JEL Classification: G10, F31, C20
Keywords: stock and foreign exchange markets, structural breaks in variance, volatility spillovers, causality in mean and variance

Testing for Causality in Mean and Variance between the Stock Market and the Foreign Exchange Market:
An Application to the Major Central and Eastern European Countries

Sinem Derindere KOSEOGLU—Istanbul University, School of Transportation and Logistics (sderin@istanbul.edu.tr)
Emrah Ismail CEVIK—Bulent Ecevit University, Department of Econometrics (emrahic@yahoo.com), *corresponding author*

Abstract
The aim of this paper is to investigate the presence of a causality relationship between the stock market and the foreign exchange market in the Czech Republic, Hungary, Poland, and Turkey. We first analyze the existence of structural breaks in the variance of stock and foreign exchange rate returns series. Then, we employ the causality-in-mean/variance test proposed by Hong (2001). Our empirical results suggest that stock markets Granger-cause foreign exchange markets for all the countries in both mean and variance. Therefore, it can be said that the stock market has an important role in the price discovery process for the foreign exchange market for the countries in question

1. Introduction

The relationship between the exchange rate and stock prices has been widely examined in the literature according to two main exchange rate determination models —the Balance of Trade Model (Flow Oriented Model, Goods Market Model) and the Portfolio Balance Model (Stock Oriented Model, Asset Market Model). The former model assumes that changes in the exchange rate affect the trade balance (Dornbusch and Fischer, 1980). This also has an impact on the outcomes and income of a company, which means they affect stock prices, as the latter are defined as the present value of the future cash flows of a company. The effect changes depending on whether the company is export or import-oriented. A depreciation (appreciation) of the domestic currency of a country affects the competitiveness of an export-oriented company by leading to an increase (decrease) in the demand for its export goods. However, the weak domestic currency has a negative effect on an importing company because of the increase in the cost of imported goods. In this way, exchange rate movements affect stock prices depending on the direction of the aforementioned effects on the company's income. In other words, with the Balance of Trade Model, a change in the exchange rate is completely reflected in traded goods prices, i.e., there is a full pass-through, and so the cash flows of a company are affected. Nevertheless, the exchange rate pass-through might sometimes be incomplete. It has been widely remarked that traded goods prices may not fully reflect changes in exchange rates all the time. Krugman (1987) explains this situation with the pricing-to-market phenomenon. This is the practice where a company fixes its traded goods prices in the market when exchange rates change because it does not want to lose its market share. Instead, it prefers to lose income for a short period. When pricing-to-market is being practiced, exchange rate changes do not affect companies' cash flows, which means that stock prices are not affected by exchange rates. Therefore, in the case of the Balance of Trade model empirical results should be considered in relation to pricing to market.

The latter model assumes that currencies are assets (Frankel, 1983) and currency prices are therefore determined in the same manner as the prices of assets such as stocks, bonds, gold or real estate (Shapiro, 1994). Thus, an increase in a country's stock prices results in growth of wealth, which leads investors to increase their demand for money, which in turn raises the country's interest rates. As the main characteristic of short-term capital inflows is their dependence on short-term interest rates, higher interest rates attract mainly short-term capital inflows and initiate an increase in foreign demand for the country's currency. This leads to an appreciation of the currency. In this way there is an expected relationship between stock prices and exchange rates in the Portfolio Balance Model.

The theoretical literature shows that there is no consensus on the relationship between stock prices and exchange rates. The empirical literature has similarly failed to reach a consensus. Thus, there is a need to conduct further empirical tests of this issue. In our study, we investigate the relationship between stock returns and exchange rates in three major Central European countries, namely, the Czech Republic, Hungary, and Poland, as well as Turkey as a South-Eastern European country, for the period July 30, 2002–July 28, 2011. These countries (except for Turkey), which are generally referred to as transition economies, have experienced many changes in their economic policies since the collapse of communism. For instance, portfolio capital inflows to these countries have significantly increased recently. These capital inflows might have impacted on stock prices and exchange rates. Thus, we analyze the relationship between the stock and exchange rate markets for these countries. Most of the studies that examine the relationship between stock prices and exchange rates have focused on the U.S. or South-Eastern and South Asian countries. There have been a limited number of studies that focus on the Central and Eastern European countries (Murinde and Poshakwale, 2004; Islami, 2008), and these studies generally investigate the causality-in-mean relation between stock prices and exchange rates. However, causality-in-variance (or the volatility spillover effect) is especially important for financial variables because the rapid rise in the globalization of financial markets is increasing the transmission of returns and volatilities between these markets. Thus, in this study we analyze the causality not only in the mean, but also in the variance to reveal this relationship between stock prices and exchange rates for the countries in question. Fedorova and Saleem (2010) examine the relationships between stock and foreign exchange markets in terms of both mean and volatility for Eastern European countries. However, our study additionally examines the relationship between the stock and foreign exchange markets for Turkey and is different from this previous study in many aspects, which are mentioned in detail in the empirical results section.

In our study, we test for the presence of a causal link between stock prices and exchange rates for the Czech Republic, Hungary, Poland, and Turkey by using the two-step methodology proposed by Cheung and Ng (1996) and Hong (2001). In addition, we investigate the existence of regime shifts in the stock and foreign exchange markets of these countries due to the fact that the global financial crisis in 2008 and the recent European debt crisis may have led to structural changes in the volatility of stock and foreign exchange markets. All the empirical results indicate the existence of at least one structural break in the variance of the stock and exchange rate returns series in all countries, except for the foreign exchange market of Turkey. Secondly, the causality-in-mean test results show the presence of a unidirectional causal link going from the stock market to the foreign exchange market in all the countries. Finally, we observe the presence of causality-in-variance for all the countries, running from stock markets to foreign exchange markets. These results generally suggest that stock markets play a more dominant role in the price discovery process in all the countries.

The rest of the paper is organized as follows. Section 2 provides a brief review of the literature on the stock market and foreign exchange market. Section 3 discusses the econometric methodology. Section 4 presents the empirical results of the causality-in-mean and causality-in-variance tests. Section 5 brings the study to a conclusion.

2. Literature Review

The relationship between stock returns and exchange rates has been a subject of extensive research for a decade. These studies have used various different methodologies for different countries to examine this relationship, but the literature as a whole has failed to reach a consensus.

Early studies, in general, investigated the relations between stock returns and exchange rates by simple correlation and regression methods and dealt with unidirectional causality (e.g., Franck and Young, 1972; Aggarwal, 1981; Giovannini and Jorion, 1987; Smith, 1992; Solnik, 1987; Jorion 1990; Bondar and Gentry, 1993). More recent studies, however, have used more complicated econometric techniques. For instance, Bahmani-Oskooee and Sohrabian (1992) used cointegration and Granger causality tests to analyze the relations between stock prices and exchange rates in the U.S. The evidence showed that there is no long-run relationship between the variables. However, Roll (1992) found a positive relationship between stock prices and exchange rates in the U.S. Mok (1993) determined the existence of causality between the daily exchange rate and stock prices in Hong Kong for the period from 1986 to 1991. Ajayi and Mougoue (1996) examined the same relations for eight developed countries using daily data and found a significant relationship between the foreign exchange and stock markets.

Yu (1997) conducted the same study for Hong Kong, Tokyo, and Singapore using the Granger causality test and daily data for the period 1983–1994. The empirical results suggested that in Hong Kong stock prices are caused by exchange rates. There is no causality for Singapore and finally there is bidirectional causality in Tokyo. In addition, he also analyzed and found long-run relations between the variables in all countries. Abdalla and Murinde (1997) examined the long-run relation between stock prices and the real effective exchange rate for Pakistan, Korea, India, and the Philippines using cointegration tests for the period 1985–1994. The empirical evidence showed that there is no long-run relationship for Pakistan and Korea, but there is a long-run relationship for India and the Philippines.

Nieh and Lee (2001) conducted the same study for G7 countries using daily data for the period 1993–1996. They found no long-run relation between stock prices and exchange rates for any of the G-7 countries. Mansor (2000) also found no long-run relation between the variables for Malaysian markets. Granger (2000) investigated the causality of the variables for Asian countries and found a positive relation between the variables in the Japanese and Thai markets, a negative relation in the Taiwanese market, and no relation in the other markets. Kim (2003) investigated relations between stock prices and exchange rates for U.S. markets using a multivariate cointegration and error correction model for the period 1974–1998. The results revealed that the stock price is negatively related to the exchange rate.

