# No. 44-2009

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This version: 10 November 2009

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# News and Correlations of CEEC-3 Financial Markets

## Abstract

We investigate conditional correlations between six CEEC-3 financial markets estimated by DCC-MGARCH models. In general, the highest correlations exist between Hungary and Poland in foreign exchange and stock markets. Short-term money markets are rather isolated from each other. We find that the associations of CEEC-3 exchange rates versus the euro are weaker than those versus the US dollar. The persistence of the effect of shocks on the time-varying correlations is strongest for foreign exchange and stock markets, indicating a tendency toward contagion. In searching for the origins of financial market volatility in the CEEC-3, we uncover some evidence of Granger-causality on the foreign exchange markets. Finally, using a pool model, we investigate the impact of euro area, US, and CEEC-3 news on the correlations. Apart from ECB monetary policy news, we observe no broad effects of international news on correlations; instead, local news exerts an influence, which suggests a dominance of country- or market-specific circumstances.

## I. Introduction

In today's world of integrated financial markets, local news seldom has a merely local effect but often also causes financial market reactions in neighbouring countries. This will be even more likely to occur if the countries in question share some key characteristics, as do the CEEC-3 (Czech Republic, Hungary, and Poland), which are all emerging transition economies. If, in making investment decisions, economic agents do not distinguish between individual countries but treat them as a homogenous region, it could result in contagion. The question of whether there is contagion among interrelated financial markets is of great concern to financial investors, as its existence would imply that in the case of a shock diversification becomes ineffective.

We investigate the CEEC-3 financial markets for several reasons. First, they represent some of the largest financial markets in the region in terms of liquidity and market capitalisation.[1] Second, the three economies are closely interrelated in terms of trade relations and geographic proximity. Third, as emerging economies, they are particularly prone to financial crisis, as witnessed during the 1990s. Finally, they are in the process of integrating into the European Union. To establish whether the CEEC-3 exhibit signs of contagion, we use a DCC-MGARCH (dynamic conditional correlation multivariate generalized autoregressive conditional heteroskedasticity) model to estimate the cross-country conditional correlations of returns of six financial markets. We then test the reaction of these correlations to local (political and macroeconomic) news and EU as well as US macroeconomic news. If these events result in a strengthening of financial market interdependence, we interpret it as evidence of contagion.

Even though our study does not cover a period of crisis, it is closely related to the literature examining financial market contagion. This branch of research flourished after the financial market crises of the 1990s (such as the Mexican, Asian, Russian, Argentine, and Brazilian financial market crises). However, despite the popularity of the term, 'contagion' has no unanimous definition, nor has a common measure of detecting it been established.

We use the definition of contagion proposed by Forbes and Rigobon (2001): 'a significant increase in cross-market linkages after a shock to an individual country (or group of countries)'.[2] Thus, a co-movement of markets (which some authors define as contagion) or merely a high interdependence of two markets is not sufficient, under our definition, to

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[1] For cross-country data on financial market developments, see the EBRD country database.
[2] Forbes and Rigobon (2001) call this kind of contagion 'shift-contagion'.

constitute contagion.[3] Contagion has also been defined as an increase in the probability of a crisis in one country given a crisis in another country (Eichengreen et al., 1996).

To empirically detect instances of contagion, two main approaches are employed. The first approach is related to the Eichengreen et al. (1996) definition and consists of investigating whether the likelihood of a crisis in one country depends on local fundamentals, events in another country, or some common factors shared by these countries (e.g., Haile and Pozo, 2008; Fazio, 2007; Eichengreen et al., 1996).

The second approach is based on the Forbes and Rigobon (2001) definition of contagion and empirically examines the developments of cross-country correlation coefficients between financial markets. The studies in this branch of the literature examine cross-country correlation coefficients during periods of crisis and/or they estimate the impact of a specific type of event (not necessarily during a time of crisis) on their development (e.g., Chiang et al., 2007).

There are several channels through which contagion can occur but, again, no consensus as to what these channels are.[4] Three of the most important (and agreed upon) channels are the following (see also Didier et al., 2008; Fazio, 2007). First, a financial crisis can be transmitted via trade (Glick and Rose, 1999). Thus, increased trade integration makes countries vulnerable to contagion. Second, contagion can be caused by financial markets themselves. An important aspect of this channel for emerging markets is that financial investors might treat seemingly similar countries as equal due to a lack of information. Third, similar macroeconomic weaknesses in different countries may imply that all countries will be treated the same if one of them faces a crisis (Fazio, 2007).

We examine the impact of specific local (political and macroeconomic) and international macroeconomic news (shocks) on CEEC-3 conditional correlations between six financial markets. Time-varying correlation coefficients are estimated by DCC-MGARCH models for each financial market, an approach that overcomes the problem noted by Forbes and Rigobon (2001, 2002) of a bias toward finding that contagion exists when using unconditional correlations, as these increase during crises as a result of higher volatility. The DCC-MGARCH model addresses this problem and thus eliminates this potential bias.[5] By analysing the influence of local news, we can test whether real linkages between the CEEC-3 and/or lack of information cause investors to treat them equally and thus create contagion.

