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Time-Varying Comovements in Developed and Emerging European Stock Markets: Evidence from Intraday Data

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Time-Varying Comovements in Developed and Emerging European Stock Markets: Evidence from Intraday Data

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Abstract

We study comovements between three developed (France, Germany, the United Kingdom) and three emerging (the Czech Republic, Hungary and Poland) European stock markets. The novelty of our paper is that we apply the Dynamic Conditional Correlation GARCH models proposed by Engle (2002) to five-minute tick intraday stock price data for the period from June 2003 to January 2006. We find a strong correlation between the German and French markets and also between these two markets and the UK stock market. By contrast, very little systematic positive correlation can be detected between the Western European stock markets and the three stock markets of Central and Eastern Europe, as well as within the latter group.

JEL codes: F37 G15

Keywords: stock markets, intraday data, comovements, bi-variate GARCH, European integration

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1. Introduction

The ongoing process of globalization and European integration has entailed large cross-border capital flows and resulted in stronger real economic linkages between old and new EU member states. Portfolio capital flows accompanied by a deepening of the financial systems in Central and Eastern Europe (CEE) may also have promoted financial market integration in Europe. Greater real and financial integration may imply higher synchronization between developed and emerging European stock markets, as well as among the CEE markets as a group. The hypotheses of higher synchronization are an issue we aim to address in this paper. Our results offer less evidence than might be expected.

Earlier studies that investigated short-run and long-run comovements and contagions[1] between CEE markets and their Western European counterparts did not produce very strong evidence. For instance, Gilmore and McManus (2002, 2003) and Černý and Koblas (2005) did not establish any long-term relationship between the three CEE markets Hungary, the Czech Republic and Poland and the German stock markets for daily or intraday data. Voronkova (2004) shows the presence of long-run links using daily stock market data, on the condition that structural changes are properly accounted for. In a similar vein, Syriopoulos (2004) finds that the CEE markets tend to display stronger linkages with their mature counterparts than with their neighbors. Furthermore, Scheicher (2001) finds evidence of limited interaction between some of the CEE markets and the major markets for daily stock market volatility. There is also little evidence of contagion effects in the CEE stock markets, and CEE stock markets are not more prone to contagion than more developed stock markets (Tse, Wu, and Young, 2003; Serwa and Bohl, 2005).

The above-listed literature uses conventional econometric techniques including cointegration, causality tests and univariate GARCH models. The (G)ARCH revolution entailed the emergence of a number of multivariate GARCH models that provide more efficient tools for analyzing comovements and volatility spillovers between financial assets than the other methods. Still, the first class of multivariate GARCH models implied substantial computing requirements (Kearney and Poti, 2006). A solution for circumventing this problem is the Dynamic Conditional Correlation (DCC) GARCH model of Engle and Sheppard (2001) and

[1] The literature distinguishes cross-market comovements during calm periods from those in periods before and after a crisis. Interdependence defines how strong the interlinkage between two markets is during normal times. We speak of contagion if the interlinkage becomes stronger in the aftermath of a crisis than before it (Forbes and Rigobon, 2002).

Engle (2002), which proved to give a better description of the data than the Constant Conditional Correlation GARCH model (see e.g. Cappiello, Engle and Sheppard, 2003).[2]

Multivariate GARCH applications are largely absent in analyzing intraday data. To the best of our knowledge, the DCC GARCH model has not been applied to emerging European stock markets at this frequency.[3] We aim to fill this gap by investigating the dynamic correlation of time-varying volatilities between three CEE stock markets and also between them and three Western European counterparts over the period from June 2003 to January 2006 on the basis of intraday data recorded in five-minute intervals. The limited evidence of intraday comovements between the CEE and the Western European markets indicate that stock market integration is less than complete.

The outline of the paper is as follows: Section 2 deals with data issues, section 3 focuses on the testing procedure. Section 4 presents the estimation results and section 5 provides concluding remarks.

2. Description of the Intraday Dataset

Our dataset consists of intraday data for European stock markets. We consider three emerging CEE markets (Hungary, the Czech Republic and Poland) and three developed markets (Germany, France and the United Kingdom). Stock exchange indices quoted by Bloomberg are available in five-minute intervals (ticks) for the stock markets in Budapest (BUX), Prague (PX 50), Warsaw (WIG 20), Frankfurt (DAX 30), Paris (CAC 40) and London (UKX).