Tsoukalas (2003) examined the same relations for the Cypriot equity market. The results of the study showed strong relations between the variables. Murinde and Poshakwale (2004) investigated this relationship for Hungary, the Czech Republic, and Poland for two specific periods: the pre-euro period and the euro period. They found that for the pre-euro period, stock prices Granger-cause exchange rates in Hungary only; in the Czech Republic and Poland, bi-directional causality between exchange rates and stock prices seems to exist. During the euro period, exchange rates unidirectional-Granger-cause stock prices in all the three economies. Pan et al. (2007) examined the dynamic linkages between exchange rates and stock prices for East Asian countries over the period 1988 to 1998. The results showed that there is a bidirectional causal relation for Hong Kong before the 1997 Asian crises. Also, there is a unidirectional causal relation from exchange rates to stock prices for Japan, Malaysia, and Thailand and from stock prices to exchange rates for Korea and Singapore. During the Asian crisis, there is only a causal relation from exchange rates to stock prices for all the countries except Malaysia. Sundaram (2009) found that there is no long-run or causal relationship between stock returns and exchange rates for India.

Rahman and Uddin (2009) investigated the interactions between stock prices and exchange rates in three emerging countries of South Asia, namely, Bangladesh, India, and Pakistan. They also found no long-run or causal relations between the variables. Hamrita and Trifi (2011) examined the relationships between the variables using a wavelet transform. They found a bidirectional relation between exchange rate returns and stock index returns in the U.S. As is apparent from the literature given above, the empirical results about price or return relationships between stock and exchange rate markets change depending on the countries analyzed, the time period taken, and the methodologies used. Hence, all the empirical results are inconclusive.

When we also look at studies investigating volatility spillovers and causality in variance between these markets, the results are still inconclusive. For instance, Kanas (2000) analyzed volatility spillovers between stock and exchange rate markets in the U.S., the UK, Japan, Germany, France, and Canada. He found evidence of spillovers from stock returns to exchange rate changes for all the countries except for Germany. There is no volatility spillover from exchange rate markets to stock markets. Similarly, Caporali et al. (2002) investigated the causal relationship between stock prices and exchange rate volatility for four East Asian countries. They found that in the pre-crisis sample stock prices affect exchange rates negatively in Japan and South Korea and positively in Indonesia and Thailand. In the latter two countries after the onset of the 1997 East Asian crisis the spillover effects are found to be bidirectional.

Yang and Doong (2004) investigated the same relations for the G7 countries. The results indicated that there is significant volatility spillover and an asymmetric effect from the stock market to the foreign exchange market for France, Italy, Japan, and the U.S. In line with our study, Fedorova and Saleem (2010) examined the dynamic relations between stock exchange and foreign exchange markets in Poland, Hungary, Russia, and the Czech Republic using weekly data for the period of 1995–2008. They provided evidence of unidirectional volatility spillovers from currency markets to stock markets in all the countries with the exception of the Czech Republic, where changes in the stock market influenced the currency market.

3. Econometric Framework

In this study, we employ the two-step methodology proposed by Cheng and Ng (1996) and Hong (2001) to determine the causal relation between stock and foreign exchange markets. Although in the literature the most common approach to obtaining the causal link between variables is the traditional Granger causality test, the Granger causality test procedure is very sensitive to the choice of lag length (Jones, 1989; Urbain, 1989) and relies on some distributional and time series assumptions (e.g., normality, homoscedasticity, stationarity, and cointegration). On the other hand, it is well known that most financial returns series exhibit non-normality and the ARCH effect.[1] In addition, the Granger causality test focuses only on changes in the mean of the two variables. However, causality in variance is as important as causality in mean for financial variables such as stock returns and exchange rates, because it implies a general pattern in volatility transmission between financial markets. In this context, Mantalos and Shukur (2010) determined that the Wald test based on the VAR model over-rejects the null hypothesis of noncausality when there are volatility spillover effects between the financial variables. Moreover, Cheung and Ng (1996) indicated that changes in variance are said to reflect the arrival of information and the extent to which the market evaluates and assimilates the new information. In addition, the causation pattern in variance provides an insight into the characteristics and dynamics of economic and financial prices, and such information can be used to construct better econometric models describing the temporal dynamics of the time series. Therefore, we examine the presence of a causal link in mean and variance between stock returns and foreign exchange rates in this study.

Two approaches have been widely used in the literature for testing causality in variance. One of them is the two-step methodology proposed by Cheung and Ng (1996) and Hong (2001), which is based on the cross correlation function (CCF) of squared GARCH residuals. The other approach depends on a dynamic specification of a multivariate GARCH (MGARCH) model, where causality in variance can be represented in terms of specific parameter restrictions.

Hafner and Herwartz (2006) indicated that likelihood-based tests within multivariate dynamic models typically suffer from the curse of dimensionality. Also, multivariate GARCH models are argued in the literature because they require

[1] Note that the Granger causality test can be employed by using consistent variance-covariance matrices when series exhibit heteroskedasticity or autoregressive conditional heteroscedasticity (ARCH) properties.

the imposition of a large amount of parameter constraints to ensure covariance stationary in the estimation procedure. In this context, it can be said that the Cheung and Ng (1996) and Hong (2001) causality tests are more practical and yet still have a powerful fit with large data series, leptokurtic series, and cases of noncorrelated errors (see, e.g., Caporale et al., 2006; Pardo and Torro, 2007; Quadan and Yagil, 2012). Therefore, we examine the presence of a causal link between stock and exchange markets based on the estimation of univariate GARCH models.

The first step of the causality-in-mean/variance test is to estimate a univariate GARCH model for the stock market index and foreign exchange rates. Although different GARCH model specifications can be considered for testing for causal relations between variables, we employ the exponential GARCH (EGARCH) model proposed by Nelson (1991) to examine the presence of a leverage effect in the volatility of stock price and exchange returns series.[2] The causality-in-variance test can be defined as follows.

Let the stock market (s_t) and the foreign exchange rate (fx_t) be two stationary and ergodic time series described by the following EGARCH processes:

$$\begin{aligned} s_t &= \mu_{s,t} + \varepsilon_t \\ \varepsilon_t \backslash &\left(\varepsilon_{t-1}, \varepsilon_{t-2}, ..., s_{t-1}, s_{t-2}, ...\right) \sim GED\left(0, h_{s,t}\right) \\ h_{s,t} &= \omega + \beta \log\left(h^2_{s,t-j}\right) + \alpha \left|\frac{\varepsilon_{t-1}}{\sigma_{t-1}}\right| + \gamma \frac{\varepsilon_{t-1}}{\sigma_{t-1}} \end{aligned} \tag{1}$$

$$\begin{aligned} fx_t &= \mu_{fx,t} + \zeta_t \\ \zeta_t \backslash &\left(\zeta_{t-1}, \zeta_{t-2}, ..., fx_{t-1}, fx_{t-2}, ...\right) \sim GED\left(0, h_{fx,t}\right) \\ h_{fx,t} &= \omega + \beta \log\left(h^2_{fx,t-j}\right) + \alpha \left|\frac{\zeta_{t-1}}{\sigma_{t-1}}\right| + \gamma \frac{\zeta_{t-1}}{\sigma_{t-1}} \end{aligned} \tag{2}$$

where $\mu_{s,t}$ and $\mu_{fx,t}$ are the means of s_t and fx_t, and ε_t and ζ_t are the innovation processes for s_t and fx_t, respectively.

Let I_t and J_t be two information sets defined by $I_t = \left\{s_{t-j}; j \geq 0\right\}$ and $J_t = \left\{s_{t-j}, fx_{t-j}; j \geq 0\right\}$. fx_t is said to cause s_{t+1} in variance if:

$$E\left\{\left(s_{t+1} - \mu_{s,t+1}\right)^2 | I_t\right\} \neq E\left\{\left(s_{t+1} - \mu_{s,t+1}\right)^2 | J_t\right\} \tag{3}$$

Equation (3) indicates that the information set for the foreign exchange returns series will improve the forecast of the stock returns series, hence the foreign exchange returns series is referred to as the Granger cause of the stock returns series. Cheung and Ng (1996) operationalize the notion of causality in variance defined in equation (3). First, the squares of the standardized innovations ε_t and ζ_t in equations

[2] We also employ other univariate GARCH specifications, such as the GARCH-in-mean model, and find similar causality results.

