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The Financial Crisis and the Stock Markets of the CEE Countries*

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Abstract

Stock markets in Central and Eastern European (CEE) countries significantly collapsed during the financial crisis of 2008. We studied whether the collapse of stock markets in CEE countries was due to international linkages of deteriorating fundamentals or international spillovers of speculative bubbles. To this end, we estimated a state-space model to decompose the stock market indexes of three large CEE countries (Czech Republic, Hungary, and Poland) into fundamentals and speculative bubbles. We then used the techniques of cointegration analysis to study the long-run linkages of fundamentals and speculative bubbles. Our results suggest that international long-run linkages varied over time. The long-run linkages with the U.S. stock market strengthened in terms of both fundamentals and speculative bubbles during the market jitters caused by the financial crisis of 2008.

1. Introduction

Modern portfolio theory implies that international linkages of stock markets are a key determinant of the benefits of international portfolio diversification. A thorough understanding of the international linkages of stock markets is, therefore, of key importance for international investors. A question of particular interest for international investors is whether international linkages of stock markets strengthen in times of financial crises. If financial crises result in stronger international linkages of stock markets, the benefits of international portfolio diversification may liquefy when they are most urgently needed. In times of financial crises, international linkages of stock markets may increase because fundamentals deteriorate more or less simultaneously across countries, or because of "contagion" effects. While there is no consensus on how exactly contagion should be defined (Forbes and Rigobon, 2001; Rigobon, 2002), the empirical framework we lay out in this paper makes it a natural choice to define contagion in terms of significant international linkages of speculative bubbles in times of financial crises. A speculative bubble, in our research, is defined as the extent to which the actual stock price stochastically deviates from the fundamentals-based stock price. Our perspective of contagion is, in economic terms, in line with that proposed by Bekaert et al. (2005, p. 40), who define "contagion as excess correlation, that is, correlation over and above what one would expect from economic fundamentals".

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The significant collapse of the stock markets in Central and Eastern European (CEE) countries during the recent financial crisis of 2008 provides a natural experiment to study how international linkages of stock markets, fundamentals, and speculative bubbles change in times of crisis.[1] We studied whether the collapse of stock markets in CEE countries was due to international linkages caused by a correlated cross-country deterioration of fundamentals or by contagion effects reflecting international spillovers of speculative bubbles. To this end, we estimated a state-space model similar to the one suggested by Wu (1995, 1997) to decompose the stock market indexes of three large CEE countries (Czech Republic, Hungary, and Poland) into fundamentals and speculative bubbles. Studying these three large CEE countries is interesting because they are member countries of the European Union and, conditional on sustainable convergence as specified in the EC Treaty, they wish to adopt the euro. We then used the techniques of cointegration analysis to analyze the international cointegration linkages of fundamentals and speculative bubbles (Bhar and Hamori, 2005). In doing this, we accounted for the potential instability of the long--term international linkages of the stock markets of the CEE countries. We used monthly data for the sample period from 1995 to 2008 in our empirical analysis.

Our results suggest that the international cointegration linkages of the stock markets of the CEE countries varied over time, and that cointegration linkages of both fundamentals and speculative bubbles strengthened during the market jitters caused by the financial crisis of 2008. *Transatlantic* cointegration linkages with U.S. fundamentals and U.S. speculative bubbles strengthened to a much more significant extent than *continental* cointegration linkages with fundamentals and speculative bubbles estimated for Germany and the United Kingdom. *Intraregional* cointegration linkages of speculative bubbles among the CEE countries also became stronger during the recent financial crisis, but there is hardly evidence that the crisis triggered stronger intraregional cointegration linkages of fundamentals. Taken together, our results imply that the collapse of the stock markets of the CEE countries during the recent financial crisis is likely to reflect both a correlated transatlantic deterioration of fundamentals and contagion effects due to international spillover of speculative bubbles that originated in the U.S. stock market.

The focus of our analysis is on the long-term international cointegration linkages of the stock markets of the CEE countries. Our focus on long-term linkages implies that our analysis of linkages of stock markets differs from the analyses that are characteristic of the literature on contagion. This literature typically focuses on excess short-term linkages of financial markets in times of financial crises. We also define contagion in terms of "excess linkages" of financial markets, but our analysis focuses on the potential effects of the recent financial crisis on the long-term equilibrium linkages of stock markets. Depending on the investment horizon of international investors, the existence and nature of long-term international cointegration linkages of stock markets can be a major determinant of the benefits of international portfolio diversification. We used the cointegration methodology developed by Johansen (1988, 1991) to analyze international cointegration linkages of the stock markets of the CEE countries. In the empirical literature, Kasa (1992), Francis and Leachman (1998), Masih and Masih (2001), and many others use the techniques of cointe-

[1] For a detailed account of the current global financial crisis, see Bartram and Bodnar (2009).

gration analysis to study international cointegration linkages of leading developed stock markets. Phylaktis (1999) and Manning (2002) use the cointegration methodology to study international cointegration linkages of Asian stock markets, while Choudhry (1997) and Chen et al. (2002) focus on Latin American stock markets.

Concerning the international cointegration linkages of the CEE countries, MacDonald (2001) finds evidence of cointegration of the stock markets of the CEE countries with the stock markets of Germany, the United Kingdom, and the United States. Syriopoulos (2006) reports that the international cointegration linkages of the stock markets of the CEE countries with developed stock markets are stronger than the intraregional linkages of the stock markets of the CEE countries. Syllignakis and Kouretas (2006) report that the stock markets of the CEE countries are partially integrated with the German and U.S. stock markets. Gilmore and McManus (2002), in contrast, find that the stock markets of the CEE countries are not cointegrated, either individually or as a group, with the U.S. stock market. Similarly, Yuca and Simga-Mugan (2000) do not find evidence of cointegration between CEE countries, and limited evidence of international cointegration linkages with developed stock markets. The conflicting evidence reported in the earlier literature might be the result of structural instability of the long-term international cointegration linkages caused by financial crises (Jochum et al., 1999; Voronkova, 2004). The recent financial crisis may have brought about just another significant change in the international cointegration linkages of the stock markets of the CEE countries.

We organize the remainder of this paper as follows. In Section 2, we describe the data that we used in our empirical analysis. In Section 3, we derive the state-space model that we used to decompose stock market indexes into fundamentals and speculative bubbles. We also report the estimation results for the state-space model. In Section 4, we report the results of our cointegration analysis. In Section 5, we offer some concluding remarks.

