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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 correlation between stock markets has increased from 2001 to 2007. In particular, the Czech and Polish markets show a higher correlation to the Euro-zone. However, this is not a broad-based phenomenon across Eastern Europe. We also find that the increase in correlations is not a reflection of a world-wide phenomenon of financial integration but appears to be specific to the European market.

JEL classifications: C32; C51; F36; G15

Keywords: Multivariate GARCH; Smooth Transition Conditional Correlation; Stock Return Comovement; New EU Members.

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 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

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

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 synchronization establishing that many of these new EU members have achieved a high degree of cycle correlation.[3] However, less is known about the progress of these countries towards financial integration which is another important aspect of economic integration.[4] As new EU members these countries will eventually join the EMU and thus it is important to monitor the development of economic and financial links between them and the Euro-zone.

We contribute to fill this gap by addressing two important questions regarding stock market integration between five CEECs (Hungary, Czech Republic, Slovakia, Slovenia and Poland) and the Euro-zone. First, has stock market integration increased following the accession to the EU? 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

[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).

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 (Silvennoinen and Teräsvirta, 2005, and Berben and Jansen, 2005). 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 Balázs and Kočenda (2007), who apply the dynamic conditional correlation GARCH (DCC-GARCH) model of Engle (2002). In this model, however, the underlying unconditional correlation is assumed to be constant over time and, therefore, may not reliably capture the progress toward financial integration experienced by Eastern Europe. The long-run dynamics of financial integration are better captured by our model where the unconditional correlation is allowed to change over time. The idea of a smooth transition in correlations is also explored by Chelley-Steeley (2005), 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 seven years that includes the accession of the CEECs to the EU. Moreover, unlike the previous papers we follow a systematic testing procedure to determine the number of changes in correlations.[5]

[4] Balázs and Kočenda (2007), Cappiello, Gérard, Kadareja and Manganelli (2006) and Chelley-Steeley (2005) constitute notable exceptions.
[5] Cappiello, Gérard, Kadareja and Manganelli (2006) also address this issue although they simply test for the constancy in correlations by using a regression quantile approach.

Our results are summarised as follows. We find that the correlation between the CEECs and the Euro-zone stock markets increased between 2001 and 2007. However, this is not a broad-based phenomenon across the five CEECs. Also, the increase in correlations is not a reflection of a world-wide phenomenon of financial integration but appears to be specific to the European market.

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, Section 4 concludes and discusses further extensions.

2. Methodology

2.1 The model

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

$$y_t = \mu + r_t \qquad t = 1,...,T \tag{1}$$

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

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

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. 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).[6] 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

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

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, $s_t=t/T$. Another way of writing Eq. (5) is

$$\rho_t=\begin{cases}\rho_1 & ,G_t=0\\ \rho_2 & ,G_t=1\end{cases}$$

Values of zero of the transition function identify regime one and values of unity identify the alternative regime. Also, values of G_t between 0 and 1 define situations where the correlation is a mixture of the two regimes (during the transition period).

[6] 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.

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

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

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 $\gamma \to \infty$, $G_t(s_t;\gamma,c)$ becomes a step function ($G_t(s_t;\gamma,c)=0$ if $s_t \le c$ 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 (LM_{CCC}) 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 (LM_{STCC}) of Silvennoinen and Teräsvirta (2007). Under the null hypothesis a single STCC (one change in correlations) is adequate whereas the alternative supports a double STCC (two changes in correlations).[7] 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. 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 (7). 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 robust standard errors are used for the parameter estimates (Bollerslev and Wooldridge, 1992). Furthermore, the log-likelihood is maximized with respect to all parameters simultaneously.[8]

3. Empirical results

The data we use consists of weekly data, denominated in Euro,[9] of the following stock market indices: Budapest (BUX), PX Global Index in the Czech Republic, Slovakia SAX 16, Slovenian PIX, Warsaw General Index 20, Dow Jones STOXX 50 and S&P 500 Composite. The Dow Jones STOXX 50 provides a blue-chip representation of 50 super-sector leaders in the Euro-zone. 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

[7] For analytical expressions of the test statistics and the required derivatives, the reader is referred to Silvennoinen and Teräsvirta (2005, 2007).
[8] All computations are carried out using GAUSS.