(1) and (2) are taken: $u_t = \left\{ \left(s_t - \mu_{s,t} \right)^2 / h_{s,t} \right\} = \varepsilon_t^2$ and $v_t = \left\{ \left(fx_t - \mu_{fx,t} \right)^2 / h_{fx,t} \right\} = \zeta_t^2$. Then, an *S*-statistic for testing the causal relationship at a specified lag M is computed:

$$S = T \sum_{j=i}^{M} \hat{\rho}_{uv}^2 (j) \tag{4}$$

In equation (4), $\hat{\rho}_{uv}^2\,(j)$ indicates the sample cross-correlation at lag *j*, which is calculated from $\hat{\rho}_{uv}^2 (j) = \left\{ \hat{C}_{uu}(0) \hat{C}_{vv}(0) \right\}^{-1/2} \hat{C}_{uv}(j)$, where the sample cross-covariance function is given by:

$$\hat{C}_{uv}(j) = \begin{cases} T^{-1} \sum_{t=j+1}^{T} \hat{u}_t \hat{v}_{t-j}, & j \geq 0 \\ T^{-1} \sum_{t=-j+1}^{T} \hat{u}_{t+j} \hat{v}_t, & j < 0 \end{cases}$$

with $\hat{C}_{uu}(0) = T^{-1} \sum_{t=1}^{T} \hat{u}_t^2$, $\hat{C}_{vv}(0) = T^{-1} \sum_{t=1}^{T} \hat{v}_t^2$. Note that $\hat{u}_t$ and $\hat{v}_t$ are squared standardized residuals derived from the GARCH models expressed in equations (1) and (2).

Cheung and Ng's (1996) *S*-statistic has a chi-square distribution with (M–i+1) degrees of freedom. The null hypothesis is that of no causality at all lags from *i* to *M*. The alternative hypothesis is the presence of causality at some lag *j*. Note that the presence of a causal relationship in the mean of the variables of interest can also be tested for by employing Cheung and Ng's (1996) methodology. In this case, the *S*-statistic in equation (4) is calculated using the respective GARCH innovations ε_t and ζ_t (instead of their squares) from equations (1) and (2) in the sample cross-correlation and cross-covariance functions.

On the other hand, one criticism of the *S* test statistic is that it may not be fully efficient when a large *M* is used because it gives equal weighting to each of the *M* sample cross-correlations. However, empirical studies show that the cross-correlation between financial assets decays to zero when the lag order *l* is increased. In this context, Hong (2001) introduced the modified *S* statistic by using the non-uniform kernels weighting function. He indicates that his test statistic, in which the null hypothesis shows that there is no causality, outperforms the *S* statistic in Monte Carlo simulation studies. Hong's (2001) test statistic is defined as:

$$Q = \frac{T \sum_{l=1}^{T-1} k^2 \left(\frac{l}{M} \right) \hat{\rho}_{uv}^2 (l) - C_{1T}(k)}{\sqrt{2 D_{1T}(k)}} \tag{5}$$

In equation (5), $k(l / M)$ is a weight function, for which we use the Bartlett kernel

$$k(l/M)=\begin{cases}1-|l/(M+1)| & if\ k/(M+1)\le 1\\ 0 & otherwise\end{cases} \quad (6)$$

where

$C_{1T}(k)=\sum_{l=1}^{T-1}(1-|l|/T)k^2(l/M)$ and $D_{1T}(k)=\sum_{l=1}^{T-1}(1-|l|/T)\{1-(|l|+1)/T\}k^4(l/M)$.

The test Q-statistic is a one-sided test, and upper tailed normal distribution critical values should be used. For example, the asymptotic critical value at the 5% level is 1.645. The test procedure, summarized by Hong (2001), is given as:

- Estimate univariate GARCH (p, q) models for the time series and save the standardized residuals.
- Compute the sample cross-correlation function $\hat{\rho}_{uv}(l)$ between the centered standardized residuals.
- Choose an integer M and compute $C_{1T}(k)$ and $D_{1T}(k)$.

Then compute the test Q-statistic and compare it to the upper-tailed critical value of the normal distribution at an appropriate level. If Q is larger than the critical value, there is no causality, and accordingly the null hypothesis is rejected.

However, an extensive literature on the estimation of GARCH models has argued that the presence of structural breaks in the unconditional variance of series leads to overestimation of GARCH parameters. For instance, Hillebrand (2005) showed that parameter regime changes in GARCH models that are not accounted for in global estimations cause the sum of the estimated GARCH parameters to converge to one via Monte Carlo simulations. He referred to this effect as "spurious almost-integration".

These findings are very important for testing causality in variance, because Javed and Mantalos (2011) determined that causality-in-variance test results are very sensitive to the GARCH parameters. Therefore, biased GARCH model results can generate misleading causality results. These findings are confirmed by Van Dijk et al. (2005) and Rodrigues and Rubia (2007), because they determined that the causality-in-variance test suffers from severe size distortions when there are structural breaks in the variance of series. Accordingly, we examined the presence of structural breaks in the unconditional variance of both returns series before testing for causality in variance.[3]

Inclan and Tiao (1994) proposed a test procedure that is based on ICSS (Iterative Cumulative Sum of Squares) to detect structural breaks in the unconditional variance of a stochastic process. In order to test the null hypothesis of constant unconditional variance against the alternative hypothesis of a break in the unconditional variance, Inclan and Tiao (1994) proposed using the statistic given by:

$$IT=\sqrt{T/2}D_k \quad (7)$$

where $D_k=(C_k/C_T)-(k/T)$, C_T is the sum of the squared residuals from

[3] Note that the presence of structural breaks in the mean of series may affect the causality-in-mean test results. Therefore, we also implement the multiple structural breaks test proposed by Bai and Perron (1998, 2003). According to the test results, we do not determine the presence of structural breaks in the mean of the series. The results are available on request.

the whole sample period, and $C_k = \sum_{t=1}^{k} r_t^2$ is the cumulative sum of squares of a series of uncorrelated random variables with mean 0 and variance σ_t^2 , $t = 1,2,...T$. The value of k ($k = 1,...T$) that maximizes $\left|\sqrt{T/2}D_k\right|$ is the estimate of the structural break date. Under variance homogeneity, the *IT*-statistic behaves like a Brownian bridge asymptotically. At the 5% significance level, the critical value computed by Inclan and Tiao (1994) is $C_{0.05} = 1.358$.

The most serious drawback of the *IT*-test statistic is that it is designed for independently and identically distributed random variables. However, Andreuo and Ghysels (2002) and Sanso et al. (2004) determined that the test statistic generates oversized results when the dependent variable exhibits a conditional heteroscedasticity process. In this context, Fernandez (2008) determined that the *IT* test statistic fails to find an effect of the September 11 terrorist attacks on the volatility of world stock markets. Sanso et al. (2004) modified the *IT* test statistic for a GARCH process in the dependent variable and showed that the modified test statistic outperforms the *IT*-test statistic by means of Monte Carlo simulation. In this study, the modified *IT*-test statistic was used to detect break points in the variance of returns series as in Arago-Manzana and Fernandez-Izquierdo (2007), Rapach and Strauss (2008), and Ewing and Malik (2010). The modified *IT* test statistic is given by

$$\kappa = \sup_k \left|T^{-1/2} G_k\right| \tag{8}$$

where $G_k = \hat{\omega}_4^{-1/2}\left(C_k - \frac{k}{T}C_T\right)$ and $\hat{\omega}_4$ is a consistent estimator of ω_4. The non-parametric estimator of ω_4 is

$$\hat{\omega}_4 = \frac{1}{T}\sum_{t=1}^{T}\left(r_t^2 - \hat{\sigma}^2\right)^2 + \frac{2}{T}\sum_{l=1}^{m}\omega(l,m)\sum_{t=l+1}^{T}\left(r_t^2 - \hat{\sigma}^2\right)\left(r_{t-1}^2 - \hat{\sigma}^2\right) \tag{9}$$

where $\omega(l,m)$ is a lag window, such as the Bartlett, defined as $\omega(l,m) = 1 - l/(m+1)$, or the quadratic spectral.

In the test procedure, if we were looking for only the possibility of a single point change, then the G_k function would provide a satisfactory procedure. But when we are interested in finding multiple change points on an observed series, the usefulness of the G_k function becomes questionable because of the masking effect. A solution is an iterative scheme based on successive application of G_k to pieces of the series, dividing consecutively after a possible change point is found (see Inclan and Tiao, 1994, for ICSS procedure details).

