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Abstract

We study the stock market integration of emerging CEE-3 stock markets (namely, the Czech, Hungarian, and Polish markets) and hypothesize that this process has been gradual over time. As a proxy for integration, co-movements with three stock market indices that represent the developed markets (i.e., MSCI Germany, the Dow Jones Euro Stoxx 50, and MSCI World) are estimated using the standard, asymmetric, and corrected DCC-GARCH model. A smooth transition logistic trend model is then fitted to the dynamic correlations to examine the integration process. Evidence of strengthening relationships among the markets under study is provided.

Keywords: stock market integration; dynamic conditional correlations; CEE-3 countries; smooth transition model

JEL codes: C32, G01, G15

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Research area: finance

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Introduction

Over the last 40 years, studies investigating stock market integration have been related to portfolio theory and effective international diversification (for earlier works, see, e.g., Grubel, 1968; Ripley, 1973; Lessard 1974, 1976; Panton et al., 1976; Hilliard, 1979). Changes in the covariance structure of cross-country stock returns affect diversification opportunities, the attractiveness of foreign stock markets, sensitivity to returns and volatility spillovers and the cost of capital; thus, these changes affect investment decisions per se. As stock markets may predict the development of the real economy, stock market integration is of particular interest to policy makers.

An increase in cross-market linkages was observed in the 1980s (Jaffe and Westerfield, 1985; Schöllhammer and Sand, 1985; Asprem, 1989; Eun and Shim, 1989), especially after the US stock market crash in October 1987. This event shed light on the emerging markets, where investors might still benefit from their low interdependence with developed and well-examined markets. However, emerging markets are unique due to the potential existence of barriers that may discourage foreign investors (see Bekaert, 1995, or Bekaert and Harvey, 2002). Since the 1980s, many emerging markets throughout the world have implemented liberalization policies to render themselves more attractive to international investors, and as a consequence, these markets have become more integrated with developed markets. Of course, it cannot be expected that all of the barriers will be eliminated at once, and therefore, the transition process from segmented to integrated markets should occur gradually.

During the last 15 years, the CEE-3 countries have witnessed several institutional and economic events that should support the above hypothesis of progressive co-movements. For example: (i) European Union accession negotiations have gradually improved the regulatory and legal framework to offer higher investor protection and, thus, to relax investment barriers; (ii) the trading platforms on the CEE-3 stock exchanges have improved; (iii) in May 2004, the CEE-3 countries became members of the EU; (iv) there has been enduring interest from foreign investors (see Private Capital Flows in Appendix B); (v) currencies have entered floating exchange rate regimes; (vi) events such as financial crises; and specifically, (vii) EU debt crisis, may also increase the co-movements.

In this paper, to verify the hypothesis of gradual stock market integration, we apply the smooth transition logistic trend model, which is fitted to time-varying correlations between the emerging European markets of the CEE-3 (namely, the Czech Republic, Hungary, and Poland) and three stock market indices that represent the developed markets (i.e., MSCI Germany, the Dow Jones Euro Stoxx 50, and MSCI World). First, we show that during the recent financial crisis, long-run unconditional correlations between the CEE-3 and developed markets have increased. Next, the dynamic conditional correlations are estimated using the standard DCC-GARCH model (henceforth, DCC), the asymmetric DCC-GARCH model (ADCC), and the corrected DCC-GARCH model (cDCC). These correlations represent the degree of stock market co-movements, which we use as a proxy for stock market integration. Using the non-linear smooth transition logistic model, we provide evidence for the hypothesis of gradual stock market integration between the Czech and Polish stock markets and the developed markets. The evidence of a smooth transition for the Hungarian stock market is less convincing.

The remainder of the paper is organized as follows: in Section 1, we briefly discuss the history and peculiarities of the CEE-3 markets. Section 2 discusses the related literature of empirical studies examining CEE-3 stock markets with an emphasis on those that use a methodology similar to ours. Section 3 describes the data and Section 4 explains the applied methodology. In Section 5, the estimated correlations are presented along with the results from the smooth transition model. Finally, Section 6 provides concluding remarks.

1. The history and peculiarities of the CEE-3 stock markets

Since the early 1990s, the Czech Republic[1], Hungary, and Poland (which are denoted as the CEE-3 countries) underwent a drastic transformation from centrally planned to market-oriented economies. Prior to 1990, all financial transactions were part of the central plan. Banks were responsible for both keeping records of the transactions among economic entities and collecting household deposits (Hermes and Lensink, 2000; Bonin and Wachtel, 2003). For example, in Czechoslovakia, the financial intermediation toward private subjects was limited or virtually non-existent. Therefore, one of the earlier tasks of the first democratically elected governments (after the dissolution of the Eastern Bloc in 1989) was to design a new financial system.

In many developed countries, stock markets are integral parts of the existing financial system. The establishment of stock markets in the CEE-3 was not a market-driven process. Instead, these stock markets were established by the governments, as they were needed to support the various forms of privatization of state-owned enterprises (SOEs). This privatization

[1] This country was known as Czechoslovakia until its federal parliament dissolved it into the Czech Republic and Slovakia in January 1, 1993. Note that we are not dealing with the Slovak stock market due to its small size, the small number of actively traded stocks, the low level of liquidity, a near absence of initial public offerings, and related factors. This market is practically inefficient and reports very low correlations with other markets (both within the CEE region and with the developed markets); see Baumöhl (2013).

makes the stock markets of the former Eastern Bloc distinct from other emerging markets around the world.[2]

Different CEE-3 countries took different approaches to privatization, and these approaches had different impacts on the formation and role of the stock market. In contrast to Poland, where the agricultural sector was partially in private ownership during the 1980s, and Hungary, where small businesses were allowed, Czechoslovakia was an extreme with virtually no private ownership before the 1990s (see Shafik, 1995). Furthermore, compared to the book value of state-owned assets offered for privatization, the private savings in Czechoslovakia were low. Therefore, it was feared that foreign investors could easily purchase assets at the expense of domestic investors (e.g., households). At that time, this possibility was politically inadmissible (see Mejstřík and Burger, 1994; Bornstein, 1997; Megginson and Netter, 2001). Consequently, the government in Czechoslovakia (in the first wave of privatization) and later in the Czech Republic (in the second wave of privatization) carried out a unique mass privatization based on vouchers, which were sold (with a discount) to the adult population (see Brada, 1996; Bornstein, 1997).[3] The privatized companies were required to be listed on the newly established stock market(s) (Bonin and Wachtel, 2003): "(Mass) *Privatization led to extensive equity ownership of firms but often the stakes were quite small and the holdings were dispersed. Thus, all of a sudden some transition economies* (Czechoslovakia) *found themselves with extensive equity markets.*" In 1993, 955 securities from the first wave of privatization were listed on the PSE, and an additional 674 securities emerged from the second wave of privatization in 1995.[4] Although this mass privatization resulted in a large number of listed companies, there was insufficient liquidity. Indeed, mass privatization led to a mass "tunneling," in which firm managers, investment funds (which owned most of the vouchers), and other majority shareholders expropriated the resources of the companies at the expense of the other (minority) shareholders (Johnson et al., 2000). As a consequence, from 1996 to 1997, 83% of the local companies were delisted from the PSE (Caviglia et al., 2002). More importantly, during the 1990s, no new capital was raised on the PSE (see Kominek, 2004; Glaeser et al., 2001), which is historically the primary function of stock markets. In a well-functioning stock market, companies need to have good corporate governance to attract capital. However, this requirement did not exist in the Czech Republic, as the stock market was not established to raise capital; instead, the capital was

[2] The Warsaw stock exchange (WSE) was established on April 16, 1991; the Budapest stock exchange (BSE) was established on June 21, 1990; and the Prague stock exchange (PSE) was established on November 24, 1992. Note that these dates do not necessarily correspond to the first trading days.
[3] The CEE-3 countries carried out several forms of privatization, but we focus on those forms that determined the establishment and subsequent development of their stock markets.
[4] The data were obtained from the homepage of the PSE: www.pse.cz.

merely redistributed (see Stringham et al., 2008). Although the regulatory framework in the Czech Republic improved during the 2000s (see Groh and von Liechtenstein, 2009), IPOs are still quite uncommon, which is most likely due to trust-related issues.