2. The Data

We used monthly data for the sample period from January 1995 to December 2008 in our empirical analysis. The choice of the sample period was mainly governed by the availability of data. Our data source is Thompson Financial Datastream. We retrieved from Datastream data on stock market indexes and Datastream-estimated dividend yields for the Czech Republic, Hungary, and Poland. In addition, we retrieved data for Germany, the United Kingdom, and the United States. We converted all national stock market indexes into U.S. dollars. We used the official conversion rate of the European Central Bank to compute a continuous exchange rate of the euro vis-à-vis the U.S. dollar for the entire sample period. In order to compute the real stock market indexes and real dividends, we used the U.S. consumer price index as a deflator. *Figure 1* shows the real stock market indexes. A cross-country comparison of the stock market indexes is possible because we scaled the indexes such that they assume the value 100 at the beginning of the sample period. The stock market indexes for the Western countries clearly show the run-up and eventual collapse of the dot-com bubble in 2000. Moreover, at the end of the sample period, the stock market indexes of the Western countries and the CEE countries collapsed during the recent financial crisis.

Figure 1 Stock-Market

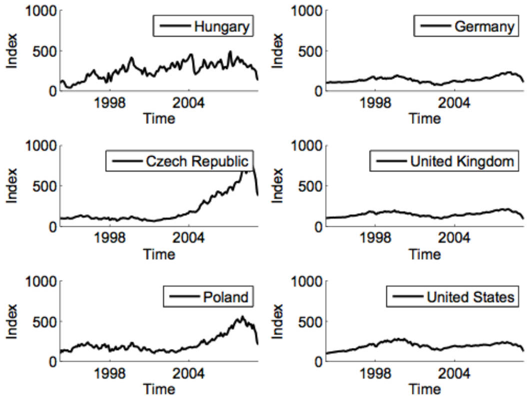


Figure 1 suggests that the stock markets of the CEE countries were hit particularly hard by the recent financial crisis. Specifically, during the last three months of 2008, the stock market indexes showed a clear tendency to decrease. The observed plunge in the CEE stock markets, compounded by a depreciation of local currencies, resulted in large negative U.S. dollar real returns. For example, the most severe collapse in the CEE stock markets, which occurred in November 2008, right after the collapse of Lehman Brothers, resulted in U.S. dollar real monthly returns (in monthly percentage) of -45.33%, -31.67%, and -39.67%, in Hungary, the Czech Republic, and Poland, respectively. Of the same order of magnitude, but less acute, was the collapse of the stock markets in Germany, the UK, and the U.S., that in November 2008 saw real returns of -21.25%, -21.67%, and -16.83%. It is, therefore, interesting to study whether the widespread collapse of stock markets reflects contagion effects or a simultaneous deterioration of fundamentals.

Table 1 contains the results of tests for a unit root in the real stock price indexes and the real dividends. We applied the unit root tests to the natural logarithms of the real stock price indexes and real dividends. The table presents the results of the DFGLS test for a unit root developed by Elliott et al. (1996). The null hypothesis of the test is that the time series being analyzed features a unit root. The DFGLS test does not require specification of the deterministic components of the unit-root regression equation insofar as the data are detrended before testing for a unit root. The results yield evidence that both the real stock price indexes and the real dividends have a unit root.

The evidence in favor of the unit-root property of the real stock price indexes and the real dividends gives rise to the question whether both series feature a common stochastic trend and are, thus, cointegrated. Absence of cointegration provides evidence for the presence of speculative bubbles in stock markets (Diba and Grossman, 1988). The intuition for this argument is that the expected discounted value of the stream of real dividends can be interpreted as a measure of the "fundamental" value of a stock. Rational speculative bubbles should drive a wedge between fundamentals and stock prices. As a result, speculative bubbles may result in noncointegration between fundamentals and stock prices. The column "Dividend-price ratio"

Table 1 Results of Unit Root Tests and of a Test for Speculative Bubbles

	Real stock price index	Real dividends	Dividend-price ratio
Czech Republic	0.1724	0.2088	-1.1201
Hungary	-0.6908	-0.3169	-2.1481
Poland	-1.2507	1.0300	0.1700
Germany	-0.8896	0.8641	1.2063
United Kingdom	-0.9502	0.0615	-0.9503
United States	-0.3687	1.1188	-0.6570

Notes: The table reports test statistics of the DFGLS test for a unit root developed by Elliott et al. (1996). The critical values of the test are -1.6152 and -1.9432 at the 10 percent and 5 percent level of significance. The column "Dividend-price ratio" summarizes the results of DFGLS tests for a unit root in the (natural logarithm) of the dividend-price ratio. If the null hypothesis of a unit root in the dividend--price ratio is rejected, there is evidence of cointegration between real stock market indexes and real dividends. Speculative bubbles cannot be ruled out if real stock price indexes and real dividends are not cointegrated.

in *Table 1* summarizes the results of tests for a unit root in the (natural logarithm) of the dividend-price ratio. The results indicate that it is in general not possible to reject the null hypothesis of noncointegration, implying that a more detailed analysis of the presence of speculative bubbles in the stock markets under investigation is warranted.

3. Fundamentals and Speculative Bubbles

We describe in Section 3.1 the state-space model that we used to estimate fundamentals and speculative bubbles. We summarize the estimation results in Section 3.2.

3.1 The State-Space Model

The framework for our analysis is the standard present-value model of stock price determination. We assumed that the current real stock price can be expressed as the present value of next period's expected real stock price and real dividends:

$$P_t = E_t\left(P_{t+1} + D_t\right)/\left(1+R\right) \tag{1}$$

where E_t denotes expectations conditional on information available up to and including time t, P_t denotes the real stock price at time t, D_t denotes the real dividends paid between time t and time $t+1$, and R denotes the constant required real rate of return. Using lowercase letters to denote the natural logarithm of a variable, the linear approximation of Equation (1) can be written as follows (Campbell et al. 1997, Chapter 7):

$$p_t = \kappa - r + \phi E_t\left(p_{t+1}\right) + \left(1-\phi\right)d_t \tag{2}$$

where

$\kappa = -\log\left(\phi\right) - \left(1-\phi\right)\log\left(1/\phi - 1\right)$, $\phi = 1\Big/\left(1+\exp\left(\overline{d-p}\right)\right)$, and $\overline{d-p}$ denotes the average log dividend-price ratio. Invoking the transversality condition, $\lim_{j\to\infty}\phi^j E_t p_{t+j} = 0$, the unique fundamental forward-looking no-bubble solution, p_t^f, can be written as

$$p_t^f = \left(\kappa - r\right)/\left(1-\phi\right) + \left(1-\phi\right)E_t\sum_{j=0}^{\infty}\phi^j d_{t+j} \tag{3}$$

We refer to p_t^f as "fundamentals". If the transversality condition does not hold, a rational speculative bubble may exist and the general solution for the stock price is given by

$$p_t = p_t^f + b_t \tag{4}$$

where the speculative bubble, b_t, satisfies the following difference equation:

$$E_t b_{t+j} = \left(1/\phi\right)^j b_t \tag{5}$$

where $j = 1, 2, ...$. In our empirical work, we parameterized Equation (5) as follows:

$$b_t = \left(1/\phi\right) b_{t-1} + \varepsilon_t \tag{6}$$

where $0 < \phi < 1$. The disturbance term, ε_t, is normally distributed with mean zero and variance σ_ε^2.

