The U.S. Subprime Mortgage Crisis and the Stock Markets of the CEE Countries

September 2009

Abstract

Stock markets in Central and Eastern European (CEE) countries significantly collapsed during the recent U.S. subprime mortage crisis. 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. In order to answer this question, 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 analyze the international cointegration linkages of fundamentals and speculative bubbles. Our results suggest that international cointegration linkages varied over time, and that cointegration linkages with the U.S. stock market strengthened in terms of of both fundamentals and speculative bubbles during the market jitters caused by the U.S. subprime mortgage crisis.

JEL classification: C32, F37, G15
Keywords: Stock markets, Fundamentals, Speculative bubbles, Cointegration Analysis, CEE countries, Kalman filter

Addresses:

Christian Pierdzioch, Saarland University, Department of Economics, P.O.B. 15 11 50, 66041 Saarbruecken, Germany, c.pierdzioch@mx.uni-saarland.de.

Renatas Kizys, Instituto Tecnologico y de Estudios Superiores, Departamento de Economia, C.P. 64849, Monterrey, Nuevo Leon, Mexico, renatas.kizys@itesm.mx.

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 crisis. If financial crisis 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 crisis, 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, Bekaert et al. 2005), the empirical framework we lay out in this paper makes it a natural choice to define contagion in terms of a strengthening of international linkages of speculative bubbles in times of financial crisis.

The significant collapse of the stock markets in Central and Eastern European (CEE) countries during the recent U.S. subprime mortage crisis provides a natural experiment to study how international linkages of stock markets, fundamentals, and speculative bubbles change in times of crisis (for a detailed account of the current global financial crisis, see Bartram and Bodnar 2009). 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 U.S. subprime mortgage crisis. *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 U.S. subprime mortgage 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 stock markets of the CEE countries during the U.S. subprime mortgage crisis is likely to reflect both a correlated transatlantic deterioration of fundamentals and contagion effects due to international spillover effects 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. 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 the international cointegration linkages of the stock markets of the CEE countries. In the recent empirical literature, Kasa (1992), Francis and Leachman (1998), Masih and Masih (2001), and many others use the techniques of cointegration analysis to study the international cointegration linkages of leading developed stock markets. Phylaktis (1999) and Manning (2002) use the cointegration methodology to study the 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 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 among 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) does 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 U.S. subprime mortgage 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 to 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 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. 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 for the Western countries and the CEE countries collapsed during the U.S. sub-prime loan crisis.

– Insert Figure 1 about here.–

Figure 1 suggests that the stock markets in the CEE countries were hit particularly hard by the ensuing U.S. subprime mortgage 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 that occurred in November 2008, resulted in U.S. Dollar real monthly returns (in annualized percentage) of -544%, -380%, and -476%, in Hungary, Czech Republic, and Poland, respectively. Of the same order of magnitude, but less acute, was the collapse of the stock markets in Germany, UK, and U.S., that in November 2008 saw the real returns of -255%, -260%, and 202%. 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 has the property that there is no need to specify which deterministic components one wants to include in the unit-root regression equation because 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.

– Insert Table 1 about here.–

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 results summarized in Table 1 indicate that the log real stock market indexes and the log real dividends are not cointegrated, implying that the presence of speculative bubbles cannot be ruled out.

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:

$$P_t = \mathbb{E}_t\left(P_{t+1} + D_t\right) / \left(1 + R\right), \tag{1}$$

where $\mathbb{E}_t$ 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):

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

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

$$p_t^f = (\kappa - r)/(1-\phi) + (1-\phi)\mathbb{E}_t\sum_{j=0}^{\infty}\phi^j d_{t+j}. \tag{3}$$