4. Data and Empirical Results

In this study, we employ daily data for all countries covering the period from July 30, 2002 to July 28, 2011 for a total of 2348 observations. The stock price indices are obtained from the MSCI-Barra database and represent all share indices for the local stock market. The foreign exchange rates are the local currencies against

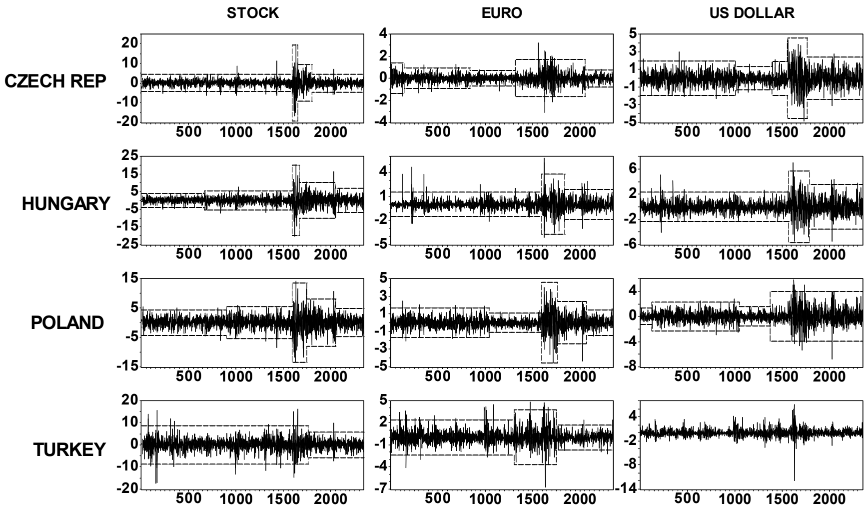
Table 1 Descriptive Statistics for Stock Index and Exchange Rates Returns Series

	CZECH REPUBLIC			HUNGARY		
	Stock	***Euro***	***US Dollar***	***Stock***	***Euro***	***US Dollar***
Mean	0.050	0.004	-0.012	0.052	0.004	-0.012
Std. Dev.	2.394	0.646	1.028	2.395	0.646	1.028
Skewness	-0.055	0.805	0.369	-0.056	0.806	0.369
Kurtosis	12.040	12.933	7.058	12.031	12.927	7.051
Jarque-Bera	7995.4 [0.000]	9907.2 [0.000]	1663.9 [0.000]	7964.1 [0.000]	9873.1 [0.000]	1655.1 [0.000]
ARCH (10)	63.226 [0.000]	39.951 [0.000]	40.943 [0.000]	63.226 [0.000]	39.951 [0.000]	40.493 [0.000]
Q (50)	151.933 [0.000]	85.690 [0.001]	76.531 [0.009]	151.933 [0.000]	85.690 [0.001]	76.531 [0.009]
Q_s (50)	2846.24 [0.000]	1073.98 [0.000]	2162.99 [0.000]	2846.24 [0.000]	1073.98 [0.000]	2162.99 [0.000]
ADF	-22.371***	-50.761***	-48.865***	-22.375***	-50.693***	-48.827***
PP	-44.341***	-51.571***	-48.877***	-44.317***	-51.501***	-48.833***
KPSS	0.228***	0.030***	0.062***	0.065***	0.029***	0.035***

	POLAND			TURKEY		
	Stock	***Euro***	***US Dollar***	***Stock***	***Euro***	***US Dollar***
Mean	0.055	-0.001	-0.017	0.067	0.017	0.001
Std. Dev.	2.102	0.664	0.994	2.656	0.862	0.896
Skewness	-0.198	0.131	0.264	-0.259	0.450	-0.100
Kurtosis	7.648	9.084	7.656	8.017	8.315	20.964
Jarque-Bera	2124.6 [0.000]	3620.7 [0.000]	2143.9 [0.000]	2483.1 [0.000]	2836.8 [0.000]	31507.8 [0.000]
ARCH (10)	54.017 [0.000]	50.286 [0.000]	55.216 [0.000]	23.570 [0.000]	29.782 [0.000]	36.708 [0.000]
Q (50)	90.166 [0.000]	134.407 [0.000]	84.818 [0.001]	89.739 [0.000]	74.346 [0.014]	87.963 [0.000]
Q_s (50)	2665.98 [0.000]	3723.48 [0.000]	3615.05 [0.000]	760.651 [0.000]	944.392 [0.000]	753.231 [0.000]
ADF	-45.083***	-30.121***	-47.967***	-46.900***	-47.501***	-46.437***
PP	-44.968***	-50.668***	-47.972***	-46.891***	-47.492***	-46.437***
KPSS	0.076***	0.079***	0.044***	0.046***	0.030***	0.028***

Notes: The figures in square brackets show the probability (p-values) of rejecting the null hypothesis. ARCH (10) indicates LM conditional variance test. Q(50) and Q_s(50) indicates Ljung-Box serial correlation test for returns and squared returns series respectively. *** indicate that the series in question is stationary at the 1% significance level.

Figure 1 Daily Returns Series for Stock Index and Exchange Rates



Notes: Dashed line indicates ± 3 standard deviations.

the U.S. dollar and the euro collected from the national central banks of the countries in question. The logarithmic stock and foreign exchange returns series are used in the empirical analysis.

The descriptive statistics for the stock index and exchange rate returns series are presented in *Table 1*. The daily mean of all the stock index returns series varies between 0.050 percent and 0.067 percent. The highest mean return occurs in the Turkish stock market. The Czech stock market, on the other hand, yields the lowest mean returns during the sample period. Furthermore, the Turkish stock returns series exhibit higher volatility according to the standard deviation statistic. When we look at the exchange rate series, the highest mean return is obtained for the Turkish foreign exchange market, which is again more volatile. All the returns series show evidence of strong skewness and excess kurtosis, which indicates that they are leptokurtic. The Jarque-Bera normality test results show that the distributions of the returns series are not normal. Box-Pierce *Q*-statistics strongly indicate the presence of serial correlation in the returns and squared returns series. Finally, we examine the existence of the unit root by means of the augmented Dickey-Fuller (ADF), Phillips-Perron (PP), and Kwiatkowski, Phillips, Schmidt and Shin (KPSS) unit root tests. The results of all three unit root tests suggest that all the returns series are stationary.

We started our empirical analysis by testing for the presence of sudden changes in the variance of the stock index and exchange rate returns series by means of the modified *IT*-statistic. *Figure 1* illustrates the returns for each series with the points of sudden change and ± 3 standard deviations. In addition to this, *Table 2* indicates the time periods of sudden changes in volatility, as identified by the ICSS algorithm.

According to the test results, we observe three sudden change points for the stock index returns series in the Czech Republic, and the dates of the structural

Table 2 Variance Structural Breaks Test Results

Countries	Series	Number of Breaks for *IT* Statistics	Number of Breaks for Modified *IT* Statistics	Break Dates			
Czech Rep.	Stock	18	3	04.09.08	25.11.08	25.06.09	-
	Euro	7	4	13.02.03	13.10.05	15.08.07	11.06.10
	US Dollar	4	4	12.06.06	05.12.07	22.07.08	07.05.09
Hungary	Stock	7	4	12.05.05	04.09.08	17.12.08	04.06.10
	Euro	6	2	04.09.08	17.08.09	-	-
	US Dollar	14	2	07.08.08	10.06.09	-	-
Poland	Stock	13	4	02.01.06	04.09.08	09.04.09	14.06.10
	Euro	14	4	25.07.06	04.09.08	30.04.09	29.06.10
	US Dollar	9	3	21.03.03	07.08.06	05.11.07	-
Turkey	Stock	20	1	30.04.09	-	-	-
	Euro	24	2	26.07.07	14.04.09	-	-
	US Dollar	13	0	-	-	-	-

breaks are related to the global financial crisis. Also, four structural break dates are detected in the variance of the foreign exchange returns series in the Czech Republic. The Hungarian stock market has five regime shifts and the break dates are observed in the period of 2005–2010. On the other hand, we observe two sudden change points in the Hungarian foreign exchange market, occurring in 2008 and 2009. It should be noted that although the U.S. dollar and euro-denominated returns series exhibit the same regime shifts in the Hungarian foreign exchange market, the test results are a little bit different for the Polish foreign exchange market. Correspondingly, although the euro-denominated returns series has five regime shifts, four regime shifts are obtained for the U.S. dollar-denominated returns series in the Polish foreign exchange market. We also observe four sudden change points in the variance of the Polish stock market.

On the other hand, we observe lower regime shifts for the stock and foreign exchange returns series in Turkey than in the other countries in question. We found only one structural break date in the variance of the stock indices in 2009. While we observe two regime shifts in the variance of the euro-denominated returns series, we do not determine any structural break for the U.S. dollar-denominated returns series in the Turkish foreign exchange market. Note that the *IT*-test results suggest more sudden change points than the modified *IT*-statistics for all countries. These results are consistent with the empirical results in the literature (see, e.g., Andreuo and Ghysels, 2002, and Sanso et al., 2002).