Because the real stock market indexes and real dividends are nonstationary (Section 2), we formulated Equation (4) in first differences. Upon defining the first--difference operator as Δ, the result is

$$\Delta p_t = \Delta p_t^f + \Delta b_t \tag{7}$$

where $\Delta p_t^f = \left(1-\phi\right)\sum_{j=0}^{\infty} \phi^j \left(E_t d_{t+j} - E_{t-1} d_{t+j-1}\right)$. Equation (7) shows that changes in fundamentals, Δp_t^f, reflect changes in expectations regarding the future stream of real dividends. We assumed that an ARIMA(n,1,0) model captures the dynamics of the demeaned real dividends. It follows

$$\Delta d_t = \sum_{j=1}^{n} \varphi_j \Delta d_{t-j} + u_t \tag{8}$$

where $n = 1, 2, ...$ and u_t is a normally distributed disturbance term with mean zero and variance σ_u^2. We assumed that the disturbance terms, ε_t and u_t, are mutually independent. Following Wu (1995), we expressed Equation (8) in its companion form as follows:

$$\boldsymbol{y}_t = \begin{pmatrix} \boldsymbol{\varphi} & \varphi_n \\ \boldsymbol{I}_{(n-1)\times(n-1)} & \boldsymbol{0}_{(n-1)\times 1} \end{pmatrix} \boldsymbol{y}_{t-1} + \boldsymbol{v}_t = \boldsymbol{A}\boldsymbol{y}_{t-1} + \boldsymbol{v}_t \tag{9}$$

where we defined $\boldsymbol{y}_t = \left(\Delta d_t, \Delta d_{t-1}, ..., \Delta d_{t-n+1}\right)'$, $\boldsymbol{v}_t = \left(u_t, \boldsymbol{0}_{1\times(n-1)}\right)'$, $\boldsymbol{\varphi} = \left(\varphi_1, \varphi_2, ..., \varphi_{n-1}\right)$, and $\boldsymbol{0}$ $\left(\boldsymbol{I}\right)$ denotes a matrix of zeros (an identity matrix) with dimension given in the index. Equation (9) implies that Equation (7) can be rearranged to obtain

$$\Delta p_t = \Delta d_t + \boldsymbol{m}\Delta \boldsymbol{y}_t + \Delta b_t \tag{10}$$

where $\boldsymbol{m} = \boldsymbol{g}\left(\boldsymbol{I}_{(n)\times(n)} - \boldsymbol{A}\right)^{-1} \boldsymbol{A}\left[\boldsymbol{I}_{(n)\times(n)} - \left(1-\phi\right)\left(\boldsymbol{I}_{(n)\times(n)} - \phi\boldsymbol{A}\right)^{-1}\right]$, and $\boldsymbol{g} = \left(1, \boldsymbol{0}_{1\times(n-1)}\right)$ denotes a selection vector.

Because the speculative bubble is not directly observable, we expressed Equation (10) in the form of a state-space model. A state-space model consists of a measurement (observation) equation and a transition (state) equation. Following Bhar and Hamori (2005), we combined Equations (6) and (8) in the state equation. The result is

$$\left(\boldsymbol{y}_t', \Delta d_{t-n}, b_t, b_{t-1}\right)' = \begin{pmatrix} \boldsymbol{A} & \boldsymbol{0}_{(n)\times 3} \\ \boldsymbol{a} & \boldsymbol{B} \end{pmatrix} \left(\boldsymbol{y}_{t-1}', \Delta d_{t-n-1}, b_{t-1}, b_{t-2}\right)' + \boldsymbol{u}_t \tag{11}$$

where $\boldsymbol{a} = \left(\boldsymbol{0}_{3\times(n-1)}, (1,0,0)'\right)$, and the other matrices are defined as

$$\boldsymbol{B} = \begin{pmatrix} 0 & 0 & 0 \\ 0 & 1/\phi & 0 \\ 0 & 1 & 0 \end{pmatrix}, \text{ and } \boldsymbol{u}_t = \begin{pmatrix} u_t \\ \boldsymbol{0}_{(n)\times 1} \\ (\varepsilon_t, 0)' \end{pmatrix}$$

The measurement equation is given by

$$\left(\Delta p_t, \Delta d_t\right)' = \begin{pmatrix} 1+m_1 & m_2 - m_1 & \dots & -m_n & 1 & -1 \\ 1 & 0 & \dots & 0 & 0 & 0 \end{pmatrix} \left(\boldsymbol{y}_t', \Delta d_{t-n}, b_t, b_{t-1}\right)' \tag{12}$$

where $m_j, j = 1,2,...,n$ denote the elements of the row vector $\boldsymbol{m}$.

3.2 Estimation Results

We used the Kalman-filter methodology as described by Kim and Nelson (1999) to estimate the structural parameters of the state-space model given in Equations (11) and (12) by the maximum-likelihood-estimation approach. The structural parameters are ϕ, φ_1, $\varphi_2,...,\varphi_n$, σ_ε^2, and σ_u^2. Equipped with estimates of the structural parameters, we computed time-series of the Kalman-filtered estimates of fundamentals and speculative bubbles, which approximate the information available to an investor at the time an investment decision must be reached. In order to retain symmetry across countries, and to economize on the number of parameters to be estimated, we used for all countries a parsimonious ARIMA(2,1,0) model to capture the dynamics of real dividends.