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:

$$\mathbb{E}_t b_{t+j} = (1/\phi)^j \, b_t, \tag{5}$$

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

$$b_t = (1/\phi) \, b_{t-1} + \epsilon_t, \tag{6}$$

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 $\Delta p_t^f = (1 - \phi) \sum_{j=0}^{\infty} \phi^j \left(\mathbb{E}_t d_{t+j} - \mathbb{E}_{t-1} d_{t+j-1} \right)$. Equation (7) shows that changes in fundamentals, Δp^f, reflect changes in expectations regarding the 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:

$$\mathbf{Y}_t = \begin{pmatrix} \varphi & \varphi_n \\ \mathbf{1}_{(n-1)\times(n-1)} & \mathbf{0}_{(n-1)\times 1} \end{pmatrix} \mathbf{Y}_{t-1} + \nu_t = \mathbf{A}\mathbf{Y}_{t-1} + \nu_t, \tag{9}$$

where we defined $\mathbf{Y}_t = (\Delta d_t, \Delta d_{t-1}, ..., \Delta d_{t-n+1})'$, $\nu_t = \left(u_t, \mathbf{0}_{1\times(n-1)}\right)'$, $\varphi = (\varphi_1, \varphi_2, ..., \varphi_{n-1})$, and $\mathbf{0}$ ($\mathbf{1}$) 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

$$\Delta p_t = \Delta d_t + \mathbf{M}\Delta\mathbf{Y}_t + \Delta b_t, \tag{10}$$

where $\mathbf{M} = \mathbf{g}\left(\mathbf{1}_{n\times n} - \mathbf{A}\right)^{-1}\mathbf{A}\left[\mathbf{1}_{n\times n} - (1-\phi)\left(\mathbf{1}_{n\times n} - \phi\mathbf{A}\right)^{-1}\right]$, and $\mathbf{g} = \left(1, \mathbf{0}_{1\times(n-1)}\right)$ 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

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

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

$$\mathbf{B} = \begin{pmatrix} 0 & 0 & 0 \\ 0 & 1/\phi & 0 \\ 0 & 1 & 0 \end{pmatrix}, \quad \text{and} \quad \mathbf{U}_t = \begin{pmatrix} u_t \\ \mathbf{0}_{n\times 1} \\ (\epsilon_t, 0)' \end{pmatrix}.$$

The measurement equation is given by

$$(\Delta p_t, \Delta d_t)' = \begin{pmatrix} 1+m_1 & m_2 - m_1 & \cdots & -m_n & 1 & -1 \\ 1 & 0 & \cdots & 0 & 0 & 0 \end{pmatrix} \Big(\mathbf{Y}_t', \Delta d_{t-n}, b_t, b_{t-1}\Big)', \tag{12}$$

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

3.2 Estimation Results

We used the Kalman-filter methodology as described by Kim and Nelson (2000) to estimate the structural parameters of the state-space model given in Equations (11) and (12) by maximum likelihood. The structural parameters are $\phi, \varphi_1, \varphi_2, ..., \varphi_n, \sigma_\epsilon^2$, and σ_u^2. Once equipped with estimates of the structural parameters, we computed time-series of the Kalman-filtered estimates of fundamentals and the 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.

– Insert Table 2 about here.–

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 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. 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 $\phi = 1/(1 + \exp(-\mu))$, where μ denotes an auxiliary parameter to be estimated. The restricted parameter ϕ is always close to unity. We used the estimation results for the restricted version of the model in our empirical analysis.

–Insert Figure 2 about here. –

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 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 U.S. subprime mortgage crisis, the speculative bubbles in the CEE countries significantly collapsed at the end of the sample period.

4 Cointegration Analysis

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 a robustness check.

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:

$$\Delta x_t = \mathbf{K} + \sum_{j=1}^{w} \mathbf{L}_j \Delta x_{t-j} + \mathbf{F} x_{t-1} + \epsilon_{t,x}, \tag{13}$$

where x_t denotes a $n \times 1$ vector of variables being analyzed, $\epsilon_{t,x}$ denotes a vector of white noise disturbance terms, $\mathbf{K}$, denotes a vector of deterministic terms, and the bold letters $\mathbf{L}_j$ and $\mathbf{F}$ denote matrices of coefficients to be estimated.

