Volatility Spillover Between Stock Prices and Exchange Rates: New Evidence Across the Recent Financial Crisis Period

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ABSTRACT

We employ an Exponential Generalised Autoregressive Conditional Heteroskedasticity (EGARCH) model to examine the volatility spillover effects between stock prices and exchange rates in three developed and three emerging countries, across the recent pre-financial-crisis, crisis and post-crisis periods. The evidence indicates asymmetric volatility spillover effects between stock prices and exchange rates in both developed and emerging economies during the financial crisis. The findings of the significant volatility spillover effects between exchange rates and stock prices imply that the markets are informationally inefficient, and one market has significant predictive power on the other.

1. INTRODUCTION

OVER THE PAST THREE DECADES, the relationship between stock prices and exchange rates has received considerable attention in the literature (e.g., Aggarwal, 1981; Bahmani-Oskooee and Sohrabian, 1992; Ajayi and Mougoue, 1996; Ajayi *et al.*, 1998; Granger *et al.*, 2000; Apte, 2001; Nieh and Lee, 2001; Phylaktis and Ravazzolo, 2005; Antell and Vaihekoski, 2007; Lin, 2012; Moore and Wang, 2014; Walid and Duc Khuong, 2014). Much of this literature — which concentrated on the 'first moment' co-movement, the cointegration properties, and the direction of causality of the variables in question — confirmed a significant relationship between stock prices and exchange rates.

In more recent years, as a result of increased cross border capital movements, greater financial integration and various financial and foreign exchange crises around the globe, research attention has focused on the investigation of the 'volatility' spillover effect (the second moment) between stock and foreign exchange markets (see, inter alia, Kanas, 2000; Yang and Doong, 2004; Wu, 2005; Antell and Vaihekoski, 2007; Mishra *et al.*, 2007; Raghavan and Dark, 2008; Giannellis *et al.*, 2010; Walid *et al.*, 2011; Kumar, 2013). The analysis of volatility spillovers is recognised to provide particularly useful insights into how information is transmitted from stock markets to foreign exchange markets and vice versa, and hence on the risk of 'contagion' effects between stock and exchange rate markets (Mishra *et al.*, 2007; Moore and Wang, 2014). Existing research generally supports the hypothesis of interdependence in volatility of stock and foreign exchange markets.

Research efforts aimed at a better understanding of the transmission mechanism of the 'volatility' in stock and foreign exchange returns, and of how information is transmitted from one market to the other, are understandable given the importance of these phenomena for option pricing, hedging, risk management, portfolio selection and regulatory policy formulation. For example, if the volatility transmits from one asset to another, both assets cannot be included in the same portfolio in order to diversify risk. Improved knowledge about the quality and quantity of the volatility spillover across stock and foreign exchange markets, in turn, can aid better decision making for portfolio managers, multinational firms, investors and policy makers alike. The importance of investigating volatility spillovers is, therefore, self-evident.

Volatility is typically defined as a measure of dispersion of returns of an asset or market index. Generally, the higher the volatility, the riskier the asset. A number of studies (e.g., Giannellis *et al.*, 2010) find that the conditional volatility of asset prices responds asymmetrically to innovations (good or bad news). In the context of volatility transmission between exchange rates and stock prices, Kanas (2000), Apte (2001), Yang and Doong (2004), and Walid *et al.* (2011) find significant asymmetric volatility spillovers between asset markets.

Many economic theories such as the *Balance of Payment Approach* (Dornbusch and Fischer, 1980) and the *Portfolio Approach* (Branson, 1983; Frankel, 1983) suggest that there is interdependence between the volatility of stock and foreign exchange markets. Moore and Wang (2014) argue that the sign, size and direction of this interdependence depends upon the financial integration and market efficiency of the country, and whether the economy is export or import dominant. Most of the existing research on the inter-temporal dependence of the volatility of asset markets tends to focus on developed countries (e.g., Yang and Doong, 2004; Raghavan and Dark, 2008), or on individual developing countries (e.g., Mishra *et al.*, 2007), with less attention being paid to the relationship in question from the perspective of both emerging and developed countries. This is striking when considering that country specific variables such as the degree of capital mobility, trade volumes and the interrelationship between macroeconomic variables may also impact the volatility transmission between foreign exchange and stock markets.

Motivated by these arguments, this study examines the interrelationship between volatility of stock and foreign exchange markets within three developed countries (Ireland, the Netherlands and Spain) and three emerging economies (Brazil, South Africa and Turkey). The specific questions we address are whether volatility in one market increases volatility in the other (within each country), and whether the impact is the same for both positive and negative shocks of the same magnitude.

Ireland, the Netherlands and Spain are selected because whilst they are members of the eurozone — the biggest economic zone in the world — they have been overlooked in previous research examining volatility spillovers between stock and foreign exchange markets. Brazil, South Africa and Turkey are chosen because they are three of the largest emerging and developing economies (in three different continents, hence less likely to suffer from cross-border ‘contagion’ effects stemming from shocks in economic fundamentals) by either nominal or inflation-adjusted GDP (IMF, 2012), have growing financial markets, and — unlike other large emerging economies such as China and India — their foreign exchange markets (and currencies) are not ‘managed’, at least officially.

The latter justification for our selection of countries is particularly important given the scope of our study, since despite growing interest towards BRICS economies (see, for example, Walid and Duc Khuong, 2014), countries such as China and India have, for decades, managed foreign exchange market pressures and episodes of exchange rate volatility through a combination of monetary and regulatory measures, along with direct and indirect interventions in line with their declared exchange rate regime.[2]

Another contribution of our study lies in the adoption of a disaggregated framework that distinguishes between pre-crisis, crisis and post-crisis periods, in order to ascertain the extent to which the recent financial crisis affected the relationship in question, an aspect which has not been given any attention in prior work for the countries considered by the present study.

2. A brief review of the literature

Theoretical models examining the relationship between exchange rates and stock prices draw from both Traditional and Portfolio approaches. Traditional or 'flow-oriented' models of exchange rates (e.g., Dornbusch and Fischer, 1980) focus on the current account or trade balances. These models posit that movements in exchange rates affect the competitiveness and profitability of firms, resulting in consequent changes in stock prices. For example, a depreciation of the exchange rate results in an increase in local firms' exports, leading to an increase in stock prices. By contrast, an exchange rate appreciation causes a negative change in the balance of trade which, in turn, leads to a decrease in stock prices. According to such models, therefore, the exchange rate is the lead variable, determining a positive correlation between exchange rates and stock prices.