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[3] In contrast, Didier et al. (2008) define contagion as a price movement in one market resulting from a shock in another market.

[4] For instance, Masson (1998) discusses monsoonal effects, spillovers, and pure contagion effects.

[5] Other solutions were proposed, inter alia, by Bonfiglioli and Favero (2005) and Corsetti et al. (2005), who also provide a more comprehensive review of other approaches.

International news can be viewed as a common shock to all three economies and is expected to increase cross-country correlations.

Our specific research questions are: (i) How can we characterise the cross-country correlations and, in particular, study whether there are differences between financial markets and/or among countries? (ii) Is there a clear direction of (Granger-) causality of volatility from one country to another and can we identify one of the CEEC-3 as the principal source of volatility? (iii) Does news originating from the CEEC-3, the European Union, and the United States affect the time-varying correlations? Can we identify categories of news that might be important sources of contagion?

The paper proceeds as follows. In Section II, we provide a brief overview of the existing literature on news and financial markets with a focus on contagion. In Section III, we describe the construction of the news events and the data sources. Section IV introduces the econometric methodology. We present our results in Section V and conclude in Section VI.

## II. Literature Overview

Only a few recent studies examine the connection between certain events and market interactions, and those that do usually concentrate on a period of financial turmoil. Chiang et al. (2007) investigate the cross-country correlations of nine Asian stock markets' daily returns by means of a DCC-MGARCH model. Within the time span of 1990–2003, they identify a period of contagion (defined as an increase in correlation) and a period of herding (defined as continued high correlation). The authors find that rating agency decisions on the creditworthiness of one of the sample countries have a statistically significant impact on the time-varying correlations between all the countries.

Filleti et al. (2008) study financial crisis contagion among several emerging markets using alternative estimation methods, including DCC-MGARCH. The authors identify six crises during the observation period 1995–2004: Asia, Russia, Brazil, NASDAQ, 11 September, and Argentina. In general, conditional correlations among the emerging markets increased during the crises. Interestingly, no contagion effects between Argentina and Brazil were detected during their respective crises. The authors argue that this finding might be because both crises were somewhat expected.[6] Alternatively, one can interpret these crises as evidence that contagion does not necessarily have to occur because financial markets perceived these shocks to be country-specific.

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[6] In regard to the Argentinean crisis, Didier et al. (2008) argue along the same line.

Albuquerque and Vega (2008) study the effect of news about economic fundamentals (macroeconomic and earning news) on the correlation of US and Portuguese stock markets using a GARCH model. They find that US news and Portuguese earning news have no affect on the cross-country stock market return correlation, but that Portuguese macroeconomic news tends to lower it. This indicates a 'common shock' effect of news released in a large economy versus idiosyncratic shocks originating from a small economy.

Our study has a bearing on this literature as one focus is on how financial markets react to CEEC-3 macroeconomic and political news compared to how they react to macroeconomic news about the euro area and the United States. The latter type of news can be considered as global shocks, which might have positive effects on the CEEC-3 cross-country correlations, whereas the former, more local-type, of news could initiate asymmetric developments. Our approach is unique in several aspects. (i) We examine a broad number of financial markets in countries not previously studied in this context. (ii) We consider a different set of news variables than do other studies. (iii) The inclusion of local political news allows an evaluation of the extent to which local news affects the CEEC-3 as a whole. (iv) We investigate whether international news plays a role, thereby also addressing the question of whether euro area or US news predominates in influencing CEEC-3 financial market interactions.

## III. Data Description

Table A1 in the Appendix presents descriptive statistics of the financial market variables. These are computed as daily changes in CEEC-3 3-month and 12-month interbank interest rates and in 10-year government bond yields. Moreover, we include the returns of CEEC-3 exchange rates (versus the euro and the US dollar) and of CEEC-3 stock indices. These transformations remove the unit roots from the series.[7] The sample consists of 687 daily observations from 2004–2006.[8] Most financial market time series display excess kurtosis and skewness, indicating non-normal distributions and the presence of ARCH.

Political news dummies were constructed using the Interfax Business Reports for the CEEC-3. Over the observation period, these daily newsletters add up to more than 20,000 pages, consisting of articles by Interfax staff writers and summaries of national newspaper reports. For this study, we selected only those events or decisions that are deemed to have a lasting effect and at least an indirect impact on the timing and/or likelihood of euro adoption,

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[7] We performed augmented Dickey-Fuller tests and all but one time series (3-month Hungarian interest rates) had unit roots. The results are available upon request.

which is of great importance for CEEC-3 financial markets. We employed both automatic and manual search procedures to identify potentially relevant passages in the electronic newsletters. Each of these passages was then checked and interpreted independently by two persons. In cases of disagreement as to proper categorisation, the opinion of a third party was sought, thus ensuring a high reliability of the coding.

