions throughout the sample (see Figure 8 in Appendix C).[22] The final estimates are summarised in Table 7.

Table 7: Final model in the pre-accession period

Likelihood ratio test of the restricted model:					$\chi^2(9) = 8.14$, p-value = 0.52		
	LCZ	LHN	LPO	LRU	LWE	LUS	trend
β_1'	1	-0.27	-0.73	0	0	0	-0.00
	(.)	(-2.98)	(-8.16)	(.)	(.)	(.)	(-2.57)
	Δ LCZ	Δ LHN	Δ LPO	Δ LRU	Δ LWE	Δ LUS	
α_1'	-0.28	0	0	0	0	0	
	(-5.75)	(.)	(.)	(.)	(.)	(.)	

Notes: The table reports the likelihood ratio test for the joint restrictions on $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ and the resulting coefficients. The corresponding t-statistics are shown in parentheses, and some of them are missing due to the imposed restrictions. LCZ, LHN, LPO, LWE, LUS and LRU stands for the logarithms of the Czech, Hungarian, Polish, Western European, US and Russian indices, respectively.

5.2. The Accession Period

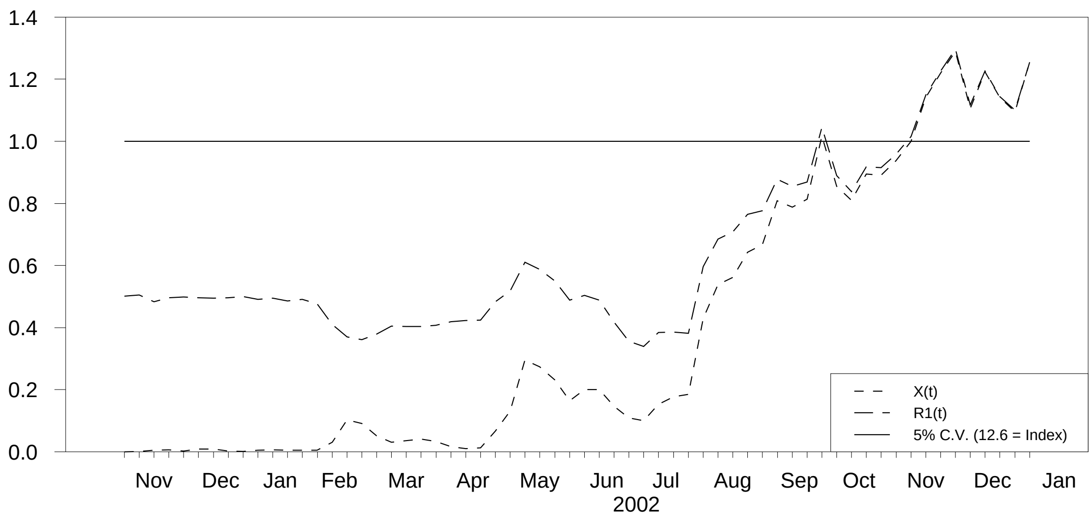
Having a reasonable model with one cointegration relation for the pre-accession period, the applicability of the model can be tested also for the following period, the accession period. For this, I conduct a recursive test of $H_0 : \beta =$ "known β". This test provides evidence for whether the cointegration relation found in the pre-accession period remains similar in the accession period as well. Therefore, the "known β" refers to the estimated β based on the pre-accession period and, consequently, its similarity to β coefficients estimated for longer periods is tested. Figure 2 shows that the cointegration relation from the pre-accession period persisted for approximately one year in the accession period and that later on, around November 2002, the relation changed permanently. 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.

5.3. The Post-accession Period

Similarly to the pre-accession period, the residuals of the basic VAR(3) model for the post-accession period do not meet all the required assumptions. The difficulties can be solved by the inclusion of the following unrestricted blip dummy variables, which account for the largest residuals in the model.

[22] The rejection of the restrictions for the shortest subsamples using the "X-form" appears to be caused by the instability of the short-run coefficients (e.g., Γ_1 and Γ_2) and thus does not indicate a serious distortion of the results for the long run.

Figure 2: Test of β = "known β" in the accession period



Notes: 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.

- Blip dummies for three weeks (March 18, October 14, November 11) in 2005 correct for the temporal instability in the Czech and Hungarian market.
- Four blip dummies (May 19, June 9, 16, 30) in 2006 account for the high volatility in the emerging markets. The volatility appears to be a result of a sharp correction in the price of riskier assets after almost three years of significant gains.
- One blip dummy in March 2, 2007, captures a global downturn of all the markets following a sharp fall of China's domestic stock markets.

Having included these dummies, the lag length 1 in the VAR model turns out to be sufficient for modelling the short-run dynamics and the resulting model appears to be reasonably specified (for more details see Appendix A.2).

The simulated values for the trace test in Table 8 indicate rank 2 at a 5 per cent confidence level and rank 3 at a 10 per cent confidence level. Moreover, considering a very small difference in magnitude between the second (0.17) and third eigenvalue roots (0.16), it seems reasonable to prefer rank 3 to rank 2. Furthermore, Figure 4 in Appendix B suggests stationarity of the third long-run equilibrium relation. Therefore, rank 3 appears to be the right choice.

