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ORIGINAL PAPER

Stock market integration between new EU member states and the Euro-zone

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Abstract This paper measures the degree in stock market integration between five Eastern European countries and the Euro-zone. A potentially gradual transition in correlations is accommodated by smooth transition conditional correlation models. We find that the Czech, Slovenian and Polish markets have increased their correlation to the Euro-zone from 1997 to 2008. However, this is not a broad-based phenomenon across Eastern Europe. The results also show that the increase in correlations is not a reflection of a world-wide phenomenon of financial integration but is mainly driven by EU-related developments.

Keywords Multivariate GARCH · Smooth transition conditional correlation · Stock return comovement · New EU members

JEL Classification C32 · C51 · F36 · G15

1 Introduction

It is a well established fact in the literature on international financial markets that the links between international stock markets tend to be strong although they change

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over time.[1] An accurate assessment of the degree of comovements among international stock markets is of interest for a number of reasons. From the point of view of investors, the optimal design of a well-diversified portfolio depends on a proper understanding of stock market correlations. Changes in comovement patterns call for an adjustment of portfolios. Policy makers are also interested in the links between stock markets because of their implications for the stability of the financial system. Particularly in the European Economic and Monetary Union (EMU), the process of financial market integration is central to monetary policy making.

Most empirical studies on comovements have focused on developed markets with only fewer papers analysing emerging markets.[2] In the case of the Central and Eastern European countries (CEECs) there has been a burgeoning literature on business cycle synchronisation establishing that many of these new EU members have achieved a high degree of cycle correlation.[3] The progress of these markets towards financial integration is however subject to some debate, with Égert and Kočenda (2007) arguing for relatively low integration in equity markets, and Cappiello et al. (2006b) and Chelley-Steeley (2005) documenting increasingly strong comovements.

This paper computes measures of the extent of stock market integration between five CEECs (Hungary, Czech Republic, Slovakia, Slovenia and Poland) and the Euro-zone. Two important questions are addressed in this analysis. First, has stock market integration increased in the run-up period to the accession to the EU in 2004 and the period afterwards? And if so, is this increase part of a world-wide phenomenon of financial integration or is it mainly driven by EU-related developments?

To answer these questions, we focus on the dominant trends in the evolution of stock market integration, which we measure by the conditional correlation between weekly returns. More specifically, we consider time-varying correlations in the stock markets using the recently developed smooth transition conditional correlation GARCH (STCC-GARCH) models: see Berben and Jansen (2005), Silvennoinen and Teräsvirta (2005) and recently Silvennoinen and Teräsvirta (2007) who extend this to a double STCC-GARCH with two transition variables. These models allow for the correlation of a constant conditional correlation GARCH (CCC-GARCH) to change smoothly over time, which seems particularly appropriate to analyse the increasing integration between the CEECs and the Euro-zone stock markets over the recent years.

Our work is close to Égert and Kočenda (2007), who apply cointegration, VAR and Granger causality techniques to stock returns. Instead, the long-run dynamics of financial integration may be better captured by our model where stock market correlations are allowed to change smoothly over time. Moreover, Égert and Kočenda find very little evidence of long-run comovements for the Hungarian, the Czech and Polish markets. They only find signs of short-run relationships. This result may be due to the

[1] Important contributions to understanding the nature of this phenomenon include Ang and Bekaert (2002), Baele (2005), Cappiello et al. (2006a), Goetzmann et al. (2005), King et al. (1994), Longin and Solnik (1995), Longin and Solnik (2001), Ramchand and Susmel (1998) among others.

[2] For papers focusing on emerging markets see for example Bekaert and Harvey (1995) and Goetzmann and Jorion (1999).

[3] For a comprehensive survey see Fidrmuc and Korhonen (2006). Also, for economic convergence among the CEECs see Kočenda (2001) and Kutan and Yigit (2004).

use of very high frequency data (5-min tick intraday data) as well as the short period under consideration (June 2003–February 2005).[4] Instead, we use weekly data, which are less noisy and, therefore, may more reliably capture the progress toward financial integration experienced by Eastern Europe.

The idea of a smooth transition in correlations is also explored by Chelley-Steeley (2005) for Hungary, the Czech Republic, Poland and Russia, although with a different procedure: she first estimates monthly correlations from daily data, and then fits a smooth transition regression (STR) model to the previously estimated correlations. Instead, we model the conditional correlations directly. Another important difference with Chelley-Steeley (2005) is that her analysis is confined to the period from 1994 to 1999, while our study looks at the more recent and relevant period of the last twelve years that includes the accession of the CEECs to the EU. Regarding the results, Chelley-Steeley finds that the Hungarian market is the most integrated among the Eastern European markets.

Moreover, unlike the previous papers we follow a systematic testing procedure to determine the number of changes in correlations. Cappiello et al. (2006b) also address this issue although they simply test for the constancy in correlations by using a regression quantile approach. In general, they find evidence of stock market integration for Hungary, the Czech Republic, Poland, Slovenia, Estonia, Latvia and Cyprus using daily data over January 1994–November 2005.