Initially, Hungary and Poland employed a small-scale privatization that utilized a case-by-case approach (this form of privatization proved to be more successful than mass privatization; see Megginson and Netter, 2001). Some of the larger SOEs were sold using initial public offerings (IPOs) (Bornstein, 1997; Caviglia et al., 2002), which resulted in a much lower number of listed companies (see Appendix B). In Poland, a distinctive form of mass privatization started as late as 1995, and it combined different forms of privatization methods. For example, whereas stakes of the SOE were offered to citizens and employees, a large portion of the shares was transferred to national investment funds that could sell their shares to foreign investors. A certain stake remained in state-ownership as well. The national investment funds later became listed on the WSE (Caviglia et al., 2002; Bonin and Wachtel, 2003).

The situation of the BSE might not seem very different from the situation of the PSE, as after the first decade of its existence, the BSE had experienced only a limited number of IPOs. However, the current number of issuers in the BSE is larger than in the PSE, and the turnover ratio of the BSE is the highest among all of the CEE-3 markets (Appendix B). A higher turnover ratio might be associated with more transparent corporate governance and more careful selection of the enterprises that enter the stock market, which might increase the credibility of the market as a whole and, thus, help the BSE in the future. The business of IPOs is flourishing exclusively on the WSE. In addition, the number of listed companies has steadily increased, and furthermore, listings and investors from countries other than Poland have been attracted.

Appendix B summarizes the basic statistics of the CEE-3 economies and their stock markets. To facilitate comparison, we supplemented these statistics with data on other European countries (stock markets) where the banking sector dominates the financial system. We also included data from the United States of America and Great Britain.

The peculiarities of the CEE-3 stock markets can be summarized as follows: (i) the financial sectors in the CEE-3 are dominated by banks; (ii) the stock markets are the by-product of the privatization of SOEs; (iii) due to the lower number of listed companies in the PSE and BSE, these markets have high concentrations (according to market capitalization); (iv) as the first issuers were SOEs, enterprises that are not from the financial and energy sectors are underrepresented in CEE-3 markets; (v) since the beginning of the 2000s, the process of raising capital through IPOs is stagnating for the PSE, is very limited in the BSE, but is growing in the WSE; (vi) the relative liquidity is comparable to several other European stock markets; (vii) in nominal values, the market capitalization of the CEE-3 stock markets is low, but compared to these countries nominal GDPs, these markets' capitalizations are similar to some smaller stock markets in more developed economies of Europe.

2. Related literature

Since the beginning of the negotiations for the accession of the CEE countries to the EU, the literature that focuses on CEE markets has become more extensive. After the expansion of the EU in 2004, several empirical works found significant increases in the stock market co-movements of the new member states from the CEE region.

Cappiello et al. (2006b) found evidence of increased integration for Hungary, the Czech Republic, and Poland using daily data from January 1994 to November 2005. In addition, Wang and Moore (2008) found an increasing level of integration of the CEE-3 toward EU markets (based on the aggregate Euro-zone index of the 12 EMU markets) over the sample period 1994 – 2006, and the dynamic conditional correlations (DCCs) at the end of the examined period were approximately 0.3 – 0.5. Using weekly data, Savva and Aslanidis (2010) found that from 1997 to 2008, the DCCs between the Polish and Czech stock market and the Eurozone increased, and the DCCs for the Hungarian stock market remained constant but high.[5] Similar results were found by Syllignakis and Kouretas (2011) using weekly data from October 1997 to February 2009. Additionally, Syllignakis and Kouretas (2011) found that during the financial crisis of 2007 – 2009, the correlation increased from approximately 0.5 to 0.75 (between the CEE-3 countries and the German and US stock markets). Subsequently, a positive relationship was found between the levels of the DCCs and the conditional volatility. Moreover, Gjika and Horvath (2012) used daily data from December 2001 to October 2011 to examine the time-varying correlations between the CEE-3 and the Eurozone (Stoxx 50), which was performed using the ADCC model framework. The conditional correlations increased significantly after the CEE-3 countries entered the EU (May 2004) and remained at these high levels (approximately 0.6 – 0.7) during the recent financial crisis. Again, Gjika and Horvath (2012) showed that the levels of the DCCs can be linked to the levels of conditional volatility on the corresponding markets. Baumöhl et al. (2011) used a different approach: using daily data from January 1998 to March 2010, they linked the levels of DCCs (among the CEE-3 and developed stock markets in Germany and the US) with regimes of the unconditional volatility of stock market returns. In general, an increase in

[5] They use a Smooth Transition Conditional Correlation GARCH model, which is similar to the methodology employed in this paper in that it allows one to model the time-varying correlation by means of one or two transition variables.

volatility was associated with higher levels of DCCs (and vice versa), which provides evidence for the presence of a shift-contagion effect. The estimated correlations found by Baumöhl et al. (2011) ranged from 0.5 to 0.7 at the end of the examined sample, and a sharp peak was detected during the recent financial crisis. Horvath and Petrovski (2013) compared stock market correlations (which were estimated in the BEKK-GARCH framework) between Western Europe (Stoxx Europe 600) and the markets of countries in the CEE-3 and Southeastern Europe (Croatia, Macedonia, and Serbia) over the period 2006 – 2011. After comparing these two groups of emerging European markets, the authors concluded that the degree of integration is much higher in the CEE-3 countries (as the conditional correlations vary but remain at approximately 0.6 with no visible pattern).