Several recursive tests are again conducted to check the assumption of constant parameters. The tests for constancy of the log-likelihood function and of the $\boldsymbol{\beta}$ parameters (see Figures 9 and 11 in Appendix C) do not indicate a violation of the constancy assumption. However, the development of eigenvalues in Figure 10 in Appendix C clearly detects non-constancy, particularly in the two largest eigenvalues corresponding to the first two stationary relations. This in turn indicates non-constant $\boldsymbol{\alpha}$ parameters.[23] Therefore, I cannot

[23] Since the eigenvalues are linear functions of the corresponding $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ parameters and the constancy of $\boldsymbol{\beta}$ parameters is not violated, the rejection occurs due to the non-constant $\boldsymbol{\alpha}$ parameters.

Table 8: Trace test for the post-accession period

H_0 :	Eigenvalue	Trace	Trace*	Frac95	P-Value	P-Value*
r=0	0.22	125.9	123.5	108.1	0.00	0.00
r=1	0.17	87.7	86.3	81.4	0.02	0.02
r=2	0.16	58.4	57.8	59.6	0.06	0.07
r=3	0.10	31.7	31.5	40.5	0.27	0.29

Notes: The results of the asymptotic trace test and the corresponding eigenvalues are reported. The simulation framework is the same as for the pre-accession period, i.e., a length of 158 random walks (the same as the length of the sample) and 5000 replications were used. Frac95 denotes the 95% quantile from the simulated trace test distribution. Trace* and P-Value* denotes the results of a small sample Bartlett correction introduced in Johansen (2002).

consider the estimates of $\boldsymbol{\alpha}$ to be reliable and concentrate only on the examination of the $\boldsymbol{\beta}$ coefficients.

Compared to the pre-accession period, the higher number of stationary relations suggests stronger integration among the six markets in general. However, it is not clear whether new relations between the Central European and the Western European markets emerged or whether the linkages among the Central European markets strengthened, and it is not clear which role is played by the other markets, the US and the Russian markets. To learn more about this, the long-run structure needs to be identified.

As a starting point, the tests of exclusion in Table 9 show that no market can be excluded from all three cointegration relations simultaneously, meaning that each market is involved in at least one cointegration relation.

Table 9: Tests of exclusion restrictions on β in the post-accession period

Test	DF	CV	LCZ	LHN	LPO	LRU	LWE	LUS
exclusion	3	7.81	10.81	9.33	17.19	14.44	14.93	16.68
			[0.01]	[0.03]	[0.00]	[0.00]	[0.00]	[0.00]

Notes: The table reports likelihood ratio test statistics of exclusion restrictions on $\boldsymbol{\beta}$ under the rank 3 assumption. The null hypothesis is $\beta_{1,i} = \beta_{2,i} = \beta_{3,i} = 0$ for a particular market i. LCZ, LHN, LPO, LWE, LUS and LRU stands for the logarithms of the Czech, Hungarian, Polish, Western European, US and Russian indices, respectively. The test statistics are χ^2 distributed, DF denotes the degree of freedoms and CV the corresponding 5% critical value. P-values are reported in the parentheses.

An interesting question arises in relation to the pre-accession period, namely, whether the same or a similar cointegration relation can be found among the Central European markets in the post-accession period as well. Therefore, I test for joint exclusion of all the non-Central European markets from a single relation, and it is not rejected, as shown in Table 10). The homogeneity of the Central European markets under $\mathcal{H}_2$ is also not rejected, thought the estimated coefficients are very different from those in the pre-accession period. The change in $\boldsymbol{\beta}$ coefficients is in line with the finding for the accession period that the estimated $\boldsymbol{\beta}$ from the pre-accession period do not remain constant. In particular, the importance of the Hungarian market has increased at the expense of the Polish market, and there is even some evidence for spread stationarity between the Czech and Hungarian markets ($\mathcal{H}_3$), provided that the time trend is included. This indicates, on the one side, a very strong link between the Czech and Hungarian markets that significantly limits, for instance, the diversification benefits in the post-accession period, as the two markets share one common driving trend. On the other side, the diversification possibilities between the Czech and Polish market have increased compared to the pre-accession period, as their spread is not found to be stationary anymore, and the markets follow different paths.[24]

I prefer the initial cointegration relation under $\mathcal{H}_1$ to the restricted homogeneity ($\mathcal{H}_2$) and spread ($\mathcal{H}_3$) relations due to the highest p-value. I regard this relation as the first identified cointegration relation in the model for the post-accession period and label it the "Central European relation". This relation is irreducible in the sense that it is not a linear combination of two "smaller" separate stationary relations and that, consequently, both of the two remaining cointegration relations involve also the 0non-Central European markets. In fact, the test of joint exclusion of the three Central European markets under $\mathcal{H}_4$ indicates that the two remaining cointegration relations bridge the two groups of markets, the non-Central European and the Central European markets, which is a novelty arising in the post-accession period.