Our results are summarised as follows. We find that the correlation of three CEECs (the Czech Republic, Slovenia and Poland) and the Euro-zone stock markets increased between 1997 and 2008. However, this is not a broad-based phenomenon across Eastern Europe. Also, the increase in correlations is not a reflection of a world-wide phenomenon of financial integration but is mainly driven by EU-related developments.

The paper is organised as follows. Section 2 presents the smooth transition conditional correlation model as well as the discussion of the tests to determine the number of changes in correlations. Section 3 discusses the data and presents the results. Finally, Sect. 4 concludes and discusses further extensions.

2 Methodology

2.1 The model

Consider the following 2-dimensional vector process of stock returns (y_t)

$$y_t = \mu + r_t \tag{1}$$

where $t = 1, \ldots, T$ and $\mu \neq 0$ denotes the vector of mean returns. The conditional covariances of the shocks in (1) are time-varying, such that

$$r_t | \Im_{t-1} \sim N(0, H_t) \tag{2}$$

[4] We thank an anonymous referee for pointing this out.

where $\Im_{t-1}$ is the information set at time t and N denotes the bivariate normal distribution. Each of the univariate error processes has the specification

$$r_{i,t} = h_{i,t}^{1/2} \varepsilon_{i,t}$$

where the errors $\varepsilon_{i,t}$ form a sequence of independent random variables with mean zero and variance one, for each of the stock returns $i = 1, 2$. The conditional covariance matrix of r_t, H_t, is time-varying as follows: each conditional variance $E(r_{i,t}^2/\Im_{t-1}) = h_{i,t}$ follows a univariate GARCH(1,1) process

$$h_{i,t} = \alpha_{i0} + \alpha_{i1} r_{i,t-1}^2 + \beta_{i1} h_{i,t-1} \tag{3}$$

with the non-negativity and stationarity restrictions imposed.

Rather than modelling the off-diagonal elements of H_t directly, the definition

$$\rho_t = h_{12,t}(h_{11,t} h_{22,t})^{-1/2} \tag{4}$$

allows the focus to be placed on the conditional correlations ρ_t. $h_{12,t}$ is the conditional covariance between stock returns. We allow the conditional correlations to be time-varying by considering the smooth transition conditional correlation GARCH (STCC-GARCH) specification proposed in Silvennoinen and Teräsvirta (2005) and Berben and Jansen (2005).[5] This model assumes two regimes with regime-specific constant correlations, and allows for a smooth change over time between correlation regimes. More specifically, the correlation ρ_t follows

$$\rho_t = \rho_1 \left(1 - G_t \left(s_t; \gamma, c\right)\right) + \rho_2 G_t \left(s_t; \gamma, c\right) \tag{5}$$

where ρ_1 and ρ_2 are the regime-specific correlations. The function $G_t\left(s_t; \gamma, c\right)$ is the transition function, assumed to be continuous and bounded between zero and unity, γ and c are its parameters, whereas s_t is the transition variable. As our focus is on dominant, long-run trends in correlations, there is one change in correlation regime and the transition variable is specified as a linear function of time, $s_t = t/T$.

To capture integration we need the transition function to be monotonically increasing in t, which is achieved by using the logistic function

$$G_t\left(s_t; \gamma, c\right) = \frac{1}{1 + \exp\left(-\gamma\left(s_t - c\right)\right)}, \quad \gamma > 0 \tag{6}$$

where c is the threshold parameter and locates the midpoint between the two regimes. The slope parameter γ determines the smoothness of the change in the transition function and shows the versatility of the model. In particular, when $\gamma \to \infty$, $G_t\left(s_t; \gamma, c\right)$

[5] The model of Berben and Jansen (2005) is bivariate with a time trend as the transition variable, while the framework in Silvennoinen and Teräsvirta (2005) is multivariate and their transition variable can be deterministic or stochastic.

becomes a step function ($G_t(s_t; \gamma, c) = 0$ if $s_t < c$ and $G_t(s_t; \gamma, c) = 1$ if $s_t > c$), and the transition between the two extreme correlation states becomes abrupt. In that case, the model with time transition approaches a structural break model in conditional correlations.

Before considering the STCC model it is important to determine whether the change in correlation is statistically significant. To that purpose, we perform the Lagrange Multiplier test of Silvennoinen and Teräsvirta (2005). Under the null hypothesis the model is a constant conditional correlation (CCC, Bollerslev 1990) model, whereas the alternative model is a STCC with $s_t = t/T$. Only in the case we reject the hypothesis of constant correlation, we proceed with the estimation of the STCC model.