Johansen's approach amounts to determining the rank of the matrix $\mathbf{F}$. The rank of the matrix $\mathbf{F}$ is equal to the number of independent cointegration vectors. If the matrix $\mathbf{F}$ has full rank or zero rank, there are no cointegration vectors. If the matrix $\mathbf{F}$ has reduced rank, the number of cointegration vectors can be determined by testing for the significance of the eigenvalues of $\mathbf{F}$. To this end, the λ_{trace} statistic and the λ_{max} statistics can be used.

The λ_{trace} statistic tests the null hypothesis that there are at most r distinct cointegration vectors. The λ_{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 length five years to compute the λ_{trace} and the λ_{max} statistics. A rolling estimation window renders it possible to account for changes in the cointegration vectors over time. 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 λ_{max} statistics. In order to assess the statistical significance of our results, we formed the ratios of the λ_{trace} and the λ_{max} statistics and their respective 95 percent critical values (Rangvid 2001, Pascual 2003). If the ratios exceed one, the null hypothesis will be rejected.

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.

4.2 Results of the Cointegration Analysis

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. The evidence of cointegration between speculative bubbles becomes even weaker in the second half of 2003, but regains strength again in 2007/2008. There is also evidence of cointegration between fundamentals in 2001. Thereafter, cointegration between fundamentals becomes weaker. Cointegration between fundamentals, however, regains strength in 2003 and in early 2004. Thereafter, cointegration between fundamentals significantly weakens and, at the end of the sample period, is thus overturned by cointegration between speculative bubbles.

– Insert Figure 3 about here. –

Figure 4 summarizes the results for a model that features the stock market indexes (fundamentals, speculative bubbles) of the stock markets of the three CEE countries and of one of the Western stock markets. 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 links between the stock market indexes of the CEE countries and the stock market indexes of the Western countries have become tighter over time. There is no evidence that the recent U.S. subprime mortgage crisis has had a significant effect on the international cointegration linkages of stock market indexes.

– Insert Figure 4 about here. –

With regard to fundamentals, our results suggest that the cointegration linkages between 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. The significant cointegration among fundamentals lost in significance in 2004. At the end of the sample period, however, the cointegration linkages between the fundamentals again have become stronger, especially when the vector error correction model features the fundamentals of the U.S. stock market. Transatlantic cointegration linkages thus strengthened in terms of fundamentals during the recent U.S. subprime mortgage 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. The evidence of cointegration between the speculative bubbles becomes weaker in the second half of the sample period, but becomes strongly significant again in 2007. Evidence of contagion effects reflecting international cointegration linkages of speculative bubbles is particulary strong if the Western country is represented by the United States in the vector-error-correction model. The contagion effects are weaker when one considers the vector-error-correction models that contain the United Kingdom and Germany.

4.3 Robustness Check

In order to analyze the robustness of our results, we replaced the speculative bubbles in our vector-error correction model with the (log) 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-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 then the stock price index, or (ii) the speculative bubbles decrease at a lower rate than the stock price index. In both cases, fundamentals decrease.

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, there is 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 have grown stronger again. One can argue that the stronger evidence of cointegration of speculative bubbles in the last two years of our sample period may be an indication of the increased incidence of intraregional contagion. With regard to the transatlantic cointegration linkages, there is evidence of contagion in the last two years of our sample (the Western country is represented by the United States). Evidence of continental cointegration linkages is somewhat weaker (the Western country is represented by the United Kingdom or Germany). In sum, the results of our robustness check confirm our main result that cointegration linkages have strengthened since the outbreak of the recent U.S. subprime mortgage crisis.

5 Concluding Remarks

Our results suggest that the intraregional linkages among the stock markets of the three CEE countries 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 of both fundamentals and speculative bubbles during the market jitters caused by the U.S. subprime mortgage crisis.

