What Drives the Stock Market Integration in the CEE-3?[1]

Tomáš VÝROST* – Eduard BAUMÖHL** – Štefan LYÓCSA***

Abstract

In this article, we study the possible explanatory power of macroeconomic factors that may drive the stock market integration between the Czech Republic, Poland and Hungary (CEE-3) and developed countries, using Germany as a benchmark. Our findings suggest that the recent global financial crisis has affected time-varying correlations between certain stock markets more substantially than the entry of the CEE-3 into the EU. The results of our analysis of the effects of these macroeconomic factors were inconclusive. Only our proxy of exchange rate risk was significant in all cases, with positive effects on integration, thus supporting the presence of contagion among different markets.

Keywords: *stock market integration, CEE-3, time-varying correlations, DCC MV-GARCH model, macroeconomic factors*

JEL Classification: G01, G15, C32

Introduction

Stock market integration has been studied for several decades with respect to international portfolio diversification (see, inter alia, Grubel, 1968; Ripley, 1973; Lessard, 1974). Formerly, because the correlation between equity returns and international markets was low and because equity returns were attributed to national factors, diversification among these markets was advisable. As Gjerde

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and Saettem (1995) explained, the segmentation of stock markets before 1980 can be attributed to barriers preventing capital mobility and foreign exchange transactions.

In the 1980s, various empirical studies reported a substantial increase in the interdependence of national stock markets (e.g., Jaffe and Westerfield, 1985; Schöllhammer and Sand, 1985; Asprem, 1989; Eun and Shim, 1989; Grinold, Rudd and Stefek, 1989; Meric and Meric, 1989).[2]

This progression naturally shed light on emerging markets in which investors could gain higher returns and benefit from international diversification. This premise may have been valid until the co-movements of emerging markets vis-à--vis developed markets were weak. Unfortunately for investors, after the effects of globalization, many emerging markets dropped all barriers to foreign participants in their local capital markets. However, as noted by Bekaert and Harvey (2002), regulatory liberalizations do not necessarily lead to market integration: "First, the market might have been integrated before the regulatory liberalization. That is, foreigners might have had the ability to access the market through other means, such as country funds and depository receipts. Second, the liberalization might have little or no effect because either foreign investors do not believe the regulatory reforms will be long lasting or other market imperfections exist that keep them out of the market." Because of the difficulty of evaluating the degree of integration, it is more convenient from an empirical point of view to focus on the co-movements of stock markets that may, in fact, be interpreted as the result of integration.

A convenient way of capturing the degree of stock market integration is to estimate dynamic conditional correlations (DCCs) with the DCC MV-GARCH model proposed by Engle and Sheppard (2001) and Engle (2002). Using DCCs allows the evolution of the relationships between stock markets to be assessed.

In this paper, we estimated DCCs between the CEE-3 (Central and Eastern European countries – Czech Republic, Poland, and Hungary) stock markets and the German stock market. We then study these time-varying correlations with regard to local and global events (such as entry to the EU and the global financial crisis), and explaining DCCs by applying various macroeconomic variables.

Our findings suggest that the global financial crisis increased stock market integration for the CEE-3 countries and, as shown in previous studies, we did not find strong evidence that monetary convergence or the real economic activity of the CEE-3 countries is associated with stock market integration. The contribution

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[2] Later, the methodologies applied in this field of study varied from basic correlation analyses to Granger causalities, co-integration techniques and then to more advanced techniques, such as regime-switching models and dynamic copulas.

of this paper is therefore twofold. First, we show by utilizing new data that the recent global crisis had a significant effect on stock market integration. Second, we further confirm that key macroeconomic variables do not explain the evolution of stock market integration.

The remainder of this paper is organized as follows. Section 1 provides a brief discussion of related empirical research. Section 2 describes the data utilized, and Section 3 specifies the methodology applied. Section 4 presents the results obtained, and the last Section concludes.

1. Related literature

After the US stock market crash in October 1987, co-movements between national stock markets were shown to have arisen (see, King and Wadhwani, 1990; Arshanapalli and Doukas, 1993), which is the so-called contagion effect, leading to a wide area of research within the framework of stock market integration.

The discussion about financial contagion notably expanded after the work of Forbes and Rigobon (2002), in which the contagion effect was defined as "a significant increase in cross-market linkages after a shock to one country (or group of countries)". Alternatively, continued market dependence at high levels is considered to be "no contagion, only interdependence".