The STCC model allows for a monotonic change in correlations. In practice, this might be restrictive and, therefore, it would be of interest to extend the model to allow for non-monotonic correlation patterns. This possibility is investigated by using the Lagrange Multiplier test of Silvennoinen and Teräsvirta (2007). In this test, under the null hypothesis a single STCC (one change in correlations) is adequate whereas the alternative supports a double STCC (two changes in correlations).[6] If evidence of a second change in correlations is found, then we estimate the double smooth transition conditional correlation (DSTCC) given by the following equation

$$\begin{aligned} \rho_t = {} & \rho_1(1 - G_{1t}(s_t; \gamma_1, c_1)) + \rho_2 G_{1t}(s_t; \gamma_1, c_1)(1 - G_{2t}(s_t; \gamma_2, c_2)) \\ & + \rho_3 G_{1t}(s_t; \gamma_1, c_1) G_{2t}(s_t; \gamma_2, c_2) \end{aligned} \tag{7}$$

Notice that the second transition variable is also a function of time ($s_t = t/T$), and hence (7) allows for the possibility of a non-monotonic change in correlation over the sample. This is a special case of Silvennoinen and Teräsvirta (2007) as the transition variables are the same. The transition functions $G_{1t}(s_t; \gamma_1, c_1)$ and $G_{2t}(s_t; \gamma_2, c_2)$ are logistic functions as defined in (6). The parameters γ_i and $c_i (i = 1, 2)$ are interpreted in the same manner as in the STCC model, but in order to ensure identification we require $c_1 < c_2$ and hence that the two correlation transitions occur at different points of time.

We estimate the (D)STCC-GARCH models by quasi-maximum likelihood (QML), where standard errors that are robust to the violation of the assumption of normality in (2) are used for the parameter estimates (Bollerslev and Wooldridge 1992). Furthermore, the log-likelihood is maximised with respect to all parameters simultaneously.[7] All computations are carried out using GAUSS.

[6] For analytical expressions of the test statistics and the required derivatives, the reader is referred to Silvennoinen and Teräsvirta (2005, 2007).

[7] In addition to the one-step estimation, the models were also estimated in two steps, namely the mean and volatility equations in a first step, followed by the (D)STCC parameters in a second step. Qualitatively, both approaches led to the same results in terms of coefficient estimates. However, the standard errors of the (D)STCC coefficients were smaller when the one-step estimation was employed. So, in line with increased efficiency we reported the single-step estimation results.

Table 1 Summary statistics of the stock returns

	Abbr.	Min	Max	Mean	SD	Skewness	Kurtosis
Budapest (BUX), Hungary	HU	−22.81	19.86	0.138	4.468	−0.848	7.036
PX Global Index, Czech Rep.	CZ	−19.21	11.27	0.117	3.542	−0.895	6.099
Slovakia SAX 16, Slovakia	SK	−14.27	13.47	0.161	3.002	0.054	5.759
Slovenian (PIX), Slovenia	SL	−10.82	11.37	0.219	2.495	−0.075	6.775
Warsaw Gen. Index 20, Poland	PL	−22.74	15.46	0.025	4.668	−0.561	5.065
DJ Euro Stoxx 50, Euro-zone	EURO	−14.53	16.76	0.042	3.256	−0.477	6.462
S&P 500 Composite, USA	US	−13.87	12.32	0.025	2.897	−0.391	5.207

Source is DataStream

3 Empirical results

The data we use consists of weekly data, denominated in Euro,[8] of the following stock market indices: Hungarian BUX (HU), PX Global Index in the Czech Republic (CZ), Slovakia SAX 16 (SK), Slovenian PIX (SL), Warsaw General Index 20 (PL), Dow Jones STOXX 50 (EURO) and S&P 500 Composite (US). The Dow Jones STOXX 50 provides a blue-chip representation of 50 super-sector leaders in the Euro-zone.[9] On the other hand, given its size the US market is a natural proxy for the world stock market. The inclusion of the S&P 500 index also allows capturing international influence since the US is home to many of the world's largest companies. All data is obtained from DataStream, and we refer to the manual *DataStream Global Equity Indices* for further details. The sample period starts on January 6, 1997 and ends on November 24, 2008, which yields 621 observations. This time span includes the run-up period to the accession to the EU in 2004 and the period afterwards and thus allows us to assess whether the stock markets of Eastern Europe and the Euro-zone have become more integrated since joining the EU.[10] Descriptive statistics of the data are presented in Table 1.

In most cases, the results for the volatility models are very close to those found elsewhere in the literature. For example, in the GARCH equations the betas are normally between 0.80 and 0.95, except for a few cases where they are estimated in the range 0.50–0.60. Similarly, the alphas are estimated to be between 0.07 and 0.15. We also refined our basic GARCH(1,1) and tested for asymmetry in volatility by considering the GJR-GARCH(1,1) model. The results showed that for this dataset the asymmetry effect was not statistically significant. Furthermore, we examined volatility linkages in the above specifications by calculating the correlation between the estimated variances of the two assets and found that the conditional variances are weekly correlated.

8 Results remain qualitatively the same for data denominated in local currency.

9 We have also used the DAX index as a representative of the EU market with the results remaining qualitatively similar to those presented here.

10 Note that the announcement of EU enlargement was on December 12–13, 2002 with the actual membership happening on May 1, 2004.