If we summarize the structural break test results, the important thing to note for both the stock and foreign exchange markets is that all countries experienced a significant drop in their stock and foreign exchange markets in the period of 2007–2010 due to the global financial crisis. This leads to a finding of the presence of structural breaks in the variance of the series. Both the currency values and the stock exchange indices of these countries decreased significantly in the aftermath of the crisis.

It is of interest to relate structural breaks to the exchange rate policy of the country in question. For instance, the Czech Republic adopted a floating exchange rate regime in 1997, whereas Poland and Hungary have been using this regime since 2000 and 2008, respectively. Hungary was the last country to start practicing this regime among these three Central European countries; before February 2008 it followed a crawling corridor or target zone system. Thus, the structural break test results for the foreign exchange market are consistent with the theoretical expectations, because we do not observe the presence of a structural break before 2008.

In order to eliminate the effects of the structural breaks, we construct dummy variables with regard to the time periods of sudden changes, as in Lamoureux and Lastrapes (1990), Aggarwal et al. (1999), Arago-Manzana and Fernandez-Izquierdo (2007), Wang and Thi (2007), and Ewing and Malik (2010).

Next, we estimated the univariate EGARCH model with and without dummy variables for the stock index and exchange rate returns series. In order to control for the possible day-of-the-week effect, we consider four dummy variables (Tuesday excluded) in the estimation method. The EGARCH (1,1) model was found to be sufficient for adequate model volatility for the stock index returns series.[4] The parameter *gamma* (γ), which indicates the presence of the leverage effect in the conditional volatility, was not found to be negative and statistically significant for the exchange rate returns series for any country, and hence we consider the GARCH model for the exchange rate returns series. The EGARCH model results are presented in *Table 3*.

According to the results in *Table 3*, the *alpha* (α) and *beta* (β) parameters are found to be statistically significant for the stock index and exchange rate returns series of all the countries. For all the stock index returns series, the *gamma* parameter is determined as statistically significant at the 1% level. Hence, this result provides evidence of the existence of the leverage effect in the stock market. Therefore, we can say that good news and bad news have a different impact on the volatility of stock index returns.

In addition, it can be seen in *Table 3* that the structural breaks in the variance of the series lead to an increase in the sum of the *alpha* and *beta* parameters. In particular, we observe a dramatic decrease in the *beta* parameter in all the returns series. These findings are consistent with the empirical results in the literature (see, e.g., Lamoureux and Lastrapes, 1990; Aggarwal et al., 1999; Arago-Manzana and Fernandez-Izquierdo, 2007; Wang and Thi, 2007; and Ewing and Malik, 2010).

The log likelihood values in *Table 3* indicate that the GARCH model with dummy variables gives a better fit for all the returns series. In addition, we employed a likelihood ratio (LR) test to determine the significance of the dummy variables in the volatility process. The LR test can be calculated by using $LR = 2[L(M_d) - L(M)]$, where $L(M_d)$ and $L(M)$ are the maximum log likelihood values derived from the GARCH models with and without dummy variables, respectively. The test statistic is asymptotically χ^2 distributed with degrees of freedom equal to the number of restrictions (or the number of dummy variables).

The LR test results in *Table 3* strongly indicate that the null hypothesis of no change in variance is rejected at the 1 percent level. Therefore, it can be said that

[4] We consider the Akaike information criteria in selecting the number of autoregressive parameters in the ARMA model.

Table 3 (E)GARCH Model Results

	CZECH REPUBLIC						HUNGARY					
	Stock		US Dollar		Euro		Stock		US Dollar		Euro	
	Without dummies	With dummies	Without dummies	With dummies	Without dummies	With dummies	Without dummies	With dummies	Without dummies	With dummies	Without dummies	With dummies
Constant	-0.056	-0.072	-0.054*	-0.049	0.001	0.001	0.107	0.079	-0.075**	-0.074**	-0.006	-0.006
AR(1)	0.733***	0.666**	0.496**	0.911**	0.101**	-0.259***	-1.196***	-0.636***	0.745***	0.794***	0.367	0.339
AR(2)	0.264***	0.332	0.154	-0.170	1.339***	0.369***	-1.234***	-0.602***	-1.089***	-1.646***	-0.267	-0.322
AR(3)	-	-	-	-	0.152***	-0.449***	-1.248***	-0.668***	0.453***	0.757***	-0.097	-0.112
AR(4)	-	-	-	-	0.914***	-0.833***	-0.628***	0.272***	-0.546***	-0.858***	0.115	0.121
AR(5)	-		-	-	0.016	0.047**	0.031	-0.024	-	-	-0.029	-0.127
MA(1)	-0.682***	-0.619**	-0.451**	-0.865*	-0.064	0.304***	1.241***	0.685***	-0.726***	-0.774***	-0.332	-0.301
MA(2)	-0.315***	-0.384	-0.141	0.160	-1.330***	-0.350***	1.287***	0.611***	1.066***	1.653***	0.181	0.239
MA(3)	-		-	-	0.204***	0.430***	1.306***	0.663***	-0.459***	-0.758***	0.078	0.088
MA(4)	-		-	-	0.893***	0.861***	0.700***	- 0.261***	0.523***	0.854***	-0.162	-0.169
MA(5)	-		-	-	-	-	-	-	-	-	-	-
ω	-0.141***	-0.083***	0.002*	0.255**	0.001***	0.009**	-0.116***	-0.057**	0.014***	0.044***	0.007***	0.012***
α	0.218***	0.188***	0.031***	0.038*	0.066***	0.072***	0.18***	0.152***	0.072***	0.080***	0.184***	0.196***
β	0.967***	0.891***	0.964***	0.372	0.920***	0.882***	0.982***	0.903***	0.914***	0.841***	0.816***	0.763***
γ	-0.066***	-0.103***	-	-	-	-	-0.041***	-0.076***	-	-	-	-
v	1.503***	1.564***	1.451***	1.578***	1.277***	1.291***	1.572***	1.625***	1.444***	1.484***	1.088***	1.099***
$\alpha + \beta$	1.185	1.079	0.995	0.41	0.986	0.954	1.162	1.055	0.986	0.923	1.000	0.959
R^2	0.006	0.006	0.003	0.004	0.016	0.011	0.017	0.011	0.002	0.005	-0.005	-0.007
Ln(*L*)	-4264.08	-4239.322	-2475.646	-2449.108	-739.709	-727.829	-4856.29	-4840.06	-3043.05	-3026.44	-1665.94	-1652.95
Q (50)	52.899 [0.225]	54.135 [0.192]	39.608 [0.699]	12.968 [0.605]	16.221 [0.133]	21.448 [0.029]	37.063 [0.646]	38.021 [0.604]	49.580 [0.197]	43.876 [0.392]	58.751 [0.036]	58.073 [0.032]
Q_s (50)	62.528 [0.053]	55.556 [0.158]	49.167 [0.310]	15.180 [0.439]	10.594 [0.478]	9.504 [0.575]	50.901 [0.138]	47.697 [0.219]	54.352 [0.096]	61.093 [0.029]	3.429 [0.999]	4.156 [0.999]
ARCH(5)	1.016 [0.406]	1.230 [0.291]	0.180 [0.970]	0.673 [0.643]	0.321 [0.900]	0.198 [0.963]	0.841 [0.520]	1.152 [0.330]	2.085 [0.064]	2.478 [0.030]	0.095 [0.992]	0.174 [0.972]
LR	49.516 [0.000]		53.076 [0.000]		23.760 [0.000]		32.458 [0.000]		33.212 [0.000]		25.978 [0.000]	

Notes: The figures in square brackets show the *p*-values. *v* is GED parameter. *Q*(20) and Q_s(20) indicates Ljung-Box serial correlation test values for the returns and the squared returns series respectively. ARCH (5) shows heteroskedasticity test results. ***, ** and * indicates statistical significance at the 1%, 5% and 10% level, respectively

Table 3 (E)GARCH Model Results (*continued*)