Several studies have recently found evidence of the contagion effect. For example, using the sample of the CEE-3 indices, the German DAX and the U.S. S&P500, Baumöhl, Lyócsa and Výrost (2011) showed that endogenously detected volatility breaks in stock market returns are significantly associated with DCCs. When breaks are linked to a decrease in volatility, the correlations between the indices decrease as well. A sudden increase in volatility is similarly accompanied by an increase in DCCs and thus provides evidence for the presence of a shift contagion effect.[3] Kenourgios, Samitas and Paltalidis (2011) also provide evidence of contagion in a sample consisting of four emerging markets (BRIC countries) and two developed markets (UK and US) during the period 1995 – 2006, using a multivariate regime-switching Gaussian copula model and the asymmetric time-varying framework (AG-DCC).

Studies on stock market integration evolved further, from identification of and search for contagion among markets to exploratory studies that focused on finding the driving forces behind stock market integration, see Hardouvelis, Malliaropulos and Priestley (2006), Wang and Moore (2008) or Büttner and Hayo (2011).

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[3] Shift contagion can be defined as a significant change in the co-movement of stock markets between consecutive regimes of volatility (see a more complete definition in Rigobon, 2002).

With respect to analyses of the integration of CEE-3 stock markets with developed stock markets, several works have concluded that these markets present a substantial degree of co-movement (for a DCC-GARCH approach, see, e.g., Baumöhl, Lyócsa and Výrost, 2011, and Gjika and Horváth, 2012, for BEKK--GARCH, see, Horváth and Petrovski, 2012 and for applied sigma- and beta--convergence concepts, see Babecký, Komárek and Komárková, 2010).[4] Although the integration of the CEE-3 stock markets vis-à-vis developed markets may be interpreted as strong, Wang and Moore (2008) concluded that "… financial market integration seems to be a largely self-fuelling process, depending on existing levels of financial sector development…". The dominance of country- or market-specific news on correlations between the CEE-3 financial markets (stock, exchange, and money markets) was found by Büttner and Hayo (2011).

2. Data

To estimate the DCCs, we apply the daily closing prices of the stock market indices of the Czech Republic (PX), Hungary (BUX), Poland (WIG) and Germany (DAX)[5] from January 2000 to the end of August 2011. From these prices we compute weekly returns as log differences on a Wednesday-to-Wednesday basis to avoid the possibility of day-of-the-week effects. When Wednesdays were not active trading days, the closing prices were chosen from the next date with valid prices from the sequence of the nearest days (i.e., Tuesday, Thursday, Monday, and Friday).[6] Because the macroeconomic variables utilized are available with monthly frequency, estimated DCCs were averaged for one month. The DCCs were explained by a set of the following macroeconomic and market data.[7]

Oil prices (Brent) are long believed to be related to stock markets (e.g., Aslanidis, Osborn and Sensier, 2008). However, their relationship to stock market integration is not well understood. For example, Aslanidis, Osborn and Sensier (2008) were unable to explain the increase in DCCs utilizing changes in oil prices. However, because the relationship between oil prices and stock markets is

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[4] Surprisingly, Égert and Kočenda (2011) using intraday data on the sample period from June 2003 to January 2006 found low (around zero) conditional correlations among CEE-3 stock markets. A possible explanation is provided by Büttner and Hayo (2011): "One explanation of this noteworthy difference from our results could be that markets in the CEEC-3 are too slow in their reaction, possibly because of low liquidity and less advanced trading platforms." For an analysis of asset prices of CEE-3 stock markets reacting to macroeconomic news and for which spillover effects are present, see Hanousek, Kočenda and Kutan (2009) and Hanousek and Kočenda (2011).

[5] The German DAX is set as a proxy of stock market returns in the Eurozone, as in Hanousek and Kočenda (2011), because Germany is regionally and economically close to CEE countries.

[6] More days were unnecessary; this ensured that there were no missing observations.

[7] A detailed description of data sources are presented in the Appendix 1.

believed to be negative and because oil prices are a global economic factor, large swings in oil prices (especially increase in oil prices) may correspond to the increase of integration between markets.