Table 2 Tests of CCC- against STCC-GARCH

Panel a	
HU–EURO	0.440
CZ–EURO	0.000**
SK–EURO	0.836
SL–EURO	0.004**
PL–EURO	0.044*
Panel b	
HU–CZ	0.000**
HU–SK	0.541
HU–SL	0.275
HU–PL	0.000**
CZ–SK	0.955
CZ–SL	0.016*
CZ–PL	0.000**
SK–SL	0.317
SK–PL	0.731
SL–PL	0.167
Panel c	
HU–US	0.635
CZ–US	0.014*
SK–US	0.126
SL–US	0.247
PL–US	0.755
EURO–US	0.375

Results (p values) from tests of constant correlation against STCC-GARCH models
*, ** Significance at the 5 and 1% level, respectively

Finally, we tried models using the bivariate stud-t distribution for the errors with the results remaining qualitatively the same.[11]

Our next step is to test whether there is a statistically significant change in the stock market correlations between Eastern Europe and the Euro-zone by performing the constancy test of Silvennoinen and Teräsvirta (2005). Table 2 shows that the null hypothesis of constant correlation is rejected for the Czech, Slovenian and Polish markets with the first two implying strong rejections. For the other two markets, the test shows no evidence of changing correlation.

As to the magnitude of the correlations, the constant conditional correlation (CCC) estimates in Table 3 (panel a) show that for the Hungarian, Czech and Polish markets the correlation is considerably higher than for the smaller markets. For instance, for the first three markets the correlation is around 0.5, for the Slovenian market close to 0.2, while for the Slovakian one about 0.06 (statistically insignificant).

[11] All these results are available from the authors upon request.

Table 3 CCC-GARCH models

	ρ
Panel a	
HU–EURO	0.534 (0.036)
CZ–EURO	0.452 (0.038)
SK–EURO	0.064 (0.044)
SL–EURO	0.197 (0.047)
PL–EURO	0.561 (0.031)
Panel b	
HU–CZ	0.564 (0.032)
HU–SK	0.111 (0.042)
HU–SL	0.274 (0.041)
HU–PL	0.628 (0.030)
CZ–SK	0.147 (0.039)
CZ–SL	0.241 (0.040)
CZ–PL	0.562 (0.034)
SK–SL	0.035 (0.041)
SK–PL	0.132 (0.040)
SL–PL	0.212 (0.038)
Panel c	
HU–US	0.430 (0.039)
CZ–US	0.318 (0.041)
SK–US	0.087 (0.046)
SL–US	0.124 (0.048)
PL–US	0.449 (0.038)
EURO–US	0.689 (0.024)

The table presents quasi maximum likelihood estimates of the correlation parameters of CCC-GARCH models; remaining parameter estimates are available upon request; values in parentheses are robust standard errors

Table 4 (panel a) reports the estimated STCC-GARCH for those models where constancy is rejected. The parameter c defines the middle of the transition period and is expressed as a fraction of the sample size. The heading ‘Date’ reports the week corresponding to c. As observed, for the Czech Republic and Slovenia the estimates imply a considerable increase in correlation. This can also be seen clearly in Fig. 1, which plots the correlations implied by all three models. For Slovenia, the level of correlation is lower compared to the other two countries, even after the shift in late 2002. As for Poland, the increase in correlation is relatively small, possibly reflecting the mild significance at 4.4% of the constancy results in Table 2.

As stressed before, when using a STCC model it is of interest to test whether there is a second change in correlations. We tested this by performing the Silvennoinen and Teräsvirta (2007) LM test for an additional transition in STCC. The results (reported in Table 5, panel a) support a second change in correlation only for the Czech market

Table 4 STCC-GARCH models

	ρ_1	ρ_2	γ	c	Date
Panel a					
CZ–EURO	0.357 (0.049)	0.714 (0.039)	8.864 (6.672)	0.738 (0.059)	17 Oct 05
SL–EURO	0.077 (0.054)	0.341 (0.051)	500 (.)	0.485 (0.010)	21 Oct 02
PL–EURO	0.518 (0.049)	0.628 (0.058)	15.00 (43.26)	0.558 (0.163)	1 Sep 03
Panel b					
HU–CZ	0.534 (0.032)	0.666 (0.042)	500 (.)	0.791 (0.001)	05 Jun 06
HU–PL	0.546 (0.043)	0.782 (0.026)	14.70 (13.98)	0.640 (0.105)	23 Aug 04
CZ–SL	0.173 (0.045)	0.422 (0.061)	500 (.)	0.787 (0.033)	22 May 06
CZ–PL	0.476 (0.036)	0.757 (0.027)	500 (.)	0.688 (0.047)	14 Mar 05
Panel c					
CZ–US	0.283 (0.041)	0.432 (0.067)	17.39 (80.22)	0.801 (0.114)	17 Jul 06

The table presents quasi maximum likelihood estimates of part of the parameters of STCC-GARCH models; remaining parameter estimates are available upon request; 'Date' is the week that corresponds to c (threshold); (.) γ is estimated large with large standard error (for more details, see Teräsvirta 1994, p. 213); values in parentheses are robust standard errors