	POLAND						TURKEY					
	Stock		US Dollar		Euro		Stock		US Dollar		Euro	
	Without dummies	With dummies	Without dummies	With dummies	Without dummies	With dummies	Without dummies	With dummies	Without dummies	With dummies	Without dummies	With dummies
Constant	0.030	-0.044	-0.022	-0.020	-0.022	-0.017	-0.084	-0.079	-0.034	-	-0.021	-0.018
AR(1)	-0.613***	0.358***	-0.771***	-0.769***	0.693***	0.352	0.090***	-0.315***	0.472*	-	-0.759***	0.114
AR(2)	0.094***	0.288***	-0.595***	-0.594***	-0.055	-0.244	-0.581***	-0.479***	-0.603***	-	-0.861***	0.231*
AR(3)	0.599***	0.587***	-0.032	-0.028	0.195***	0.133	-0.596***	-0.876***	0.545***	-	-0.638***	0.471***
AR(4)	-0.172***	-0.277***	-0.047*	-0.044*	-1.007***	-0.305**	-0.014	-0.192***	-0.241	-	-1.096***	-0.772***
AR(5)	-0.010	0.039**	-0.018	-0.014	0.583***	0.179***	-	-	-	-	-0.458**	-0.201***
MA(1)	0.633***	-0.340***	0.803***	0.803***	-0.729***	-0.385***	-0.030***	0.372***	-0.434*	-	0.825***	-0.059
MA(2)	-0.089***	-0.311***	0.610***	0.611***	0.051	0.235	0.583***	0.498***	0.540***	-	0.872***	-0.287**
MA(3)	-0.617***	-0.591***	-	-	-0.201***	-0.162	0.624***	0.891***	-0.521***	-	0.682***	-0.486***
MA(4)	0.135***	0.243***	-	-	1.012***	0.280*	0.064***	0.256***	0.204	-	1.123***	0.766***
MA(5)	-	-	-	-	-0.616***	-0.192	-	-	-	-	0.514**	0.265
ω	-0.070***	0.056	0.007***	0.010***	0.004***	0.031***	-0.070***	-0.039	0.025***	-	0.026***	0.045***
α	0.116***	0.075**	0.063***	0.068***	0.071***	0.069***	0.198***	0.194***	0.174***	-	0.158***	0.163***
β	0.984***	0.842***	0.928***	0.895***	0.916***	0.820***	0.953***	0.942***	0.803***	-	0.813***	0.769***
γ	-0.046***	-0.109***	-	-	-	-	-0.083***	-0.035***	-	-	-	-
v	1.479***	1.528***	1.488***	1.528***	1.428***	1.515***	1.36***	1.369***	1.252***	-	1.205***	1.244***
$\alpha + \beta$	1.100	0.917	0.991	0.963	0.987	0.889	1.151	1.136	0.977	-	0.968	0.932
R^2	0.008	0.006	0.000	0.000	0.009	0.007	0.002	0.001	0.001	-	0.008	0.009
$Ln(L)$	-4711.89	-4690.69	-2866.78	-2856.23	-1866.63	-1848.96	-5316.036	-5310.82	-2568.836	-	-2565.935	-2554.786
Q (50)	44.057 [0.344]	39.037 [0.556]	51.233 [0.182]	51.276 [0.181]	49.235 [0.150]	57.762 [0.085]	55.690 [0.077]	58.069 [0.051]	45.219 [0.339]	-	41.918 [0.388]	40.642 [0.442]
Q_s (50)	44.339 [0.333]	70.530 [0.003]	51.240 [0.182]	45.500 [0.368]	60.049 [0.022]	52.165 [0.094]	81.533 [0.000]	80.805 [0.005]	37.152 [0.683]	-	36.339 [0.636]	44.235 [0.297]
ARCH(5)	2.615 [0.022]	1.730 [0.140]	1.741 [0.138]	1.355 [0.246]	3.066 [0.016]	0.702 [0.590]	0.936 [0.441]	1.159 [0.326]	0.354 [0.841]	-	0.592 [0.668]	0.613 [0.653]
LR	42.400 [0.000]		21.102 [0.000]		35.336 [0.000]		10.432 [0.000]		-		22.298 [0.000]	

Notes: The figures in square brackets show the *p*-values. v is GED parameter. $Q(20)$ and $Q_s(20)$ indicates Ljung-Box serial correlation test values for the returns and the squared returns series respectively. ARCH (5) shows heteroskedasticity test results. ***, ** and * indicates statistical significance at the 1%, 5% and 10% level, respectively.

Table 4 Hong's Causality in Mean Test Results

	Causality Direction	*M* = 1	*M* = 2	*M* = 3	*M* = 4	*M* = 5
Czech Rep.	Stock → Euro	5.182*** [0.000]	4.965*** [0.000]	4.843*** [0.000]	4.749*** [0.000]	4.612*** [0.000]
	Euro → Stock	0.224 [0.411]	0.169 [0.433]	0.066 [0.474]	-0.055 [0.522]	-0.175 [0.570]
	Stock → US Dollar	36.395*** [0.000]	35.191*** [0.000]	32.862*** [0.000]	30.668*** [0.000]	28.814*** [0.000]
	US Dollar → Stock	0.379 [0.352]	0.234 [0.407]	0.178 [0.429]	0.162 [0.436]	0.122 [0.451]
Hungary	Stock → Euro	102.161*** [0.000]	98.968*** [0.000]	92.679*** [0.000]	86.823*** [0.000]	81.780*** [0.000]
	Euro → Stock	1.243 [0.107]	1.043 [0.148]	0.792 [0.214]	0.592 [0.277]	0.445 [0.328]
	Stock → US Dollar	140.354*** [0.000]	136.221*** [0.000]	127.632*** [0.000]	119.471*** [0.000]	112.417*** [0.000]
	US Dollar → Stock	-0.017 [0.507]	-0.139 [0.555]	-0.264 [0.604]	-0.387 [0.651]	-0.508 [0.694]
Poland	Stock → Euro	113.092*** [0.000]	109.707*** [0.000]	102.741*** [0.000]	96.135*** [0.000]	90.427*** [0.000]
	Euro → Stock	0.216 [0.414]	0.388 [0.349]	0.777 [0.219]	1.148 [0.126]	1.386* [0.083]
	Stock → US Dollar	122.938*** [0.000]	119.117*** [0.000]	111.455*** [0.000]	104.108*** [0.000]	97.690*** [0.000]
	US Dollar → Stock	0.510 [0.305]	0.423 [0.336]	0.335 [0.369]	0.237 [0.406]	0.169 [0.433]
Turkey	Stock → Euro	219.550*** [0.000]	229.138*** [0.000]	226.444*** [0.000]	218.398*** [0.000]	209.207*** [0.000]
	Euro → Stock	2.093** [0.018]	1.868** [0.031]	1.631* [0.051]	1.422* [0.077]	1.236 [0.108]
	Stock → US Dollar	447.565*** [0.000]	466.222*** [0.000]	460.342*** [0.000]	443.933*** [0.000]	425.304*** [0.000]
	US Dollar → Stock	-0.576 [0.718]	-0.729 [0.767]	-0.862 [0.806]	-0.951 [0.829]	-1.019 [0.846]

Notes: The figures in square brackets show the *p*-values. *, ** and *** indicates the existence of causal link at the 1%, 5% and 10% level respectively.

the inclusion of dummy variables in the GARCH model increases the explanatory power of the model.

We then employed Hong's test to determine the causal relation between the stock and foreign exchange markets. The results are presented in *Table 4*. The Hong's causality-in-mean test results indicate the presence of a causal link going from the stock market to the foreign exchange market for all countries. This finding is consistent with the Portfolio Balance Model. Thus, this relationship might be because the stock markets of these countries attract short-term capital flows from foreign investors, which affects the demand for domestic currency. The recent increase in portfolio capital inflows to these countries supports this result. On the other hand, we observe no causal link running from the foreign exchange market to the stock market in any country except Turkey, in which the euro-denominated returns series is found to be a Granger cause of the stock index returns series at the 5% level. This result shows that a full pass-through has not been observed in the exchange markets of these countries, which is consistent with Krugman's (1987) pricing-to-market phenomenon. It means that currency changes in these countries do not reflect traded goods prices completely; therefore, companies' cash flows are not affected. Exceptionally, the existence of causality running from the euro-denominated returns series to the stock market in Turkey shows that traded goods prices in Turkey have fully reflected changes in the euro exchange rate. This might be due to the recent significant increase in the amount of exports and imports between European Union (EU) countries and total trade. If trade with the EU is linked with the international competitiveness of Turkish companies, this provides a natural mechanism whereby foreign exchange movements can pass through to the Turkish stock market.