Table 2 summarizes the estimation results for the three CEE countries and for Germany, the United Kingdom, and the United States. The estimates of the variances of the disturbance terms are highly significant, and the parameters of the process governing the dynamics of real dividends are of a reasonable magnitude and are significant. The estimation results indicate that the parameter ϕ is always close to unity and precisely estimated. In conflict with the theoretical model (Section 3.1), however, the estimated parameter ϕ always exceeds unity. The parameter ϕ represents a real discount factor and should take on values between zero and one in the long run. Because the inverse of the discount factor is used in Equation (6), which thus assumes values greater than one, the speculative bubble follows an asymptotically explosive process. In order to make sure that the empirical estimates satisfy the restriction $0 < \phi < 1$, we estimated a restricted version of the model in which this restriction by construction always holds. We invoked the restriction by using the transformation $\phi = 1/\left(1+\exp(-\mu)\right)$, where μ denotes an auxiliary parameter to be estimated. The restricted parameter ϕ is always close to unity.[2]

Figure 2 Estimated Speculative Bubbles

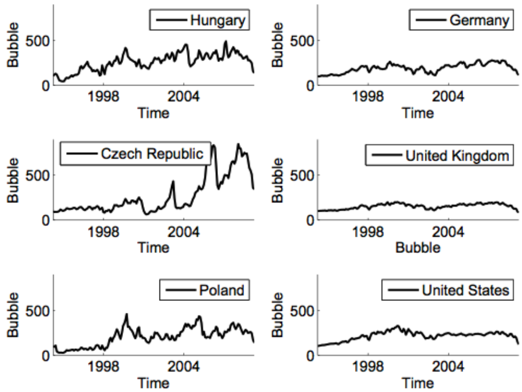


Table 2 Estimation Results for the State-Space Model

	φ_1	φ_2	Φ	σ^2_ε	σ^2_u
			Czech Republic		
Coefficient	-0.4256	-0.3335	1.0153 [0.9751]	252.6730	191.1269
Std. Error	0.0681	0.0684	0.0141	15.7369	11.8986
			Hungary		
Coefficient	-0.5846	-0.5569	1.0142 [0.9523]	221.7298	181.3189
Std. Error	0.0684	0.0696	0.0118	13.6610	11.1742
			Poland		
Coefficient	-0.4652	-0.3140	1.0201 [0.9960]	174.6303	181.1756
Std. Error	0.0724	0.0723	0.0149	10.8020	11.2025
			Germany		
Coefficient	-0.6578	-0.3697	1.0123 [0.9738]	109.7067	109.0153
Std. Error	0.0721	0.0713	0.0110	6.7046	6.6611
			United Kingdom		
Coefficient	-0.7334	-0.4729	1.0093 0.9656]	78.1637	75.3744
Std. Error	0.0690	0.0692	0.0095	4.6829	4.5158
			United States		
Coefficient	-0.6613	-0.4727	1.0044 [0.9852]	63.5884	82.4322
Std. Error	0.0717	0.0718	0.0067	3.8192	4.9508

Note: The numbers in brackets denote the estimation results for a restricted model with $0<\Phi<1$.

Figure 2 shows the estimated speculative bubbles, all scaled to assume the value 100 at the beginning of the sample period. Upon comparing *Figure 2* with *Figure 1*, it becomes apparent that the speculative bubbles account for a substantial

[2] We used the estimation results for the restricted version of the model in our empirical analysis. We used Matlab for estimation of the state-space model, and EViews and R for the cointegration and correlation analysis outlined in Section 4.

proportion of the stock market indexes, both in the CEE countries and in the Western countries. Moreover, the magnitudes of the speculative bubbles have undergone substantial changes over time. Reflecting the financial crisis of 2008, the speculative bubbles in the CEE countries significantly collapsed at the end of the sample period.

4. Empirical Results on Intermarket Linkages

In Section 4.1, we briefly describe how we used the approach advanced by Johansen (1988, 1991) to test for cointegration. In Section 4.2, we report the results of our cointegration analysis. In Section 4.3, we report the results of robustness checks.

4.1 Testing for Cointegration

The approach developed by Johansen (1988, 1991) has been widely used in empirical research to test for cointegration. A key advantage of Johansen's approach is that it can be used to test for cointegration in a multivariate setting. In order to illustrate the approach, it is useful to examine the following vector error correction model:

$$\Delta \boldsymbol{x}_t = \boldsymbol{k} + \sum_{j=1}^{w} \boldsymbol{L}_j \Delta \boldsymbol{x}_{t-j} + \boldsymbol{F}\boldsymbol{x}_{t-1} + \boldsymbol{\varepsilon}_{t,x} \tag{13}$$

where $\boldsymbol{x}_t$ denotes an $n \times 1$ vector of variables being analyzed, $\boldsymbol{\varepsilon}_{t,x}$ denotes a vector of white noise disturbance terms, $\boldsymbol{k}$ denotes a vector of deterministic terms, and the bold letters $\boldsymbol{L}_j$ and $\boldsymbol{F}$ denote matrices of coefficients to be estimated.

Johansen's approach amounts to determining the rank of the matrix $\boldsymbol{F}$. The rank of the matrix $\boldsymbol{F}$ is equal to the number of independent cointegration vectors. If the matrix $\boldsymbol{F}$ has full rank or zero rank, there are no cointegration vectors. If the matrix $\boldsymbol{F}$ has reduced rank, the number of cointegration vectors can be determined by testing for the significance of the eigenvalues of $\boldsymbol{F}$. To this end, the λ_{trace} and the $\lambda_{\max}$ statistics can be used. The λ_{trace} statistic tests the null hypothesis that there are at most r distinct cointegration vectors. The $\lambda_{\max}$ statistic tests the null hypothesis that the number of distinct cointegrating vectors is at most r against the alternative that the number of distinct cointegration vectors is $r + 1$.

We used a rolling sample window of five years to compute the λ_{trace} and the $\lambda_{\max}$ statistics. A rolling estimation window renders it possible to account for changes in the cointegration vectors over time. A rolling estimation window should help to detect such changes even if they take place gradually and even if there are multiple changes. A recent application of the rolling-estimation-window approach to cointegration analysis can be found in López Andión et al. (2010), who use a rolling estimation window to analyze changes in cointegration linkages of the mortgage markets in the European Monetary Union. Along similar lines, Brada et al. (2005) use a rolling cointegration technique to measure real and monetary convergence between Germany, France, and the CEE countries.