Macroeconomic news in the CEEC-3 was taken from preliminary publications of statistical offices and central banks, which encompass announcements on GDP growth rates, inflation rates, and the balance of payments (current account and/or trade balance shocks). In line with the efficient market hypothesis, an announcement is considered news if it deviates from the expectations of financial markets and consensus forecasts. Due to data limitations, we have news on the balance of payments only in some cases.

For the United States and the euro area, we derive the surprise component of official macroeconomic announcements by calculating the difference between the announcements and consensus forecasts published by Bloomberg.[9] Since we are interested in macroeconomic shocks, we usually cannot use the euro area announcements, as these are based on an aggregation of non-synchronously published national data. Instead, we employ macroeconomic news related to Germany, which is the largest EU economy, except in the case of the business climate indicator and consumer confidence, where appropriate European values are available. We cover nominal indicators (consumer price index, producer price index) and real indicators (gross domestic product, retail sales, industrial production, trade balance, unemployment rate), as well as forward-looking indicators (such as consumer confidence, business climate, IFO index, ISM manufacturing index).

Also included are monetary policy decisions on interest rates in the CEEC-3, as well as by the ECB and the FED. An overview of the news variables is given in Table A2. Following the literature, we do not distinguish between positive and negative news. For the impact on cross-country correlations, it only matters whether a news item triggered global, regional, or country-specific effects. The different types of news are incorporated into the model as impulse dummies and are aggregated in political (Pol), macroeconomic (Macro), and monetary policy (IR) news categories.

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[8] A more detailed description of the data is given in Büttner and Hayo (2008). Note that the number of trading days here differs markedly from the number in that paper, due to the estimation of a multivariate model.
[9] A more detailed description of the data is given in Büttner et al. (2009).

## IV. Econometric Model

Our modelling approach consists of two steps.[10] First, we estimate a DCC(1,1)-MGARCH(1,1) model (Engle, 2002) of the financial market returns and extract the time-varying conditional correlations between CEEC-3 markets. We include five autoregressive lags to account for autocorrelation.[11] Second, we estimate the effects of local and international news on these conditional correlations.

The advantage of MGARCH models is that they allow for time-varying conditional volatilities in the univariate series as well as for time-varying conditional correlations of two or more time series. The first feature is a more realistic representation of financial time series displaying volatility clusters, while the second takes the integration of financial markets into account. The DCC class of these models adds a further dimension by explicitly modelling the time-varying correlations across financial markets and thereby allowing us to discover potential contagion.[12]

In a first stage, we estimate the conditional variance by univariate GARCH(1,1) models of daily returns, specified as

$$h_{it} = \omega_i + \alpha_i u_{it-1}^2 + \beta_i h_{it-1}, \tag{1}$$

where $h_{it}$ and $u_{it}$ are the conditional variance of an asset in country $i$ and the error term, respectively. The sum of $\alpha_i$ and $\beta_i$ determines the persistence of innovations on the conditional variance. If this sum equals 1, we have an integrated GARCH process.[13]

As a starting point for the second (the multivariate) stage we have:

$$u_t | F_{t-1} \approx N(0, H_t) \tag{2}$$

where $F_{t-1}$ captures all information up to $t$–1.

The covariance matrix $H_t$ is defined as:

$$H_t \equiv D_t R_t D_t. \tag{3}$$

$R_t$ is the time-varying correlation matrix and the diagonal matrix $D_t$ contains, on its main diagonal, the time-varying conditional standard deviations $\sqrt{h_{it}}$ of an asset in country $i$

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[10] Chandra (2006) and Chiang et al. (2007) use a similar two-step approach. This is necessary, as DCC-MGARCH models are sensitive to including a large number of dummy variables.

[11] We considered using a richer set of exogenous variables in the univariate GARCH equations, i.e., more than five autoregressive lags, but convergence problems of the estimators suggested employing a parsimonious model.

[12] DCC-MGARCH models assume a uniform dynamic structure for all cross-correlations. Richer dynamics can be modelled using the BEKK model but the number of parameters to be estimated increases substantially. For an overview of MGARCH models, see Silvennoinen and Teräsvirta (2008) or Bauwens et al. (2006).