It should be noted that the causal link between the second moment of the stock and foreign exchange returns series indicates the presence of a volatility spillover effect. In this context, in order to examine the existence of a volatility spillover effect, we obtain the squares of the standardized GARCH residuals and employ the causality test. However, Cheung and Ng (1996) and Pantelidis and Pittis (2004) indicated that the presence of causality in mean leads to severe size distortions in testing for causality in variance if such effects are not filtered out. In our study, we remove the causality-in-mean effects by including the lagged values of the stock return series, which cause the exchange returns series in-mean in the mean equation of the GARCH model.

The causality-in-variance test results are presented in *Table 5*. They strongly indicate the existence of a causality relation going from the stock markets to the foreign exchange markets in all countries. The causality test results show that the stock market Granger-causes the foreign exchange market in mean and variance for all countries. These results are generally consistent with the Portfolio Balance Model. The findings are also important from a risk management perspective. Import or export-oriented companies whose costs and revenues are denominated in more than one currency face exchange rate risk due to the volatility of stock returns.

In comparison with our results, Fedorova and Saleem (2010) found a different direction of volatility spillover from the currency market to the stock market in Poland and Hungary. However, our study is different from this previous study in many respects. First, we examine the presence of a causal link between the stock and foreign exchange markets using the two-step methodology proposed by Hong (2001). Secondly, although they examined the presence of a causality relation for the period of 1995–2008, we do not consider the transition period, hence our sample covers the period of 2002–2011. In addition, our data frequency is daily, whereas Fedorova and Saleem (2010) considered weekly data in their empirical analysis. Ultimately, it can be said that these differences in methodology, sample periods, and data frequency may lead to the finding of a different causal link in Fedorova and Saleem (2010). In this context, Koutmos (1998) argued that empirical findings from different data frequencies may give different results, a finding which may have significance for investors with alternative investment periods.

More importantly, Fedorova and Saleem (2010) did not consider the effects of structural breaks on the testing procedure. It is well known that the stock and foreign exchange markets in transition economies were very volatile at the beginning of the transition period. Also, they were negatively affected by the 1998 Russian and

Table 5 Hong's Causality in Variance Test Results

	Causality Direction	*M* = 1	*M* = 2	*M* = 3	*M* = 4	*M* = 5
Czech Rep.	Stock → Euro	2.414*** [0.008]	2.219** [0.013]	1.919** [0.027]	1.628* [0.052]	1.416* [0.078]
	Euro → Stock	0.620 [0.268]	0.577 [0.282]	0.451 [0.326]	0.370 [0.356]	0.355 [0.361]
	Stock → US Dollar	6.668*** [0.000]	6.457*** [0.000]	5.973*** [0.000]	5.466*** [0.000]	5.155*** [0.000]
	US Dollar → Stock	0.248 [0.402]	0.077 [0.469]	0.218 [0.414]	0.428 [0.334]	0.543 [0.294]
Hungary	Stock → Euro	2.860*** [0.002]	2.615*** [0.004]	2.292** [0.011]	2.003** [0.023]	1.767** [0.039]
	Euro → Stock	-0.299 [0.617]	-0.411 [0.659]	-0.261 [0.603]	0.058 [0.477]	0.339 [0.367]
	Stock → US Dollar	6.020*** [0.000]	5.691*** [0.000]	5.176*** [0.000]	4.733*** [0.000]	4.416*** [0.000]
	US Dollar → Stock	-0.707 [0.760]	-0.856 [0.804]	-0.866 [0.807]	-0.731 [0.768]	-0.540 [0.706]
Poland	Stock → Euro	4.564*** [0.000]	4.273*** [0.000]	3.839*** [0.000]	3.435*** [0.000]	3.076*** [0.001]
	Euro → Stock	0.255 [0.399]	0.170 [0.432]	0.549 [0.291]	0.994 [0.160]	1.444* [0.074]
	Stock → US Dollar	7.482*** [0.000]	8.234*** [0.000]	8.517*** [0.000]	8.440*** [0.000]	8.182*** [0.000]
	US Dollar → Stock	-0.648 [0.741]	-0.698 [0.757]	-0.511 [0.695]	-0.308 [0.621]	-0.128 [0.551]
Turkey	Stock → Euro	17.737*** [0.000]	19.108*** [0.000]	20.141*** [0.000]	20.590*** [0.000]	20.545*** [0.000]
	Euro → Stock	-0.158 [0.563]	-0.276 [0.609]	-0.404 [0.657]	-0.527 [0.701]	-0.629 [0.735]
	Stock → US Dollar	6.577*** [0.000]	8.586*** [0.000]	10.290*** [0.000]	11.267*** [0.000]	11.680*** [0.000]
	US Dollar → Stock	-0.484 [0.686]	-0.568 [0.715]	-0.659 [0.745]	-0.731 [0.768]	-0.792 [0.786]

Notes: The figures in square brackets show the *p*-values. *, ** and *** indicates the existence of causal link at the 1%, 5% and 10% level respectively.

1997 Czech crises. These events might have caused structural breaks in their stock and foreign exchange returns series. In this context, Bensafta (2010) indicates that when there are structural breaks in the variance of series, the multivariate GARCH-BEKK model overestimates the volatility persistence, leading to biased causality-in-variance results. In addition, Arogo and Salvador (2011) showed that multivariate GARCH models including sudden changes in volatility outperform alternative models in terms of in-sample and out-of-sample forecasts. Therefore, we additionally examine the existence of structural changes in the countries' stock and foreign exchange markets to get more accurate causality results. This is very important for our time period due to the global financial crisis and the European debt crisis. Finally, we also investigate the presence of causal relations for Turkey as a South-Eastern European country.

5. Conclusion

In this study, we examined the dynamic relation between the stock market and the foreign exchange market for the Czech Republic, Hungary, Poland, and Turkey. In order to determine the presence of a causal link between the stock and foreign exchange rate returns series, we employed the two-step test procedure proposed by Hong (2001). In addition, empirical studies in the literature showed that structural breaks in the variance of series lead to size distortions in testing for causality in variance. Therefore, we first examine the existence of structural breaks in the variance of the stock and foreign exchange returns series by means of the modified *IT*-test. The modified *IT*-test statistic results indicate the presence of at least one structural break in the variance of the stock and foreign exchange returns series in all the countries. In order to eliminate regime shifts, we construct dummy variables and then we estimate a GARCH model with the dummy variables.

The causality test results strongly indicate that the stock market Granger-causes the foreign exchange market in mean and variance in all countries. These results are consistent with the earlier findings of Kanas (2000), Caporali et al. (2002), and Yang and Doong (2004), who found evidence in favor of the Portfolio Balance Model. This finding is also important from a risk management perspective, because external trade-oriented companies whose costs and revenues are denominated in more than one currency face exchange rate risk depending on the volatility of the stock returns. Moreover, these results are particularly important for the Czech Republic, Hungary, and Poland, because these countries are part of the European Union and aspire to adopt the euro. Hence, it is important to understand the interactions between the stock exchange and foreign exchange markets in these countries. In addition, these findings may help financial managers and hedgers better understand the dynamic relationship between stock and foreign exchange markets, leading to the construction of better portfolio management strategies.

REFERENCES

Abdalla I, Murinde V (1997): Exchange rates and stock price interactions in emerging financial markets: evidence on India, Korea, Pakistan and the Philippines. *Applied Financial Economics*, 7:25–35.

Aggarwal R (1981): Exchange rates and stock prices: A study of the United States capital markets under floating exchange rates. *Akron Business and Economic Review*, 12:7–12.

Aggarwal R, Inclan C, Leal R (1999): Volatility in emerging markets. *Journal of Financial and Quantitative Analysis*, 34:33–55.

Ajayi RA, Mougoue M (1996): On the dynamic relation between stock prices and exchange rates. *Journal of Financial Research*, 19:193–207.

Andreou E, Ghysels E (2002): Detecting multiple breaks in financial market volatility dynamics. *Journal of Applied Econometrics*, 17:579–600.

Arago V, Salvador E (2011): Sudden changes in variance and time varying hedge ratios. *European Journal of Operational Research*, 215:393–403.

Arago-Manzana V, Fernandez-Izquierdo MA (2007): Influence of structural changes in transmission of information between stock markets: A European empirical study. *Journal of Multinational Financial Management*, 17:112–124.

Bai J, Perron P (1998): Estimating and testing linear models with multiple structural changes. *Econometrica*, 66:47–78.

Bai J, Perron P (2003): Computation and analysis of multiple structural change models. *Journal of Applied Econometrics*, 18:1–22.

Bahmani-Oskooee M, Sohrabian A (1992): Stock prices and the effective exchange rate of the Dollar. *Applied Economics*, 24:459–464.