We used a model that features as a deterministic component a constant in the cointegration vectors. Concerning the number of lags in the vector error correction model, we varied the lag length between one, $w = 1$, and four, $w = 4$, and selected, in every estimation, the lag length based on a likelihood ratio test. We used the small-sample correction suggested by Reimers (1992) to compute the λ_{trace} and the $\lambda_{\max}$ statistics. In order to assess the statistical significance of our results, we formed the ratios of the λ_{trace} and the $\lambda_{\max}$ statistics and their respective 95 percent critical values (Rangvid, 2001; Pascual, 2003). If the ratios exceed one, the null hypothesis will be rejected.[3]

4.2 Results of the Cointegration Analysis

In order to set the stage for our analysis, we tested for evidence of cointegration using data for the full sample period. A potential difficulty with this approach, however, is that it assumes constancy of the cointegration linkages and, thus, cannot account for potential changes in the cointegrating vectors against the evidence of a number of economic and financial events that occurred in the CEE countries during the sample period.[4]

Tables 3 and *4* summarize the results of the cointegration analysis for the full sample period using the λ_{trace} and the $\lambda_{\max}$ statistics. We present results for stock market indexes, fundamentals, speculative bubbles, and the bubble-price ratios.[5] The overall evidence of cointegration in the full sample of data is rather moderate. The weak evidence of cointegration, however, perhaps is not surprising. Brada et al. (2005, p. 255) argue that tests for cointegration tend to rarely reject the hypothesis of no cointegration if the cointegrating vector changes over time. They also argue that tests for structural breaks tend to reject the hypothesis of a structural break in case cointegration linkages change gradually over time. A rolling-sample-window cointegration analysis can be used to inspect the potential variation over time in the cointegrating vectors.[6]

Figure 3 summarizes the results for a model that features as variables in the vector x_t the stock market indexes (fundamentals, speculative bubbles) of the stock markets of the three CEE countries. With regard to the stock market indexes, there is weak evidence of perhaps one cointegration vector in 2001/2002. There is also weak evidence of cointegration between speculative bubbles in 2002 and the first half of 2003. Cointegration between speculative bubbles became weaker in the second half of 2003, but regained strength in 2008. There is also evidence of cointegration be-

[3] An issue that may be important concerning the interpretation of the results of our cointegration analysis is that speculative bubbles obey an asymptotically explosive data-generating process. In order to deal with this issue, we shall present the results of a robustness check in Section 4.3.

[4] The list of significant economic and financial events includes the Asian crisis (1997), the Russian crisis (1998), accession to the EU (2004), and the global financial crisis (since 2007), to name just a few.

[5] In the context of our model, the bubble-price ratio is defined as the proportion of the speculative bubble in the overall stock price index, $100\times\left(P_t-P_t^F\right)\big/P_t$, where capital letters represent the variables in levels. The bubble-price ratio, therefore, does not share the asymptotically explosive property of speculative bubbles, but is by construction strictly first-order integrated. We present results for the bubble-price ratio as a robustness check.

[6] As for model diagnostics, we studied autocorrelations and cross-correlations of the residuals implied by the full-sample VAR and VEC models (the results are not reported but are available upon request). It turned out that four lags suffice to eliminate the predictable components from the residuals. It should be noted that we did not restrict the coefficients of the short-run dynamics of the model to be constant over time.

Table 3 Cointegration Results (VECM(4), λ_{trace})

Null hypothesis	Stock market index	Fundamentals	Speculative bubbles	Bubble-price ratio
		Germany – CEE Cointegration		
$r = 0$	0.0814	0.2631	0.7043	0.1944
$r \leq 1$	0.4163	0.8799	0.8075	0.4907
$r \leq 2$	0.5491	0.7780	0.6969	0.5539
$r \leq 3$	0.7013	0.5565	0.3759	0.2639
		UK – CEE Cointegration		
$r = 0$	0.1837	0.0366	0.8562	0.3242
$r \leq 1$	0.6883	0.3979	0.9297	0.5091
$r \leq 2$	0.6977	0.7009	0.8951	0.8847
$r \leq 3$	0.4894	0.8261	0.8410	0.9214
		US – CEE Cointegration		
$r = 0$	0.2237	0.0384	0.3337	0.0009
$r \leq 1$	0.7243	0.3635	0.4917	0.1253
$r \leq 2$	0.7179	0.5712	0.8207	0.6767
$r \leq 3$	0.6692	0.7174	0.6849	0.7800
		CEE Intraregional Cointegration		
$r = 0$	0.7325	0.0400	0.4957	0.1866
$r \leq 1$	0.7496	0.5935	0.5376	0.2662
$r \leq 2$	0.6521	0.6424	0.2875	0.2309

Notes: This table summarizes MacKinnon-Haug-Michelis *p*-values for Johansen's (1988) λ_{trace} test for cointegration. r = number of cointegration vectors. The null hypothesis in the case of $r = 0$ ($r \leq 1$, $r \leq 2$, $r \leq 3$) is that there is no cointegration vector (one cointegration vector, two cointegration vectors, three cointegration vectors). The alternative hypothesis stipulates $r > 0$ ($r > 1$, $r > 2$, $r > 3$) cointegration vectors. In the first stage, the null hypothesis $r = 0$ is tested. If the null hypothesis of noncointegration cannot be rejected, there is no evidence of one cointegration vector. If the null hypothesis can be rejected, in the second stage, the null hypothesis $r \leq 1$ is tested. If the null hypothesis cannot be rejected, there is evidence of one cointegration vector. If the null hypothesis can be rejected, in the third stage, the null hypothesis $r \leq 2$ is tested. If the null hypothesis cannot be rejected, there is evidence of two cointegration vectors. If the null hypothesis can be rejected, in the fourth stage, the null hypothesis $r \leq 3$ is tested. If the null hypothesis cannot be rejected, there is evidence of three cointegration vectors. If the null hypothesis can be rejected, there is evidence of four cointegration vectors, implying that the four series being analyzed are stationary. The sample period runs from 04/1995 to 12/2008. The table summarizes the results for a vector error correction model (VECM) with four lags. The results are based on estimates of speculative bubbles derived from a state-space model in the presence of the restriction $0<\Phi<1$.

tween fundamentals in 2001. Thereafter, cointegration between fundamentals became weaker. Cointegration between fundamentals, however, regained strength in 2003 and in early 2004. Thereafter, cointegration between fundamentals significantly weakened and, at the end of the sample period, was thus overturned by cointegration between speculative bubbles.