Industrial production (IP) and its correlations have been applied by Wang and Moore (2008) to capture the convergence between economies, while Büttner and Hayo (2011) used IP to measure the extent of business cycle synchronization between two economies. In both cases, the variables were unable to satisfactorily explain the evolution of stock market integration. We use IP as a proxy for the development of real economy in the CEE-3 countries.

Inflation differential (HICPd) is calculated as the difference between the inflation rate of the CEE-3 countries and the average inflation rate of the three lowest inflation EU countries.[8] Inflation differential is used as a measure of monetary convergence between the CEE-3 economies and their more developed counterparts in EU.[9]

Short-term interest rate differential (STId) is also a measure of monetary convergence. The STId is calculated from 1-month interest rates of applicable money markets as a logarithmic difference between the interest rates of a CEE-3 country and Germany. As noted by Hardouvelis, Malliaropulos and Priestley (2006, p. 377): "… a convergence toward the low levels of Germany would show that the country's monetary authority was under no pressure to follow an unusually strict policy to satisfy the Maastricht criteria."

Exchange rate risk (GKA) is included as it is often used to explain stock market integration (Hardouvelis, Malliaropulos and Priestley, 2006; Wang and Moore, 2008; Büttner and Hayo, 2011). We apply the Garman and Klass (1980) range-based unconditional volatility ($\hat{\sigma}^2_{GK,t}$) with jump adjustment j_t (Molnár, 2012). If O, H, L, and C denote the opening, highest, lowest, and closing prices in a given week, respectively, then $h_t = ln(H_t) - ln(O_t)$, $l_t = ln(L_t) - ln(O_t)$, $c_t = ln(C_t) - ln(O_t)$, $j_t = ln(O_t) - ln(C_{t-1})$ and volatility is estimated as:

$$\hat{\sigma}^2_{GK,t} = 0.5(h_t - l_t)^2 - (2\ln 2 - 1)c_t^2 + j_t^2 \tag{1}$$

In the particular case of the CEE-3 countries, accession into the EU and the EMU (*Economic and Monetary Union*) are both motivating factors for the convergence of those economies (although, during the recent Eurozone crisis, early accession into the EMU does not seem to be a particularly important agenda item for the CEE-3 countries). Later, higher currency volatility increases the costs of diversification and therefore decreases stock market integration (Hardouvelis,

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[8] By adding inflation points, we would obtain one of the Maastricht criteria.

[9] As Hardouvelis, Malliaropulos and Priestley (2006) suggested, stock market integration may be explained by several economic channels that are related mostly to the monetary union.

Malliaropulos and Priestley, 2006). However, this effect may arguably change with respect to the degree of stock market integration. For example, if stock markets are integrated, higher volatility in currency markets may trigger changes in portfolios, thus leading to increasing integration. Conversely, if stock markets are less integrated, higher volatility could decrease the level of integration, as suggested by Hardouvelis, Malliaropulos and Priestley (2006).

Finally, we used *market capitalization as a percentage of GDP* (MC) (as in previous studies) because it captures the level of stock market development of a country. Countries with longer stock market histories tend to have a higher MC than countries with shorter histories, as is the case with the CEE-3 countries. Moreover, a higher MC suggests that the stock market more plausibly represents the development of the real economy.

3. Methodology

To estimate time-varying correlations, we used the DCC MV-GARCH model proposed by Engle and Sheppard (2001) and Engle (2002). It is a two-step procedure in which univariate GARCH models are fitted in the first step, followed by the DCCs themselves. As a mean equation we fitted the appropriate ARMA models, in which autocorrelation was checked by the Ljung-Box Q test up to int[$0.05T$] lags. Between the two models (or more) in which autocorrelation was not present, we chose according to the Akaike information criterion. As a variance equation, we fitted standard univariate GARCH models and an asymmetric EGARCH model proposed by Nelson (1991), taking into account the different effects of positive or negative shocks on the conditional variance. We tested again for remaining ARCH effects in the squares of residuals with the Ljung-Box Q^2 test (up to int[$0.05T$] lags). The DCC MV-GARCH model takes the following form:

$$\mathbf{r}_t \mid \Omega_{t-1} \sim N(\mathbf{0}, \mathbf{H}_t) \tag{2}$$

$$\mathbf{H}_t = \mathbf{D}_t \mathbf{C}_t \mathbf{D}_t \tag{3}$$

$$\mathbf{C}_t = diag\{\mathbf{Q}_t^*\}^{-1} \mathbf{Q}_t \, diag\{\mathbf{Q}_t^*\}^{-1} \tag{4}$$