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. This data is available from 1996. However, as the period 1997-2000 is characterized by financial market turbulence (Asian-Latin American-Russian crises) and the burst of the dotcom bubble we confine our analysis to the period 2001-2007. More precisely, the sample period starts on January 8, 2001 and ends on July 30, 2007, which yields 343 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. We find that the conditional variances are weekly correlated with an average correlation of 0.086. Finally, we tried models using the bivariate stud-*t* distribution for the errors with the results remaining qualitatively the same.[11]

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[9] Results remain qualitatively the same for data denominated in local currency.
[10] Note that the announcement of EU enlargement was on December 12–13, 2002 with the actual membership happening on May 1, 2004.
[11] All these results are available from the authors upon request.

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 (panel a) shows that the null hypothesis of constant correlation is rejected only for the Czech and Polish markets. For the other three 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 in the Hungarian, Czech and Polish markets the correlation is considerably higher than in the Slovakian and Slovenian markets. In particular, for the first three markets correlations are about 0.5, while for the last two they are about 0.1.

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 both the Czech Republic and Poland the estimates imply a considerable increase in correlation. This can also be seen clearly in Figure 1, which plots the correlations implied by the models. The dates of change and the length of the transition period differ across the two countries. In particular, for the Czech market the estimates point to an instantaneous increase in correlation, from 0.51 to 0.81, in early December 2006. On the other hand, the Polish results show a gradual rise in correlations from 0.41 in mid 2002 to 0.65 in mid 2004. So, the increase happened within a time span of two years. The difference between these patterns may relate to the different approaches taken to financial market development – while the Czech Republic started with large scale privatizations, Poland followed a more gradual process reforming first the legal system and then allowing for the subsequent listing of stocks.

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 are reported in Table 5 (panel a). As the test indicates, there is no second change in the correlations for either the Czech or the Polish markets.

The above findings constitute evidence of increased correlation between the Eastern European and the Euro-zone stock markets. In particular, the Czech and Polish markets show a higher correlation to the Euro-zone over the last seven years.[12] The timing and speed of the increases in correlations differ across the two markets. While for the Czech Republic it occurs as an abrupt break in early December 2006, in the case of Poland, it starts in the period before the accession and gradually continues until mid 2004. More precisely, the increase in correlation observed for Poland starts around the date of announcement of EU enlargement on December 12–13, 2002. Furthermore, the higher degree of correlation is not a broad-based phenomenon across the five CEECs. For example, the correlation for the Hungarian market has remained unchanged. The same is true for the Slovakian and Slovenian markets, which show very low and statistically insignificant correlations (particularly in the case of Slovakia).

The evidence of increased correlation between some Eastern European and the Euro-zone stock markets is in line with Kim, Moshirian and Wu´s (2005) findings of rapid increase in stock market integration for most old EU-15 members already since

[12] Interestingly, Balázs and Kočenda (2007) find very little evidence of stock market integration for these countries. We believe this is because their DCC model does not allow the level of unconditional correlation to change over time and, therefore, may not reliably capture the progress toward financial integration experienced by these countries.

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.[13] 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.

Another interesting issue is whether the correlations among the five CEECs markets considered here 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 the Hungary-Poland and Czech Republic-Poland cases, for which the hypothesis of constant correlation is rejected (panel b of Table 2). In both cases, the estimates imply a very gradual rise in correlation from about 0.35 in 2001 to about 0.8 in 2007, as shown in Figure 2 (solid line). The above conclusion is also supported by the DSTCC models (Table 6) with the time-varying correlations plotted in Figure 2 (dotted lines).[14] Overall, the evidence of increased correlation between the Czech and Polish markets strengthens our earlier finding of strong comovement between these two Eastern European markets and the Euro-zone.

We now turn to our second question: was the increase in correlations found in Czech Republic- 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

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

answer this question we estimate correlations of stock returns between the two Eastern European markets and the US. In particular, we choose the S&P 500 as the most dominant market in the world. The results are presented in Tables 2-3 (panel c) and can be summarized as follows. The CCC correlations in Table 3 show that the Czech and Poland 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) do not reject the null hypothesis of constant correlation in these two markets. Therefore, we conclude that the increase in correlations found with the Euro-zone is not a reflection of a world-wide phenomenon of financial integration but appears to be specific to the European market. [15]

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 STCC-GARCH methodology to analyze the degree of comovements between the stock markets of five CEECs (Hungary, Czech Republic, Slovakia, Slovenia and Poland) 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 Polish and Czech stock markets and those of the Euro-zone has increased substantially from 2001 to 2007.

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[14] Before obtaining the DSTCC estimates, we performed additional transition tests and reported the results in Table 5 (panel b).