Figure 4 summarizes the results for a model that features the stock market indexes (fundamentals, speculative bubbles) of the three CEE countries and of one of the Western countries. As regards the stock market indexes, there is evidence of temporary cointegration in 2001 and in 2002/2003. The results further suggest the temporary presence of one cointegration vector in 2005. The results, however, do not suggest that the cointegration linkages of the stock market indexes of the CEE countries and the stock market indexes of the Western countries became tighter over time.

Table 4 Cointegration Results (VECM(4), λ_{max})

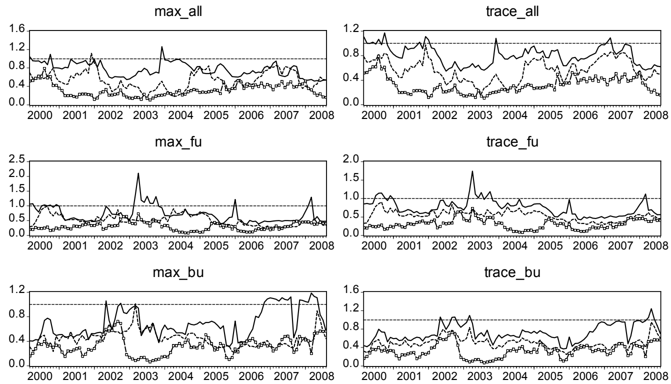
Null hypothesis	Stock market index	Fundamentals	Speculative bubbles	Bubble-price ratio
		Germany – CEE Cointegration		
$r = 0$	0.0812	0.0686	0.6809	0.2159
$r = 1$	0.4795	0.9256	0.8974	0.5869
$r = 2$	0.4839	0.8501	0.8692	0.8212
$r = 3$	0.7013	0.5565	0.3759	0.2639
		UK – CEE Cointegration		
$r = 0$	0.0929	0.0279	0.7440	0.4252
$r = 1$	0.7320	0.3227	0.9171	0.2914
$r = 2$	0.7997	0.5935	0.8554	0.7902
$r = 3$	0.4894	0.8261	0.8410	0.9214
		US – CEE Cointegration		
$r = 0$	0.1120	0.0357	0.4637	0.0016
$r = 1$	0.7641	0.3816	0.3345	0.0619
$r = 2$	0.7116	0.5009	0.8330	0.5918
$r = 3$	0.6692	0.7174	0.6849	0.7800
		CEE Intraregional Cointegration		
$r = 0$	0.7454	0.0166	0.6110	0.3758
$r = 1$	0.7620	0.5725	0.7785	0.4529
$r = 2$	0.6521	0.6424	0.2875	0.2309

Notes: This table summarizes MacKinnon-Haug-Michelis *p*-values for Johansen's (1988) λ_{max} test for cointegration. r = number of cointegration vectors. The null hypothesis in the case of $r = 0$ ($r = 1$, $r = 2$, $r = 3$) is that there is no cointegration vector (one cointegration vector, two cointegration vectors, three cointegration vectors). The alternative hypothesis stipulates $r = 1$ ($r = 2$, $r = 3$, $r = 4$) cointegration vectors. In the first stage, the null hypothesis $r = 0$ is tested. If the null hypothesis of noncointegration cannot be rejected, there is no evidence of one cointegration vector. If the null hypothesis can be rejected, in the second stage, the null hypothesis $r = 1$ is tested. If the null hypothesis cannot be rejected, there is evidence of one cointegration vector. If the null hypothesis can be rejected, in the third stage, the null hypothesis $r = 2$ is tested. If the null hypothesis cannot be rejected, there is evidence of two cointegration vectors. If the null hypothesis can be rejected, in the fourth stage, the null hypothesis $r = 3$ is tested. If the null hypothesis cannot be rejected, there is evidence of three cointegration vectors. If the null hypothesis is rejected, there is evidence of four cointegration vectors, implying that the four series being analyzed are stationary. The sample period runs from 04/1995 to 12/2008. The table summarizes the results for a vector error correction model (VECM) with four lags. The results are based on estimates of speculative bubbles derived from a state-space model in the presence of the restriction $0<\Phi<1$.

There is no evidence that the recent financial crisis had a significant effect on the international cointegration linkages of stock market indexes.

With regard to fundamentals, our results suggest that the cointegration linkages of the fundamentals of the stock markets of the CEE countries and the stock markets of Germany, the United Kingdom, and the United States were significant in late 2002 and in 2003. Efforts to meet the Maastricht criteria for EU accession were perhaps an important determinant of the strengthening in cointegration linkages of the fundamentals of the stock markets of the CEE countries and those of the stock markets of Germany and the United Kingdom. The significant cointegration among fundamentals lost in significance in 2004. At the end of the sample period, however, the cointegration linkages of the fundamentals again became stronger, especially when the vector error correction model features the fundamentals of the U.S. stock market.

Figure 3 Cointegration among the Stock Markets of the CEE Countries



Notes: The tests for cointegration are based on models that feature the stock markets of the three CEE countries. The solid (dotted, squared) lines represent the results of tests for cointegration based on the λ_{trace} (0) ($\lambda_{trace}(1)$, $\lambda_{trace}(2)$) and the λ_{max} (0) ($\lambda_{max}(1)$, $\lambda_{max}(2)$) statistics. The dashed vertical line represents the 95 percent critical value. The statistics are scaled by their critical values. The critical values are taken from Osterwald-Lenum (1992). *all* = test for cointegration between stock market indexes, *fu* = test for cointegration between fundamentals, *bu* = test for cointegration between speculative bubbles.

Transatlantic cointegration linkages, thus, strengthened in terms of fundamentals during the recent financial crisis.

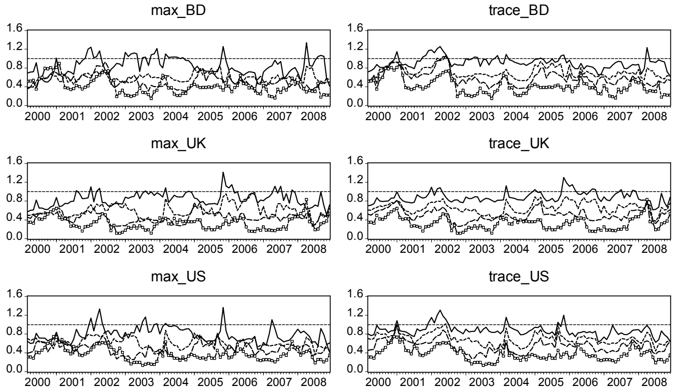
Our results suggest that, with regard to speculative bubbles, cointegration linkages were relatively strong in 2002–2003. Perhaps investors' expectations concerning the EU accession of the CEE countries brought about a greater degree of economic uncertainty and distorted investors' perceptions regarding future fundamentals of the CEE countries. Cointegration between the speculative bubbles weakened in the second half of the sample period, but became strongly significant again in 2007. We found particularly strong evidence of contagion effects reflecting international cointegration linkages of speculative bubbles when we used the United States to represent the Western country in the vector-error-correction model. We found weaker contagion effects when we analyzed the vector-error-correction models that contain the United Kingdom and Germany as the Western country.