Table 10: Estimates of β under different restrictions in the post-accession period

Test	LCZ	LHN	LPO	LRU	LWE	LUS	trend	DF	χ^2	p-value
$\mathcal{H}_1$	1	-0.78	-0.57	0	0	0	0.00	1	0.59	0.44
$\mathcal{H}_2$	1	-0.81	-0.19	0	0	0	-0.00	2	2.25	0.32
$\mathcal{H}_3$	1	-1	0	0	0	0	-0.00	3	5.02	0.17
$\mathcal{H}_4$	0	0	0	-0.08	1	-0.82	-0.00	1	5.01	0.03
$\mathcal{H}_5$	-1.05	0.87	0	0	1	0	-0.00	1	0.01	0.92
$\mathcal{H}_6$	-1	0.82	0	0	1	0	-0.00	2	0.01	0.99

Notes: The table presents the estimated coefficients under different restrictions on $\boldsymbol{\beta}$ and the results of likelihood ratio test statistics of these restrictions. LCZ, LHN, LPO, LWE, LUS and LRU stands for the logarithms of the Czech, Hungarian, Polish, Western European, US and the Russian indices, respectively. All the test statistics are χ^2 distributed. DF denotes the degree of freedoms.

Since I expect that the Central European markets are particularly linked to the Western European markets after the EU accession, I look for a stationary relation involving some of these markets. One convenient candidate for such a relation consists of that between the Czech, Hungarian and Western European markets ($\mathcal{H}_5$) because of its high p-value.[25] The p-value even increases by the additional restriction $\beta_{2,LCZ} = -\beta_{2,LWE}$ under $\mathcal{H}_6$. Since the joint hypothesis for $\mathcal{H}_6$ and the "Central European relation" in $\mathcal{H}_1$ is also not rejected ($\chi^2(3) = 2.26$ with a p-value of 0.52), the relation under $\mathcal{H}_6$ is included in the cointegration space and labelled as the "new EU relation". It clearly captures a new linkage between the Central European and the Western European markets that emerged in the post-accession period and could not be detected in the pre-accession period. The fact that the relation includes the Czech and Hungarian markets but not the Polish one can be explained by the two smaller countries' higher openness towards foreign capital and stronger trade integration with the EU. For instance, Czech and Hungarian exports to the EU in the post-accession period account for more than 50 and 40 per cent of their GDPs, respectively, compared to less than 25 per cent reported for Poland (IMF, 2008a).

As there is no economic prior for the third relation, it is just-identified by the exclusion of the Czech and the Hungarian market and links the Polish market to the non-Central European markets.[26] The resulting

[24] 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.

[25] Note that the exclusion of LRU and LUS from a single relation is trivial and leads to a just-identified relation in the model achieved by rotation of the cointegration space, but not by testable over-identifying restrictions.

[26] Note that just-identification of the third cointegration relation can generally be achieved by the exclusion of any market pair from the group of the Czech, Hungarian, Polish and the Western European market. Hence, the relation represents a linkage among all the markets.

estimate of the whole $\boldsymbol{\beta}$ matrix is reported in Table 11. Although the third cointegration relation also contains the Russian market, and although the market cannot be excluded according to the t-test, the estimated coefficient of $\beta_{3,LRU}$ is relatively small.[27] This suggests a rather minor importance of the Russian market in the long-run structure. 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 movements would not be very large.

Table 11: Final model in the post-accession period

Likelihood ratio test of the restricted model:					$\chi^2(3) = 2.27$, p-value = 0.52		
	LCZ	LHN	LPO	LRU	LWE	LUS	trend
β'_1	1	-0.77	-0.63	0	0	0	0.00
	(.)	(-13.50)	(-7.90)	(.)	(.)	(.)	(2.00)
β'_2	-1	0.82	0	0	1	0	-0.00
	(.)	(13.94)	(.)	(.)	(.)	(.)	(-4.03)
β'_3	0	0	-0.36	-0.05	1	-0.35	-0.00
	(.)	(.)	(-5.61)	(-5.12)	(.)	(-7.90)	(-1.54)

Notes: The table reports the likelihood ratio test for the over-identifying restrictions on $\boldsymbol{\beta}$, the resulting coefficients as well as their t-statistics in the parentheses. Some t-statistics are missing due to the imposed restrictions. LCZ, LHN, LPO, LWE, LUS and LRU stands for the logarithms of the Czech, Hungarian, Polish, Western European, US and Russian indices, respectively.

To check the stability of the imposed over-identifying restrictions, I again conduct a recursive likelihood ratio test. The joint restrictions in Table 11 seem to be plausible for most of the subsamples, as can be seen in Figure 12 in Appendix C. The only problematic period appears in August 2005, when the restrictions are rejected. Since this is only borderline and temporary, the detected instability is not serious.