On the other hand, Portfolio or Asset Market approaches (Branson, 1983; Frankel, 1983) postulate that it is capital flows, not trade flows, which determine exchange rates. In these models, exchange rates are determined by demand and supply of domestic assets (stocks and bonds). An increase in stock prices leads to an increase in demand for domestic assets, resulting in an appreciation of the domestic currency. By contrast, a decrease in stock prices causes exchange rates to depreciate. Here, the stock price is the lead variable and the postulated relationship is negative.

Turning to the empirical literature, early work on the relationship between exchange rates and stock prices tended to focus on the first moment (mean) of the distribution of asset returns. Ajayi and Mougoue (1996) find a significant short- and long-term relationship between the variables in eight developed countries. They suggest that an increase in stock prices results in an increase in exchange rates in the USA and the UK. The relationship is explained in terms of economic agents treating rising stock prices as a sign of higher inflation expectations which, in turn, leads to a lower demand for local currency, with the latter depreciating as a result. In stark contrast, Aggarwal (1981) finds a negative correlation between US stock prices and trade-weighted US dollars. This result is rationalised in terms of a decline in stock prices pushing foreign investors to sell their assets denominated in local currency, leading to a depreciation of the currency. Solnik (1987) too finds a negative relationship between stock prices and exchange rates. However, Bahmani-Oskooee and Sohrabian (1992), Granger *et al.* (2000) and Nieh and Lee (2001) argue in favour of the existence of bi-directional causality in the relationship.

A few studies have examined the second moment (variance) distribution of the variables. For example, Kanas (2000), Yang and Doong (2004), and Raghavan and Dark (2008), analyse the volatility spillover effects between stock and foreign exchange markets in developed countries, while Apte (2001), Wu (2005), Walid *et al.* (2011) and Kumar (2013) concentrate on emerging countries.

It is noteworthy that there is a dearth of literature that has investigated volatility spillover effects from both the developed and emerging countries' perspective using the same methodology and time span. This is an important shortcoming because, as argued by Moore and Wang (2014), the sources and nature of the linkage between exchange rates and stock prices in developed economies are different from those in emerging economies.

Kanas (2000) examines the volatility spillover between stock and foreign exchange rate markets in the USA, the UK, Canada, Japan, Germany and France, using daily data and employing a bivariate Exponential Generalised Autoregressive Conditional Heteroskedasticity (EGARCH) model. The study concludes that there is a significant symmetric volatility spillover from stock returns to exchange rate changes for all countries except Germany. However, the volatility spillovers from exchange rate changes to stock returns were not significant for any of the countries examined. One possible explanation for the result of an insignificant spillover is that the use of daily data cannot capture the effect of trade flows on exchange rate changes.

Apte (2001) examines the volatility spillover effect between stock and foreign exchange markets in India using a bivariate EGARCH model and daily data from 02 January 1991 to 24 April 2000. Contrary to the findings of Kanas (2000), the study finds a significant volatility spillover from the foreign exchange market to the stock market and an insignificant volatility spillover from the stock market to the foreign exchange market. Mishra *et al.* (2007) point out that the main limitation of this study is that it generated the data on stock indices from 1991 to 1994 by simulation. By using an EGARCH model and by extending the sample period, Mishra *et al.* (2007) find a bi-directional volatility spillover between stock and foreign exchange markets in India.

Yang and Doong (2004) investigate the mean and volatility spillover from one market to another in G-7 countries using weekly (Friday closing) exchange rates and stock market indices. Employing a bivariate EGARCH model, they find an asymmetric spillover effect from stock markets to foreign exchange markets in France, Italy, Japan and the USA. However, the volatility in exchange rates is found to have a less pronounced impact on the volatility in stock prices.

Wu (2005) examines the volatility spillover effects between stock and foreign exchange markets in Japan, South Korea, Indonesia, Philippines, Singapore, Thailand and Taiwan for the period 1997-2000, splitting the data into pre-Asian financial crisis and post-crisis periods. The study finds a bi-directional spillover effect between the variables during the recovery period in almost all countries.

Raghavan and Dark (2008) find a significant unidirectional return and volatility spillover from the USD/AUD exchange rates to Australian All Ordinaries Index (AOI) by using daily data on the USD/AUD and the Australian All Ordinaries Index (AOI) from 2 January 1995 to 31 December 2004. The major drawback of this study is that it ignores the possibility of volatility asymmetry in financial asset prices, i.e. that positive and negative shocks may induce a different degree of volatility.

Walid *et al.* (2011) examine the impact of exchange rate changes on stock market volatility in four emerging markets (Hong Kong, Singapore, Malaysia and Mexico). Using weekly data and employing a two regime Markov-Switching EGARCH model, they find evidence of regime switching behaviour in volatility of emerging countries' stock markets, and that the relationship between exchange rates and stock prices is regime dependent. Their evidence also indicates that the volatility in foreign exchange markets spills over to stock markets asymmetrically.

The most recent results by Kumar (2013) are of particular interest since his bivariate GARCH model estimations suggest — in addition to the existence of bi-directional volatility spillover between stock and foreign exchange markets in India, Brazil and South Africa (IBSA countries) — that stock markets play a relatively more important role than foreign exchange markets in the first and second moment interactions and spillovers.

Moore and Wang (2014) focus on investigating the sources of the dynamic linkage between real exchange rates and stock returns in relation to the US market for developed and emerging Asian markets, and find that the trade balance is the main determinant of the dynamic correlation for Asian markets, whereas the interest rate differential is the driving force for the developed markets.

It is clear from the above review of relevant literature that the findings on volatility spillover effects between stock prices and exchange rates are rather mixed, with no clear consensus from which to discern a conventional wisdom. The results vary from country to country, also depending on the methodology adopted as well as the quantity, quality and time span of data employed. There is, therefore, value in investigating further the effects of volatility spillovers between exchange rates and stock prices, particularly by comparing the experience of a sample of both developed and emerging markets within the same framework, across pre-crisis, crisis and post-crisis periods using the same methodology and time span, and an up-to-date dataset.