The time period starts on June 2, 2003, at 1:30 p.m. and ends on January 24, 2006. The time difference between the markets is accounted for by using Central European Daylight Time (CEDT) for all indices, which eliminates the time difference between London and continental Europe. Table 1 gives an overview of the trading hours at the six stock markets.

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[2] The use of multivariate ARCH specifications to model the conditional mean and volatility of stock prices is still not as widespread as the use of conventional univariate models. The methodology is usually used in two strands of financial modeling. One is the modeling of the behavior of stock prices, related financial instruments or stock indices in order to exploit the effect of conditional variance and covariance. Ledoit, Santa-Clara, and Wolf (2003), Bystrom (2004), Hutson and Kearney (2005), McKenzie and Kim (2007) and Kearney and Muckley (2007) are examples of such applications. Testing the validity of the CAPM model is another line of research where Engle and Rodrigues (1989) and Clare et al. (1998) can serve as examples in which the CAPM model with time-varying covariances is rejected.

[3] Lucey and Voronkova (2005) used this method for daily Russian stock market returns. Crespo-Cuaresma and Wójcik (2006) made use of this technique to analyze interest rate data.

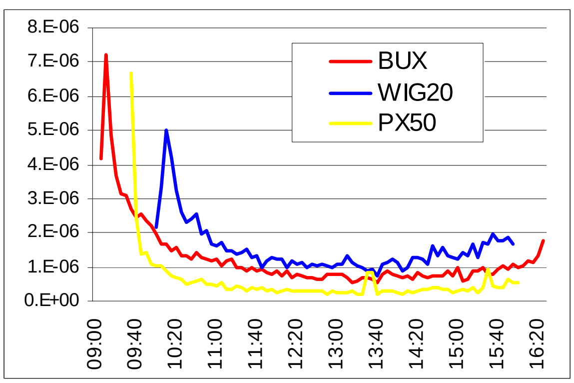
Table 1. Overview of Trading Hours

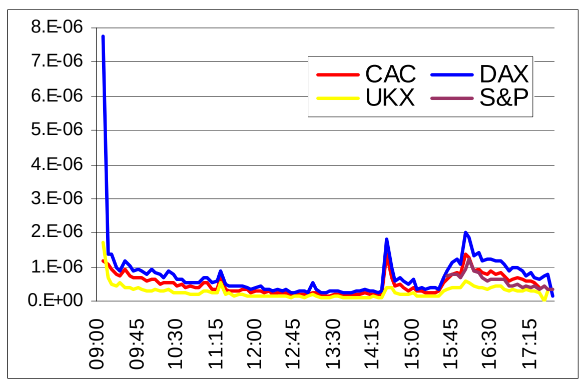
	Start	End	Ticks
BUX	9:00 a.m.	4:25 p.m.	90
PX 50	9:30 a.m.	4:00 p.m.	79
WIG 20	10:00 a.m.	3:55 p.m.	72
DAX	9:00 a.m.	8:10/5:40 p.m.*	135/105
CAC	9:05 a.m.	5:25 p.m.	101
UKX	9:00 a.m.	5:35 p.m.	104

* From November 2003, trading ends at 5:40 p.m.

A big advantage of using intraday data is that the estimates are more robust with respect to structural breaks (Terzi, 2003) given the relatively short time horizon (2 years) as compared to studies employing daily data (up to 10 years). Yet there are two problems that need to be addressed. The first one relates to the fact that trading hours are longer in Western Europe than in the CEE markets. In order to make our analysis fully comparable and executable, we need a common denominator. This could be, for instance, the shortest window, i.e. the one for the WIG 20 running from 10:00 a.m. to 3.55 p.m.

Figure 1. Average Squared Returns and the Intraday *U*-Shaped Pattern





Source: Authors' calculations.