4.3 Robustness Checks

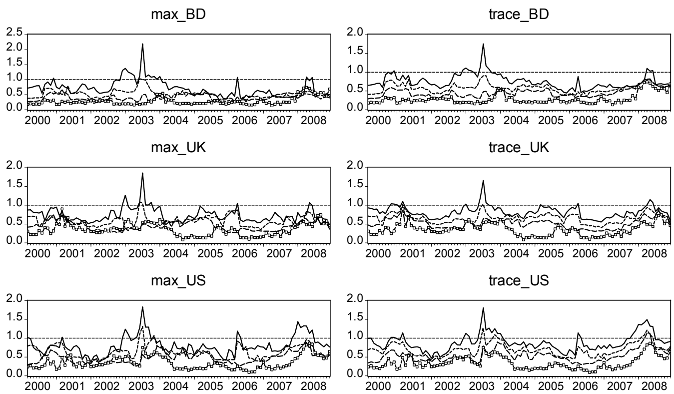
In order to analyze the robustness of our results, we replaced the speculative bubbles in our vector-error correction model with the bubble-price ratios estimated for the CEE and the Western countries. While the speculative bubbles have a root outside the unit circle (that is, $1/\phi > 1$), the bubble-price ratio is by construction difference stationary. The bubble-price ratio is difference stationary because the stock price, in our model of speculative bubbles, is equal to the sum of fundamentals and speculative bubbles. As a result, the bubble-price ratio is the inverse of fundamentals, which are difference stationary according to Equation (8). The bubble-price ratio increases when (i) the speculative bubbles increase at a higher rate than the stock price index, or (ii) the speculative bubbles decrease at a lower rate than the stock price index. In both cases, fundamentals decrease.

Figure 4 Cointegration between CEE Countries and Western Countries

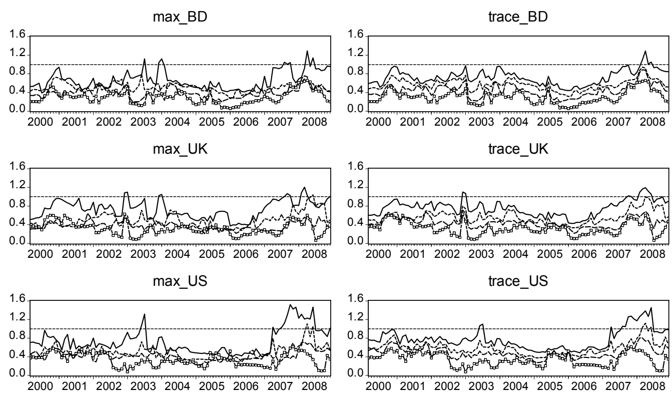
Panel (a): Cointegration between Stock Market Indexes



Panel (b): Cointegration between Fundamentals



Panel (c): Cointegration between Speculative Bubbles



Notes: The tests for cointegration are based on models that feature the stock markets of the three CEE countries plus one of the stock markets of the Western countries. The solid (dotted, dashed, squared) lines represent the results of tests for cointegration based on the λ_{trace} (0) ($\lambda_{trace}(1)$, $\lambda_{trace}(2)$, $\lambda_{trace}(3)$) and the λ_{max} (0) ($\lambda_{max}(1)$, $\lambda_{max}(2)$, $\lambda_{max}(3)$) statistics. The dashed vertical line represents the 95 percent critical value. The statistics are scaled by their critical values. The critical values are taken from Osterwald-Lenum (1992). *BD* = Germany, *UK* = United Kingdom, *US* = United States.

For the sake of brevity, we only summarize the results of the robustness check (the results are available upon request). With regard to the intraregional cointegration linkages, we found evidence of temporary cointegration of the bubble-price ratios in 2000 and in 2002/2003. At the end of the sample period, from the beginning of 2007 onwards, the intraregional cointegration linkages of the bubble-price ratios strengthened again. The stronger evidence of cointegration of speculative bubbles in the last two years of our sample period may thus reflect, to some extent, stronger incidence of intraregional contagion. With regard to the transatlantic cointegration linkages, we found evidence of contagion in the last two years of our sample (the Western country is represented by the United States). In contrast, evidence of continental cointegration linkages turned out to be somewhat weaker (the Western country is represented by the United Kingdom or Germany). In sum, the results of our robustness check confirmed our main result that cointegration linkages (in terms of fundamentals and speculative bubbles) strengthened following the outbreak of the recent financial crisis.

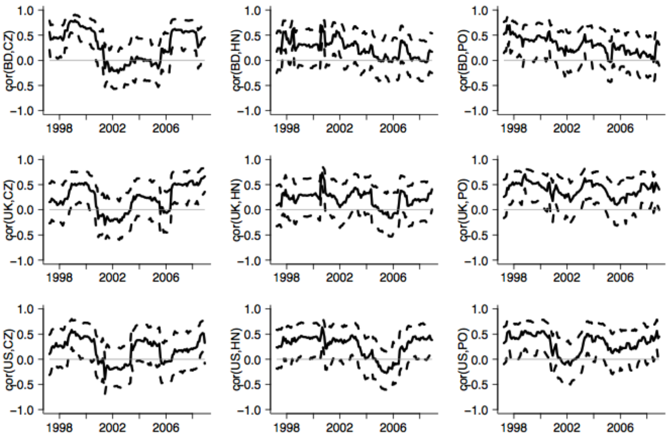
As another robustness check, we used stock market indexes (returns, dividends) denominated in local currency to estimate fundamentals and speculative bubbles in the CEE countries (the results are not reported, but are available from the authors upon request). To this end, we used national consumer-price indexes to deflate the data. Using stock market indexes in local currency renders it possible to integrate out valuation effects stemming from movements in exchange rates. As in Section 4.2, we found that the financial crisis led to a strengthening of cointegration links between fundamentals and speculative bubbles. As concerns intraregional dynamics, cointegration linkages of speculative bubbles strengthened during the financial crisis, indicating the presence of intraregional contagion effects. We also found evidence of stronger intraregional cointegration linkages of fundamentals, which were, however, rather short-lived. Moreover, during the financial crisis, cointegration linkages of fundamentals strengthened in the model featuring U.S. fundamentals. Cointegration linkages of speculative bubbles strengthened in the models featuring Germany and the U.S. as Western countries. Furthermore, the scaled cointegration statistics, while being statistically significant, turned out to be somewhat smaller than the cointegration statistics summarized in Section 4.2. The relative magnitude of the cointegration statistics can be interpreted to reflect the presence of a valuation effect triggered by the widespread depreciation of the exchange rates of the CEE countries against the U.S. dollar by the end of the year 2008.