In order to develop a better understanding of the economic forces that are responsible for how financial crisis 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. One could draw, for example, on results reported by Bracker et al. (1999), who document that international linkages of stock markets are influenced by bilateral import dependence, size differentials across markets, and real interest rate differentials. Forbes and Chinn (2004) find that bilateral trade flows are an important determinant of international linkages of stock markets. Pretorius (2002) analyze 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 international linkages of speculative bubbles. The results of such research would be interesting for economists, but also for investors who try 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 U.S. subprime mortgage crisis, our results may be helpful to develop more efficient early-warning indicators that forecast contagious financial crisis. In the earlier literature, many authors have reported that it is notoriously difficult to forecast financial crisis. 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 co-movement of fundamentals and speculative bubbles as financial crises gather steam.

References

Bekaert, G., C.R. Harvey, and A. Ng (2005). Market Integration and Contagion. Journal of Business 79: 39−69.

Bhar, R., and S. Hamori (2005). Empirical Techniques in Finance. Berlin: Springer Verlag.

Bracker, K., D.S. Docking, and P.D. Koch (1999). Economic Determinants of Evolution of International Stock Market Integration. Journal of Empirical Finance 6: 1−27.

Campbell, J.Y., A.W. Lo, and A.C. MacKinlay (1997). The Econometrics of Financial Markets. Princeton, New Jersey: Princeton University Press.

Chen, G.M., M. Firth, and O.M. Rui (2002). Stock Market Linkages: Evidence from Latin America. Journal of Banking and Finance 26: 1113−1141.

Choudhry, T. (1997). Stochastic Trends in Stock Prices: Evidence from Latin American Markets. Journal of Macroeconomics 19: 285−304.

Conroy R.M., K.M. Eades, and R.S. Harris (2000). Effects of Dividends and Earnings: Evidence from Simultaneous Announcements in Japan. Journal of Finance 55: 1199−1227.

Diba, B.T., and H. Grossman (1988). Explosive Rational Bubbles in Stock Prices? American Economic Review 78: 520−530.

Elliott, G., T.J. Rothenberg, and J.H. Stock (1996). Efficient Tests for an Autoregressive Unit Root. Econometrica 64: 813−836.

Engle, R.F., and C.W.J. Granger (1987). Cointegration and Error Correction: Representation, Estimation, and Testing. Econometrica 55: 251−276.

Forbes, K.J., and M.D. Chinn (2004). A Decomposition of Global Linkages in Financial Markets Over Time. Review of Economics and Statistics 86: 705−722.

Forbes, K.J., and R. Rigobon (2001). Contagion in Latin America: Definitions, Measurement and Policy Implications. Economia 1: 1−46.

Francis, B.B., and L.L. Leachman (1998). Superexogeneity and the Dynamic Linkages among International Stock Markets. Journal of International Money and Finance 17: 475−492.

Gilmore, C.G., and G.M. McManus (2002). International Portfolio Diversification: US and Central European Equity Markets. Emerging Markets Review 3: 69−83.

Grullon, G., and R. Michaely (2002). Dividends, Share Repurchases, and the Substitution Hypothesis. Journal of Finance 57: 1649−1684.

Jochum, C., G. Kirchgässner, and M. Platek (1999). A Long-Run Relationship between Eastern European Markets? Cointegration and the 1997/1998 Crisis in Emerging Markets. Weltwirtschaftliches Archiv / Review of World Economics 135: 455−479.

Johansen, S. (1988). Statistical Analysis of Cointegration Vectors. Journal of Economic Dynamics and Control 12: 231−254.

Johansen, S. (1991). Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models. Econometrica 59: 1551−1580.

Kasa, K. (1992). Common Stochastic Trends in International Stock Markets. Journal of Monetary Economics 29: 95−124.

Kim, C.J., and C.R. Nelson (2000). State-Space Models with Regime Switching. Cambridge, Massachusetts: MIT Press.

Manning, N. (2002). Common Trends and Convergence? South East Asian Equity Markets, 1988–1999. Journal of International Money and Finance 21: 183−202.