5.4. Robustness Check

To see if important information is lost by using weekly instead of daily data, I replicate the models with the Tuesday closing prices instead of the Friday closing prices. I adjust only the inclusion of dummies. For instance, since September 11, 2001 was Tuesday, the corresponding dummy is shifted 3 days ahead from Friday to Tuesday. The results show that the cointegration rank and the identified long-run relations are robust to the day of the week used in both the pre- and post-accession periods. More specifically, the same restrictions on the β coefficients are not rejected, and the magnitudes of the estimated coefficients remain similar. Certain differences are found only in the estimated α coefficients, as the Hungarian market in the pre-accession period is suggested to be adjusting to the long-run relation in addition to the Czech market. Considering the similarity of the Hungarian and Czech markets in terms of their size and openness, this finding appears to be plausible and underlines the importance of Polish market movements for the whole Central European region. The α estimates in the post-accession period are found to be unstable as in the case of the Friday data. To further investigate the choice of the data, I also use data in US dollars instead of the local currencies. My analysis confirms the findings in Yang et al. (2006), Koch and Koch (1991) and Bessler and Yang (2003) that the results do not depend substantially on the currency used. In particular, the number of detected equilibrium relations is found to be the same regarding both periods.

[27] The relatively small size of $\beta_{3,LRU}$ is also found under the other identification schemes.

As a last robustness check, I examine alternative model specifications. Including the three dummies for the instability after September 11, 2001, the characteristics of the long-run relation in the pre-accession period do not substantially depend on the lag length of the model or on additional dummies used in the model. Nevertheless, the results for the post-accession period are found to be more sensible to the choice of lag length or the inclusion of dummy variables. Under alternative model specifications, I find stronger support for preferring rank 2 to rank 3 according to the trace test. The two equilibrium relations still detected bridge the Central European and the non-Central European markets, but the "pure Central European relation" is no longer found to be stationary. Hence, the main result that new linkages between the Central European and the other markets emerge in the post-accession period remains unchanged. On the contrary, the degree of cointegration between the Central European markets seems to decrease using some of the different model specifications. However, most of these specifications seriously violate the residual or constancy assumptions and thus are not reliable. These differences in results deliver further evidence about the importance of fulfilled model assumptions for statistical inference.

6. Conclusions

The Central European countries have been the leaders of the transition process from centrally planned towards free market economies. A substantial role in the process of transition was played by developed countries, especially those in Western Europe. For instance, large capital inflows, especially in the form of FDI, represented one important channel for tightening economic relations. Consequently, the Central European countries became members of the EU in May 2004. This study attempts to answer the question whether the political, legal and broad economic integration is accompanied by financial market integration as well. The focus is on the financial integration of the three largest Central European markets with their mature counterpart, the Western European market, but the US and the Russian market are also added to the analysis as representatives of the global economy. The financial integration is studied from the point of long-run convergence towards stable equilibrium relations among the stock markets, as modelled by the Johansen cointegration method. Compared to the previous literature, particular attention is paid to the appropriate application of the technique, and the technique is soundly used to get deeper insights into the development, structure and changes in the long-run relations.

The results show that the Central European stock markets indeed became more integrated with the global economy in general and with the "old" EU in particular after the EU accession. This is evidenced by the emergence of two new equilibrium relations in the post-accession period that link the movements of the Central European markets to the movements of the Western European, US and Russian markets, whereas no such relations can be detected before the EU enlargement. In fact, one new long-run relation could be identified as the "new EU relation" because it connects the developments of the Czech and the Hungarian market to the development of the Western European market representing the "old" EU. The accepted exclusion of the Polish market from this relation can be explained by less openness towards foreign investment and weaker trade integration of the Polish market with the EU compared to the two smaller Central European markets. The other equilibrium relation mentioned represents a linkage among the Western European US and Russian markets as well as the Central European markets, though the role of the Russian market is found to be relatively limited. Hence, the existence of the relation points to the importance of the US stock market to the Central European markets, in addition to the influence from the Western European market, detected after the EU enlargement.

Considering only the three Central European markets, their degree of integration is found to be the same in the period before and after the EU enlargement. More specifically, I detect one cointegration relation linking the three Central European markets in both periods. However, its characteristics changed over time.

Between November 1998 and October 2001, this relation can be characterised by a strong adjustment of the Czech market to movements of the Polish market. Using an alternative day of the week, I also find weak evidence for adjustment of the Hungarian market. These results suggest a major importance of the Polish market, the largest stock market in the region, whose movements were likely to reflect especially local market events, because only a relatively low fraction of investment to the market originated abroad. Nevertheless, the characteristics of the relation changed permanently around November 2002, which roughly corresponds to the end of the EU admission negotiations on December 13, 2002. In particular, the importance of the Polish market is indicated to be smaller in the relation detected after the EU accession in 2004, since a strong link between the Czech and Hungarian markets is found. Considering the rich long-run structure in the post-accession period, the Polish market as an initial driving force of the "Central European relation" in the pre-accession period was likely to be substituted later on by the stochastic trends of the mature markets, in particular by the Western European market. Unfortunately, recursive tests of parameter constancy detect serious instability of the α coefficients, which impedes the confirmation of this hypothesis by the data.

This study finds evidence for a significantly stronger financial integration of the Central European markets with the global economy after the EU enlargement in 2004, particularly with the "old" EU. I have shown that new long-run linkages between the Central European markets and the developed markets in Western Europe and the US emerged after the EU accession, though no such relation could be found before the EU enlargement. From the perspective of stock market investors, the results suggest that the benefits of long-run portfolio diversification between the developed and the Central European markets were reduced.