3. Econometric approach and data description

Prior to undertaking EGARCH estimations, our preliminary testing phase entailed assessing the integration and cointegration properties of the series. For the former we employ the Exponential Smooth Transition Autoregressive (ESTAR) non-linear unit root test by Kapetanios, Shin and Snell (2003), also known as KSS unit root test. With regard to the latter, following Yau and Nieh (2009), we use the threshold cointegration test elaborated by Enders and Granger (1998) and Enders and Siklos (2001), which hinges on a residual-based two-stage estimation procedure. The first stage regression is given by:

$$Y_{1t} = \alpha + \beta Y_{2t} + \mu_t \quad (1)$$

where Y_{1t} and Y_{2t} are two I(1) series and u_t is the error. The second stage focuses on the test of stationarity of the error terms:

$$\Delta\mu_t = I_t \rho_1 \mu_{t-1} + (1 - I_t)\rho_2 \mu_{t-1} + \sum_{i=1}^{l} \gamma_i \Delta\mu_{t-1} + \varepsilon_t \quad (2)$$

where ε_t is a white-noise disturbance, and the residual u_t, is extracted from equation (1). The term I_t is the Heaviside indicator function where $I_t = 1$ if $u_{t-1} \geq \tau$ and $I_t = 0$ if $u_{t-1} \leq \tau$ and τ is the threshold value, which constitutes a Threshold Autoregressive (TAR) model of the error. Assuming the system is convergent, $u_t = 0$ can be considered as the long-run equilibrium value (threshold value) of the sequence. If u_t is above its long-run equilibrium, then the adjustment is $\rho 1 u_{t-1}$. However, if u_t is below its long-run equilibrium, then the adjustment is $\rho 2 u_{t-1}$. The null hypothesis is $\rho 1 = \rho 2 = 0$ and thus rejection implies the existence of a threshold cointegrating relationship. If a cointegrating relationship is found then the null of symmetric adjustment (Ho: $\rho 1 = \rho 2$) can be tested. Rejection of both the null hypotheses ($\rho 1 = \rho 2 = 0$ and $\rho 1 = \rho 2$) implies the existence of threshold cointegration and asymmetric adjustment.

In equation (2), the decay could also be allowed to depend on Δu_{t-1}. The Heaviside indicator could then be $I_t = 1$ if $\Delta u_{t-1} \geq \tau$ and $I_t = 0$ if $\Delta u_{t-1} \leq \tau$. This model is especially valuable when the adjustment is asymmetric, such that the series exhibits more 'momentum' in one direction than the other (Enders and Granger, 1998), hence the name Momentum Threshold Autoregressive (M-TAR) model. The Akaike Information Criterion (AIC) is used in order to select between TAR and M-TAR models.

We also employ the Momentum Threshold Error Correction Model (M-TECM) as elaborated by Enders and Granger (1998) and Enders and Siklos (2001), to examine Granger-causality. The M-TECM is expressed as:

$$\Delta Y_{it} = \alpha + \gamma_1 Z_{t-1}^{+} + \gamma_2 Z_{t-1}^{-} + \sum_{i=1}^{k_1} \delta_i \Delta Y_{1,t-i} + \sum_{i=1}^{k_2} \theta_i \Delta Y_{2,t-i} + \upsilon_t \qquad (3)$$

where Y_{it} denotes the variables concerned, $Z_{t-1}^{+} = I_t \Delta \hat{\mu}_{t-1}$ and $Z_{t-1}^{-} = (1 - I)_t \Delta \hat{\mu}_{t-1}$, and υ_t is a white-noise disturbance.

The Granger-causality test is based on a standard F-test of whether all the coefficients of $\Delta Y_{1,t-i}$ or $\Delta Y_{2,t-i}$ are jointly statistically different from zero, and/or whether the γ_j coefficients of the error-correction term are significant. Rejection of the null hypotheses Ho: $\delta_1 = \delta_2 = 0$ and Ho: $\gamma_1 = \gamma_2 = 0$ indicates a short-term causality, rejection of the null hypotheses Ho: $\delta_1 = \delta_2 = \gamma_1 = 0$, Ho: $\theta_1 = \theta_2 = \gamma_2 = 0$, indicates a long-term causality, and rejection of the null hypotheses Ho: $\gamma_2 = \gamma_1 = 0$ and Ho: $\gamma_2 = \gamma_1$ indicates an asymmetric mean reversion process. Since the Granger-causality test is very sensitive to the selection of lag length, the AIC is used to determine the appropriate lag length.

Once the non-linear integration and cointegration properties of the series are established, and the causality tests performed, we then employ the multivariate extension (specifically, the bivariate extension) of the EGARCH model proposed by Nelson (1991). This model has some advantages over other GARCH models. For example, there is no need to impose a non-negativity constraint on model parameters since the log of the conditional variance is modelled, thus allowing for the detection of asymmetric effects in volatility. The model specification is as follows.

Mean equations:

$$S_t = \theta_{s,0} + \sum_{i=1}^{k} \theta_{s,i} S_{t-i} + \sum_{i=1}^{k} \theta_{e,i} E_{t-i} + \alpha_s \lambda_{s,t-1} + \varepsilon_{s,t} \qquad (4)$$

$$E_t = \theta_{e,0} + \sum_{i=1}^{k} \theta_{e,i} E_{t-i} + \sum_{i=1}^{k} \theta_{s,i} S_{t-i} + \alpha_e \lambda_{e,t-1} + \varepsilon_{e,t} \quad (5)$$