$$\mathbf{Q}_t = \left(1 - \sum_{p=1}^{P} \alpha_p - \sum_{q=1}^{Q} \beta_q\right)\overline{\mathbf{Q}} + \sum_{p=1}^{P} \alpha_p \left(\mathbf{s}_{t-p}\mathbf{s}_{t-p}^T\right) + \sum_{q=1}^{Q} \beta_q \, \mathbf{Q}_{t-q} \tag{5}$$

$$\rho_{i,j,t} = \frac{q_{i,j,t}}{q_{i,i,t} \, q_{j,j,t}}, \; i, j = 1, \, 2 \ldots n; \; i \neq j \tag{6}$$

For each pair of residuals, $\mathbf{r}_t = (\varepsilon_{i,t}, \varepsilon_{j,t})^{\mathrm{T}}$, obtained from the ARMA models, it is assumed that, based on the information set Ω_{t-1}, they follow a multivariate normal distribution with variance-covariance matrix $\mathbf{H}_t$. This matrix can be decomposed as in equation (3), where $\mathbf{C}_t$ is the time-varying correlation matrix, $\mathbf{D}_t$ is the diagonal matrix of time-varying conditional standard deviations from univariate GARCH models, $\mathbf{s}_t$ are standardized residuals, $\overline{\mathbf{Q}}$ is the unconditional correlation matrix in dynamic correlation structure $\mathbf{Q}_t$, $\mathbf{Q}_t^*$ is a diagonal matrix with the square root of the i-th diagonal element of $\mathbf{Q}_t$ on its i-th diagonal position and a typical element of $\mathbf{C}_t$ takes the form of $\rho_{i,j,t}$, which are the DCCs.

After the bivariate dynamic correlations between the CEE-3 stock market indices and the German DAX are estimated, we run a regression of the following form on each pair (k) of DCCs:

$$DCC_{k,t} = \beta_{k,0} + \beta_{k,1}DUM_1 + \beta_{k,2}DUM_2 + \varepsilon_{k,t} \tag{7}$$

where $\beta_{k,m}$ for $m = \{0, 1, 2\}$ are the regression parameters and $\varepsilon_{k,t}$ is the error term. We have included two dummy variables (DUM_1 and DUM_2) into our model. First, we added a dummy variable for May 2004 that takes the value of 1 after and 0 before this date. This corresponds to the entry of all three CEE-3 countries into the EU. The second dummy variable seeks to account for the start of the global financial crisis. It is not straightforward to pinpoint the exact date when the global crisis started. According to Chor and Manova (2011), September 2008 was a month during which banking lending volume decreased sharply when the "collapse of Lehman Brothers and the government bailout of AIG – brought credit activity to a virtual standstill and raised the prospect of a financial sector meltdown in the US". The dummy variable is coded 0 for all months before September 2008 and 1 for all months after that date.[10]

Finally, we ran another simple regression to determine the effect of selected macroeconomic variables:

$$DCC_{k,t} = \beta_{k,0} + \sum_{s=1}^{6} \beta_{k,s}X_{s,t} + \sum_{l=1}^{L} \beta_{k,l}'DCC_{k,t-1} + \varepsilon_{k,t} \tag{8}$$

where regression parameters and error terms are denoted as in equation (7). Matrix $X_{s,t}$ consists of six macroeconomic variables described in Section 2, i.e., oil price (Brent), industrial production (IP), inflation differential (HICPd), short-term interest rate differential (STId), range-based volatility as an exchange rate risk proxy (GKA) and market capitalization as a percentage of GDP (MC). The second sum in equation (8) has been added to control for the autocorrelation

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[10] It is possible to add a third dummy variable to control for the recent EU debt crisis but, of course, more observations would be necessary.