[13] In the cases where we cannot rule out an integrated GARCH (Nelson, 1990) process in the univariate time series, we use IGARCH.

obtained in the previous stage. $R_t$ is derived as follows. Using $u_{it} = \sqrt{h_{it}}\varepsilon_{it}$, we arrive at matrix $Q_t$, defining the dynamic correlation structure:

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

where $\varepsilon_t$ is a vector with three elements, namely, the standardised errors from the country-specific GARCH(1,1) models. $\overline{Q}$ is the unconditional, time-invariant correlation matrix of $\varepsilon_t$ (or the unconditional covariance matrix of $u_t$). The structure of this time-varying covariance equation is the same as in the univariate GARCH(1,1) above, i.e., the sum of $\alpha$ and $\beta$ gives the persistence of a shock on the covariance. The off-diagonal elements of $Q_t$ are the conditional covariances of our variables of interest. Correlation matrix $R_t$ is derived as:

$$R_t = Q_t^{*-1} Q_t Q_t^{*-1},$$ (5)

where the diagonal matrix $Q_t^*$ contains the square roots of the diagonal elements of $Q_t$. In the case of the CEEC-3, we obtain:

$$R_t = \begin{pmatrix} 1 & \rho_{CZHUt} & \rho_{CZPLt} \\ \rho_{CZHUt} & 1 & \rho_{HUPLt} \\ \rho_{CZPLt} & \rho_{HUPLt} & 1 \end{pmatrix}.$$ (6)

A typical element of $R_t$ is:

$$\rho_{CZHUt} = \frac{q_{CZHUt}}{\sqrt{q_{CZCZt}\, q_{HUHUt}}}$$ (7)

$$= \frac{(1-\alpha-\beta)\overline{q}_{CZHU} + \alpha(\varepsilon_{CZt-1}\varepsilon'_{HUt-1}) + \beta q_{CZHU\,t-1}}{\sqrt{(1-\alpha-\beta)\overline{q}_{CZCZ} + \alpha\varepsilon_{CZt-1}^2 + \beta q_{CZCZt-1}}\sqrt{(1-\alpha-\beta)\overline{q}_{HUHU} + \alpha\varepsilon_{HU\,t-1}^2 + \beta q_{HUHU\,t-1}}}$$

with $q_{ij}$ (i, j = CZ, HU, PL) an element of $Q_t$ and $\overline{q}_{ij}$ of $\overline{Q}$. The parameters of the DCC-MGARCH model can be estimated by maximum likelihood.[14]

## V. Estimation Results

The estimation reveals several important findings. In the first-stage univariate GARCH equations, in all but six markets weak market efficiency is violated, as past returns can be used to predict today's returns.[15] A possible explanation is the low liquidity of the still relatively immature CEEC-3 financial markets. For no market can we reject the existence of an integrated GARCH in at least one of the three countries.

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[14] In the case of 12-month interest rates, we had to exclude an outlier to ensure convergence.

Turning to the second-stage estimates of Equation (8), we observe that the unconditional correlations are high between Hungary and Poland in all markets except 3-month interest rates and bond yields (see Table 1). This suggests that these two countries' financial markets are very integrated, perhaps perceived to be at similar stages of transition. The Czech Republic can be viewed as more advanced in terms of real and nominal convergence.[16] From a market perspective, correlations are highest on the foreign exchange and stock markets.[17] Short-term interest rates show only small correlations, which suggests dissimilar domestic monetary policies.

**Table 1: Parameters of the Correlation Equations**

| Equation | 3M | 12M | 10Y | CEEC-3/€ | CEEC-3/$ | Stocks |
|---|---|---|---|---|---|---|
| $\overline{\rho}_{CZHU}$ | 0.06* | 0.09*** | 0.18*** | 0.41*** | 0.74*** | 0.53*** |
| $\overline{\rho}_{CZPL}$ | 0.04 | 0.10*** | 0.36*** | 0.41*** | 0.68*** | 0.52*** |
| $\overline{\rho}_{HUPL}$ | 0.06 | 0.18*** | 0.34*** | 0.56*** | 0.81*** | 0.61*** |
| $\alpha$ | 0.02 | 0.02 | 0.01 | 0.03*** | 0.02*** | 0.03 |
| $\beta$ | 0.05 | 0.48*** | 0.46 | 0.95*** | 0.98*** | 0.94*** |
| df | 2.60 | 2.57 | 4.46 | 5.54 | 9.18 | 5.35 |

Note: ***, **, and * indicate significance at a 1%, 5%, and 10% level, respectively.

The correlations of exchange rates versus the euro are lower than those versus the US dollar for the entire sample. This may be because for the CEEC-3, the main reference currency is the euro. For instance, Hungary was operating an ERM-II-style hybrid exchange rate regime versus the euro during the observation period. Hence, shocks of the exchange rate versus the dollar mainly reflect changes in the euro/dollar exchange rate and are therefore symmetric for the CEEC-3. The persistence of shocks on the time-varying correlations is highest for the exchange rates and the stock markets, which may be an indication of contagion

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[15] The weakly efficient markets are: Czech 3-month interest rates, Czech and Polish government bond yields, Czech crown/euro exchange rate, and the respective exchange rates of the Polish zloty versus the euro and the US dollar. Detailed results are available upon request.