In addition, an increase in the stock market integration that is related to the entry of the CEE-3 countries into the EU, the financial crisis, or other crises, is reported in studies that use other methodologies. Using daily data from July 1995 to February 2005, Gilmore et al. (2008) found that the long-term (cointegration) relationship between CEE-3 and the German or UK stock market is disrupted by episodes dominated by short-term domestic factors. The authors conclude that there is little evidence of steady progress toward a long-term equilibrium in the relationship of the CEE equity markets with the markets of the UK and Germany. On a sample of daily data from October 2000 to April 2007, Christiansen and Ranaldo (2009) used coexceedance (the joint occurrence of extreme returns in different markets) within ten CEE countries (i.e., new EU member states, which included the CEE-3) as a measure of stock market integration. Among other results, the authors found that the coexceedance in the ten CEE countries was associated with the coexceedance in the existing EU countries, which they interpreted as a sign of increased stock market integration. Pukthuanthong and Roll (2009) asserted that stock market integration can be measured by the proportion of a country's stock market return that is explained by global factors. Their empirical analysis included the Czech, Polish, and Hungarian stock markets. Using daily data, these authors found that from the beginning of the CEE-3 stock markets to February 2008, the proportion of variance explained by global factors increased linearly (see their Table 3, p. 226). This pattern might be interpreted as evidence for a gradual increase in stock market integration. Similar logic was recently put forward by Hooy and Lim (2013, p. 32), who suggested that "*when a market becomes more globally integrated, the explanatory power of global over local factor increases, leading to a faster incorporation of global information into stock prices and hence a higher degree of informational efficiency.*" On a sample of weekly data from the beginning of 1995 to 2007, their results suggest that among the CEE-3 countries, Poland had the highest stock market efficiency and the highest level of stock market integration. Graham et al. (2012) used weekly data from January 2001 to April 2010 and a method based on wavelet squared coherency between emerging stock markets (including the CEE-3) and the US stock market. These authors found that since 2006, long-term co-movements have evolved into short-term co-movements. Finally, the studies by Lucey and Zhang (2010) and Frijns et al. (2012) provide a unique view of the stock market integration. Lucey and Zhang (2010) found that cultural distance[6] can explain the stock market integration (which is measured by unconditional and conditional correlations) both within emerging markets (including the CEE-3) and between emerging and developed markets. Similarly, Frijns et al. (2012) argued that political crises have a direct impact on the level of Foreign Direct Investments and that these crises increase the risk aversion of foreign investors. Consequently, political crises should decrease stock market integration. These authors were able to provide empirical evidence that supports their hypothesis of a negative relationship between political crises and stock market integration. Their sample included 49 countries, such as the CEE-3 and other emerging countries, as emerging countries are in general more prone to political crises.

Although a substantial increase in correlations is reported in almost all of the relevant studies, in this paper, we are interested in describing the process of such an increase. Recall from the Introduction that the underlying theory suggests that stock market integration should be a gradual process.

To our knowledge, the first time the smooth transition logistic trend model was applied within the literature on stock market integration was by Chelley-Steeley (2004). She analyzed a sample of Asia-Pacific emerging markets (Korea, Taiwan, Thailand, and Singapore) and developed markets (US, UK, Canada, France, Germany, and Japan) over the period from January 1990 to January 2002. In her second paper, Chelley-Steeley (2005) examined the integration of equity markets in the CEE-3 and Russia with respect to the US, the UK, Germany, Japan, and France during the period from July 1994 to December 1999. In both papers, she applied the smooth transition logistic trend model (which is described in Section 4) to bivariate Pearson's correlations, which were calculated for each month using the daily returns within the corresponding month.

Using the DCC-GARCH model to estimate the time-varying correlations, a smooth transition logistic trend model was similarly utilized by Lahrech and Sylwester (2011) to establish the degree of stock market integration between the US and Latin American stock

[6] This factor is measured by religious communality and using a measure proposed in Kogut and Singh (1988).

markets in the period from December 1988 to March 2004. The same approach was applied by Durai and Bhaduri (2011) on a sample of markets in the US, UK, Germany, India, Malaysia, Indonesia, Singapore, South Korea, Japan, and Taiwan over the period from July 1997 to August 2006.

Our study differs from previous studies in several ways. First, we use more recent data that incorporate not only the financial crisis but also data from after the crisis. This approach is of particular importance, as short periods of increased integration may cause only short-term contagion, which, in turn, might make the results spurious. Second, we include not only the Eurozone and US markets but also a proxy for the world's stock markets, which allows us to observe the differences in the integration between the main trading partners of the CEE-3 and the world in general. In addition, we verify the stability of our results based on different DCC models and GARCH specifications, and most importantly, as we use weekly data, we verify the sensitivity of our results with regard to the different types of construction of weekly returns.

3. Data description

We utilized the daily closing prices, which start on 1st January 1999, and end on 28th September 2012, of the stock market indices from the CEE-3 countries, namely, the Czech Republic (PX), Hungary (BUX), and Poland (WIG). Weekly returns were constructed using the log differences of averages taken from the daily prices of a given week. We used the local currencies, as using a common currency would confound the results with respect to the evolution of exchange rates. For example, it is known that a decrease in the exchange rate volatility can increase the stock market integration because it is an important source of the risk that is priced in capital markets (e.g., see Fratzscher, 2002). This position is also consistent with currency hedged international equity portfolios. Our proxy of developed markets included the following: the Dow Jones Euro Stoxx 50 (STX50), MSCI Germany (GER), and MSCI World (MSCI). Although Germany is an important trading partner of the CEE-3 countries, Europe as a whole represents an important region with which the CEE-3 economies tend to fully integrate. The MSCI World index was included to measure the tendencies of the integration of the CEE-3 with world stock markets in general. The data source and a further description of the data used in this study are provided in Appendix A.

4. Methodology

Let $\mathbf{r}_t = (r_{1,t}, \ldots, r_{k,t})^{\mathrm{T}}$, $t = 1,2,\ldots,T$ denote the excess returns[7] of k assets, with $\mathrm{E}[\mathbf{r}_t|F_{t-1}] = \mathbf{0}$ and $\mathrm{Var}[\mathbf{r}_t|F_{t-1}] = \mathbf{H}_t$, where F_{t-1} is the information set that is available at time t–1. The DCC model of Engle and Sheppard (2001) and Engle (2002) is defined as follows:

$$\mathbf{H}_t = \mathbf{D}_t \mathbf{R}_t \mathbf{D}_t \tag{1}$$

$$\mathbf{R}_t = diag\{\mathbf{Q}_t\}^{-1/2} \mathbf{Q}_t diag\{\mathbf{Q}_t\}^{-1/2} \tag{2}$$

$$\mathbf{Q}_t = (1-\varphi-\psi)\overline{\mathbf{Q}} + \varphi \boldsymbol{\varepsilon}_{t-1}\boldsymbol{\varepsilon}_{t-1}^{\mathrm{T}} + \psi \mathbf{Q}_{t-1} \tag{3}$$

where $\mathbf{D}_t$ is a diagonal matrix of the time-varying conditional standard deviations, $\boldsymbol{\varepsilon}_t$ are standardized excess returns, and $\mathbf{R}_t$ is the time-varying correlation matrix. In addition, $\overline{\mathbf{Q}}$ is estimated via the moment estimator $T^{-1}\sum_{t=1}^{T} \hat{\boldsymbol{\varepsilon}}_t \hat{\boldsymbol{\varepsilon}}_t^{\mathrm{T}}$. The following restrictions are imposed, which ensure that the matrix $\mathbf{Q}_t$ is positive definite: the scalar parameters $\varphi, \psi \geq 0$ and $\varphi + \psi < 1$. Then, the DCCs are calculated as follows:

$$\rho_{ij,t} = \frac{q_{ij,t}}{\sqrt{q_{ii,t}\, q_{jj,t}}},\ i, j = 1,2,\ldots,k; i \neq j \tag{1}$$