of residuals, while lag 3 was the maximum lag order (L) necessary to remove autocorrelation in all models. Such representation prevents our results from being spurious. Although this model specification is rarely used, with the condition that residuals should be stationary, it was recommended by Hamilton (1994, pp. 561 – 562) and, recently, by McCallum (2010), who used simulation studies to show that removing autocorrelation from residuals prevents the regression results from being spurious.[11] Later, the interpretation is straightforward (because we use level data instead of differences). Thus, it is unnecessary to test for co-integrating relationships between variables and the order of integration of variables is also of lesser importance. Nevertheless, an F-test of the joint null hypothesis has a nonstandard limiting distribution and is therefore no longer valid (see, Hamilton, 1994, p. 562). All variables have been tested for the presence of a unit root by applying panel tests (Harris and Tzavalis, 1999; Hadri, 2000 and Breitung and Das, 2005) for their better power compared to univariate tests (the results may be found in Appendix 2; for further discussion see Lyócsa, Výrost and Baumöhl, 2011). Only Brent was tested by the DF-GLS test because it is not a country-specific variable; thus, no panel has been formed. Residuals from equation (8) were subjected to the DF-GLS test as well.

4. Results

In Figure 1, we show our estimated relationships between the CEE-3 stock markets and the German DAX. At first sight it is noteworthy that when in the first step of estimation the DCC model EGARCH models are applied, the DCCs are more volatile due to non-linear response to good and bad news. From visual inspection of the charts, we observe that correlations of the Czech PX and the Polish WIG with the DAX are high in recent years (approximately 0.70), while the Hungarian BUX tends to correlate with the DAX in a slightly weaker manner in our sample.

Table 1 presents the basic descriptive statistics of the estimated DCCs. The highest correlations of the Czech stock market are noteworthy because they were reported in October 2008. This date is interesting because, on the 15th of October 2008, the American stock market reported its largest decline since the stock market crash of 1987. This sharp decline of returns and rising correlations suggests close linkage with developed markets.[12] Such a degree of co-movement supports the presence of contagion and has, in fact, implications for the effectiveness of international diversification. The highest correlations of the Polish stock

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[11] This approach was also used by Wang and Moore (2008).

[12] This can also be observable in Baumöhl, Lyócsa and Výrost (2011).

market were in January 2009, which is also attributable to the recent financial crisis. Hungary is different, however; the DCCs contain two sharp peaks (these may be not obvious from Figure 1 because all the charts are scaled to an identical range), one peak in which correlations were maximum and a second peak corresponding to November 2008, in which DCCs were slightly lower, 0.5977 (DCC-GARCH) and 0.6249 (DCC-EGARCH).

F i g u r e 1
The Estimated DCCs Using Univariate GARCH and EGARCH Models
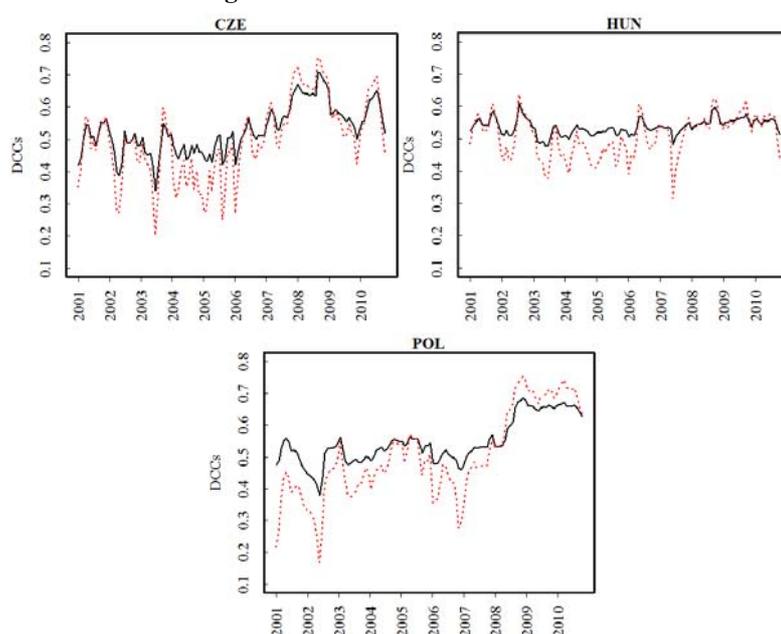
Notes: The dotted lines correspond to DCC-EGARCH models, solid lines to DCC-GARCH.
Source: Authors.