Appendix A. Model Specifications

Appendix A.1. The Pre-accession Period

As shown in Table 12, the normality of the residuals is clearly rejected for the basic VAR(3) model in the pre-accession period without any dummy variables.[28] The assumption of independent residuals implies no autocorrelation. This is violated for model with lag 2, but not with lag 3 and 1. Furthermore, the tests for ARCH effects reject the null hypothesis of no autocorrelation in second moments and detect heteroscedasticity in the residuals for every lag length reported. Moreover, several large standardised residuals (over 3.5) could be detected. Therefore, the specification of the model is not satisfactory and the situation does not improve when additional lags are included.

Table 12: Misspecification tests

		Pre-accession period		Post-accession period	
Test	DF	Basic model	Adjusted model	Basic model	Adjusted model
Normality:	12	62.3 [0.00]	**17.2 [0.14]**	29.1 [0.00]	**14.0 [0.30]**
Autocorrelation:					
LM(1):	36	44.8 [0.15]	34.1 [0.56]	39.1 [0.33]	**35.6 [0.49]**
LM(2):	36	56.3 [0.02]	34.3 [0.55]	31.4 [0.69]	36.0 [0.47]
LM(3):	36	41.1 [0.26]	**33.7 [0.58]**	31.8 [0.67]	34.0 [0.56]
ARCH effects:					
LM(1):	441	545.7 [0.00]	516.1 [0.01]	485.9 [0.07]	**364.7 [0.99]**
LM(2):	882	1038.6 [0.00]	943.4 [0.07]	1059.8 [0.00]	905.8 [0.28]
LM(3):	1323	1519.1 [0.00]	**1479.3 [0.00]**	1487.6 [0.00]	1319.2 [0.52]

Notes: The table reports misspecification tests for multivariate normality proposed in Doornik and Hansen (2008) (H_0 : normality), Lagrange Multiplier tests for autocorrelation (H_0 : no autocorrelation) as well as ARCH effects (H_0 : no autocorrelation in second moments) in the residuals for models with one to three lags (Anderson, 2003; Rao, 1973). All the test statistics are χ^2 distributed. DF denotes degree of freedom. P-values are reported in parenthesis. Adjusted model refers to the model with dummies (and specific lag length in the pre-accession period). The results of tests for the models used in the following are indicated by boldface.

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 deviations from normality (Juselius, 2007, page 128) and presence of ARCH effects (Gonzalo, 1994; Lee and Tse, 1996). However, appropriate dummy variables might mitigate autocorrelation by the elimination of large residuals and improve the skewness and kurtosis of the residual distribution. Therefore, I include several dummies, as listed in the main text.

I further adjust the lag length of the model. Generally, it is set to 3, because no autocorrelation in residuals is rejected for the VAR(2) model, but not for the VAR(3) model. A longer lag structure appears to be redundant, and a shorter structure is rejected by the tests for lag reduction. Nevertheless, the examination of the coefficients of the Γ_1 and Γ_2 matrices in the VECM specification[29] indicates that the columns $\Gamma_{1,LRU}$, $\Gamma_{2,LWE}$ and $\Gamma_{2,LUS}$ contain only insignificant coefficients and can be individually as well

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

[29]Note that the lag length of 3 in the VAR form corresponds to the lag length of 2 in the VECM form and, thus, to only two Γ matrices.

as jointly excluded. For instance, the likelihood ratio test statistic of the joint hypothesis of exclusion is $LR = 2(3333.8 - 3322.7) = 11.1$, which is smaller than the critical value $\chi^2_{0.95}(18) = 28.9$. Therefore, I use this more parsimonious structure of lags.

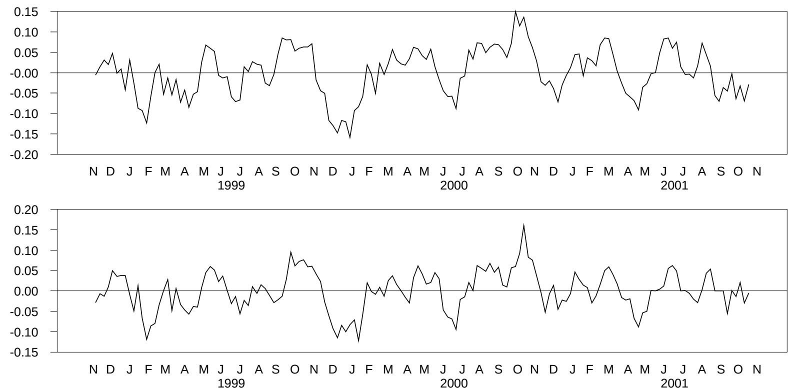
The misspecification tests for the resulting model (Adjusted model in Table 12) show that both no autocorrelation and normality of the residuals have improved substantially. Hence, the extended model is preferred to the basic one.

Appendix A.2. The Post-accession Period

The basic VAR model for the post-accession period surprisingly does not suffer from autocorrelation in residuals. However, non-normality, ARCH effects and large residuals are still detected (Table 12, Basic model). The inclusion of the dummies introduced in the main text improve the model substantially (see Adjusted model in Table 12). Furthermore, the lag length is set to 1 for two reasons. First, the likelihood ratio tests for reducing lag length from 3 to 1 ($\chi^2(72) = 82.9$ with a p-value 0.18) and from 2 to 1 ($\chi^2(36) = 39.98$ with a p-value 0.30) do not reject the hypothesis of the sub-model with lag length 1. Second, the misspecification tests for this model do not detect any residual autocorrelation, non-normality or even presence of ARCH effects and thus can be used further on.