[16] For instance, according to Eurostat, the per capita income level in comparison to the EU-27 average in 2007 was 80.2% in the Czech Republic, 62.6% in Hungary, and 53.7% in Poland.

[17] Interestingly, Égert and Kočenda (2007) find correlations among CEEC-3 stock markets of just 0.02–0.05. Although they cover a slightly different sample period (06/2003–01/2006), they apply the same type of DCC-MGARCH model to intraday data. For three western European stock markets (France, Germany, and the United Kingdom), the authors find much higher correlations. One explanation of this noteworthy difference from our results could be that markets in the CEEC-3 are too slow in their reaction, possibly due to low liquidity and less advanced trading platforms.

effects across CEEC-3 financial markets, as shocks on stock and foreign exchange markets lead to a more persistent increase in correlations.


## Granger-Causality Tests

Given that market developments appear to be correlated, it is worth taking a look at whether there is (Granger-) causality from one country to another. We conduct trivariate Granger-causality tests for the conditional volatilities derived in the first-stage estimation. The main results are as follows.[18] For the 3-month interbank interest rates, we find no Granger-causality, highlighting the domestic focus of short-term monetary policy. However, for the 12-month horizon, we find bidirectional Granger-causality between the Czech Republic and Poland. Thus, there appears to be a connection between money markets in the medium term.

We also find that the conditional variance of Polish 10-year interest rates Granger-causes changes in the Hungarian market. We could not obtain transaction volume figures, but we do know that the stock of outstanding government debt is larger in Poland than it is in Hungary. Greater market size can explain why the Polish market leads the Hungarian market with respect to volatility spillovers.

On foreign exchange markets, the conditional variance of the Hungarian forint/euro exchange rate Granger-causes both the Polish zloty/euro and the Czech crown/euro exchange rate volatility. The Polish zloty/euro exchange rate volatility also Granger-causes the Czech crown/euro conditional variance. For the exchange rates against the US dollar, we find a bidirectional relationship between the volatilities of the Czech crown and the Hungarian forint. Thus, the Hungarian forint is a major source of volatility. However, this aspect of the forint cannot be explained by the turnover data on local foreign exchange spot transactions. Figures from the CEEC central banks for April 2007[19] reveal that the sum of spot transactions of local currency versus the US dollar is the highest in the Czech market, followed by the Polish and, finally, the Hungarian markets. Moreover, the market turnover of the Hungarian forint against the euro is lower than the respective figure for the Polish zloty. Therefore, market liquidity does not explain why the Hungarian forint plays such a leading role in CEEC-3 foreign exchange volatility.

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[18] Concerning the lag order of the exogenous variables in the test, we follow the procedure suggested by Hayo (1999), where the choice of the lag length is guided by the elimination of serial correlation in the residuals. Full results are available upon request.

[19] This is the last reference month of the "Triennial Central Bank Survey of Foreign Exchange and Derivatives Market Activity" by the BIS. According to figures retrieved from the CEEC central bank homepages, in April 2007 the respective foreign exchange market turnover of local currencies against the US dollar (euro) was (in million US dollars): Czech Republic 239 (554), Hungary: 103 (718), and Poland: 190 (1,287).

Alternatively, the effect may be connected to problems arising from the stabilisation of the forint due to the so-called Hungarian twin deficit, which might affect volatilities of other currencies as well. Finally, both Czech and Hungarian stock markets Granger-cause the volatility of the Polish stock market. Again, this cannot be explained by market size, as during observation period, the Polish market had the largest market capitalisation.[20]

**POOL estimates**

Next, we test the impact of CEEC-3, EU, and US news on the conditional correlations obtained in the first stage in the framework of a stacked pool model. Due to non-stationarity, we examine the stock market and foreign exchange correlations in differences. That is, we estimate the following pooled OLS equation for each of the markets:

$$\rho_{ij,t} = c + CZHU_{SD} + CZPL_{SD} + \sum_i \sum_{r=1}^{5} \delta_r \rho_{t-r}^i + \sum_{r=0}^{2} News_{t-r} \gamma_r + \varepsilon_t \quad (8)$$

where $\rho_{ij,t}$ is the country-pairwise conditional correlation as in Equations (6) and (7), $c$ is a constant, $CZHU_{SD}$ and $CZPL_{SD}$ are step dummies that capture country fixed effects, $\delta$ and $\gamma$ are vectors of coefficients (parameters are assumed to be homogenous), and $News$ is a matrix containing the three news categories from the CEEC-3, European Union, and the United States. We employ heteroskedasticity-consistent or autocorrelation-consistent standard errors where applicable.[21] Starting with Equation (8), we apply a consistent general-to-specific modelling approach to simplify the model (Hendry, 1995).[22]

**VI. Interpretation of Estimation Results**

Table 2 provides an overview of the remaining, statistically significant variables.[23] Conditional correlations on money markets are barely affected by news, as expectations on domestic monetary policy and the domestic macroeconomic environment dominate these markets. Moreover, as shown in Table 1, these markets are rather isolated from each other. Foreign exchange and stock markets are more integrated with regional and global financial markets and, therefore, market movements following news are more frequent.