Another, and perhaps more substantial problem is the well-observed fact that absolute returns and volatility, measured for instance in terms of squared returns, exhibit a *U*-shaped pattern during the trading day both in mature and emerging markets. This means that absolute returns and volatility tend to be higher after market opening and before market closing than during the rest of the trading day.[4] This pattern is present in the data because of the arrival and incorporation of news during the beginning of the trading day, differences in intraday trading activity, and also because of the opening and closing of positions at the beginning and at the

[4] See e.g. McMillan and Speight (2002) for the UK and Fan and Lai (2006) for Taiwan.

end of the trading session. The presence of intraday volatility seasonality should be accounted for to avoid compounded results.

Bearing this in mind, we computed the average squared returns during the trading day for the six stock market indices introduced above and for the Standard & Poor's index. The results are plotted in Figure 1 and reveal some stylized facts.

First, one can indeed observe a *U*-shaped pattern for all stock indices. Noticeably, the squared returns are much higher after market opening than before closing. Especially for BUX and WIG 20, the *U*-shape is highly asymmetric as a result. For these two indices, a bump emerges during the first 15 to 30 minutes after market opening, implying that markets need some time to react and incorporate news that materialized between two trading days. The *U*-shape is actually an inverted *J* curve for the other stock indices, as squared returns before market closure do not differ on average from those observed during the day.

Second, volatility in the CEE stock markets appears to be larger during the early hours of trading than in their Western counterparts, with the exception of the tick at 9:05 a.m. of the DAX. Third, as evidenced by the developments in squared returns of the Standard & Poor's index, Western European stock markets are clearly influenced by US macroeconomic announcements at 2:30 p.m. CEDT and by the opening of the US stock markets at 3:30 p.m. CEDT. Yet the CEE markets seem to be affected by none of these effects, perhaps with the exception of PX 50.

This observed intraday behavior can clearly have an influence on the estimation results. For this reason, we take the shortest common window given by WIG 20, i.e. from 10:00 a.m. to 2:40 p.m. and account for the *U*-shaped pattern and the impact of the US event within this window. This leads us to downsize the WIG 20 window to the period running from 11:00 a.m. to 2:40 p.m.

We compute the returns as log first differences where each trading day is a separate sub-sample in order to prevent our results from being distorted by overnight returns. This means that the first return observation on each day is not based on the closing price of the previous day. However, overnight returns are eliminated already by the shortened common window that is free from the *U*-shaped pattern.[5]

[5] It should be noted that some observations are missing for some of the series; they are replaced with zeros.

Table 2 shows some descriptive statistics for the window corrected for the *U*-shaped pattern according to which the log stock returns exhibit a high degree of autocorrelation (Ljung-Box test for residuals).

Table 2. Descriptive Statistics, Common Window

	Log levels						Log differences					
	BUX	**PX-50**	**WIG20**	**CAC**	**DAX**	**UKX**	**BUX**	**PX-50**	**WIG20**	**CAC**	**DAX**	**UKX**
Mean	9.51	6.82	7.52	8.25	8.33	8.46	1.2E-05	1.5E-05	8.3E-06	1.5E-06	3.0E-06	-3.1E-06
Median	9.46	6.79	7.50	8.23	8.32	8.44	0.0E+00	0.0E+00	-3.9E-06	4.4E-06	2.6E-06	0.0E+00
Maximum	10.08	7.33	7.99	8.50	8.62	8.66	0.01	0.02	0.01	0.01	0.01	0.004
Minimum	8.95	6.27	7.08	8.01	8.01	8.29	-0.01	-0.02	-0.01	-0.02	-0.02	-0.01
Std. Dev.	0.33	0.31	0.19	0.12	0.13	0.09	0.00	0.00	0.00	0.00	0.00	0.00
Skewness	0.10	-0.05	0.22	0.18	0.16	0.36	-0.22	-0.31	0.10	-1.30	-0.69	-1.59
Kurtosis	1.67	1.72	2.92	2.25	2.50	2.04	8.46	105.71	6.45	50.64	26.39	55.73
Jarque-Bera	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
No. of Obs.	27,423	27,379	27,456	28,040	27,919	27,481	27,422	27,269	27,449	28,036	27,916	27,478

Note: P-values are reported for the Jarque-Bera normality test.