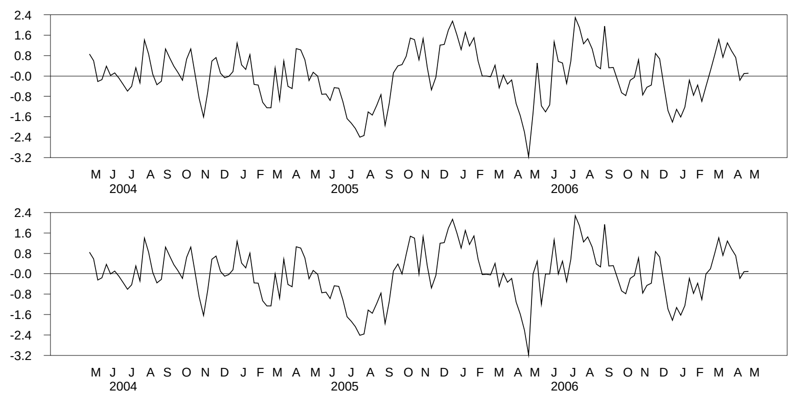
Appendix B. Cointegration Relations

Figure 3: The first cointegration relation in the pre-accession period



Notes: The figure plots the first cointegration relation in the pre-accession period in the "X-form" (upper part) and "R-form" (lower part).

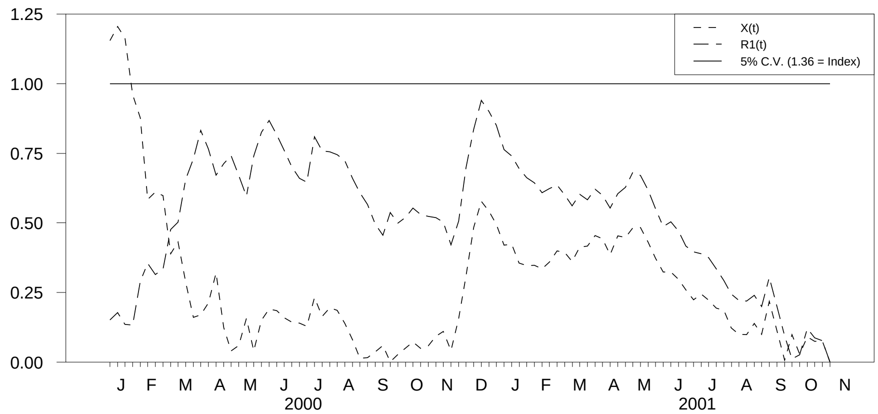
Figure 4: The third cointegration relation in the post-accession period



Notes: The figure plots the third cointegration relation in the post-accession period in the "X-form" (upper part) and "R-form" (lower part).

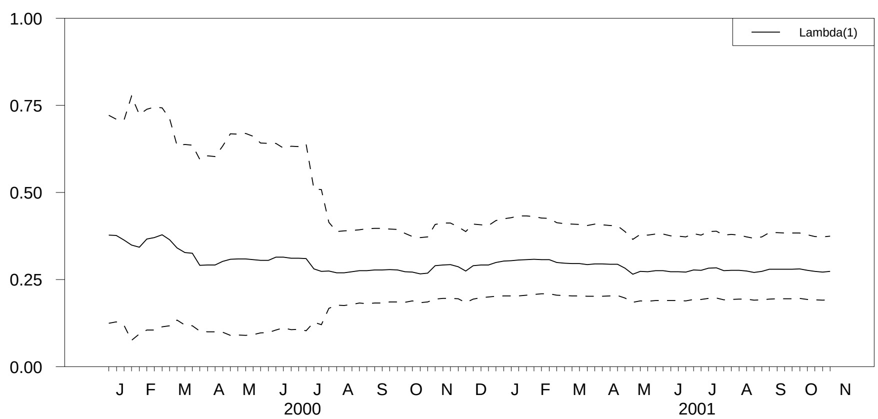
Appendix C. Constancy of Parameters

Figure 5: Test for constancy of log-likelihood function in the pre-accession period