Bensafta KM (2010): Non-stationary variance and volatility causality. *Economics Bulletin*, 30(4): 2920–2935.

Bodnar GM, Gentry WM (1993): Exchange-rate exposure and industry characteristics: evidence from Canada, Japan and the US. *Journal of International Money and Finance*, 12:29–45.

Caporale GM, Pittis N, Spagnolo N (2002): Testing for causality-in-variance: an application to the East Asian markets. *International Journal of Finance and Economics*, 7:235–245.

Caporale GM, Pittis N, Spagnolo N (2006): Volatility transmission and financial crises. *Journal of Economics and Finance*, 30:376–90.

Cheung YW, Ng LK (1996): A causality-in-variance test and its application to financial market prices. *Journal of Econometrics*, 72:33–48.

Dickey DA, Fuller WA (1979): Distribution of the estimators for autoregressive time series with a unit root. *Journal of the American Statistical Association*, 74:427–431.

Dijk VD, Osborn DR, Sensier M (2005): Testing for causality in variance in the presence of breaks. *Economics Letters*, 89:193–199.

Dornbusch R, Fischer S (1980): Exchange rates and the current account. *American Economic Review*, 12:7–12.

Ewing BT, Malik F (2010): Estimating volatility persistence in oil prices under structural breaks. The *Financial Review*, 45:1011–1023.

Fedorova E, Saleem K (2010): Volatility Spillovers between Stock and Currency Markets: Evidence from Emerging Eastern Europe. *Finance a úvěr-Czech Journal of Economics and Finance*, 60(6):519–533

Fernandez V (2006): The impact of major global events on volatility shifts: Evidence from the Asian crisis and 9/11. *Economic Systems*, 30:79–97.

Franck P, Young A (1972): Stock price reaction of multinational firms to exchange realignments. *Financial Management*, 1:66–73.

Frankel J (1983): Monetary and portfolio-balance models of exchange rate determination. In: Bhandari J, Putman B (Eds): *Economic Interdependence and Flexible Exchange Rates*. MA, MIT Press.

Giovannini A, Jorion P (1987): Interest rates and risk premia in the stock market and in the foreign Exchange market. *Journal of International Money and Finance*, 6:107–124.

Granger CWJ, Huang BN, Yang CW (2000): A Bivariate causality between stock prices and exchange rates: Evidence from recent Asian flu. *Quarterly Review of Economics and Finance*, 40:337–54.

Hafner CM, Herwartz H (2006): A Lagrange multiplier test for causality in variance. *Economics Letters*, 93:137–141.

Hamrita ME, Trifi A (2011) The Relationship between interest rate, exchange rate and stock price: A wavelet analysis. *International Journal of Economics and Financial Issues*, 1(4):220–228

Hillebrand E (2005): Neglecting parameter changes in GARCH models. *Journal of Econometrics*, 129:121–138.

Hong Y (2001): A test for volatility spillover with application to exchange rates. *Journal of Econometrics*, 103:183–224.

Inclan C, Tiao GC (1994): Use of cumulative sums of squares for retrospective detection of changes in variance. *Journal of the American Statistic Association*, 89:913–923.

Islami M (2008): Interdependence between foreign exchange markets and stock markets in selected European Countries. *Universitätsbibliothek Wuppertal*, *Schumpeter Discussion Papers*, sdp08007.

Javed F, Mantalos P (2011): *Sensitivity of the causality in variance test to the GARCH(1;1) parameters*. Available at SSRN: http://ssrn.com/abstract=1856055.

Jones JD (1989): A comparison of lag-length selection techniques in tests of Granger causality between money growth and inflation: evidence for the US, 1959–86. *Applied Economics*, 21: 809–822.

Jorion P (1990): The exchange rate exposure of US multinationals. *Journal of Business*, 63: 331–345.

Kanas A (2000): Volatility spillovers between stock returns and exchange rate changes: international evidence. *Journal of Business Finance and Accounting*, 27:447–467.

Kim K (2003): Dollar exchange rate and stock price: evidence from multivariate cointegration and error correction model. *Review of Financial Economics*, 12:301–313.

Koutmos G (1998): The volatility of interest rates across maturities and frequencies. *Journal of Fixed Income*, 8:27–31.

Krugman PR (1987): Pricing to Market when the Exchange Rate Changes. In: Arndt S, Richardson JD (eds.): *Real Financial Linkages Among Open Economies*. Cambridge, MIT Press.

Kwiatkowski D, Phillips PCB, Schmidt P, Shin Y (1992): Testing the null hypothesis of stationarity against the alternative of a unit root. *Journal of Econometrics*, 54:159–178.

Lamoureux CG, Lastrapes WD (1990): Persistence in variance, structural change and the GARCH model. *Journal of Business and Economic Statistics*, 68:225–234.

Mahmod Q, Yagil J (2012): Fear sentiments and gold price: testing causality in-mean and invariance. *Applied Economics Letters*, 19(4):363–366.

Mansor HI (2000): Cointegration and Granger causality tests of stock price and exchange rate interactions in Malaysia. *ASEAN Economic Bulletin*, 17:36–47.

Mantalos P, Shukur G (2010): The effect of spillover on the Granger causality test. *Journal of Applied Statistics*, 37:1473–1486.

Mok HMK (1993): Causality of interest rate, exchange rate and stock prices at stock market open and close in Hong Kong. *Asia Pacific Journal of Management*, 10:123–143.

Murinde V, Poshakwale S (2004): Exchange rate and stock price interactions in European emerging financial markets before and after the Euro. *SSRN Electronic Journal*, doi: 10.2139/ssrn.493543.

Nelson DB (1991): Conditional heteroskedasticity in asset returns: A new approach. *Econometrica*, 59:347–370.

Nieh CC, Lee CF (2001): Dynamic relationship between stock prices and exchange rates for G-7 Countries. *Quarterly Review of Economics and Finance*, 41:477–490.

Pan MS, Fok RCW, Liu YA (2007): Dynamic linkages between exchange rates and stock prices: Evidence from East Asian markets. *International Review of Economics and Finance*, 16:503–520.

Pantelidis T, Pittis N (2004): Testing for Granger causality in variance in the presence of causality in mean. *Economics Letters*, 85:201–207.

Pardo A, Torro H (2007): Trading with asymmetric volatility spillovers. *Journal of Business Finance & Accounting*, 34:1548–1568.

Phillips PCB, Perron P (1988): Testing for a unit root in time series regression. *Biometrika*, 75:335–346.

Rahman L, Uddin J (2009) Dynamic relationship between stock prices and exchange rates: Evidence from three South Asian Countries. *International Business Research*, 2(2):167–174.

Rapach DE, Strauss JK (2008): Structural breaks and GARCH models of exchange rate volatility. *Journal of Applied Econometrics*, 23:65–90.

Rodrigues PMM, Rubia A (2007): Testing for causality in variance under nonstationarity in variance. *Economics Letters*, 97:133–137.

Roll R (1992): Industrial structure and the comparative behavior of international stock markets indices. *Journal of Finance*, 47:3–41.

Sanso A, Arago V, Carrion JL (2004): Testing for change in the unconditional variance of financial time series. *Revista de Economia Financiera*, 4:32–53.

Shapiro AC (1994): *Foundations of Multinational Financial Management*. 2nd ed. Allyn and Bacon, International Edition.

Smith C (1992): Stock markets and the exchange rate: a multicountry approach. *Journal of Macroeconomics*, 14:607–629.

Solnik B (1987): Using financial prices to test exchange rate models: A note. *Journal of Finance*, 42:141–49.

Sundaram K (2009): *Investigating causal relationship between stock return with respect to exchange rate and FII: evidence from India.* Birla Institute of Technology & Science Pilani. Munich Personal RePEc Archive.

Tsoukalas D (2003): Macroeconomic factors and stock prices in the emerging Cypriot equity market. *Managerial Finance*, 29(4):87–92.

Urbain JP (1989): Model selection criteria and Granger causality tests: An empirical note. *Economics Letters*, 29:317–320.

Wang KM, Thi TBN (2007): Testing for contagion under asymmetric dynamics: Evidence from the stock markets between US and Taiwan. *Physica A*, 376:422–432.

Yang S, Doong S (2004): Price and volatility spillovers between stock prices and exchange rates: Empirical evidence from the G-7 Countries. *International Journal of Business and Economics*, 3:139–153.

Yu Q (1997): Stock prices and exchange rates: Experience in leading East Asian financial centers: Tokyo, Hong Kong and Singapore. *Singapore Economic Review*, 41:47–56.