3. Econometric Method – Dynamic Conditional Correlation GARCH

We aim to study the pairwise dynamic correlations for two stock market returns, Δr_1 and Δr_2, at the six markets under research. We hypothesize that the correlations between pairs of returns vary over time and later we document this to be the reality in Figures 2 to 4. Therefore, for the estimation, we opt to use the bivariate version of the Dynamic Conditional Correlation GARCH (DCC-GARCH) model developed by Engle (2002) and Engle and Sheppard (2001).[6]

The estimation of the DCC-GARCH model encompasses two stages. In the first stage, a univariate GARCH model is estimated for the individual time series. In the second stage, the standardized residuals obtained from the first stage are used to derive the conditional correlation estimator.

Following Engle (2002), the DCC-GARCH model for the bivariate vector $\Delta r_t \equiv [\Delta r_{1t}, \Delta r_{2t}]'$ is specified as follows:

$$\Delta r_t \left| \Omega_{t-1} \sim N(0, D_t R_t D_t) \right., \qquad (1)$$

$$D_t^2 = diag\{\omega_1 \omega_2\} + diag\{\kappa_1 \kappa_2\} \circ \Delta r_{t-1} \Delta r_{t-1}' + diag\{\lambda_1 \lambda_2\} \circ D_{t-1}^2, \qquad (2)$$

[6] Engle and Sheppard (2001) use the DCC-GARCH model to estimate the conditional covariance of up to 100 assets using S&P 500 Sector Indices and Dow Jones Industrial Average stocks, and conduct specification tests of the estimator using an industry standard benchmark for volatility models. They demonstrate the strong performance and easy implementation of the estimator.

$$\varepsilon_t = D_t^{-1} \Delta r_t \, , \tag{3}$$

$$Q_t = S(1-\alpha-\beta) + \alpha(\varepsilon_{t-1}\varepsilon'_{t-1}) + \beta Q_{t-1} \, , \tag{4}$$

$$R_t = diag\{Q_t\}^{-1} Q_t diag\{Q_t\}^{-1} \, , \tag{5}$$

where equation (3) represents the standardized errors, S is the unconditional correlation matrix of the errors and $\circ$ is the Hadamard product of two matrices of the same size (element-by-element multiplication). The parameters of the DCC-GARCH model can be estimated using maximum likelihood.

If $\alpha + \beta < 1$, equation (4) is mean reverting (mean reverting DCC-GARCH). On other hand, $\alpha + \beta = 1$ results in the integrated DCC-GARCH model as equation (4) collapses to equation (4'):

$$Q_t = (1-\phi)(\varepsilon_{t-1}\varepsilon'_{t-1}) + \phi Q_{t-1} \, . \tag{4'}$$

A standard Likelihood Ratio test ($LR = 2(\log L_{\alpha+\beta=1} - \log L_{\alpha+\beta<1}), LR \sim \chi^2$) can be used to discriminate between (4') and (4).

4. Empirical Findings

We first need to check the stationarity of the stock return series. We use three unit root and stationarity tests: the augmented Dickey-Fuller (ADF) and Philips-Perron (PP) unit root tests and the Kwiatkowski, Phillips, Schmidt, and Shin (KPSS) stationarity test. The results reported in the Appendix indicate clearly that the stock market return series are stationary both in levels and in first differences.

As outlined in section 3, the estimation of the DCC-GARCH model consists of two stages. In the first stage, we follow Lee (2006) and Crespo-Cuaresma and Wójcik (2006) and estimate a bivariate Vector Autoregression (VAR) model for the return series. Formal checks justify the use of a GARCH model as the null hypothesis of normality and that of homoscedasticity can be rejected for the VAR model residuals (see Table A2 in the Appendix). Hence, we then use the residuals of the VAR model as inputs for the univariate GARCH (1,1) model. The information criterion always selects the simple GARCH model with no autoregressive term in the conditional mean equation. In the second stage we use the standardized residuals from a

GARCH model to estimate the pairwise specifications of the DCC-GARCH model. Table 3 shows that the Likelihood Ratio test always selects the integrated DCC models.