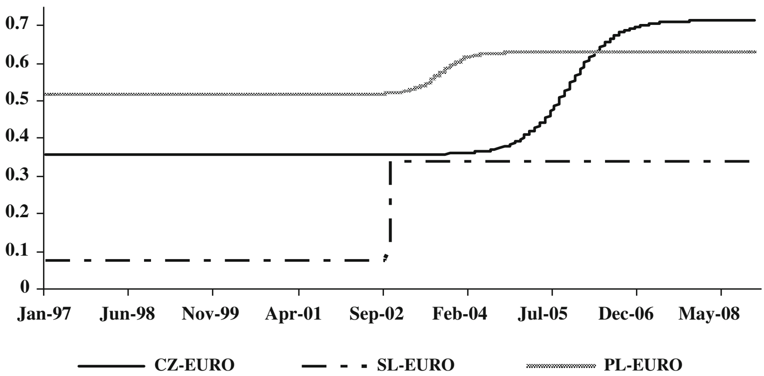

Fig. 1 Time-varying correlations (STCC): Czech Republic versus EURO, Slovenia versus EURO and Poland versus EURO

(p value is 0.017). Therefore, we proceeded with the estimation of a DSTCC model for this market. The results are reported in Table 6 (panel a) and the correlations are shown in Fig. 2. As seen, the Czech market demonstrates a twice increasing correlation pattern generating a stepwise process. Nevertheless, the time-pattern of increase in correlation is similar to that implied by the single transition STCC model in Table 4. Notice also that the threshold parameters in DSTCC are imprecisely estimated. In light of this, the DSTCC estimates seem to be just a refinement of the single STCC ones.

The above findings constitute evidence of increased correlation between three Eastern European and the Euro-zone stock markets. In particular, the Czech, Slovenian and Polish markets have increased their correlation to the Euro-zone from 1997 to 2008. Furthermore, the timing of the shifts in correlations is interesting. In particular,

Table 5 Tests of STCC- against DSTCC-GARCH

Panel a	
CZ–EURO	0.017*
SL–EURO	0.791
PL–EURO	0.778
Panel b	
HU–CZ	0.002**
HU–PL	0.600
CZ–SL	0.649
CZ–PL	0.045*
Panel c	
CZ–US	0.155

Results (p values) from tests of an additional transition in STCC-GARCH
*, ** Significance at the 5 and 1% level, respectively

Table 6 DSTCC-GARCH models

	ρ_1	ρ_2	ρ_3	γ_1	γ_2	c_1	c_2	Date1	Date2
Panel a									
CZ–EURO	0.356	0.555	0.726	500	500	0.692	0.792	04 Apr 05	12 Jun 06
	(0.037)	(0.119)	(0.037)	(.)	(.)	(0.537)	(0.437)		
Panel b									
HU–CZ	0.500	0.694	0.685	15.07	14.99	0.648	0.682	27 Sep 04	21 Feb 05
	(0.034)	(9.929)	(0.036)	(174.9)	(3059)	(1.413)	(21.44)		
CZ–PL	0.470	0.566	0.759	39.84	500	0.590	0.689	19 Jan 04	21 Mar 05
	(0.033)	(0.099)	(0.026)	(412.7)	(.)	(0.103)	(0.157)		

The table presents quasi maximum likelihood estimates of part of the parameters of DSTCC-GARCH models; remaining parameter estimates are available upon request; 'Date1' is the week that corresponds to c_1 (threshold 1) and 'Date2' is the week that corresponds to c_2 (threshold 2); (.) gammas are estimated large with large standard error (for more details, see Teräsvirta 1994, p. 213); values in parentheses are robust standard errors

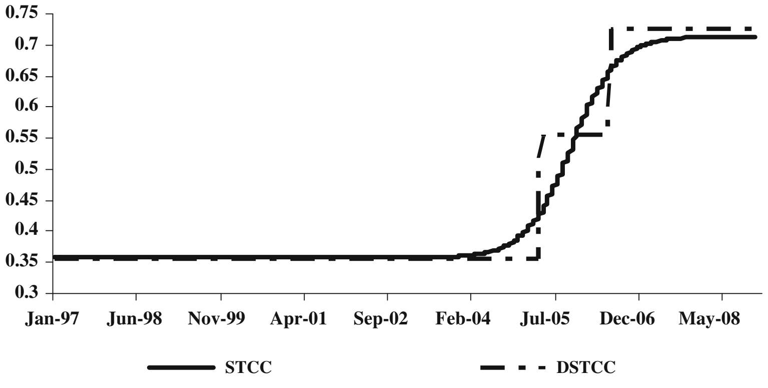

Fig. 2 Time-varying correlations (STCC vs. DSTCC): Czech Republic versus EURO

for Slovenia and Poland, the shift occurs before the EU accession, while for the Czech Republic it starts in the period before the accession and gradually continues after the accession date. The results for Slovenia are of particular interest. This country joined the ERM II in June 2004 and adopted the common currency in January 2007. These findings are in line with Kim et al. (2005) results of rapid increase in stock market integration for most old EU-15 members already since 1996–1997. As the authors argue there are three broad channels through which a currency union can affect financial market integration, namely, exchange rate risk, business cycle and monetary policy convergence. All new EU members are expected to join the Euro at some point in the near future and have achieved a high degree of business cycle correlation.[12] As stock markets move in anticipation of future events, forward looking investors may have already factored in the adoption of the Euro into the stock prices prior to its introduction.