T a b l e 1
Descriptive Statistics of DCCs

	CZE		HUN		POL	
	DCC-GARCH	DCC-EGARCH	DCC-GARCH	DCC-EGARCH	DCC-GARCH	DCC-EGARCH
Average	0.5283	0.4991	0.5370	0.5088	0.5470	0.5034
Std. dev.	0.0744	0.1208	0.0243	0.0630	0.0706	0.1369
Min.	0.3410	0.2016	0.4776	0.3159	0.3790	0.1676
(date)	2003M07	2003M07	2003M06	2007M07	2002M06	2002M06
Max.	0.7097	0.7558	0.6092	0.6378	0.6861	0.7545
(date)	2008M10	2008M10	2002M08	2002M08	2009M01	2009M01
AR(1)	0.8858***	0.8397***	0.7951***	0.7876***	0.9691***	0.9440***
	(0.0440)	*(0.0471)*	*(0.0448)*	*(0.0443)*	*(0.0290)*	*(0.0322)*

Notes: The last row presents the estimate of the coefficient in the autoregressive model of order 1. Significance levels are denoted as *, **, and *** for 10%, 5%, and 1%, respectively. Standard errors are in parentheses.
Source: Authors.

To formalize the previous discussion about the contagion of the CEE-3 countries during the recent financial crisis, we ran a simple regression (see equation 7) in which two events are captured via dummy variables – entry to the EU and global crisis.

The results from Table 2 suggest that the recent financial crisis was a significant event that influenced the degree of stock market integration between the CEE-3 and Germany. Moreover, this effect caused increasing correlations because the coefficients have positive signs that speak in favor of the presence of contagion. As in the regression, only dummies are used as explanatory variables, a high coefficient of determination is surprising (particularly in the case of Poland). Another surprising result is that entry into the EU seems to be a significant factor of stock market integration only for the Czech Republic and Poland.

T a b l e 2

Effects of Main Events

	CZE		HUN		POL	
	GARCH	EGARCH	GARCH	EGARCH	GARCH	EGARCH
Constant	0.4834^{***}	0.4498^{***}	0.5323^{***}	0.5002^{***}	0.4939^{***}	0.3912^{***}
	(0.0118)	*(0.0213)*	*(0.0080)*	*(0.0173)*	*(0.0112)*	*(0.0200)*
Entry to EU	0.0428^{**}	0.0365	-0.0022	-0.0121	0.0343^{**}	0.0897^{***}
	(0.0215)	*(0.0375)*	*(0.0086)*	*(0.0210)*	*(0.0131)*	*(0.0269)*
Global financial crisis	0.0705^{***}	0.1070^{**}	0.0263^{***}	0.0713^{***}	0.1300^{***}	0.2245^{***}
	(0.0247)	*(0.0412)*	*(0.0050)*	*(0.0148)*	*(0.0081)*	*(0.0195)*
Adjusted R^2	0.3101	0.1901	0.1845	0.1914	0.8006	0.7460

Notes: Significance levels are denoted as *, **, and *** for 10%, 5%, and 1%, respectively. Standard errors are in parentheses.

Source: Authors.

Next, our analysis determines the potential explanatory power of the effect of various macroeconomic variables on the evolution of DCCs. For this purpose, we estimated a simple regression (equation 8), but without the dummy variables for entry to the EU and for the global financial crisis in this case. The first important point about this is that our emphasis now lies in estimating the effects of the macroeconomic variables themselves. The second and perhaps more important point is that macroeconomic variables should contain identical information as our dummies. The results from these estimated regressions are shown in Table 3.

Our results have relatively high levels of R^2 (0.95 in the case of Poland). We emphasize that this is not a result of spurious regression because the residuals are stationary and are not autocorrelated (see, McCallum, 2010). However, it is apparent that the high R^2 is clearly determined by the addition of the lagged DCCs (see Table 1, in which the highest autoregressive coefficient is for Poland) and not solely by the exogenous variables.