Cappiello et al. (2006a) considered a specification that allows for the incorporation of asymmetries in the correlation dynamics:

$$\mathbf{Q}_t = (1-\varphi-\psi)\overline{\mathbf{Q}} - \xi \overline{\mathbf{N}} + \varphi \boldsymbol{\varepsilon}_{t-1}\boldsymbol{\varepsilon}_{t-1}^{\mathrm{T}} + \psi \mathbf{Q}_{t-1} + \xi \mathbf{n}_{t-1}\mathbf{n}_{t-1}^{\mathrm{T}} \tag{5}$$

where $\overline{\mathbf{N}} = T^{-1}\sum_{t=1}^{T} \mathbf{n}_t \mathbf{n}_t^{\mathrm{T}}$, $\mathbf{n}_t = I[\boldsymbol{\varepsilon}_t < 0] \circ \boldsymbol{\varepsilon}_t$, $I[.]$ is a $k \times 1$ indicator function that takes the value of 1 if the argument is true (and 0 otherwise), and "$\circ$" represents the Hadamard product. The positive definiteness of $\mathbf{Q}_t$ is ensured in a similar manner: $\varphi, \psi, \xi \geq 0$, and $\varphi + \psi + \delta\xi < 1$, where δ = the maximum eigenvalue $[\overline{\mathbf{Q}}^{-1/2}\overline{\mathbf{N}}\overline{\mathbf{Q}}^{-1/2}]$ (for more details, see Cappiello et al., 2006a). The entire analysis was conducted with the R software using the *rmgarch* (Ghalanos, 2012a) and *rugarch* (Ghalanos, 2012b) packages.

Standardized excess returns $\boldsymbol{\varepsilon}_t$ were obtained from univariate ARMA-GARCH models. As stated by Cappiello et al. (2006a), the correlation estimates are inconsistent when univariate models are not well specified, although different GARCH models usually produce similar volatility patterns. However, to minimize the risk, we chose from among several GARCH model specifications (standard GARCH, AVGARCH, NGARCH, EGARCH, GJR-GARCH, APARCH, TGARCH, and CSGARCH), to determine the one that best fit the data according to the Bayesian information criterion (BIC). We allowed up to two lags in the GARCH specifications and up to

[7] These returns were obtained from the mean equations, as described below.

five lags in the ARMA specifications to control for the autocorrelation and ARCH effects in the standardized residuals; the presence of both these issues was tested using the Ljung-Box test at a 5% significance level. Following Nelson (1991), instead of the normality condition on the distribution of excess returns, we utilized a generalized error distribution (GED).

The standard DCC model suffers from the downward bias of the estimated parameters, although this problem seems to be particularly severe in higher dimensions (Hafner and Reznikova, 2012). Indeed, Aielli (2013) shows that the moment estimator $\overline{\mathbf{Q}}$ may be biased and inconsistent. To double-check our results, we also checked their sensitivity using the corrected DCC model (cDCC) of Aielli (2013). Instead of the standard DCC model, he proposes to model the following process:

$$\mathbf{Q}_t = (1-\varphi-\psi)\overline{\mathbf{Q}}^* + \varphi\boldsymbol{\varepsilon}^*_{t-1}\boldsymbol{\varepsilon}^{*\mathrm{T}}_{t-1} + \psi\mathbf{Q}_{t-1} \tag{6}$$

where $\boldsymbol{\varepsilon}^*_t = \mathrm{diag}\{\mathbf{Q}_t\}^{1/2}$. $\boldsymbol{\varepsilon}_t$ and $\overline{\mathbf{Q}}^*$ can now be consistently estimated via a sample covariance matrix of $\boldsymbol{\varepsilon}^*_t$.

After the time-varying correlations are obtained, we verified that the stock market integration[8] can be considered a gradual process. The non-linear smooth transition logistic model suggested by Granger and Teräsvirta (1993) is fitted in the following form[9]:

$$\rho_{ij,t} = \alpha + \beta S_t(\gamma,\tau) + \delta\rho_{ij,t-1} + \upsilon_t \tag{7}$$

where $\rho_{ij,t}$ are the estimated bivariate dynamic conditional correlations, α, β, and δ are the regression parameters, and υ_t is the error term. The logistic function $S_t(\gamma,\tau)$ is defined as follows:

$$S_t(\gamma,\tau) = \left(1+\exp(-\gamma(t-\tau T))\right)^{-1}, \gamma > 0 \tag{8}$$

where T is the sample size, τ is the parameter determining the transition midpoint between the two regimes, and γ measures the speed of the transition. For small values of γ, we may consider integration from the first regime $\alpha/(1–\delta)$ to $(\alpha+\beta)/(1–\delta)$ to be slow and gradual. For larger values of γ, the shift between the two regimes occurs more quickly.[10]

[8] It is worth recalling from the Introduction that we use correlations as a proxy for real stock market integration.
[9] The subscripts i and j for the regression parameters are suppressed for the sake of clarity.
[10] Estimation of (7) is not straightforward. A grid of feasible values for parameters γ and τ was first constructed and then model (7) was estimated, using this grid, for each of the fixed γ and τ parameters using conditional least squares. With 1190 values for γ and 50 for τ, 59,500 estimates were derived for each parameter α, β, and δ. After discarding obvious outliers, we also obtained intervals for α, β, and δ. These intervals were used to conduct 20 random searches of 10,000 itterations each. This random-search approach resulted in 20 models. The parameter estimates of these models were used as starting values in a non-linear least squares estimation of model (7) employing the L-BFGS-B algorithm from Byrd et al. (1995). The resulting parameter estimates then became the starting values for a non-linear least squares estimation of model (7), using the standard Newton and Gauss approach. From the resulting 20 models, the one with the minimum sum of squared residuals was selected.

H_0: $\gamma = 0$ should be tested to determine whether there were changes in the level of integration. Model (7) has the property of being identifiable only under the alternative. A consequence is the *t*-ratio of H_0: $\gamma = 0$ having no usual interpretation. For that reason, we do not report significances (from corresponding *t*-statistics) for γ and τ, but instead use a general polynomial test of changes in the mean:

$$\rho_{ij,t} = \alpha + \delta\rho_{ij,t-1} + \sum_{k=1}^{K} \beta_k t^k + \upsilon_t \tag{9}$$

This specification involves a general test of non-linear changes in the mean, testing H_0: $\beta_1 = \ldots = \beta_K = 0$. These hypotheses are tested using a Lagrange Multiplier (LM) type test. With $K = 3$, this approach is essentially the same as was proposed by Luukkonen et al. (1988) and with $K = 4$ as in Escribano and Jordá (2001). The idea is to have the transition function $S_t(\gamma,\tau)$ approximated by a Taylor expansion under the null hypothesis. If the nonparametric unweighted bootstrap (Cribari-Neto and Zarkos, 1999) version of White's (1980) test suggests the errors in (9) are heteroskedastic, then a heteroskedasticity-robust version of the LM test would be used.

5. Results

In this chapter, we present the main results. First, a brief analysis of the long-run correlations is provided. Second, the estimation results of the DCC models and a basic description of the estimated correlations are discussed. Finally, the smooth transition logistic trend regression model is fitted on correlations to evaluate the pattern of the stock market integration.

5.1 Long-run correlations

We divided the sample into three periods and calculated the long-run correlation coefficients for different bandwidths (see Andrews, 1991).[11] The results are reported in Table 1, and they provide descriptive evidence about the overall increase of the stock market integration among the CEE-3 and developed countries.