In general, international news plays a minor role. In fact, in the case of US news, only macroeconomic news has a negative influence on correlations of 10-year bond yields and

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[20] At the end of 2006, the Polish/Czech/Hungarian market capitalisations were measured as 113/35/32 billion euros (Eurostat).

[21] In the case of heteroskedasticity, we use JHCSE, according to MacKinnon and White (1985). When we find autocorrelation, we apply HACSE, following Andrews (1991). JHCSE is preferred based on its superior small-sample performance.

[22] For the testing-down process, we use the module 'Autometrics' embedded in PcGive12 (see Doornik, 2009).

[23] The control model and the diagnostic statistics can be found in Table A3 in the Appendix.

CEEC-3/$ exchange rates. Given the importance of the United States to the world economy, this is surprising and does not support the notion that a common shock generates higher correlations. Moreover, our European macroeconomic indicators yield mixed results. German macroeconomic news causes an increase in the correlations of the exchange rates versus the dollar, but a lowering of stock market correlations. The former finding might be

**Table 2: Pool Estimates for CEEC-3 Financial Markets Correlations**

| Equation | 3M | 12M | 10Y | CEEC-3/€ | CEEC-3/$ | Stocks |
|---|---|---|---|---|---|---|
| **CZ News** | | | | | | |
| Pol lag 1 | | | 0.002* | | | |
| IR lag 1 | 0.014* | | | | -0.006* | |
| IR lag 2 | | -0.020** | -0.004* | | -0.002* | |
| Macro | | | | | 0.001* | |
| Macro lag 1 | | | -0.002** | | 0.002** | |
| Macro lag 2 | | | -0.001* | | | -0.003* |
| **HU News** | | | | | | |
| Pol | | | | | -0.002* | 0.003* |
| Pol lag 2 | | 0.002* | -0.002** | | | |
| IR | | -0.006* | | -0.007* | | |
| Macro lag 1 | | | | -0.008** | -0.003** | |
| Macro lag 2 | | | 0.002* | | -0.001 | |
| **PL News** | | | | | | |
| Pol lag 2 | 0.002** | 0.002* | | | | |
| IR lag 2 | -0.007* | | | | | |
| Macro lag 1 | | | | | | 0.004* |
| Macro lag 2 | | | | 0.006** | | |
| **EU News** | | | | | | |
| Macro | | | | | 0.001** | -0.002* |
| IR lag 2 | | 0.010* | | -0.013* | | |
| **US News** | | | | | | |
| Macro lag 2 | | | -0.001* | | -0.001** | |

Note: ** and * indicate significance at a 1% and 5% level, respectively. Where appropriate, HACSE (Andrews, 1991) or JHCSE (MacKinnon and White, 1985) were used.

Exchange rate and stock market correlations were estimated in differences, interest rates and bond yields in levels.

CEEC-3/$: CZ Macro lag 1 and CZ Macro lag 2 have been excluded for offsetting effects (Chi²(1) = 0.87).

CEEC-3/$: HU Pol and HU Pol lag 1 have been excluded for offsetting effects (Chi²(1) = 0.70).

CEEC-3/$: HU IR lag 1 and HU IR lag 2 have been excluded for offsetting effects (Chi²(1) = 0.32).

CEEC-3/$: PL Macro and PL Macro lag 2 have been excluded for offsetting effects (Chi²(1) = 0.28).

Stocks: HU Macro and HU Macro lag 1 have been excluded for offsetting effects (Chi²(1) = 0.00).

explained by the importance of the euro as an anchor for the CEEC-3; the latter result suggests an influence of diversification effects. European monetary news lowers 10-year bond correlations but increases correlations of 12-month interest rates. The last finding points to the leading role played by ECB monetary policy in the medium term.

The majority of news affecting the interaction between CEEC-3 financial markets is domestic and influences foreign exchange and stock market developments. Czech monetary policy decisions (IR) increase correlations among 3-month interest rates, but lower correlations of 12-month money markets and bond yields. Finally, Czech (Hungarian) IR news lowers the correlation between exchange rate returns against the US dollar (the euro). Hence, on the foreign exchange markets, an interest rate move leads to diversification effects. Hungarian monetary policy decisions also cause a decrease of the correlations of 12-month interest rates. Polish monetary policy decisions have the same effect on 3-month interest rates. Therefore, Hungarian and Polish (and, to a lesser extent, Czech) monetary policy cannot be considered as leading CEEC-3 monetary policy.

Czech macroeconomic surprises increase conditional correlations of exchange rates versus the US dollar, but have an adverse effect on correlations of bond yields. Hungarian macroeconomic news has the opposite effect. Polish (Czech) macroeconomic news increases (decreases) correlations of stock returns. Finally, Polish macroeconomic news also pushes up exchange rate correlations versus the euro. Hence, only Polish macroeconomic news can be considered as having contagious effects.