Conditional variance equations:

$$\sigma_{s,t}^2 = \exp\left\{\alpha_{s,o} + \sum_{i=1}^{k} \beta_{s,i} \log(\sigma_{s,t-i}^2) + \delta_{s,s}\left[\left(\left|Z_{s,t-1}\right| - E\left|Z_{s,t-1}\right|\right) + \pi_{s,s} Z_{s,t-1}\right] + \delta_{s,e}\left[\left(\left|Z_{e,t-1}\right| - E\left|Z_{e,t-1}\right|\right) + \pi_{s,e} Z_{e,t-1}\right]\right\} \quad (6)$$

$$\sigma_{e,t}^2 = \exp\left\{\alpha_{e,o} + \sum_{i=1}^{k} \beta_{e,i} \log(\sigma_{e,t-i}^2) + \delta_{e,e}\left[\left(\left|Z_{e,t-1}\right| - E\left|Z_{e,t-1}\right|\right) + \pi_{e,e} Z_{e,t-1}\right] + \delta_{e,s}\left[\left(\left|Z_{s,t-1}\right| - E\left|Z_{s,t-1}\right|\right) + \pi_{e,s} Z_{s,t-1}\right]\right\} \quad (7)$$

$$\sigma_{s,e,t} = \rho_{s,e,t} \sigma_{s,t} \sigma_{e,t} \quad (8)$$

$$\sigma_{e,s,t} = \rho_{e,s,t} \sigma_{e,t} \sigma_{s,t} \quad (9)$$

In equations (4) and (5) we examine the 'return' spillover effect between exchange rate changes (E) and stock market returns (S). Here, α is the coefficient of error correction terms (λ) and θ measures the (return) spillover effect. A lag length equal to two is selected for both E and S on the basis of the AIC.

In equations (6) and (7) the conditional variance in one market depends on its own lags and cross market standardised innovations. The persistence of volatility is measured by β. The volatility spillover effect between foreign exchange and stock markets is captured by the coefficient δ. If $\delta_{s,e}$ is significantly different from zero, then the volatility of exchange rates spills over to volatility of stock prices. Asymmetric impact is measured by the coefficient π. Asymmetry exists if π is negative and significantly different from zero. A positive and significant δ alongside a negative and significant π implies that negative shocks in one market have a larger impact on the volatility of the other market than positive shocks of the same absolute value. In other words, 'bad news' has greater impact on volatility than 'good news'.

We use weekly (Wednesday closing) spot exchange rates (local currency per US Dollar) and weekly stock price indices for Brazil, Ireland, the Netherlands, Spain, South Africa and Turkey. The data are obtained from Datastream. Weekly data are used because daily data contain too much noise. For example, some studies suggest that there is a strong Wednesday effect in stock markets, other confirm that Friday data contain weekend effects, and monthly data cannot capture the short-term dynamic relationship between exchange rates and stock prices (see Walid *et al.*, 2011).

The stock indices used in the study consist of Brazil BOVESPA, Ireland SE Overall (ISEQ), Amsterdam SE All Share, IBEX 35 (Spain), FTSE/JSE All Shares (South Africa) and ISE National All Share (Turkey). These are the main stock indices of the respective countries. The exchange rates are USD-euro, USD-Brazilian Real, USD-South African Rand and USD-Turkish Lira. Following the standard rationale provided in the relevant literature, we use nominal exchange rates, in that short-term investors are not worried about inflation effects as they do not buy goods in the basket used to construct the inflation rate.

The full sample period covers 03/01/2001 to 26/12/2012, yielding 626 observations. The period is justified by the major growth in international financial liberalisation, financial integration and foreign direct investment in the 2000s; a representative decade to measure the short time dynamic relationship between stock and foreign exchange markets. Furthermore, the sample period allows us to investigate the relationship between exchange rate movements and stock market volatility during 'good' and 'bad' times.

The global financial crisis of 2007-2008, resulted in the downturn in stock markets around the world. The active phase of the crisis can be dated from August 7, 2007, when BNP Paribas terminated withdrawals from three hedge funds recording a complete evaporation of liquidity (IMF, 2012). The downturn in stock markets around the world continued until March 2009 (De Vita and Abbott, 2011). From August 2007 to March 2009, the stock indices in Ireland, the Netherlands and Spain decreased by 75, 55 and 45 per cent, respectively. Brazilian, South African and Turkish stock indices fell by 24, 25 and 50 per cent, respectively, during the crisis period. Brazilian Real, South African Rand and Turkish Lira depreciated by 19, 32 and 28 per cent over that period, whereas the euro appreciated by 1 per cent against the US Dollar during the same period.

Following the individual split points typically adopted in the literature that also uses weekly data (e.g., Mozumder *et al.*, 2015) to examine the start of the most recent stock market crisis (01/08/2007) and its stabilisation date (01/04/2009), we divide the full sample into three sub-periods. The pre-crisis period is from 03/01/2007 to 25/07/2007, yielding 343 observations. The crisis period is from 01/08/2007 to 25/03/2009, totalling 87 observations, and the post-crisis period is from 01/04/2009 to 26/12/2012, yielding 196 observations.

The weekly return series are calculated as $Rt = \ln (P_t / P_{t-1})$, where P_t is the weekly price at time t. The plots of *stock prices* and *exchange rate* series, for all countries in our sample (not reported to conserve space), show that volatility occurs in bursts.

4. A FIRST PASS AT THE DATA

Descriptive statistics of weekly return series of stock indices for the six countries are reported in Table 1. Panel A shows descriptive statistics for the full sample (January 2001 to December 2012). The mean returns in the developed stock markets (Ireland, the Netherlands and Spain) were negative, whereas the mean returns in the emerging stock markets (Brazil, South Africa and Turkey) were positive. This reveals that stock markets in the three emerging countries performed better than in the developed countries during the full sample period. Stock market volatility in the emerging markets, especially in Turkey and Brazil, was higher than the volatility in the developed countries' stock markets. During the full sample period, the highest return and volatility was in the Turkish stock market, and the lowest return and volatility was in the Spanish stock market. This suggests that there was a trade-off between risk and return.