[11] Fixed bandwidths were used to allow for direct comparison of the long-run correlations.

Table 1: Long-run correlations

	01/1999 - 12/2002 (*T* = 208)			01/2003 – 06/2007 (*T* = 234)			07/2007 – 09/2012 (*T* = 275)		
	WIG	PX	BUX	WIG	PX	BUX	WIG	PX	BUX
Bandwidth 6									
PX	0.6108	–	–	0.6377	–	–	0.9070	–	–
BUX	0.7854	0.7233	–	0.6529	0.8164	–	0.8813	0.9031	–
GER	0.6166	0.5446	0.6879	0.6340	0.4591	0.4979	0.8813	0.8622	0.8610
STX50	0.6532	0.4580	0.6437	0.6177	0.5453	0.5521	0.8891	0.8634	0.8962
MSCI	0.5987	0.3790	0.5847	0.5848	0.5647	0.5377	0.9210	0.9224	0.9272
Bandwidth 12									
PX	0.7868	–	–	0.4649	–	–	0.9160	–	–
BUX	0.8652	0.7728	–	0.3994†	0.7434	–	0.9362	0.9392	–
GER	0.7335	0.5781	0.7697	0.7045	0.4519	0.4162†	0.8997	0.9088	0.8776
STX50	0.7170	0.5118	0.6972	0.6478	0.5187	0.4547	0.9134	0.8991	0.9077
MSCI	0.6830	0.5066	0.6615	0.5775	0.5061	0.3456‡	0.9223	0.9485	0.9094
Bandwidth 18									
PX	0.8723	–	–	0.2651‡	–	–	0.9416	–	–
BUX	0.8570	0.7970	–	0.1870‡	0.7123	–	0.9610	0.9669	–
GER	0.7098	0.5377	0.7649	0.8849	0.4916	0.4013‡	0.9240	0.9235	0.9328
STX50	0.7263	0.5355	0.7228	0.7657	0.5394	0.4591†	0.9337	0.9186	0.9216
MSCI	0.7214	0.5992	0.7357	0.7551	0.5972	0.3022‡	0.9324	0.9592	0.9749

Notes: These long-run correlations were calculated using the quadratic spectral kernel weighting scheme. All correlations are significant at least at α = 0.05 except for the correlations that are marked with the symbol †, which are significant at α = 0.1, or the symbol ‡, which are not significant at all (using the test of Panopoulou et al., 2010).

The first period ends in December 2002 and might be considered as a period in which the CEE economies were implementing legislative and economic measures that led to their accession into the EU in May 2004. December 2002 was chosen because this was the month of the Copenhagen Summit, at which 15 member states decided to admit the CEE countries as new members of the EU. The second period covers the period around the formal accession and ends before the start of the worldwide financial crisis in June 2007. Dvořák and Podpiera (2006) argued that "*the rise in stock prices in the accession countries* (before accession in May 2004) *was a result of repricing of systematic risk due to the integration of local stock markets into the world market*." Furthermore, they stated that "*it is possible that a credible announcement of the EU enlargement led to an integration of the previously segmented Central and Eastern European stock markets with the rest of the world*." Whereas our results in the first period conform to the idea of Dvořák and Podpiera (2006), our results from the second period are somewhat surprising. The long-run correlations between the Polish and Czech stock markets and the developed markets remained unchanged, whereas the correlations of the Hungarian stock market with the developed markets decreased. In addition, the long-run correlations among the CEE markets were smaller than before acession into the EU. As we increased the bandwidth from six weeks to 18 weeks, the long-run correlations among the CEE markets decreased further, which might be interpreted as a sign that the short-run relationships among the CEE markets were stronger.

These findings are in line with the idea that whereas political stability and integration into the EU might be necessary conditions for an increase in the integration within CEE markets, these conditions are not sufficient for such an increase. In addition to market-specific characteristics (the depth of the market, the number of issuers, liquidity, etc.), there might be stronger forces that drive the integration of stock markets. This possibility is clearly visible from the results in the third period, as during the crisis, the long-run correlations were high (at roughly 0.9) even though the CEE-3 countries postponed the adoption of the Euro during this period; thus, these countries reduced the extent of the expected monetary integration.

An analysis using long-run correlations with different bandwidths and periods might be interesting, but comparing non-overlapping sub-periods might mask a high variability in the stock market integration. Therefore, we continued our analysis with a time-varying correlations analysis.

5.2 The estimated dynamic conditional correlations

In Table 2, we present the types and orders of the selected mean and variance equations employed in the first step of the DCC model estimation.[12] As expected, the BIC chose GARCH models that control for asymmetry in returns; the Czech PX was the only exception. However, note that DCC models are believed to be quite robust to the specification of the variance equation.

Table 2: Fitted univariate GARCH models

Index	Mean equation	Variance equation	Q-stat	Q^2-stat	BIC
WIG	ARMA(1,1)	GJR-GARCH(1,1)	0.1954(8)	0.1675(4)	-4.5334
PX	ARMA(1,1)	GARCH(1,1)	0.1008(7)	0.7404(1)	-4.6068
BUX	ARMA(1,1)	GJR-GARCH(1,1)	0.2329(3)	0.3431(1)	-4.3353
GER	ARMA(1,1)	TGARCH(1,1)	0.3329(3)	0.0550(1)	-4.5084
STX50	ARMA(1,1)	TGARCH(1,1)	0.3179(1)	0.2790(1)	-4.6761
MSCI	ARMA(1,1)	TGARCH(1,1)	0.4996(9)	0.5829(1)	-5.1541

Notes: "Q-stat" represents the minimal p-values recorded by the Ljung-Box test of standardized residuals (which tested for autocorrelation) and the corresponding lag (in parentheses) from the entire set of 12 lags. The column "Q^2-stat" is calculated in the same manner on the squares of the standardized residuals (and tests for any remaining ARCH effects). "BIC" is the Bayesian information criterion.

[12] For the sake of brevity, detailed results have been omitted but are available upon request.

Furthermore, Table 3 presents the estimates from the DCC and ADCC bivariate models. As expected in bivariate systems, the cDCC model provides very similar outputs, and hence, we do not report results from the cDCC models because they are qualitatively the same as the results from standard the DCC models.

Table 3: The estimation results of the DCC/ADCC models

	WIG			**PX**			**BUX**		
	GER	STX50	MSCI	GER	STX50	MSCI	GER	STX50	MSCI
φ	0.0272**	0.0263***	0.038	0.0371**	0.0331**	0.0303**	0.0033	0.0402	0.0146
	(0.0113)	(0.0102)	(0.026)	(0.017)	(0.015)	(0.0123)	(0.0174)	(0.0285)	(0.0169)
ψ	0.9694***	0.9702***	0.9526***	0.9431***	0.9560***	0.9586***	0.8148***	0.8318***	0.8572***
	(0.0156)	(0.0135)	(0.039)	(0.032)	(0.023)	(0.0188)	(0.0712)	(0.1504)	(0.0619)
ξ	–	–	–	–	–	–	0.0910**	–	0.0668**
	–	–	–	–	–	–	(0.0463)	–	(0.0335)

*Notes: ** and *** denote significance at the 5% and 1% levels, respectively.*

Significant asymmetries in correlations were only found in the two cases of Hungarian BUX, and thus, for all other relationships, we estimated the standard DCC model. The dynamics of the estimated correlations are captured in Figure 1.