However, the increased degree of stock market correlation achieved by the above three countries is not a broad-based phenomenon across the five CEECs. For example, the correlation for the Hungarian market, although high, has remained unchanged, while the Slovakian market shows very low and statistically insignificant correlation.[13]

Another interesting issue is whether the correlations among the five CEECs markets change during the reporting period. The CCC models in Table 3 (panel b) reveal pronounced comovements between the Czech, Hungarian and Polish markets, which shows these three markets are strongly linked among themselves as well as vis-à-vis the Euro-zone. Table 4 (panel b) reports the STCC models for all cases, for which the hypothesis of constant correlation is rejected (Table 2). As observed, there is evidence of increased correlation among the CEECs, which strengthens our earlier finding of increased comovement between the three Eastern European markets and the Euro-zone. Notice also that, in general, the correlation among the CEECs increases after the increase in correlation between the CEECs and the Euro-zone. This can be seen clearly in Fig. 3, which plots the correlations implied by the models Therefore, it seems that the Euro-zone is the driving force behind the correlation switch among the CEECs.

Following our modelling procedure, we also performed tests for a second transition in the above STCC models and reported the results in Table 5 (panel b). There is evidence for a double transition in two cases, which is subsequently modelled by DSTCC models. The estimates are presented in Table 6 (panel b) and the correlations are plotted in Fig. 4. In the HU-CZ model, the second correlation and the two thresholds are estimated with large standard errors, which indicate that a single transition may be estimated twice. On the other hand, the CZ-PL estimates imply some evidence for a second transition, though the estimated thresholds are close. As before, the time-pattern of increase in correlations is similar to that implied by the corresponding single transition STCC models in Table 4. Hence, we conclude that the DSTCC models seem to be just a refinement of the single STCC ones, which implies that the single transition specifications are able to capture the correlation dynamics quite adequately.

[12] The literature finds that, among the Eastern European countries, Hungary, the Czech Republic, Slovenia and Poland show the highest cycle correlation to the Euro-area (e.g., Fidrmuc and Korhonen 2006; Darvas and Szapáry 2005; Savva et al., forthcoming).

[13] Chelley-Steeley (2005) and Cappiello et al. (2006b) also find a high correlation for the Hungarian market.

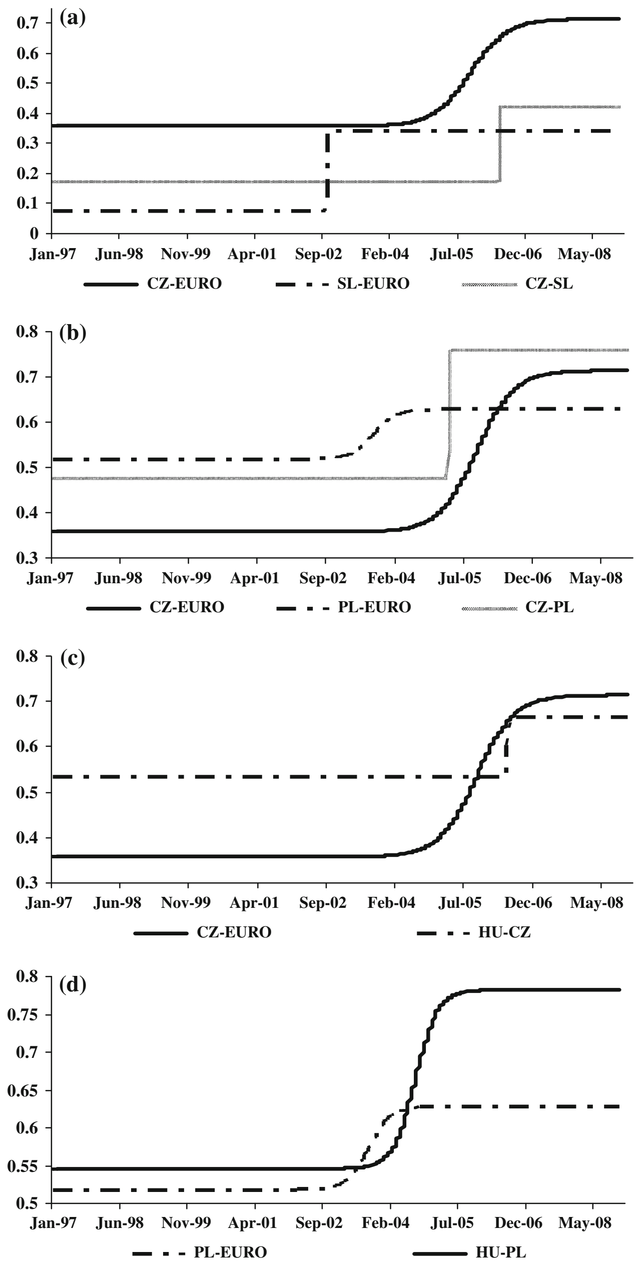

Fig. 3 Time-varying correlations (STCC): **a** Czech Republic versus EURO, Slovenia versus EURO and Czech Republic versus Slovenia; **b** Czech Republic versus EURO, Poland versus EURO and Czech Republic versus Poland; **c** Czech Republic versus EURO and Hungary versus Czech Republic; **d** Poland versus EURO and Hungary versus Poland