Table 1: Descriptive statistics of weekly return of stock indices

	Ireland	*Netherlands*	*Spain*	*Brazil*	*South Africa*	*Turkey*
	Panel A. Full sample. January 2001 - December 2012					
Mean	-0.0021	-0.0020	-1.33E-06	0.0020	0.0023	0.0026
SD	0.0335	0.0333	0.0301	0.0374	0.0293	0.0550
Skewness	-0.7580	-0.6838	-0.4181	-0.6929	-0.1352	-0.6611
Kurtosis	7.4641	8.2071	5.498013	6.8087	4.8488	7.5508
J-B	386.19**	503.61**	120.57**	427.79**	60.66**	390.21**
	Panel B. Pre-crisis period. January 2001 - July 2007					
Mean	0.0012	-0.0004	0.0014	0.0035	0.0037	0.0052
SD	0.0250	0.0348	0.0275	0.0351	0.0259	0.0560
Skewness	-0.6854	-0.3433	-0.3661	-0.2290	-0.2761	-0.5811
Kurtosis	4.7493	10.7431	5.3807	4.3779	4.4710	8.3066
J-B	70.38**	861.09**	88.41**	30.04**	35.18**	420.54**
	Panel C. Crisis period. August 2007 - March 2009					
Mean	-0.0021	-0.0098	-0.0068	-0.0030	-0.0031	-0.0081
SD	0.0607	0.0435	0.0413	0.0543	0.0425	0.0587
Skewness	0.1223	-0.5655	-0.1712	-1.1295	0.2732	-0.6117
Kurtosis	3.5947	3.8736	3.7650	6.9693	3.5599	3.7597
J-B	51.48**	57.31**	62.51**	74.74**	52.19**	7.4324**
	Panel D. Post-crisis period. April 2009 - December 2012					
Mean	0.0019	0.0022	0.0002	0.0019	0.0032	0.0056
SD	0.0300	0.0286	0.0368	0.0319	0.0233	0.0343
Skewness	-0.5478	-0.3134	-0.4181	-0.2839	-0.1507	-1.1018
Kurtosis	5.5974	4.7736	3.5134	2.9967	3.8025	8.8428
J-B	64.57**	28.75**	32.19**	22.62**	45.97**	316.83**

Note: SD and J-B denote standard error and Jarque-Bera, respectively. Jarque-Bera is the test of the null hypothesis that the residuals are normally distributed. ** denotes the level of significance at 5%. The statistic has a χ^2 distribution with two degrees of freedom. The critical value at the 5% level of significance is 5.9.

Panel B of Table 1, which presents descriptive statistics of the pre-crisis period, reveals that all countries except the Netherlands had positive stock returns and the volatility in stock markets was lower in most countries compared with the full sample period. Panel C shows that stock returns were negative and the volatility was high in all countries during the financial crisis. However, stock markets exhibited positive returns and low volatility during the recovery period, as indicated by the data presented in Panel D.

Table 2: Descriptive statistics of weekly changes in exchange rates

	Ireland	*Netherlands*	*Spain*	*Brazil*	*South Africa*	*Turkey*
	Panel A. Full sample. January 2001 - December 2012					
Mean	-0.0009	-0.0009	-0.0009	0.0001	0.0005	0.0019
SD	0.0138	0.0138	0.0138	0.0242	0.0262	0.0292
Skewness	-0.4445	-0.4445	-0.4445	1.6546	1.0295	4.0427
Kurtosis	8.9054	8.9054	8.9054	18.2200	9.3754	36.1597
J-B	619.67**	619.67**	619.67**	6317**	779.89**	20240**
	Panel B. Pre-crisis period. January 2001 - July 2007					
Mean	-0.0010	-0.0010	-0.0010	-0.0001	-0.0002	0.0018
SD	0.0125	0.0125	0.0125	0.0237	0.0231	0.0278
Skewness	0.1156	0.1156	0.1156	1.0265	0.6578	5.0391
Kurtosis	3.3265	3.3265	3.3265	10.820	5.4879	47.733
J-B	6.28**	6.28**	6.28**	931.66**	112.87**	2996**
	Panel C. Crisis period. August 2007 - March 2009					
Mean	0.0001	0.0001	0.0001	0.0020	0.0032	0.0028
SD	0.0198	0.0198	0.0198	0.0379	0.0360	0.0337
Skewness	-0.3618	-0.3618	-0.3618	2.0061	1.1948	1.5062
Kurtosis	6.0261	6.0261	6.0261	15.1398	8.2427	11.202
J-B	34.69**	34.69**	34.69**	585.78**	118.95**	273.63**
	Panel D. Post-crisis period. April 2009 - December 2012					
Mean	1.09E-05	1.09E-05	1.09E-05	-0.0005	-0.0004	0.0004
SD	0.0141	0.0141	0.0141	0.0163	0.0217	0.0139
Skewness	0.1165	0.1165	0.1165	0.3537	0.2585	-0.0406
Kurtosis	3.5630	3.5630	3.5630	4.1891	3.4195	2.9917
J-B	6.04**	6.04**	6.04**	15.55**	53.60**	54.26**

Note: See Table 1.

Table 2 exhibits descriptive statistics for weekly changes in exchange rates. Panel A of Table 2, covering the full sample period, reveals that the euro appreciated, and Brazilian Real, South African Rand and Turkish Lira depreciated against the US Dollar. However, the volatility in emerging foreign exchange markets was higher than that experienced in developed economies' markets.

As Panels B and C indicate, all exchange rates, except the Turkish Lira in the pre-crisis period, appreciated in 'good times' and depreciated in 'bad times'. Again, there was higher volatility in all foreign exchange markets during the financial crisis compared with the pre-crisis period. However, the volatility decreased gradually during the recovery period, as shown in Panel D.

It is worth highlighting that all stock return series (except for Ireland during the crisis period) are negatively skewed, indicating that the distributions have long left tails, while all the series of exchange rate changes (except for the US Dollar and the euro during the full sample and crisis period), are positively skewed, indicating that most of the distributions have long right tails. Unsurprisingly, the coefficients of kurtosis for all series are greater than three, indicating that they are leptokurtic in nature (as confirmed by the Jarque-Bera test statistics, which reject the normality hypothesis for all series at the customary five per cent significance level).

The test for an ARCH effect indicates that the squared residuals are autocorrelated in all series. Therefore, the weekly returns of both series all have common characteristics (typical of financial data) such as volatility clustering and leptokurtosis, which means that the volatility of all series can be modelled by GARCH type models.

5. Results

The ESTAR (KSS) unit root tests reveal that the time series are integrated of order (1) in (log) levels but stationary in their first differences. Hence, the series can be used to test for a cointegrating relationship.