T a b l e 3

The Effect of Macroeconomic Variables on DCCs

Dependent Variable	CZE				HUN				POL			
	DCC-GARCH		DCC-EGARCH		DCC-GARCH		DCC-EGARCH		DCC-GARCH		DCC-EGARCH	
	Coeff.	t-stat	Coeff.	t-stat	Coeff.	t-stat	Coeff.	t-stat	Coeff.	t-stat	Coeff.	t-stat
Intercept	0.0489	1.0872	−0.0243	−0.3238	0.1995	4.4100***	0.2102	3.4061***	0.0416	1.9759*	−0.0339	−1.3877
Brent	0.0002	0.9651	0.0006	1.5715	0.0001	1.7630*	0.0004	1.7673*	0.0001	1.3085	0.0004	1.6781*
IP	0.0012	1.7019*	0.0029	2.1320**	−0.0001	−0.3564	0.0001	0.2870	0.0006	1.9675*	0.0019	3.2267***
HICPd	0.0045	1.8432*	0.0079	1.8408*	0.0005	0.9076	0.0021	1.1755	0.0012	1.2237	0.0025	1.2174
STId	0.0035	0.5748	0.0089	0.8662	−0.0002	−0.1189	0.0014	0.2955	−0.0045	−0.9496	−0.0107	−1.2042
GKA	15.8356	1.9905**	27.6188	2.5286**	3.2944	2.8769***	5.2457	2.3817**	2.9895	2.0137**	3.6251	1.3870
MC	−0.0015	−1.1686	−0.0051	−1.8480*	−0.0004	−1.6093	−0.0021	−2.5525**	−0.0006	−1.8503*	−0.0024	−3.1638***
DCC$_{t-1}$	0.6999	13.4588***	0.8565	8.1259***	1.0436	10.6285***	0.9612	9.0707***	1.2632	12.8357***	1.0993	11.3582***
DCC$_{t-2}$	–	–	−0.2671	−2.8588***	−0.4080	−3.6950***	−0.3741	−3.7535***	−0.5747	−3.3881***	−0.3769	−4.9147***
DCC$_{t-3}$	–	–	–	–	–	–	–	–	0.1214	1.3064	–	–
Adj. R^2	0.8139		0.7478		0.6947		0.6697		0.9537		0.9360	
LB (Q12)	12.8810		10.2335		5.8266		7.2896		17.8750		17.2551	
$\tau^{DF\text{-}GLS}$	−2.4711**		−2.0589**		−3.4066***		−2.7336***		−5.7479***		−4.3835***	

Notes: Significance levels are denoted as *, **, and *** for 10%, 5%, and 1%, respectively. LB(Q12) stands for Ljung-Box Q statistic at lag 12, where the number of lags was selected according to Schwert's (1989) "rule of thumb" $k_{max} = \text{int}[12(T/100)^{1/4}]$. $\tau^{DF\text{-}GLS}$ denotes the test statistic for DF-GLS test of regression's residuals with critical values taken from Cook and Manning (2004) for the sample size of T = 100, i.e., −2.62 and −2.05 for 1% and 5% significance levels. "Brent" stands for oil prices, "IP" for industrial production," HICPd" for inflation differential, "STId" for short-term interest rate differential, "GKA" for exchange rate risk, "MC" for market capitalization as a percentage of GDP (see Section 2 for details).

Source: Authors.

The only factor significant for all the CEE-3 countries' integration with Germany is our proxy for exchange rate risk (GKA). This result is in sharp contrast with Wang and Moore (2008), in which exchange risk was significant at the 10% level only for Poland. However, they argue that "the volatility of exchange rates in the emerging markets tends to increase with the financial crises…, (so) the positive relationship between DCC and exchange rate risk may be plausible". Conversely, when a country exhibits stable exchange rates, it should reduce cross-currency risk premiums and, thus, supports investors' concerns about the given markets (see, Asgharian and Nossman, 2011). In our case, the positive relationship between DCCs and exchange rate risk suggests the presence of a contagion effect.

Another result that contrasts with the work of Wang and Moore (2008) is the negative sign of market capitalization as a percentage of GDP (MC), where the coefficient is even insignificant in a few cases. It seems that in recent years, an existing degree of financial market development of the CEE-3 has been sufficient to no longer contribute to strengthening the stock market integration with developed markets. One of the explanations might be, that the increase of the market capitalization is driven by rising prices and by new issuers, including those who are (i) more unique and less influenced by global economic factors and (ii) more focused on local markets.