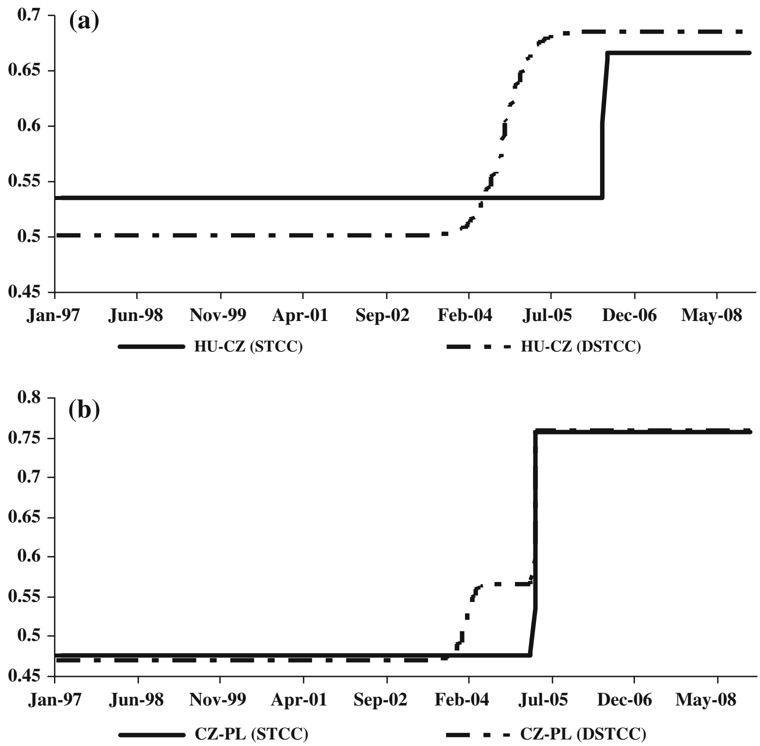

Fig. 4 Time-varying correlations (STCC vs. DSTCC): **a** Hungary versus Czech Republic, **b** Czech Republic versus Poland

We now turn to our second question: was the increase in correlations found in the Czech Republic-, Slovenia and Poland-EURO models part of a world-wide phenomenon of financial integration or was it mainly driven by EU-related developments? In order to answer this question we estimate correlations of stock returns between all five Eastern European markets and the US. In particular, we choose the S&P 500 as the most dominant market in the world. We also estimate the correlation between EURO index and the US to investigate whether the time-pattern of increase in CEE markets' correlations is possibly related to backdrop movements between the EU and the US. The results are presented in Tables 2, 3, 4 and 5 (panel c) and can be summarised as follows. The CCC correlations in Table 3 show that, apart from Slovakia, Eastern European markets are more correlated to the Euro-zone than to the US markets (panel a vs. panel c). Furthermore, the constancy tests in Table 2 (panel c) reject the null hypothesis of constant correlation only in the case of the Czech Republic. This market is subsequently modelled by a STCC model. The results are reported in Table 4 (panel c) and the correlations are shown in Fig. 5. Although, the Czech market increases its correlation with the US,[14] the shift is smaller compared to the shift versus

[14] The result for the Czech market may reflect the fact that a very significant part of the traded volume is due to foreign investors.

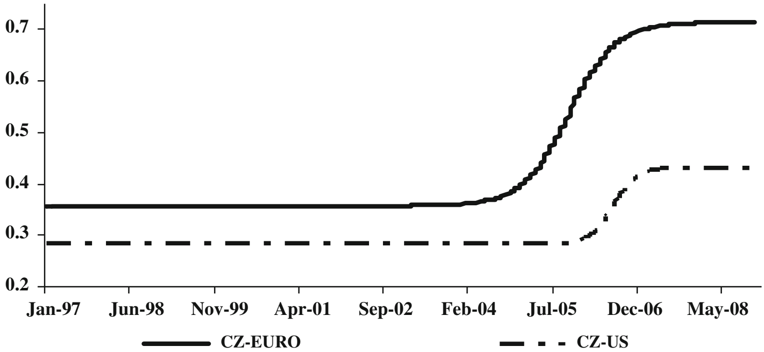

Fig. 5 Time-varying correlations (STCC): Czech Republic versus EURO/US

the Euro-zone. Furthermore, the shift versus the Euro-zone starts before the shift versus the US. Overall, we conclude that the increase in correlation found between the three CEECs and the Euro-zone is mainly driven by EU-related developments rather than by a worldwide phenomenon of financial integration.[15] This is also supported by the finding that the correlation between EURO and the US stock markets, although high, remains constant for the whole sample period, providing some more support for the claim that the increased CEECs integration is mostly due to EU-related developments.