Cointegration is a long-term concept that requires a long span of data (Hakkio and Rush, 1991). Because we are obliged to use a relatively short sample of data, we supplemented the cointegration analysis with a correlation analysis. To this end, we used the first difference of fundamentals and speculative bubbles measured in U.S. dollars to calculate rolling-window intraregional and interregional correlations.[7] As

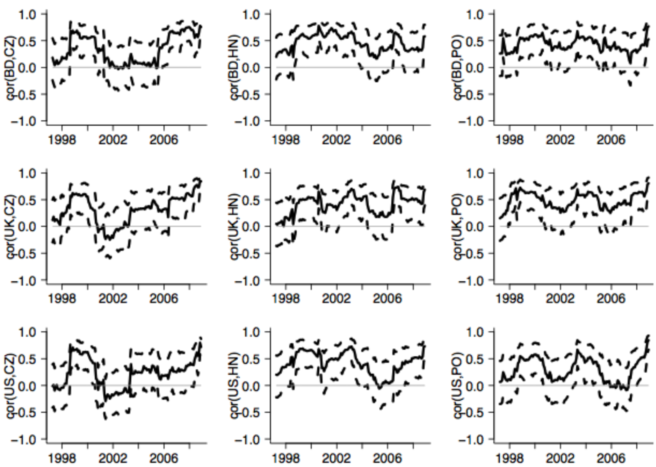
[7] Rolling-window correlations are a standard methodology to account for instability of linkages of returns in international equity markets (Longin and Solnik, 1995).

Figure 5 Correlation between CEE countries and Western Countries

Panel (a): Interregional Correlation between Fundamentals



Panel (b): Interregional Correlation between Speculative Bubbles



for the length of the rolling window, we opted for a short 24-months rolling window because the correlation analysis targets time-variation in short-run intermarket linkages. In addition, a 24-months rolling window is better suited than a longer rolling window to capture the effect of the financial crisis of 2007/2008 on intermarket linkages (the rolling window containing data for 2007–2008 covers mainly the crisis period). The results, summarized in *Figure 5*, indicate that the correlations in general increased during the last two years of the sample period, where this tendency is more visible for speculative bubbles than for fundamentals. From the three CEE countries, the Czech Republic showed the strongest correlation with the Western economies. As for the intraregional linkages (the results are not reported but are available upon request), we found that evidence of significant correlations between speculative bubbles is stronger than evidence of significant correlations between fundamentals. The results of the correlation analysis, thus, broadly corroborate the results of the cointegration analysis and help to build confidence in the results of the cointegration analysis.[8]

5. Concluding Remarks

Our results suggest that the intraregional linkages of the stock markets of the three CEE countries the Czech Republic, Hungary, and Poland have considerably changed over time. In addition, the international linkages of these three markets with the stock markets of Germany, the United Kingdom, and the United States have significantly changed over time. The cointegration linkages with the U.S. stock market strengthened in terms of both fundamentals and speculative bubbles during the market jitters caused by the financial crisis of 2008.

Our finding that the importance of fundamentals for international cointegration linkages of the stock markets of the CEE countries has increased during the recent financial crisis is in line with empirical findings recently reported by Čihák and Mitra (2009). They document that the explanatory power of regional fundamentals with respect to spreads on sovereign bonds increased during the recent financial market jitters, and that the cross-country dispersion of sovereign spreads can be explained in terms of macroeconomic performance. While the focus of our analysis differs from that of Čihák and Mitra (2009), it thus seems that macroeconomic fundamentals played a role in the transmission of the recent financial crisis to the CEE countries.

In order to develop a better understanding of the economic forces that are responsible for how financial crises become contagious and spread from one country to another, it would be interesting to study in future research the economic determinants of the time-varying intraregional and international linkages of the stock markets of the CEE countries. To this end, one could analyze to what extent time--varying intraregional and international linkages depend on the macroeconomic variables analyzed by Čihák and Mitra (2009). Empirical results recently reported by Frank and Hesse (2009) should also be useful in this respect. They find that developments of financial markets in emerging market countries during the recent financial crisis were linked to stress in the interbank market, market volatility, and default risk of financial institutions in advanced economies. Árvai et al. (2009) emphasize the importance of financial linkages through "common-lender" effects in European emerging market countries. One could also draw, for example, on the results reported by Bracker et al. (1999), who document that bilateral import dependence, size differentials across markets, and real interest rate differentials influence international linkages of stock markets. Forbes and Chinn (2004) find that bilateral trade flows are an important determinant of international linkages of stock markets. Pretorius (2002) analyzes the economic determinants of emerging stock market linkages.

Our results render it possible to study whether the impact of bilateral trade flows, bilateral capital flows, real interest rate differentials, and other economic factors on international linkages of fundamentals differs from their impact on inter-

[8] One should note, however, that potential crisis-induced shifts in the variances of the time series may bias estimates of correlations (Forbes and Rigobon, 2002). Because there is no consensus in the literature on how to control for such biases and, thus, how to estimate short-run intermarket linkages reflecting contagion, our analysis has focused on long-term cointegration linkages.

national linkages of speculative bubbles. The results of such research would be interesting for economists, but also for investors trying to figure out which economic factors determine the exposure of their portfolios to changes in the comovement of international stock markets.

While we have focused on the recent financial crisis, our results may also help to develop more efficient early-warning indicators that forecast contagious financial crises. In the earlier literature, many authors have reported that it is notoriously difficult to forecast financial crises. Policymakers who seek to prevent run ups and eventual collapses of speculative bubbles, however, need efficient and reliable early-warning indicators. It would be interesting to explore whether the performance of early-warning indicators improves once one takes into account changes in the international comovement of fundamentals and speculative bubbles as predictors of financial crises and market jitters.

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