Masih, R., and A.M.M. Masih (2001). Long and Short-Term Dynamic Causal Transmission Amongst International Stock Markets. Journal of International Money and Finance 20: 563−587.

MacDonald, R. (2001). Transformation of External Shocks and Capital Market Integration. In M. Schröder (Ed.): The New Capital Markets in Central and Eastern Europe. Springer Verlag: Berlin

Osterwald-Lenum, M. (1992). A Note with Quintiles of the Asymptotic Distribution of the ML Cointegration Rank Test Statistics. Oxford Bulletin of Economics and Statistics 54: 461−472.

Pascual, A.G. (2003). Assessing European Stock Market (Co)Integration. Economics Letters 78: 197−203.

Phylaktis, K. (1999). Capital Market Integration in the Pacific Basin Region: An Impulse Response Analysis. Journal of International Money and Finance 18: 267−287.

Pretorius, E. (2002). Economic Determinants of Emerging Stock Market Interdependence. Emerging Markets Review 3: 84−105.

Rangvid, J. (2001). Increasing Convergence among European Stock Markets? A Recursive Common Stochastic Trends Analysis. Economics Letters 71: 383−389.

Reimers, H.E. (1992). Comparisons of Tests for Multivariate Cointegration. Statistical Papers 33: 335−359.

Rigobon, R. (2002). Contagion: How to Measure It? In S. Edwards and J. Frankel (Eds.): Currency Crises Prevention. Chicago: University of Chicago Press.

Syllignakis, M.N., and G.P. Kouretas (2006). Long and Short-Run Linkages in CEE Stock Markets: Implications for Portfolio Diversification and Stock Market Integration. Working Paper 832. William Davidson Institute. Available at http://ssrn.com/abstract=910507.

Syriopoulos,T. (2006). Risk and Return Implications from Investing in Emerging European Stock Markets. Journal of International Financial Markets, Institutions, and Money 16: 28−−299.

Voronkova, S. (2004). Equity Market Integration in Central European Emerging Markets: A Cointegration Analysis with Shifting Regimes. International Review of Financial Analysis 13: 633−647.

Wu, Y. (1995). Are There Rational Bubbles in Foreign Exchange Markets? Evidence from an Alternative Test. Journal of International Money and Finance 14: 27−46.

Wu, Y. (1997). Rational Bubbles in the Stock Market: Accounting for the U.S. Stock-Price Volatility. Economic Inquiry 35: 309−319.

Yuca, A., and C. Simga-Mugan (2000). Linkages among Eastern European Stock Markets and the Major Stock Exchanges. Russian and East European Finance and Trade 36: 54−69.