Moreover, this increase seems to be specific to the European market and not driven by a worldwide phenomenon of financial integration. We also find that for the Hungarian market the correlation, although high, has remained unchanged, while for Slovenia and Slovakia we observe very low correlations. 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.

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[15] This result has also been found in Cappiello, Gérard, Kadareja and Manganelli (2006).

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Table 1: Summary statistics of the stock returns

	abbr.	min	max	mean	st.dev	skewness	kurtosis
Budapest (BUX) - Hungary	HU	-13.81	11.83	0.393	3.523	-0.376	3.858
PX Global Index - Czech Rep.	CZ	-11.65	6.998	0.460	2.570	-0.792	5.024
Slovakia SAX 16 - Slovakia	SK	-9.062	11.78	0.524	2.615	0.178	5.246
Slovenian (PIX) - Slovenia	SL	-5.938	8.661	0.365	1.869	0.445	5.309
Warsaw Gen. Index 20 - Poland	PL	-12.78	15.26	0.213	3.991	-0.196	3.838
DJ Euro Stoxx 50- Euro-zone	EURO	-12.64	12.78	-0.015	3.071	-0.304	5.710
S&P 500 Composite - USA	US	-11.34	12.13	-0.077	2.743	-0.014	5.653

Notes: Source is DataStream.

Table 2: Tests of CCC- against STCC-GARCH

	LM_{CCC}	p-value
Panel a		
HU – EURO	0.565	0.451
CZ – EURO	3.861	0.049*
SK – EURO	0.502	0.478
SL – EURO	1.610	0.204
PL – EURO	8.300	0.003**
Panel b		
HU – CZ	0.321	0.570
HU – SK	0.041	0.838
HU – SL	2.621	0.105
HU – PL	20.98	0.000**
CZ – SK	0.444	0.504
CZ – SL	0.001	0.970
CZ – PL	28.76	0.000**
SK – SL	0.759	0.383
SK – PL	0.022	0.880
SL – PL	0.028	0.865
Panel c		
CZ – US	0.184	0.667
PL – US	0.001	0.999

Notes: LM_{CCC} is the Lagrange Multiplier statistic for constant correlations; * , ** denote significance at the 5% and 1% level, respectively.

Table 3: CCC-GARCH models

	ρ
Panel a	
HU – EURO	0.480
	(0.046)
CZ – EURO	0.536
	(0.040)
SK – EURO	0.093
	(0.048)
SL – EURO	0.103
	(0.060)
PL – EURO	0.538
	(0.042)
Panel b	
HU – CZ	0.648
	(0.036)
HU – SK	0.194
	(0.048)
HU – SL	0.116
	(0.059)
HU – PL	0.638
	(0.033)
CZ – SK	0.136
	(0.047)
CZ – SL	0.111
	(0.060)
CZ – PL	0.542
	(0.040)
SK – SL	0.044
	(0.049)
SK – PL	0.146
	(0.050)
SL – PL	0.073
	(0.056)
Panel c	
CZ – US	0.417
	(0.046)
PL – US	0.473
	(0.044)

Notes: 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: STCC-GARCH models

	ρ_1	ρ_2	γ	c	Date
Panel a					
CZ - EURO	0.505	0.805	500	0.902	11 Dec 06
	(0.042)	(0.045)	(.)	(0.0001)	
PL – EURO	0.404	0.647	6.097	0.375	23 Jun 03
	(0.087)	(0.052)	(4.539)	(0.145)	
Panel b					
HU – PL	0.336	0.781	2.024	0.311	20 Jan 03
	(0.472)	(0.052)	(1.985)	(0.452)	
CZ – PL	0.281	0.999	1.250	0.669	30 May 05
	(0.248)	(0.481)	(1.680)	(0.257)	

Notes: 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.

Table 5: Tests of STCC- against DSTCC-GARCH

	LM_{STCC}	p-value
Panel a		
CZ - EURO	0.449	0.502
PL – EURO	0.093	0.759
Panel b		
HU – PL	5.506	0.018*
CZ – PL	227.5	2.05e-051**

Notes: LM_{STCC} is the Lagrange Multiplier statistic for an additional transition in STCC-GARCH; *, ** denote significance at the 5% and 1% level, respectively.