Table 3: KSS unit root test results

	Ireland	*Netherlands*	*Spain*	*Brazil*	*South Africa*	*Turkey*
			Stock indices			
		Panel A. Full sample. January 2001 - December 2012				
Level	.57	-1.23	-1.05	-1.43	-1.11	-1.80
First difference	-10.50**	-3.91**	-11.23**	-9.35**	-4.70**	-5.05**
		Panel B. Pre-crisis period. January 2001 - July 2007				
Level	-0.15	-.53	-0.73	0.92	0.39	-0.71
First difference	-21.30**	-21.12**	-10.10**	-12.52**	-10.37**	19.30**
		Panel C. Crisis period. August 2007 - March 2009				
Level	0.10	0.98	-0.72	-0.24	-0.37	-0.41
First difference	-14.44**	-7.32**	-3.78**	-2.72**	-5.20**	-3.73**
		Panel D. Post-crisis period. April 2009 - December 2012				
Level	-1.10	-1.20	-0.94	-1.58	-1.23	-1.90
First difference	-15.43**	-19.04**	-14.24**	-13.59**	-10.87**	-5.05**

cont...

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	Exchange rates					
	Panel A. Full sample. January 2001 - December 2012					
Level	-1.22	-1.00	-1.19	-1.75	-1.11	-1.17
First difference	-12.12**	-9.17**	-31.21**	-20.44**	-12.48**	-10.57**
	Panel B. Pre-crisis period. January 2001 - July 2007					
Level	-0.98	-0.79	-0.23	-1.11	-1.23	-1.90
First difference	-10.70**	-14.72**	-10.93**	-4.10**	-10.93**	-5.15**
	Panel C. Crisis period. August 2007 - March 2009					
Level	-1.24	-1.70	-1.22	-1.60	-1.06	-1.86
First difference	-19.25**	-13.07**	-17.09**	-11.22**	-11.07**	-4.79**
	Panel D. Post-crisis period. April 2009 - December 2012					
Level	-1.34	-1.41	-1.09	-1.08	-1.13	-1.80
First difference	-10.22**	-10.99**	-10.22**	-9.04**	-3.22**	-9.41**

Note: KSS denotes Kapetanios, Shin and Snell (2003). The null hypothesis is that the series has a unit root. Lag lengths are selected on the basis of the Modified Akaike Information Criterion. ** denotes the level of significance at 5%. The critical value for the KSS tests at the 5% level is -1.92.

The results of the threshold cointegration test are reported in Table 4. For the full sample period (Panel A), we find cointegrating relationships between stock prices and exchange rates for Brazil and Turkey (Figures 1 and 2 display the plots of the residuals of the cointegrating relationships in question). However, the variables do not cointegrate in the case of Ireland, the Netherlands, Spain and South Africa. The results justify the inclusion of the error correction terms in the mean equations (4) and (5) for Brazil and Turkey in the full sample period.

For the pre-crisis (Panel B), the variables are cointegrated only in the case of Brazil. We find statistically insignificant F statistics for all countries during the period of the financial crisis (Panel C) and the post-crisis (recovery) period (Panel D). Hence, error correction terms are included in mean equations (4) and (5) for Brazil in the pre-crisis period.

The preceding results show that there is no long run relationship between stock prices and exchange rates in developed countries. This finding is consistent with those by Granger *et al.* (2000), Nieh and Lee (2001) and Yang and Doong (2004). However, stock prices and exchange rates in emerging countries are found to co-move at least in one sample period. This indicates that there is a long term relationship between the variables in emerging countries, which is in line with the findings of Apte (2001) and Wu (2005). It may be that other factors, such as the trade balance, cause exchange rates and stock prices in emerging markets to move together, as suggested by Moore and Wang (2014).

Table 4: Threshold cointegration test results

	Ireland		*Netherlands*		*Spain*		*Brazil*		*South Africa*		*Turkey*	
	TAR	M-TAR	TAR	M-TAR	TAR	M-TAR	TAR	M-TAR	TAR	M-TAR	TAR	M-TAR
Panel A. Full sample. January 2001 - December 2012												
$\hat{\rho}_1$	-0.03**	-0.11**	0.08	-0.05**	-0.09	0.07**	-0.08**	0.05**	-0.01**	0.01	0.05**	-0.03**
$\hat{\rho}_2$	0.02	-0.12	-0.01**	-0.10	-0.02**	-0.03	0.01	-0.02	-0.23**	-0.03	-0.03**	-0.02
$\hat{F}_A$	1.29	0.30	1.79	0.98	4.09**	1.16	4.89**	4.75**	4.16**	1.85	5.12**	4.05**
$\hat{t}_A$	4.26**	1.35	0.24	3.35*	1.19	0.27	1.23	4.77**	1.57	0.91	0.20	5.15**
AIC	8.13	**35.20**	9.30	**40.32**	10.12	**41.50**	20.14	**39.31**	22.58	**37.80**	15.50	**35.81**
Panel B. Pre-crisis period. January 2001 - July 2007												
$\hat{\rho}_1$	0.10	-0.01**	-0.13**	-0.03**	-0.8	-0.10**	-0.04**	0.02**	-0.01**	-0.02	-0.07	-0.02**
$\hat{\rho}_2$	0.05	-0.02	0.02	-0.02	-0.05**	-0.11	-0.10	0.12	-0.11	0.04	-0.02**	-0.01
$\hat{F}_A$	4.03**	1.33	3.79**	1.06	4.22**	0.16	1.85	3.32**	1.16	1.15	1.82	0.73
$\hat{t}_A$	1.52	1.30	3.21**	1.30	1.10	4.27**	1.22	4.70**	1.21	0.54	0.25	1.43
AIC	20.12	**41.20**	8.51	**35.79**	10.37	**32.50**	12.10	**24.20**	11.20	**43.81**	10.90	**37.34**
Panel C. Crisis period. August 2007 - March 2009												
$\hat{\rho}_1$	-0.10**	0.07**	-0.01**	-0.01**	-0.05	-0.02**	-0.05	0.03	-0.02	-0.07	0.02	-0.01
$\hat{\rho}_2$	-0.05	-0.08	-0.02	0.05	-0.02**	-0.01	-0.11**	-0.02**	0.31	-0.01	-0.03**	-0.05**
$\hat{F}_A$	1.02	1.35	3.70**	1.16	4.49**	1.18	3.75**	1.89	0.81	1.23	1.02	1.07
$\hat{t}_A$	5.26**	4.35**	0.89	1.35	1.10	1.27	0.30	1.24	1.18	1.32	1.31	3.05**
AIC	8.70	**30.21**	35.50	**35.71**	10.30	**30.15**	17.31	**40.94**	27.70	**31.99**	13.10	**41.00**
Panel D. Post-crisis period. April 2009 - December 2012												
$\hat{\rho}_1$	-0.05**	-0.04**	-0.08**	-0.05**	-0.07 **	-0.05**	-0.03**	-0.11**	-0.02	-0.03**	-0.12	-0.10
$\hat{\rho}_2$	-0.04	-0.01	-0.02**	-0.03	-0.01**	-0.02	-0.05	-0.02	0.13**	-0.10	-0.13**	0.01
$\hat{F}_A$	3.03**	1.35	3.79**	3.16	1.42	1.16	1.19	1.15	0.57	1.19	1.32	1.13
$\hat{t}_A$	1.26	1.35	0.24	3.35**	1.19	3.27**	1.20	0.10	1.17	1.52	1.89	0.24
AIC	-28.72	**31.22**	25.56	**25.70**	30.34	**32.57**	12.52	**40.10**	14.19	**41.19**	10.40	**45.20**