Hungarian political news decreases conditional correlations of the CEEC-3 exchange rates versus the US dollar and of bond yields, whereas on the stock markets and on 12-month money markets, we observe an increase. The same occurs on the bond markets (on 3- and 12-month money markets) after Czech (Polish) political news, which suggests that political news is potentially contagious.

The impact of news on stock markets shows an opposite sign than it does on foreign exchange markets. For instance, national news (HU Pol and CZ Macro) causes an increase in stock market correlations and thus spillover or contagion effects. In contrast, the correlations decline on foreign exchange markets, which is indicative of diversification investments.

Finally, from a country perspective, it is interesting that in contrast to the Czech Republic and Poland, Hungarian news lowers, rather than increases, correlations. Thus investors apparently distinguish between the three countries, as Hungarian problems (the twin deficit and political crises) are not attributed to the CEEC-3 as a whole.

## VII. Robustness Checks

To check the validity of our results, we conduct a series of robustness tests. First, we compare our results to a SUR (seemingly unrelated regression) estimation of Equation (8). However, since this framework neither offers new insights nor contradicts our results, we instead report the POOL representation, which is estimated more efficiently.

Second, we vary the aggregation level of news. We separate all macro and political news according to their categories given in the news description (see Section III). However, we find no noteworthy reaction patterns in regard to specific type of news.

Third, we differentiate political news items into important and unimportant ones. This specification is based on Büttner and Hayo (2008), who find important news (defined as headline news) to have a larger impact on financial market returns than does unimportant news. This result does not carry over to the context of conditional correlations.

## VIII. Conclusion

In this paper, we obtain bivariate conditional correlations of six CEEC-3 financial markets using a DCC-MGARCH model and daily data from 2004–2006. We address three main research questions.

(i) The nature of the CEEC-3 financial market correlations. The highest conditional correlations are found between Hungary and Poland in all markets except the 3-month interest rates and 10-year bond yields. From a market perspective, correlations are highest on the foreign exchange and stock markets, whereas interbank money markets appear to be rather isolated from each other. In our sample, the correlations of CEEC-3 exchange rates versus the euro are lower than those versus the US dollar. We argue that changes in the CEEC-3/dollar exchange rate mainly reflect developments of the euro/dollar exchange rate. Among all markets, exchange rates and the stock markets show the highest persistence and thus a tendency toward contagion.

(ii) Direction of (Granger-) causality of volatility from one country to another. We find some evidence of Granger-causality on the foreign exchange markets that cannot be linked to transaction volumes on these markets but is, instead, probably related to problems of stabilising the forint due to the twin deficit. However, we are unable to identify any one of the CEEC-3 as being the principal source of volatility in the region.

(iii) The effect of news on the time-varying correlations. Combining the observations on the three countries in a stacked pool model, we investigate the impact of euro area, US, and CEEC-3 news on the conditional correlations. We observe no major effects of international news on the linkages of markets, suggesting that, at least in times without a major crisis, correlations are not determined by global or common shocks. However, local news appears to exert an influence on financial market correlations. This can be partly explained by country- or market-specific circumstances. Hungary seems to have an impact on CEEC-3 financial markets through its political, monetary, and macroeconomic news, which

generally lowers market correlations. In comparison, reaction to Polish and Czech news is less idiosyncratic.

ECB monetary policy news increases money market correlations in the medium term, indicating the leading role European monetary policy plays for CEEC-3 central banks. Evidence of contagion is found for Czech and Polish political news and for Czech, Polish, and German macroeconomic news (foreign exchange market). Polish macroeconomic news also increases correlations of stock market returns. However, the negative impact of Hungarian news on correlations shows that investors distinguish between the CEEC-3 and thus contagion does not occur following Hungarian news. Therefore, given that CEEC-3 conditional correlations of financial market returns are not highly integrated, investors can successfully engage in diversification strategies within this group of countries.