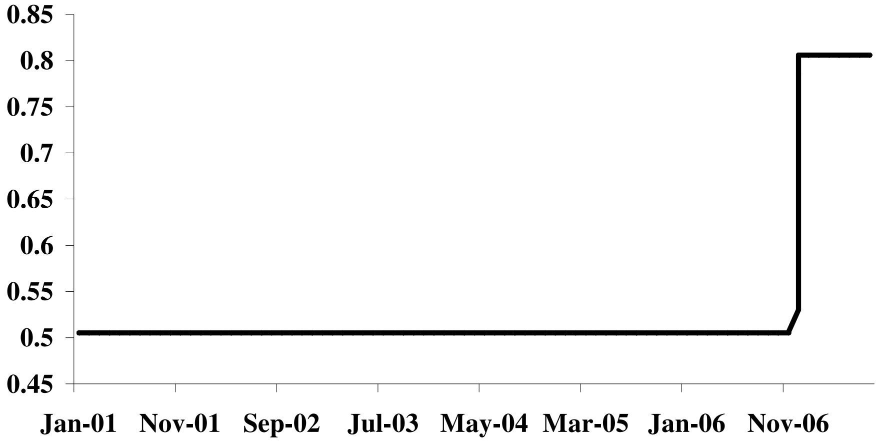
Table 6: DSTCC-GARCH models

	ρ_1	ρ_2	ρ_3	γ_1	γ_2	c_1	c_2	Date1	Date2
HU – PL	0.080	0.995	0.333	0.821	5.732	0.287	1.000	25 Nov 02	30 Jul 07
	(4.156)	(2.086)	(6.489)	(5.808)	(28.29)	(1.816)	(0.928)		
CZ – PL	0.282	0.722	0.880	2.048	500	0.429	0.878	03 Nov 03	09 Oct 06
	(0.196)	(0.128)	(0.074)	(2.381)	(.)	(0.178)	(0.0004)		

Notes: 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); (.) γ is estimated large with large standard error (for more details, see Teräsvirta, 1994, p. 213); values in parentheses are robust standard errors.

Figure 1: Time-varying correlations

(a) Chech Republic with Euro Area



(b) Poland with Euro Area

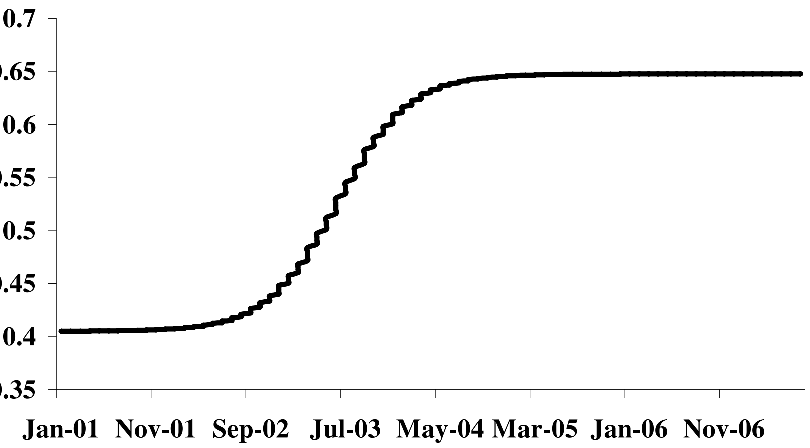
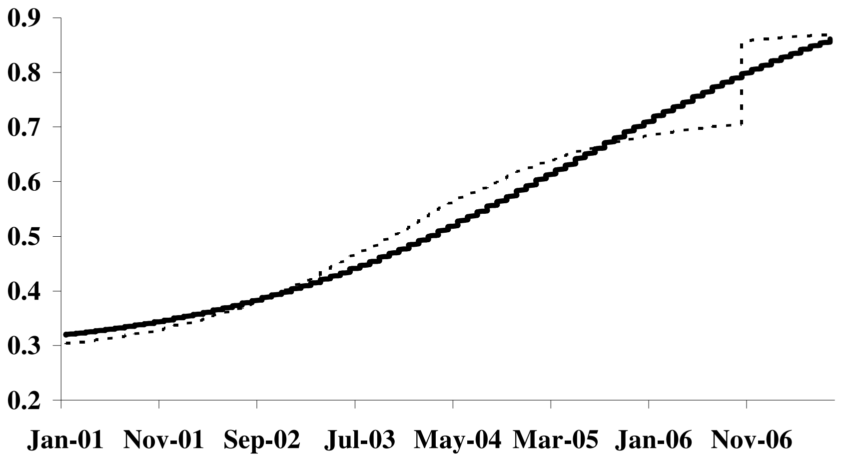


Figure 2: Time-varying correlations (STCC vs. DSTCC)

(a) Czech Republic with Poland



(b) Hungary with Poland