Note: TAR and M-TAR denote Threshold Autoregressive and the Momentum Threshold Autoregressive model, respectively. Lag-length (l = 2) is selected on the basis of the Ng and Perron (2001) unit root procedure. $\hat{F}_A$ and $\hat{t}_A$ denote the F- and t-statistics for the null hypothesis of no cointegration and symmetry, respectively. The model is specified based on the Akaike Information Criterion (AIC). ** denote the level of significance at 5%.

Figure 1: Plot of residuals of cointegrating relationship for Brazil (full sample)

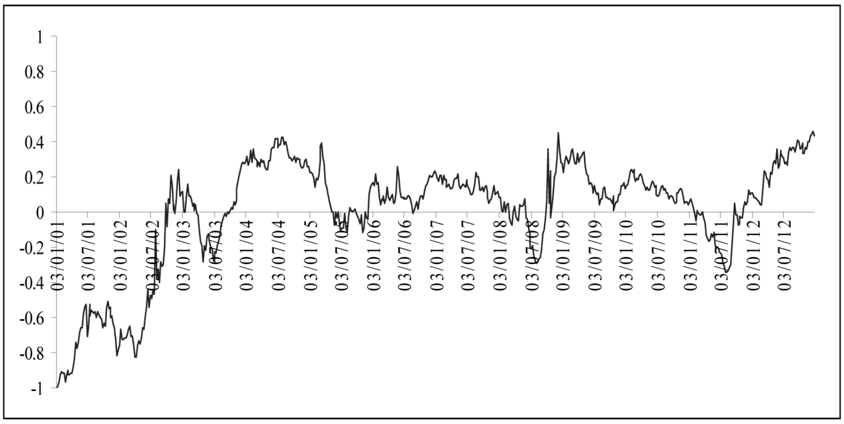

Figure 2: Plot of residuals of cointegrating relationship for Turkey (full sample)

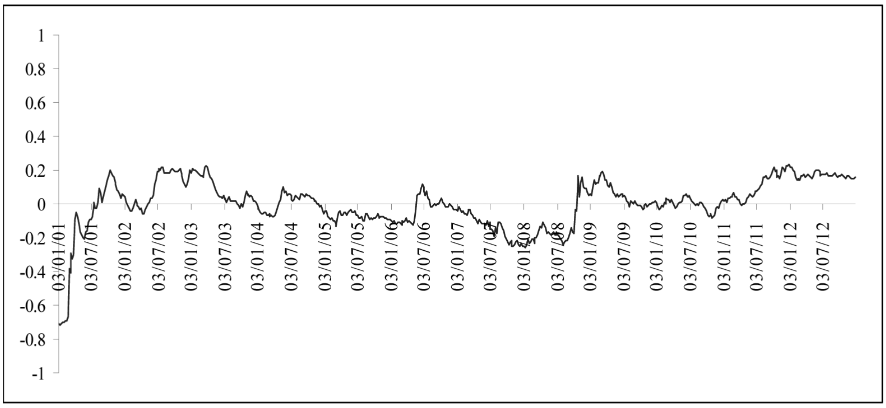

Table 5 presents the results of the non-linear Granger-causality test in a Threshold Error Correction model. The top panel shows the results of Granger-causality between stock market returns and exchange rates changes in the full sample period. The significant F-statistics for the null hypothesis that stock index returns do not Granger-cause exchange rate changes (H0: $\delta_1 = \delta_2 = 0$), indicate unidirectional short-run causality running from stock market returns to exchange rate changes in the case of Ireland and the Netherlands. Causality runs in the opposite direction (H0: $\theta_1 = \theta_2 = 0$) in the case of Turkey and there is evidence of bi-directional causality between markets in Brazil. For the pre-crisis period (Panel B), we find causal relationships between the markets only for Brazil and Turkey. Stock market returns