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**Appendix**

**Table A1: Descriptive Statistics of Daily Returns (growth rates in %)**

|  | 3M | 12M | 10Y | CEEC-3/€ | CEEC-3/$ | Stocks |
|---|---|---|---|---|---|---|
| | | | Czech Republic | | | |
| Obs. | 687 | 687 | 687 | 687 | 687 | 687 |
| Mean | 0.0007 | 0.0007 | -0.0016 | -0.0002 | -0.0003 | 0.0013 |
| Std.Devn. | 0.022 | 0.029 | 0.041 | 0.003 | 0.007 | 0.012 |
| Skewness | -2.1 | -2.0 | -0.3 | 0.0 | 0.1 | -1.0 |
| Exc. Kurt. | 66.3 | 49.2 | 5.1 | 1.6 | 1.0 | 9.0 |
| Min. | -0.25 | -0.37 | -0.27 | -0.01 | -0.03 | -0.09 |
| Max. | 0.23 | 0.23 | 0.21 | 0.01 | 0.03 | 0.07 |
| | | | Hungary | | | |
| Obs. | 687 | 687 | 687 | 687 | 687 | 687 |
| Mean | -0.0062 | -0.0043 | -0.0020 | -0.0001 | -0.0001 | 0.0014 |
| Std.Devn. | 0.078 | 0.108 | 0.084 | 0.005 | 0.008 | 0.015 |
| Skewness | 1.1 | 0.4 | 0.6 | 0.5 | 0.2 | -0.2 |
| Exc. Kurt. | 31.8 | 22.7 | 4.0 | 2.6 | 0.5 | 0.9 |
| Min. | -0.62 | -0.86 | -0.34 | -0.02 | -0.03 | -0.06 |
| Max. | 0.83 | 0.84 | 0.54 | 0.03 | 0.03 | 0.05 |
| | | | Poland | | | |
| Obs. | 687 | 687 | 687 | 687 | 687 | 687 |
| Mean | -0.0062 | -0.0043 | -0.0020 | -0.0001 | -0.0001 | 0.0014 |
| Std.Devn. | 0.078 | 0.108 | 0.084 | 0.005 | 0.008 | 0.015 |
| Skewness | 1.1 | 0.4 | 0.6 | 0.5 | 0.2 | -0.2 |
| Exc. Kurt. | 31.8 | 22.7 | 4.0 | 2.6 | 0.5 | 0.9 |
| Min. | -0.62 | -0.86 | -0.34 | -0.02 | -0.03 | -0.06 |
| Max. | 0.83 | 0.84 | 0.54 | 0.03 | 0.03 | 0.05 |

Notes: PX-50, BUX, and WIG20 are the main local stock indices.
Sources: National central banks, Eurostat, www.money.pl, www.pse.cz, Yahoo.com, www.cbonds.info.

**Table A2: Summary of News**

|  | Czech Republic | Hungary | Poland | EU | US |
|---|---|---|---|---|---|
| Political news | 41 | 99 | 91 | – | – |
| Central bank news | 8 | 21 | 10 | 6 | 17 |
| Macroeconomic news | 90 | 79 | 71 | 220 | 185 |

**Table A3: Control Model of the Pool Estimates for CEEC-3 Financial Markets Correlations**

| Equation | 3M | 12M | 10Y | CEEC-3/€ | CEEC-3/$ | Stocks |
|---|---|---|---|---|---|---|
| CORR1 | 0.110** | 0.596** | 0.468** | | | |
| CORR4 | | -0.042* | | | | |
| Constant | 0.050** | 0.080** | 0.180** | | | |
| CZHU | 0.005** | -0.038** | -0.081** | | | |
| CZPL | -0.018** | -0.033** | 0.011** | | | |
| Observations | 2034 | 2031 | 2034 | 2034 | 2034 | 2034 |
| Log-likelihood | 5834.6 | 5168.7 | 6594.5 | 4513.1 | 6915.0 | 5000.8 |
| AR | $F_{(2,2025)} = 0.9$ | $F_{(2,2018)} = 2.5$ | $F_{(2,2020)} = 1.0$ | $F_{(2,2024)} = 1.5$ | $F_{(2,2018)} = 1.8$ | $F_{(2,2026)} = 1.1$ |
| ARCH | $F_{(1,2025)} = 0.1$ | $F_{(1,2018)} = 0.0$ | $F_{(1,2020)} = 0.0$ | $F_{(1,2024)} = 0.1$ | $F_{(1,2018)} = 1.1$ | $F_{(1,2026)} = 0.3$ |
| Normality | $Chi^2(2) = 9666$** | $Chi^2(2) = 4102$** | $Chi^2(2) = 3767$** | $Chi^2(2) = 4604$** | $Chi^2(2) = 5111$** | $Chi^2(2) = 3130$** |
| Hetero | $F_{(7,2019)} = 2.6$* | $F_{(12,2007)} = 1.7$ | $F_{(12,2009)} = 1.7$ | $F_{(8,2017)} = 1.7$ | $F_{(15,2004)} = 2.5$** | $F_{(6,2021)} = 0.5$ |
| Hetero-X | $F_{(20,2006)} = 1.5$ | $F_{(52,1967)} = 1.0$ | $F_{(62,1959)} = 2.0$** | $F_{(30,1995)} = 1.6$* | $F_{(80,1939)} = 3.7$** | $F_{(21,2006)} = 0.6$ |

Note: ** and * indicate significance at a 1% and 5% level, respectively. Where appropriate, HACSE (Andrews, 1991) or JHCSE (MacKinnon and White, 1985) were used. Exchange rate and stock market correlations were estimated in differences, interest rates and bond yields in levels.