To connect these results with our analysis of the long-run correlations, we first provide simple descriptive statistics of the dynamic correlations. As before, the sample is divided into three subsamples.

Table 4: Descriptive statistics of the estimated DCCs

	WIG			**PX**			**BUX**		
Subsamples	GER	STX50	MSCI	GER	STX50	MSCI	GER	STX50	MSCI
	Mean correlations								
01/1999 – 12/2002	0.48251	0.47002	0.48680	0.43166	0.39421	0.35859	0.57181	0.55224	0.56424
01/2003 – 06/2007	0.50312	0.51542	0.54584	0.48223	0.48022	0.46199	0.54711	0.54183	0.55397
07/2007 – 09/2012	0.69418	0.69950	0.70898	0.63018	0.67053	0.68350	0.59370	0.58970	0.60992
	Minimal correlations								
01/1999 – 12/2002	0.35396	0.32484	0.33476	0.11264	0.06208	0.18190	0.41760	0.30676	0.40023
01/2003 – 06/2007	0.21825	0.28400	0.33828	0.27089	0.25907	0.27042	0.41612	0.37911	0.34092
07/2007 – 09/2012	0.52882	0.46713	0.50863	0.44143	0.52745	0.53454	0.40292	0.27432	0.39373
	Maximal correlations								
01/1999 – 12/2002	0.61480	0.60428	0.63445	0.59900	0.59697	0.58468	0.74652	0.68510	0.71141
01/2003 – 06/2007	0.65073	0.65800	0.70013	0.64069	0.65075	0.65930	0.81771	0.70629	0.81305
07/2007 – 09/2012	0.79512	0.80763	0.80622	0.74594	0.75681	0.76689	0.78740	0.75145	0.76502

Supporting evidence of an increase in the modeled correlations is clearly visible from the basic descriptive statistics in the case of the Czech PX. Furthermore, the mean, minimal, and maximal correlations are higher in each subsequent period. Similar results are found in the case of the Polish WIG, although the minimal correlations are not increasing, as there were several drops in the DCCs around the year 2006. From Table 4 (and Figure 1), it is apparent that the estimated correlations of the Hungarian BUX do not exhibit any notable pattern. The correlations oscillate around the mean of 0.5 – 0.6 with a few sharp peaks and troughs.

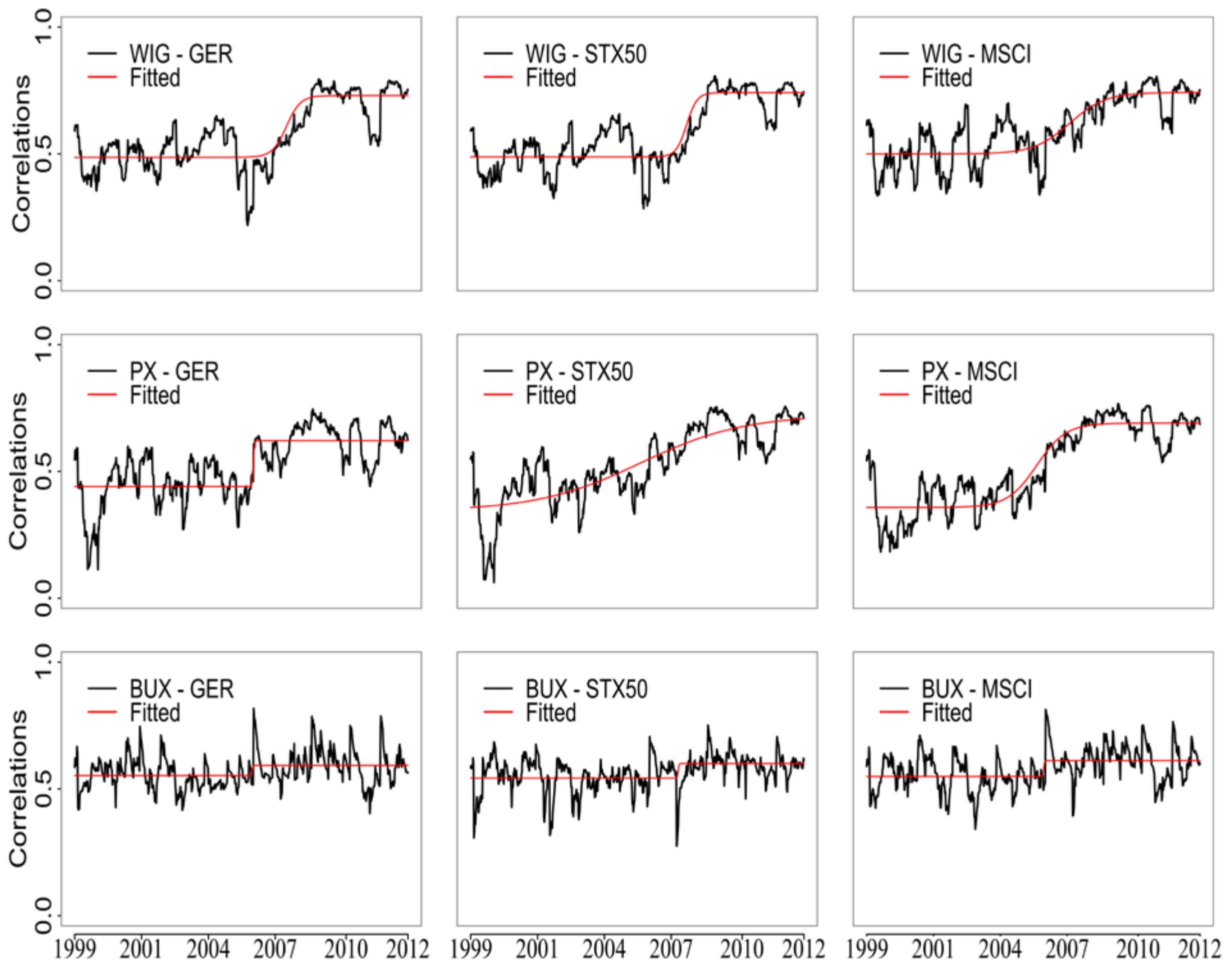

Figure 1: The estimated correlations and a fitted smooth transition model

5.3 The smooth transition model

Based on the underlying theory of stock market integration, we hypothesized that the transition process should occur gradually. Our proxy of the overall integration is the stock market co-movement, which is measured by the dynamic conditional correlations estimated in the previous subsection. The smooth transition logistic trend model was fitted on the bivariate correlations to verify whether the increase in correlations can be described as a gradual process.

The results from estimated smooth transition models are presented in Table 5, and the trend component along with the time-varying correlations are illustrated in Figure 1.