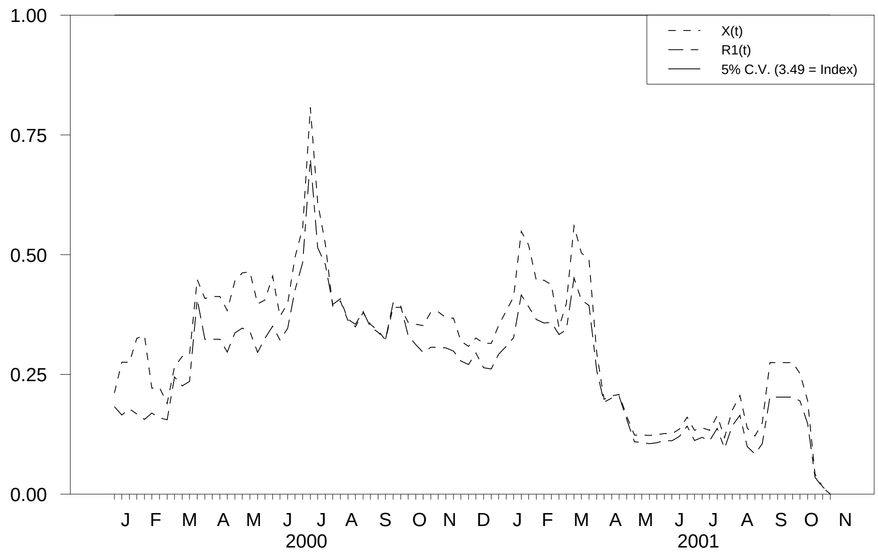
Notes: The scaling of the test is consistent, with 1 being the 5 % rejection line. The values under the rejection line indicate no violation of the constancy assumption. The line of R1(t) refers to the concentrated model in the "R-form".

Figure 6: Development of the first eigenvalue in the pre-accession period



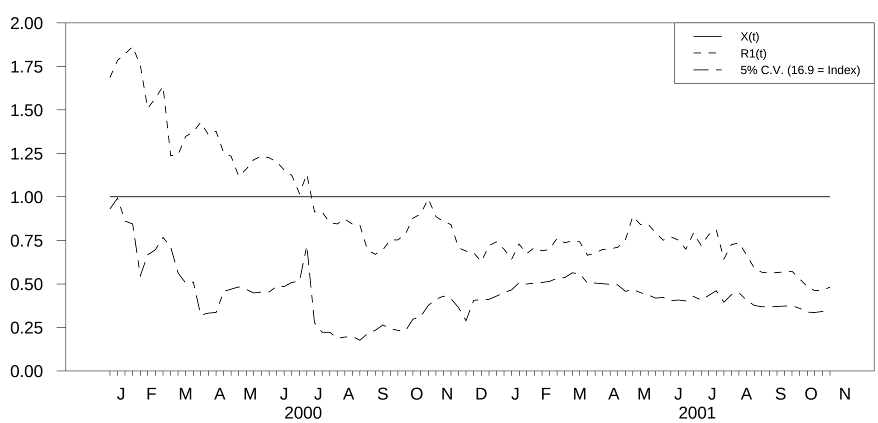
Notes: The dashed lines refer to 5 % confidence bounds. If the eigenvalue lies within the narrowest confidence bounds, the assumption of the eigenvalue's constancy is not violated.

Figure 7: Test of β constancy in the pre-accession period



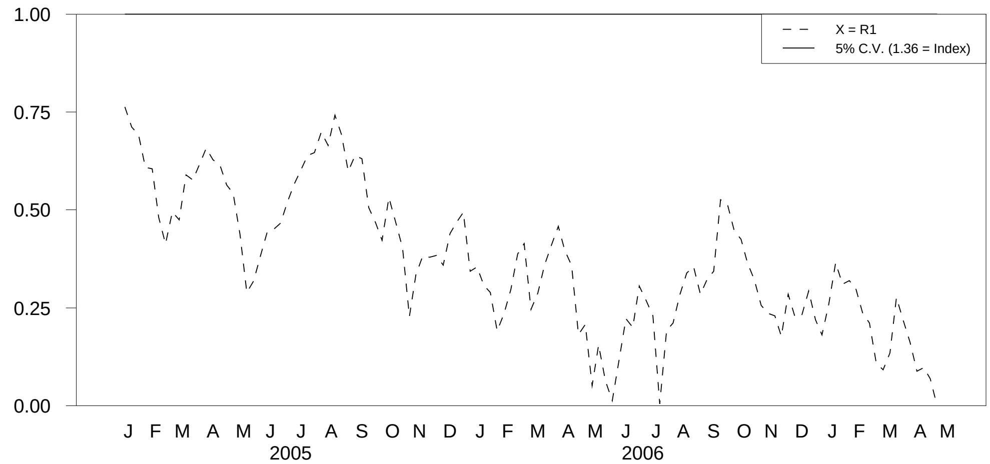
Notes: The scaling of the test is consistent, with 1 being the 5 % rejection line. The values under the rejection line indicate no violation of the constancy assumption. The line of R1(t) refers to the concentrated model in the "R-form".