Table 3. Log Likelihood Ratio Test for the Integrated and Mean Reverting DCC models

	logLR		LR test
	Integrated	**Mean reverting**	
DAX-CAC	-38663.85	-38496.72	-334.26
CAC-UKX	-45030.72	-44972.71	-116.02
DAX-UKX	-45256.54	-45226.31	-60.46
BUX-PX50	-48615.51	-48614.41	-2.20
BUX-WIG20	-48618.44	-48619.05	1.22
WIG20-PX50	-45972.53	-45972.77	0.48
CAC-BUX	-38663.85	-38496.72	-5.02
CAC-PX50	-48619.57	-48619.9	0.66
CAC-WIG20	-48617.74	-48618.42	1.36

Note: * indicates that the null of an integrated DCC is rejected in favor of the mean reverting DCC model at the 5% significance level.

The Dynamic Conditional Correlations obtained from the DCC-GARCH models are plotted in Figures 2, 3 and 4. All three figures exhibit varying patterns in the correlation dynamic path, which justifies the use of the DCC-GARCH modeling strategy. The French and German stock market indices exhibit the highest correlation for returns in general. The plotted DCC ranges in a corridor of 0.5 and 0.9 between June 2003 and January 2006. These two stock markets seem to be less correlated with the UK market, where the DCC typically varies between 0.3 and 0.6. Nevertheless, the weakening of the correlation between the French and German markets during the period under study broadly coincides with changes in the DCC between those two markets and the UK stock market, indicating a rising integration of the three markets. Further, despite that the degree of correlation between German and UK markets slightly weakens over the researched period, our results support those reported by Berben and Jansen (2005) for an earlier period.

Figure 2. Dynamic Conditional Correlation: CAC, DAX and UKX
June 2003 to January 2006

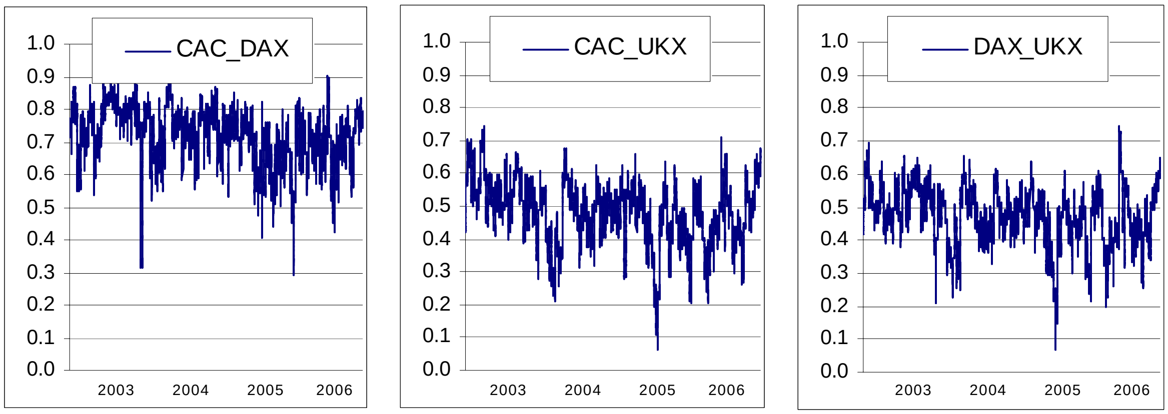


Figure 3. Dynamic Conditional Correlation Between the UK and the CEE Stock Markets:
CAC, BUX, PX50 and WIG20, June 2003 to January 2006