As correlations from DCC estimation are highly persistent, it is not surprising to see all $\hat{\delta}$ coefficients to be highly significant and above 0.9. The inclusion of autoregressive terms improved the autocorrelation structure of the errors and the inferences regarding the coefficients α and β. For all market co-movement models of the Polish WIG, both α and β are significant. The correlations in the first regime ($\hat{\alpha}/1-\hat{\delta}$) are approximately 0.5 in all cases. In addition, the increase in the second regime's correlations ($\hat{\beta}/1-\hat{\delta}$) is quite high at approximately 0.2 and, furthermore, the overall correlations at the end of our sample are above 0.7. Moreover, the general polynomial LM tests suggests the correlations were not stable, as at least one of the tests rejected the null of no change. Transition speed is determined by γ, which is the highest in the WIG-STX50 relation. The "smoothest" increase was reported with WIG-MSCI. Transition midpoints are dated around the recent financial crisis, i.e., between May and December 2007.

The results for the Czech PX are very similar. For the PX-GER correlations, the increase in the second regime is lower (0.206) and the speed parameter is quite large, suggesting a sudden increase in the correlations (see Figure 1). In contrast, the relation between the PX and the European STX50 can almost be described by a linear trend. While the speed adjustment parameter is very small (0.009), there is a steady and over time large increase in the correlations, which start at 0.344 in the first regime and end at 0.723 in the second. Such a strong increase in the second regime's correlations is also observable with the PX-MSCI. All changes in the correlations appear to be significant and, similar to before, at least one of the LM tests rejects the null hypothesis of no change in correlations. The transition midpoints are dated from January to March 2006, which appears to be earlier than in the WIG relationships.

Within our sample, the results in the case of the Hungarian BUX do not support the hypothesis of increasing market integration. At the same time, at 0.553, 0.543, and 0.549, the correlations between BUX and developed markets were already high in the first regime, while the second regime's increase in correlations is rather small, ranging from 0.040 to 0.063. The null of no changes in the mean of the correlations was not rejected by the LM test for two relationships: BUX-GER and BUX-MSCI. Figure 1 also makes clear that the smooth transition model may not be a suitable choice to characterize the estimated stock market co-movements between BUX and the developed markets.

Table 5: The fitted smooth transition model

	WIG			PX			BUX		
	GER	STX50	MSCI	GER	STX50	MSCI	GER	STX50	MSCI
α	0.021***	0.022***	0.030***	0.018***	0.015*	0.020***	0.072***	0.075***	0.058***
δ	0.956***	0.954***	0.939***	0.957***	0.957***	0.945***	0.870***	0.862***	0.895***
β	0.011**	0.012***	0.015***	0.009***	0.016*	0.018***	0.005**	0.008***	0.007**
γ	0.066	0.088	0.026	0.909	0.009	0.032	0.945	0.974	0.986
τ	0.636	0.648	0.610	0.521	0.511	0.511	0.533	0.623	0.532
Midpoint	7.10.2007	2.12.2007	27.5.2007	5.3.2006	15.1.2006	15.1.2006	7.5.2006	29.7.2007	30.4.2006
1st regime	0.487	0.488	0.500	0.428	0.344	0.358	0.553	0.543	0.549
2nd regime	0.730	0.742	0.742	0.634	0.723	0.690	0.593	0.601	0.612
LM-4	10.410**	8.004*	11.114**	8.062*	6.721	18.256***	5.054	8.904*	6.188
LM-3	8.084**	5.304	9.600**	7.679*	6.716*	17.858***	2.774	6.385*	5.377

*Notes: *, **, and *** denote significance at the 1%, 5%, and 10% levels, respectively. Regression coefficient significances were derived from the t-ratio, where the standard errors were calculated using a variance-covariance matrix with a quadratic spectral weighting scheme and an automatic bandwidth selection procedure, as in Newey and West (1994). LM-4 and LM-3 denote the χ^2 statistics of the 4th and 3rd order polynomial LM test of changes in the mean.*

As Baumöhl and Lyócsa (2012) have noted, weekly returns are commonly used to avoid non-synchronous trading effects. However, the method for constructing the weekly returns changes the statistical properties of the resulting series and, thus, has the potential to influence our results. We checked the robustness of our results by re-estimating both the DCCs and the smooth transition models using different forms of weekly returns. The general conclusions were not altered when Wednesday-to-Wednesday returns were used.

Larger differences in the estimated smooth transition models were seen when we analyzed Friday-to-Friday weekly returns. Although differences in the smooth transition model were observable for the PX and WIG correlations, our general conclusions remained the same as we observed a smooth transition from a period of lower correlations to a period of higher correlations. Qualitatively different conclusions were observed for the model involving BUX correlations. Although transition speed remained high, the first regimes were lower at 0.521, 0.468, and 0.470 for the GER, PX, and MSCI relationships respectively, while second regimes (except GER) were higher at 0.588, 0.662, and 0.655. This resulted in a larger sudden shift in market co-movements, all dated 23rd April 2006. Although there seems to be no general conclusion about the correct approach for constructing weekly returns, we adhere to the weekly averages because they are less volatile, less prone to the problems of non-synchronous trading effects, and not contaminated by day-of-the-week effects.

6. Concluding remarks

Since the early 1990s, the economies and financial systems of the CEE-3 countries have been transformed and liberalized. We hypothesized that as a consequence, there should have been a gradual increase in the integration between the CEE-3 countries and developed stock markets. To test our hypothesis, we used a sample of weekly returns constructed from daily data that spans a period from 1st January 1999 to 28th September 2012. We calculated the DCCs, which were modeled using a smooth transition logistic model.

In the case of the Czech PX, the results suggest that the transition began prior to the financial crisis of 2007. Whereas the transition midpoints for the Czech PX occurred during 2006, for the Polish WIG, these midpoints occurred at the end of 2007. For both markets, we found evidence that supports the hypothesized gradual increase in stock market integration. The "smoothest" relationship with the PX was found in the European aggregate index STX50. The Polish WIG seems to be the most integrated, as the sum of the correlations in the first and second regime is the highest. This result is not surprising, as the large number of actively traded stocks on the Polish stock market offers more diversification opportunities than the Hungarian and Czech stock markets combined.

Although the correlations were sustained at higher levels even four years after the culmination of the crisis, we are unable to rigorously rule out the possibility that the observed increases of the correlations are simply another manifestation of a contagion effect. For example, from Figure 1, it is apparent that the correlations decreased in 2010, but at the end of our sample, these correlations returned to their previous higher values. It is possible that this sudden increase was a consequence of another contagion stemming from the fiscal crisis in Europe.

The results for the Hungarian BUX are less clear. Although coefficients of the smooth transition model were significant, the increase in the correlations seems to be negligible (the maximal increase is 0.063, which corresponds to the case of BUX-MSCI). As shown in Figure 1, there seems to be no visible pattern between the two regimes. A smooth transition was only found when Friday-to-Friday weekly returns were used, but in this case, it is difficult to rule out the end-of-the-week effects.