4 Conclusions

In the last decades the Central and Eastern European countries have undertaken important processes of financial reform and stock market development. In order to design optimal investment portfolios as well as from a policy point of view it is of particular interest to assess the degree of financial integration between these countries and other stock markets.

In this paper we use the newly developed (D)STCC-GARCH methodology to analyse the degree of comovements between the stock markets of five CEECs and the Euro zone. By allowing for time-varying correlations, these models are particularly suited to capture the process, if any, of financial integration.

Our results show that the correlation between the Czech, Slovenian and Polish stock markets and those of the Euro-zone has increased substantially from 1997 to 2008. Moreover, this increase is mainly driven by EU-related developments rather than by a worldwide phenomenon of financial integration. We also find that for the Hungarian market the correlation, although high, has remained unchanged, while for Slovakia we observe very low correlation. These results constitute an important first step to further investigate the process of financial integration of Eastern Europe. In future research it would be interesting to identify the particular factors driving the process of stock market integration. In this direction, the results in Hanousek et al. (2008) are interesting as they document that the three largest CEE markets (Hungary, the Czech Republic and Poland) react to macroeconomic shocks, especially those originating from the EU.

[15] This result has also been found in Cappiello et al. (2006b).

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References

Ang A, Bekaert G (2002) International asset allocation with regime shifts. J Financial Stud 15:1137–1187

Baele L (2005) Volatility spillover effects in European equity markets. J Financial Quant Anal 40:373–401

Bekaert G, Harvey C (1995) Time varying world market integration. J Finance 50:403–444

Berben RP, Jansen WJ (2005) Comovement in international equity markets: a sectoral view. J Int Money Finance 24:832–857

Bollerslev T (1990) Modeling the coherence in short-run nominal exchange rates: a multivariate generalized ARCH model. Rev Econ Stat 72:498–505

Bollerslev T, Wooldridge JM (1992) Quasi-maximum likelihood estimation and inference in dynamic models with time-varying covariances. Econom Rev 11:143–172

Cappiello L, Engle RF, Sheppard K (2006a) Asymmetric dynamics in the correlations of global equity and bond returns. J Financial Econom 4:537–572

Cappiello L, Gérard B, Kadareja A, Manganelli S (2006b) Financial integration of new EU member states. Working Paper Series No. 683, ECB

Chelley-Steeley PL (2005) Modeling equity market integration using smooth transition analysis: A study of Eastern European stock markets. J Int Money Finance 24:818–831

Darvas Z, Szapáry G (2005) Business cycle synchronization in the enlarged EU: comovements in the new and old members. Discussion Paper No. 5179, CEPR, London

Égert B, Kočenda E (2007) Interdependence between Eastern and Western European stock markets: evidence from intraday data. Econ Syst 31:184–203

Fidrmuc J, Korhonen I (2006) A meta-analysis of the business cycle correlation between the Euro-area and CEECs: what do we know and who cares? J Comp Econ 34:518–537

Goetzmann WN, Jorion P (1999) Re-emerging markets. J Financial Quant Anal 34:1–32

Goetzmann WN, Li L, Rouwenhorst KG (2005) Long-term global market correlations. J Bus 78:1–38

Hanousek J, Kočenda E, Kutan AM (2008) The reaction of asset prices to macroeconomic announcements in new EU Markets: evidence from intraday data. Working Papers Series No. 349, CERGE-EI, Charles University, Prague

Kočenda E (2001) Macroeconomic convergence in transition economies. J Comp Econ 29:1–3

Kutan AM, Yigit TM (2004) Nominal and real stochastic convergence of transition economies. J Comp Econ 32:23–36

Kim SJ, Moshirian F, Wu E (2005) Dynamic stock market integration driven by the European Monetary Union: an empirical analysis. J Bank Finance 29:2475–2502

King M, Sentana E, Wadhwani S (1994) Volatility and links between national stock markets. Econometrica 62:901–933.1

Longin F, Solnik B (1995) Is the correlation in international equity returns constant: 1960-1990? J Int Money Finance 14:3–26

Longin F, Solnik B (2001) Extreme correlation and international equity markets. J Finance 56:649–676

Ramchand L, Susmel R (1998) Volatility and cross correlation across major stock markets. J Empir Finance 5:397–416

Savva CS, Neanidis K, Osborn DR (forthcoming) Business cycle synchronization of the Euro area with the new and negotiating member countries. Int J Finance Econ

Silvennoinen A, Teräsvirta T (2005) Multivariate autoregressive conditional heteroskedasticity with smooth transitions in conditional correlations. Working Paper Series in Economics and Finance No. 577, SSE/EFI

Silvennoinen A, Teräsvirta T (2007) Modelling multivariate conditional heteroskedasticity with the double smooth transition conditional correlation GARCH model. Working Paper Series in Economics and Finance No. 652, SSE/EFI

Teräsvirta T (1994) Specification, estimation and evaluation of smooth transition autoregressive models. J Am Stat Assoc 89:208–218