Other variables are significant only in a few cases, making general conclusions harder to infer. The higher level of underlying economic output (IP) seems to have a positive effect on the integration of stock markets in the Czech Republic and Poland but not in Hungary. Our conjecture regarding the effect of oil prices seems to be supported by our data and analysis. The coefficients were positive and significant for the stock markets in Hungary and Poland. Oil prices are a common global factor, and their increase should therefore lead to lower stock prices and higher levels of stock market co-movement. In general, monetary variables were not significant. This might not be surprising because the three countries are not aiming for the EMU in the near future. Finally, two lags of the DCCs are significant in all models, which provide evidence of substantial persistency.

Conclusion

The analysis conducted in this paper is, in many aspects, a follow-up of the research of Wang and Moore (2008). We attempted to shed more light onto the integration of the stock markets of the CEE-3 countries with those of developed countries, for which we chose Germany as a benchmark because of its regional and economical closeness.

This paper contributes to the existing literature by estimating the time-varying correlations in the recent data and by examining possible macroeconomic integration-driving factors. Except for exchange rate risk, however, our results were mixed, and no general conclusion or inference could be made about the key macroeconomic indicators of stock market integration in the CEE-3. It is also possible that integration of the CEE-3 markets is already substantial and may thus be considered a self-driven process.

As correlations play a crucial role in portfolio diversification, it is beneficial to study them more closely. We have found evidence that the recent global financial crisis can be considered a major event that has influenced the evolution of mutual relationships between stock markets under study. Because these correlations have been quite high in the last few years, the benefits of international diversification into the CEE-3 markets may be questionable.

A p p e n d i x 1 – Dataset Description

Variable	Source and Description
Oil prices	Weekly Europe (UK) Brent Blend Spot Price FOB (WEPCBRENT), <http://www.eia.gov/dnav/pet/hist/LeafHandler.ashx?n=PET&s=WEPCBRENT&f=W>. For each month, the weekly prices have been averaged over the corresponding month.
Industrial production index	OECD MEI database, production of total industry, seasonally adjusted index with 2005 as a base year.
HICP	Eurostat database, monthly data with annual rate of change, [prc_hicp_manr].
Short term interest rates	Eurostat database, money market 1-month interest rate, monthly data, [irt_st_m]. For Hungary, interest rates for the following data were not available: may 2004, June 2004, July 2007, August 2007, October 2007, and April 2006. For these months, the corresponding data were interpolated using simple linear interpolation.
Exchange rates to EUR	Daily data from www.oanda.com.
Market capitalization as a percentage of GDP	Eurostat database, [mny_stk_mcp_m].

Source: Authors.

A p p e n d i x 2 – Panel Unit Root Tests

	Breitung and Das (2005)		Harris and Tzavalis (1999)		Hadri (2000)	
	level	differences	level	differences	level	differences
DCC-GARCH	-2.2457^{**}	-3.4111^{***}	-4.2248^{***}	-58.0217^{***}	50.7632^{***}	-1.3225
DCC-EGARCH	-1.6894^{**}	-1.6894^{**}	-5.7067^{***}	-5.7067^{***}	43.9892^{***}	43.9892^{***}
IP	0.4925	-3.7089^{***}	-2.6668^{***}	-95.1763^{***}	90.7538^{***}	-1.1756
HICPd	-1.5178	-6.3774^{***}	-2.0208^{**}	-50.3987^{***}	16.4204^{***}	-0.648
STId	-0.6278	-6.0776^{***}	-0.993	-51.6022^{***}	36.9898^{***}	2.8126^{***}
GKA	-7.8824^{***}	-17.0453^{***}	-53.6185^{***}	-92.2548^{***}	3.0640^{***}	-1.6241
MC	-0.1994	-10.7071^{***}	-1.196	-75.0067^{***}	64.8717^{***}	-0.9944

Notes: Significance levels are denoted as *, **, and *** for 10%, 5%, and 1%, respectively. The Breitung – Das (2005) test takes into account cross-sectional dependence in the panel as the so called 2nd generation test. Together with Harris and Tzavalis (1999) test, the null hypothesis is H_0: All time series have a unit root, against an alternative H_1: All time series are stationary. Hadri (2000) is a stationarity test because of the reverse null H_0: All time series are stationary, H_1: Some series have a unit root.
Variable Brent was tested by DF-GLS test with critical values from Cook and Manning (2004) for sample size 100, which are -2.62, -2.05, -1.77 for 1%, 5%, and 10% significance level. The test statistic was -0.160 for levels and -4.680 for differences, suggesting that the variable Brent is stationary only in differences.

Source: Authors.

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