Note that although the CEE-3 markets are close culturally, geographically, and economically, and although they share much common history, the transitions of their stock markets from segmented to integrated markets were rather diverse. There might be several reasons for these distinctions, which could be addressed in future studies. For example, investors who enter the Czech or Hungarian stock markets might have different motives than those who enter the Polish stock market. Because the markets of IPOs on the Czech and Hungarian stock markets are virtually non-existent, it might be that foreign investors are attracted to the Czech and Hungarian stock market only to diversify. In contrast, foreign investors who enter the Polish stock market might be attracted not only to diversification opportunities but also to new investments (IPOs). The differences between the integration processes of the Czech and

Hungarian stock markets might be attributed to the turbulent history of the Czech stock market (i.e., its mass privatization). However, the Hungarian stock market seems to have been known to foreign investors for some time, as in the early 1990s, SOEs were privatized by foreign investors. In fact, the average conditional correlations between BUX and the developed stock markets for the whole period are similar[13], and in some instances even higher, than the average correlations for the PX or WIG.[14]

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Appendix A Data definitions

Name	**Abbreviation**	**Source**	**Code**	**Short description**
Market capitalization	MC	World DataBank	CM.MKT.LCAP.GD.ZS	The share price times the number of outstanding shares of listed domestic companies at the end of the year. Expressed as a % of the nominal GDP.
Turnover ratio	TR	World DataBank	CM.MKT.TRNR	The total value of the shares that were traded during the year divided by the average MC of the given year, where the average MC is calculated from end of the year data. Expressed as a % of the market capitalization.
Listed domestic companies	#L	World DataBank	CM.MKT.LDOM.NO	The domestically incorporated companies listed on the country's stock exchanges at the end of the year; this does not include investment companies, mutual funds, or other collective investment vehicles.
Domestic credit to private sector	DCPS	World DataBank	FS.AST.PRVT.GD.ZS	The financial resources provided to the private sector, e.g., through loans, purchases of non-equity securities, and trade credits and other accounts receivable that establish a claim for repayment. Expressed as a % of the nominal GDP.
Domestic credit to private sector to Market Capitalization	DCPS/MC	Calculated		Measures the size of the stock market relative to the financial institutions that provide credit.
GDP deflator	DEF	World DataBank	NY.GDP.DEFL.KD.ZG	The annual growth rate of the implicit GDP deflator.
Private capital flows	PCF	World DataBank	BN.KLT.PRVT.GD.ZS	The private capital flows consist of the net foreign direct investment and the portfolio investment. Expressed as a % of the nominal GDP.
Gross domestic product per capita	GDP_PC	World DataBank	NY.GDP.PCAP.PP.KD	The GDP per capita. The data are expressed in constant 2005 international dollars.
Czech stock market index	PX	Thomson Reuters Datastream	CZPXIDX	Daily prices in the local currency: Czech Koruna.
Hungarian stock market index	BUX	Thomson Reuters Datastream	HNBUX	Daily prices in the local currency: Hungarian Forint.
Polish stock market index	WIG	Thomson Reuters Datastream	POLWIGI	Daily prices in the local currency: Polish Zloty.
German stock market index	GER	Thomson Reuters Datastream	MSGERML	Daily prices of the MSCI Germany index.
World stock market index	MSCI	Thomson Reuters Datastream	MSWRLD$	Daily prices of the MSCI World index.
European stock market index	STX50	Thomson Reuters Datastream	DJES50I	Daily prices of the EURO STOXX 50 index.

Appendix B Basic statistics of the CEE-3 and other selected economies

		CZE	***POL***	***HUN***	***AUT***	***DNK***	***PRT***	***IRL***	***DEU***	***GBR***	***USA***
MC/GDP	1994 - 1998	21.7%	6.2%	16.4%	15.3%	43.4%	27.5%	55.5%	32.2%	137.4%	113.3%
	1999 - 2002	18.1%	16.0%	24.8%	14.9%	56.8%	43.8%	68.9%	56.5%	158.4%	142.9%
	2003 - 2007	29.7%	33.7%	30.0%	42.4%	72.3%	43.5%	59.8%	50.4%	137.1%	138.2%
	2008 - 2011	21.9%	29.0%	17.3%	17.3%	56.6%	32.9%	16.2%	36.6%	113.6%	103.6%
TR	1994 - 1998	42.0%	93.7%	53.5%	57.0%	57.6%	50.7%	50.8%	96.5%	42.6%	91.6%
	1999 - 2002	43.4%	35.0%	68.6%	28.0%	70.4%	61.3%	42.8%	101.2%	73.0%	182.1%
	2003 - 2007	77.6%	37.1%	76.1%	42.1%	82.6%	73.3%	61.6%	150.6%	155.7%	155.5%
	2008 - 2011	44.6%	50.3%	95.5%	60.3%	84.9%	55.5%	45.8%	134.5%	153.4%	282.4%
# L	1994 - 1998	956.8	106.6	46.2	104.6	236.2	161.0	79.8	643.4	2112.6	8228.6
	1999 - 2002	116.8	223.0	57.8	99.8	214.8	98.5	72.5	854.8	2044.3	6803.8
	2003 - 2007	42.8	254.2	44.4	95.0	188.6	49.8	55.6	661.2	2611.4	5186.4
	2008 - 2011	16.3	507.3	46.0	89.8	206.3	47.5	52.8	620.0	2205.0	4613.5
DCPS/MC	1994 - 1998	3.3	3.9	2.9	5.1	0.8	3.0	1.5	3.6	0.8	1.3
	1999 - 2002	2.4	1.7	1.4	7.0	2.1	3.0	1.6	2.3	0.9	1.3
	2003 - 2007	1.3	1.0	1.8	3.0	2.4	3.4	2.6	2.3	1.2	1.4
	2008 - 2011	2.5	2.0	4.2	7.2	4.0	5.9	13.8	3.0	1.9	2.0
PCF	1994 - 1998	4.9%	3.3%	9.2%	2.0%	-0.7%	-0.2%	-1.2%	-1.0%	-1.9%	1.4%
	1999 - 2002	7.9%	4.6%	7.6%	0.4%	-3.1%	1.7%	-4.4%	0.6%	0.8%	3.4%
	2003 - 2007	3.5%	4.6%	6.0%	0.9%	-6.4%	1.0%	-4.1%	0.3%	2.0%	4.3%
	2008 - 2011	3.7%	4.8%	1.3%	-0.6%	-0.2%	3.9%	17.9%	-2.9%	4.6%	1.8%
DEF	1994 - 1998	9.8%	21.2%	20.4%	1.0%	1.6%	4.1%	3.4%	1.2%	2.7%	1.9%
	1999 - 2002	2.8%	4.7%	9.3%	1.1%	2.4%	3.5%	5.8%	0.5%	1.7%	1.9%
	2003 - 2007	1.7%	2.5%	4.4%	1.7%	2.2%	2.7%	2.7%	0.9%	2.5%	2.9%
	2008 - 2011	0.5%	2.8%	3.9%	1.7%	2.4%	1.1%	-2.3%	0.9%	2.4%	1.6%
GDP_PC	1994 - 1998	15946.5	9584.0	11937.6	28131.4	28718.9	18216.9	24510.6	28099.6	25728.9	34991.7
	1999 - 2002	17572.4	11765.6	13967.3	31683.5	31474.5	21122.8	33694.2	30274.3	29374.3	39362.3
	2003 - 2007	21362.0	14000.5	16831.7	33928.1	33265.9	21488.0	38909.0	31645.8	32987.0	42322.7
	2008 - 2011	23779.4	17150.4	17216.1	35587.2	32775.2	21595.5	36926.3	33543.3	33051.3	42205.5

Notes: Averaged values over the given period are presented. CZE – the Czech Republic, POL – Poland, HUN – Hungary, AUT – Austria, DNK – Denmark, PRT – Portugal, IRL – Ireland, DEU – Germany, GBR – the United Kingdom, USA – the United States of America.

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