Figure 1: Stock-Market Indexes

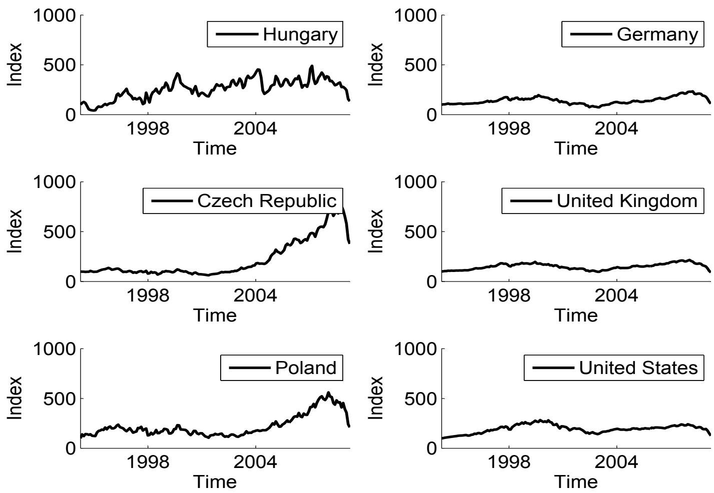


Figure 2: Estimation Results

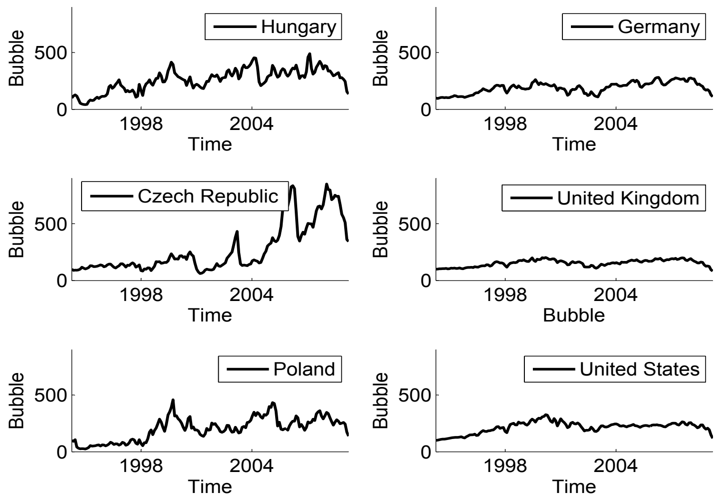
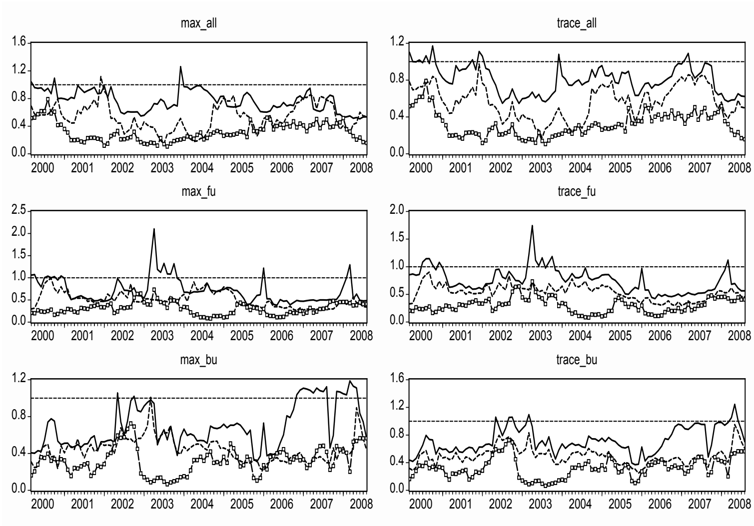


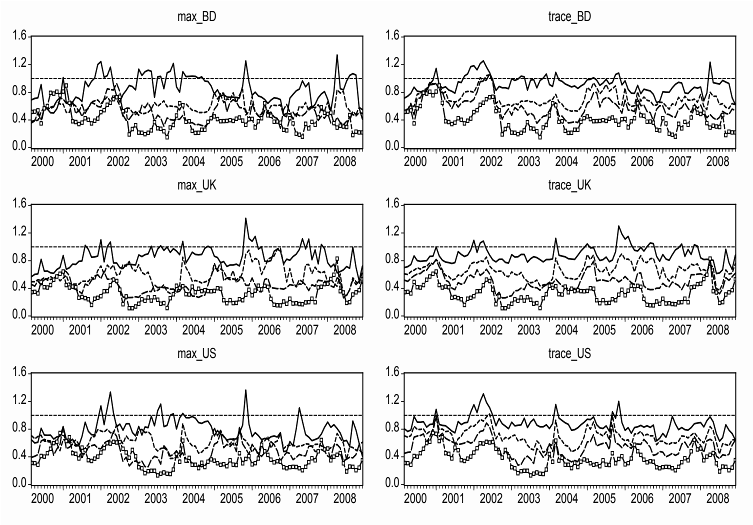
Figure 3: Cointegration among the Stock Markets of the CEE Countries