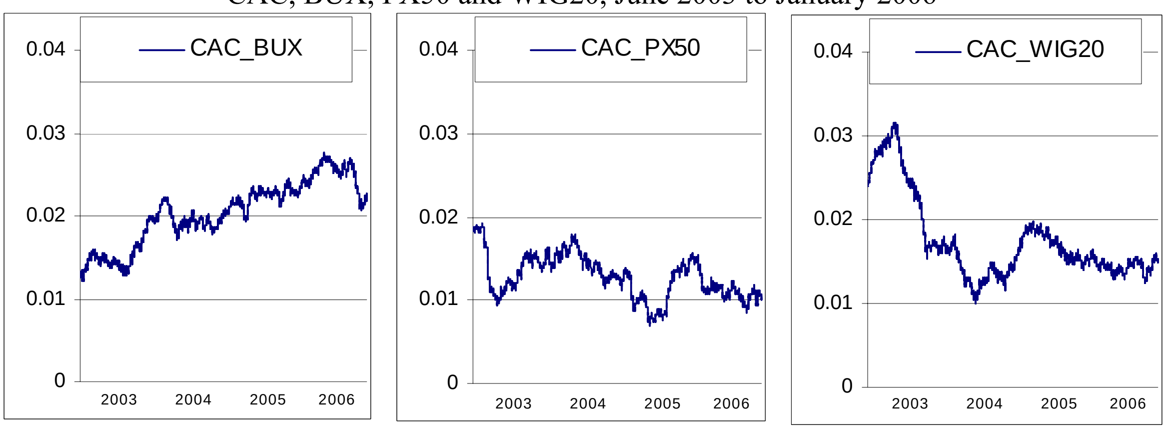
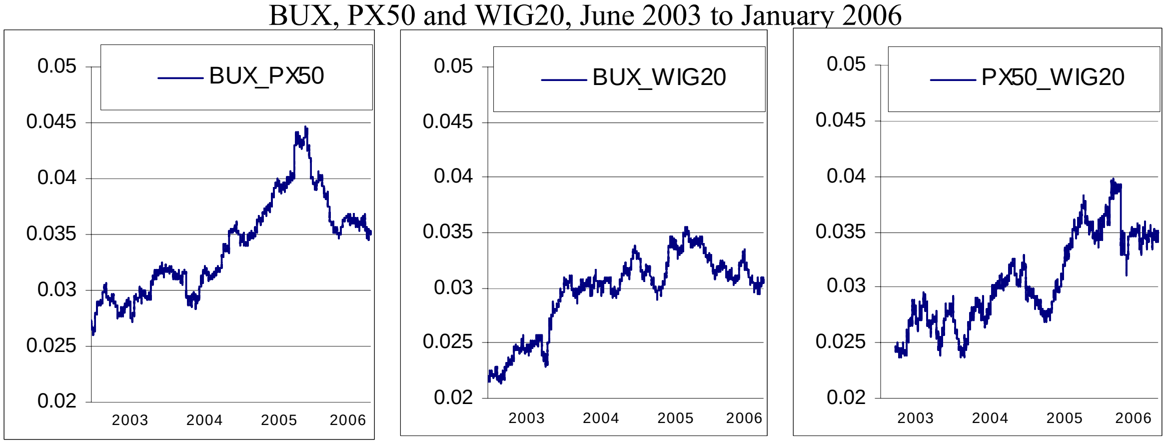


Figure 4. Dynamic Conditional Correlation Between CEE Stock Markets:
BUX, PX50 and WIG20, June 2003 to January 2006



Next, in order to assess our hypothesis of the higher synchronization between developed and emerging European stock markets, we observe the comovement of returns between the three Central and Eastern European stock markets and the French index CAC that we take as a benchmark for Western Europe. The French index is used because the market capitalization of the Paris stock exchange has been recently more than double that of Frankfurt and close to that of London and also registered the largest increase among the three markets (WFE, 2007). Even more importantly, the Paris stock exchange operates in the same time zone as the CEE markets, which eliminates data losses due to the time zone difference, which is the case for London. The comovements studied here reveal a completely different picture from the one between pairs of the developed EU markets. While all three CEE stock markets are positively correlated with the return of the French market, the correlation is quantitatively negligible, ranging between 0.01 and 0.03 (Figure 3). The low correlation goes against the higher synchronization hypothesis and hints at an existing potential for portfolio diversification. Still, the pattern of the varying correlations is different for each market pair. Budapest exhibits a mild increase in correlation with Paris, Prague seems to be quite level and Warsaw levels off after a sharp decrease in correlation.

Finally, we assess the hypothesis of the higher synchronization among the CEE markets as a group. The overall pattern looks similar to the previous account when it comes to the DCC within the group of CEE markets (Figure 4). The time-varying correlation coefficient is moving in a band of 0.02 to 0.05 for the country pairs BUX-PX50, BUX-WIG20 and PX50-WIG20. The magnitude of varying correlations is about double of that between individual CEE markets and the Western European benchmark. It does not support high synchronization hypothesis but still warrants plausible portfolio diversification among the three markets. Notwithstanding the low magnitude of the correlation, it started to increase during the second half of our sample. This might be a sign of the effect possibly brought after the three countries joined the EU in May 2004. Any stronger statement on the subject would be premature, though.