Figure 8: Likelihood ratio test of restrictions in the pre-accession period



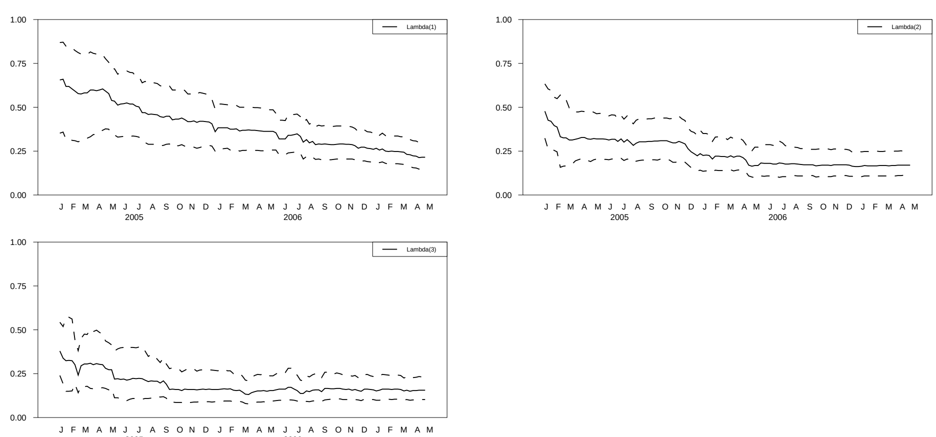
Notes: The scaling of the test is consistent, with 1 being the 5 % rejection line. The values under the rejection line indicate that the imposed restrictions are not rejected. The line of R1(t) refers to the concentrated model in the "R-form".

Figure 9: Test for constancy of log-likelihood function in the post-accession period



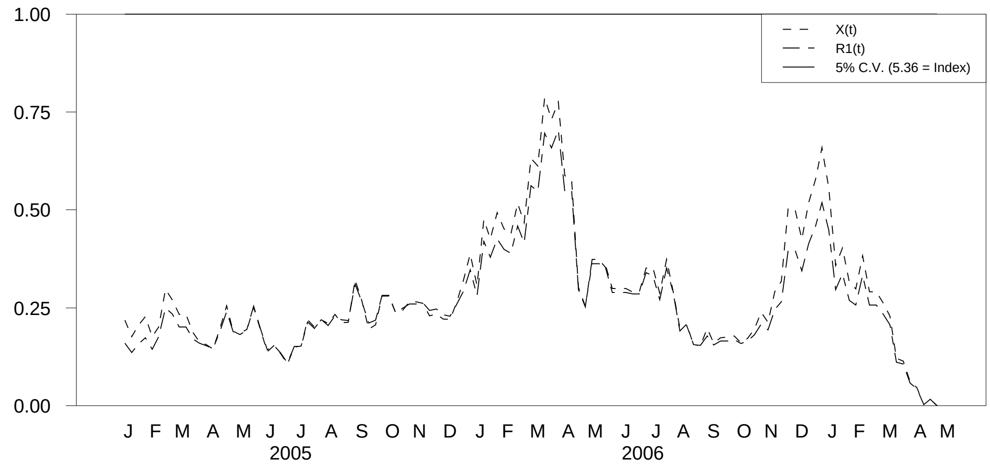
Notes: The scaling of the test is consistent, with 1 being the 5 % rejection line. The values under the rejection line indicate no violation of the constancy assumption. Due to lag length of 1 in the VAR model (i.e. zero lag in the VECM form), the test statistics for the "X-" and "R-form" are the same.

Figure 10: Development of the three eigenvalues in the post-accession period



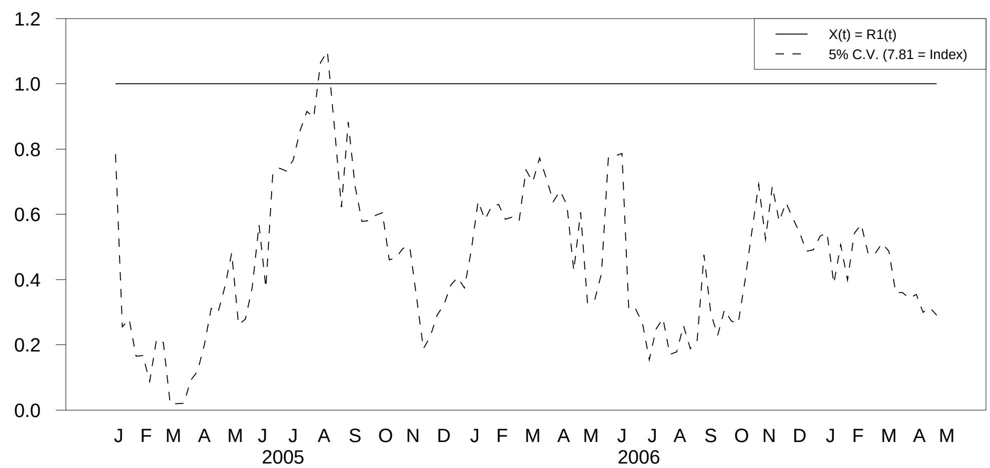
Notes: The dashed lines refer to 5 % confidence bounds. If the eigenvalue lies within the narrowest confidence bounds, the assumption of the eigenvalue's constancy is not violated.

Figure 11: Test of β constancy in the post-accession period



Notes: The scaling of the test is consistent, with 1 being the 5 % rejection line. The values under the rejection line indicate no violation of the constancy assumption. The line of R1(t) refers to the concentrated model in the "R-form".

Figure 12: Likelihood ratio test of restrictions in the post-accession period



Notes: The scaling of the test is consistent, with 1 being the 5 % rejection line. The values under the rejection line indicate that the imposed restrictions are not rejected. Due to lag length of 1 in the VAR model (i.e., zero lag in the VECM form), the test statistics for the "X-" and "R-form" are the same.

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