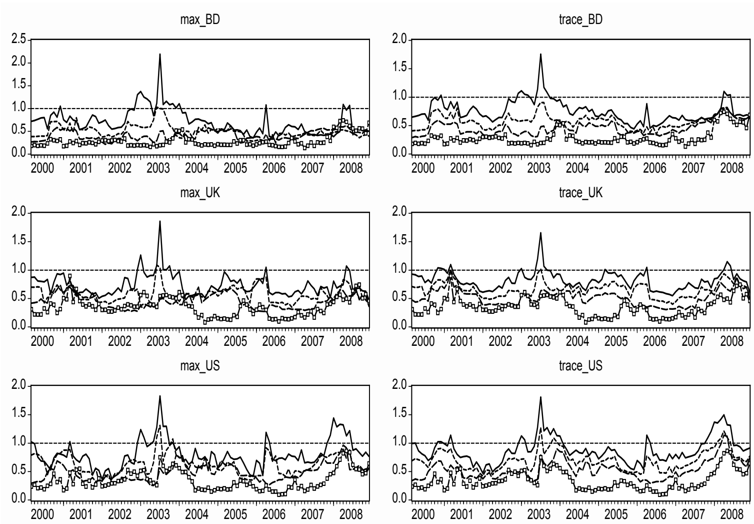
Note: The tests for cointegration are based on models that feature the stock markets of the three CEE countries. The solid (dashed, circled) lines represent the results of tests for cointegration based on the $\lambda_{trace}(0)(\lambda_{trace}(1), \lambda_{trace}(2))$ and the $\lambda_{max}(0)(\lambda_{max}(1), \lambda_{max}(2))$ statistics. The dashed vertical line represents the 95 percent critical value. The statistics have been scaled by their critical values. The critical values were 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.

Figure 4: Cointegration between CEE countries and Western Countries

Panel (a) Cointegration between Stock Market Indexes



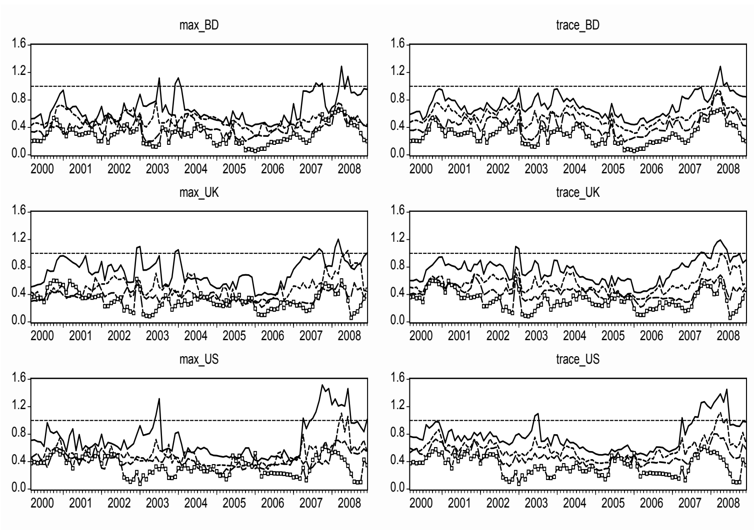
Panel (b) Cointegration between Fundamentals



to be continued

Figure 4 continued

Panel (c) Cointegration between Speculative Bubbles



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

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

	Real stock price index	Real dividends	Cointegration test
Czech Republic	0.1724	0.2088	-1.1201
Hungary	-0.6908	-0.3169	-2.1481
Poland	-1.2507	1.030	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

Note: The table reports p-values 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 five percent level of significance. The cointegration test is is a test for cointegration between real stock price indexes and real dividends. Speculative bubbles cannot be ruled out if real stock price indexes and real dividends are not cointegrated.

Table 2: Estimation Results for the State-Space Model

	φ_1	φ_2	ϕ	σ_ϵ^2	σ_u^2
Czech Republic					
Coefficient	-0.4156	-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	1.01424 [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 $\phi < 1$.