5. Conclusions

In this paper, we analyzed the time-varying correlation of intraday stock market volatilities for three Western European stock markets (CAC, DAX and UKX) and for three CEE markets (BUX, PX-50, WIG-20). The bivariate version of the Dynamic Conditional Correlation GARCH (DCC-GARCH) model shed light on the strong correlation between the German and

French markets and also between these two and the UK stock market for a common daily window adjusted for the observed *U*-shaped pattern for the period from June 2003 to January 2006. By contrast, very little systematic positive correlation can be detected between the French index (which was used as a benchmark for Western European stock markets) and the three CEE stock markets. Perhaps even more surprising is the finding that the CEE markets among themselves are not very well integrated in terms of comovements in stock market returns. The finding indicates that volatility in these specific CEE markets is apparently driven by local innovation and does not reflect transferred swings in asset prices at other markets.

Our research bears the following implications: The fact that we found very little comovement for stock market returns between the stock markets of CEE and Western Europe on the one hand and among the CEE countries on the other hand may be of importance for international portfolio diversification into the CEE. Nevertheless, the situation may be changing because of two reasons. First, the process of deepening in the CEE capital markets is advancing, and second, the degree of the CEE markets' economic integration with Western Europe is increasing as a result of the European integration process. Thus, missing or weak linkages found today may emerge or become stronger in the future.

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Appendix

Table A1. Unit Root Tests: Stock Market Indices

	ADF trend	ADF constant	PP Trend	PP constant	KPSS trend	KPSS Constant
Log stock returns						
BUX	-158.12**	-158.12**	-158.65**	-158.65**	0.09	0.09
PX-50	-99.76**	-99.75**	-147.63**	-147.66**	0.06	0.24
WIG-20	-121.12**	-121.12**	-167.42**	-167.42**	0.06	0.06
CAC	-157.15**	-157.14**	-157.17**	-157.17**	0.03	0.19
DAX	-117.56**	-117.56**	-164.31**	-164.26**	0.05	0.18
UKX	-155.36**	-155.36**	-155.69**	-155.69**	0.04	0.1
1st differences						
BUX	-40.23**	-40.23**	-4811.9**	-4811.7**	0.5**	0.5*
PX-50	-40.37**	-40.37**	-4403.3**	-4404.1**	0.02	0.02
WIG-20	-41.89**	-41.89**	-5218.8**	-5218.7**	0.02	0.02
CAC	-41.16**	-41.16**	-4956.0**	-4956.2**	0.03	0.03
DAX	-42.03**	-42.03**	-11414.7**	-11173.0**	0.05	0.05
UKX	-39.79**	-39.79**	-7147.5**	-7146.7**	0.08	0.17

Notes: ADF, PP and KPPS are the Augmented Dickey-Fuller, the Phillips-Perron, and the Kwiatowski-Phillips-Schmidt-Shin unit root tests, respectively, for the case including only a constant. In parentheses is the lag length chosen using the Schwarz information criterion for the ADF test, and the Newey West kernel estimator for the PP and KPSS tests. * and ** denote the rejection of the null hypothesis at the 5% and 1% levels, respectively. For the ADF and PP tests, the null hypothesis is the presence of a unit root, whereas for the KPSS tests, the null hypothesis is stationarity.

Table A2. Residual checks on the VAR (p-values)

Index pair	VAR lag	Breusch-Godfrey LM for serial correlation H0=no serial correlation lag10	lag 20	White Heteroscedasticity H0=homoscedasticity	Jarque-Bera H0=normality
CAC-DAX	2	0.883	0.848	0.000	0.000
CAC-UKX	2	0.783	0.408	0.000	0.000
DAX-UKX	3	0.536	0.808	0.000	0.000
BUX-PX50	2	0.143	0.174	0.000	0.000
BUX-WIG20	3	0.509	0.722	0.000	0.000
PX50-WIG20	2	0.215	0.050	0.000	0.000
CAC-BUX	1	0.347	0.680	0.000	0.000
CAC-PX50	2	0.246	0.158	0.000	0.000
CAC_WIG20	2	0.999	0.766	0.000	0.000

Note: p-values lower than 0.1, 0.05 and 0.01 indicate that the null hypothesis is rejected at the 10%, 5% and 1% levels, respectively.

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