Table 5: Threshold error correction (causality) test results

	Ireland		*Netherlands*		*Spain*		*Brazil*		*South Africa*		*Turkey*	
	RST	REX	RST	REX	RST	REX	RST	REX	RST	REX	RST	REX
Panel A. Full sample. January 2001 - December 2012												
$H0{:}\theta_1=\theta_2=0$	0.29		1.37		0.81		3.53**		0.44		4.04**	
$H0{:}\ \delta_1=\delta_2=0$		4.33**		3.53**		1.39		4.10**		0.90		1.15
$H0{:}\theta_1=\theta_2=\gamma_1=0$	0.97		0.80		1.33		1.37		0.97		0.13	
$H0{:}\theta_1=\theta_2=\gamma_2=0$	0.08		3.71**		1.01		3.59**		3.55**		0.52	
$H0{:}\delta_1=\delta_2=\gamma_1=0$		1.02		1.13		0.70		0.90		1.17		3.69**
$H0{:}\delta_1=\delta_2=\gamma_2=0$		3.18**		0.53		1.28		0.34		3.53**		0.39
$H0{:}\ \gamma_1=\gamma_2=0$	1.07	0.97	1.24	3.70**	3.78**	0.54	4.08**	1.10	0.78	1.35	0.79	0.04
$H0{:}\ \gamma_1=\gamma_2$	1.70	1.71	0.79	1.84	0.44	1.31	5.78**	1.33	1.01	0.09	0.78	0.08
Panel B. Pre-crisis period. January 2001 - July 2007												
$H0{:}\theta_1=\theta_2=0$	0.33		1.04		0.89		1.11		1.03		0.89	
$H0{:}\ \delta_1=\delta_2=0$		1.02		1.50		0.10		4.13**		1.09		3.13**
$H0{:}\theta_1=\theta_2=\gamma_1=0$	1.30		0.88		0.20		1.37		1.35		0.70	
$H0{:}\theta_1=\theta_2=\gamma_2=0$	0.72		3.91**		1.21		0.50		3.71**		5.51**	
$H0{:}\delta_1=\delta_2=\gamma_1=0$		0.90		1.01		0.63		3.90**		0.10		1.60
$H0{:}\delta_1=\delta_2=\gamma_2=0$		3.39**		0.77		1.09		0.34		0.53		1.03
$H0{:}\ \gamma_1=\gamma_2=0$	1.20	1.02	1.07	0.93	3.20**	0.79	1.50	5.82**	1.00	0.81	3.15**	0.90
$H0{:}\ \gamma_1=\gamma_2$	0.08	1.10	1.89	0.39	0.70	1.29	1.01	1.17	1.04	1.30	2.78**	1.22
Panel C. Crisis period. August 2007 - March 2009												
$H0{:}\theta_1=\theta_2=0$	0.97		1.10		0.70		0.99		1.01		0.59	
$H0{:}\ \delta_1=\delta_2=0$		4.13**		4.50**		0.33		1.57		1.50		0.11
$H0{:}\theta_1=\theta_2=\gamma_1=0$	1.30		1.10		1.13		1.18		0.93		3.10**	
$H0{:}\theta_1=\theta_2=\gamma_2=0$	1.59		1.12		0.51		3.45**		1.00		0.57	
$H0{:}\delta_1=\delta_2=\gamma_1=0$		0.97		0.45		0.69		0.71		1.10		0.60
$H0{:}\delta_1=\delta_2=\gamma_2=0$		3.34**		3.50**		5.31**		4.01**		0.13		1.29
$H0{:}\ \gamma_1=\gamma_2=0$	1.52	5.82**	3.07**	5.93**	0.15	0.74	1.50	1.82	1.00	1.03	3.10**	0.74
$H0{:}\ \gamma_1=\gamma_2$	1.07	3.35**	1.08	4.30**	1.08**	1.03	1.22	1.09	1.08	1.05	0.70	1.04
Panel D. Post-crisis period. April 2009 - December 2012												
$H0{:}\theta_1=\theta_2=0$	1.10		1.10		0.59		1.51		0.99		2.89**	
$H0{:}\ \delta_1=\delta_2=0$		0.95		0.23		1.31		1.13		0.50		0.13
$H0{:}\theta_1=\theta_2=\gamma_1=0$	1.55		0.74		1.05		1.07		0.91		1.03	
$H0{:}\theta_1=\theta_2=\gamma_2=0$	1.32		1.40		0.11		0.55		0.10		1.14	
$H0{:}\delta_1=\delta_2=\gamma_1=0$		0.97		3.13**		0.55		0.99		1.13		0.01
$H0{:}\delta_1=\delta_2=\gamma_2=0$		3.34		0.71		1.30		1.35		0.53		1.55
$H0{:}\ \gamma_1=\gamma_2=0$	0.52	1.02	1.04	3.20**	0.15	0.74	1.50	0.81	3.07**	0.90	1.10	0.70
$H0{:}\ \gamma_1=\gamma_2$	1.70	1.07	1.79	1.30	2.78**	1.39	0.70	4.30**	1.41	1.22	0.90	1.14

Note: RST and REX denote Return on Stock Indices and Return on Exchange Rates, respectively. The model's lag length specification equals two on the basis of the Akaike Information Criterion (AIC). ** denotes the level of significance at 5%.

Granger-cause exchange rate changes in Ireland and the Netherlands during the financial crisis (Panel C). In the post-crisis period (Panel D), the variables are found to Granger-cause each other only in Turkey. Overall, therefore, the study finds mixed evidence that corroborates the mixed findings reported in the existing literature reviewed earlier. Our evidence from non-linear causality tests, however, would suggest that causality is more likely to run from stock prices to exchange rates in developed economies, and in the opposite direction in emerging countries.

Table 6 presents the EGARCH estimations for both the mean and the conditional variance equations. Panel A reports the results for the full sample period. For the first moment interdependence we find significant mean spillovers running from the stock to the currency market in Ireland and the Netherlands, in the opposite direction for South Africa and Turkey, and a bi-directional spillover for Brazil.

With regard to the second moment interdependence, volatility persistence of stock market returns and exchange rate changes, measured by β, is common across all countries for the full sample period, as the βs are all statistically significant. There are volatility spillover effects from stock prices to exchange rates for the Netherlands and Turkey, as δ in the conditional variance equation of exchange rates is significant, and from exchange rates to stock prices for Brazil, as δ in the conditional variance equation of stock prices is significant. The negative sign of the significant π for Brazil indicates that unexpected 'bad news' has a greater impact on volatility than unexpected 'good news'. However, the insignificant π for the Netherlands and Turkey suggests that spillover effects are symmetric, which means that good and bad news have an equal impact on volatility.

The results for the pre-crisis period (Panel B) display an insignificant δ for all countries, indicating that there is no volatility spillover between the markets at all in 'good times'. Nevertheless, the volatility persistence of stock prices and exchange rates are found to be common for all countries during the pre-crisis period.

The EGARCH results for the period of the financial crisis are reported in Panel C. Again, the results indicate that the volatility in stock prices and exchange rates are persistent during the financial crisis. There are asymmetric volatility spillover effects from stock prices to exchange rates for Ireland, the Netherlands and Turkey, and a symmetric spillover effect (in the same direction) for South Africa during the crisis. However, the spillover effect for Brazil is bi-directional and asymmetric. The results also indicate that the sign of the asymmetric spillover effects is negative for all countries, which means that 'bad news' has a greater impact on volatility than 'good news' during the financial crisis.

Table 6: EGARCH estimations

	Ireland		*Netherlands*		*Spain*		*Brazil*		*South Africa*		*Turkey*	
	S	E	S	E	S	E	S	E	S	E	S	E
Panel A. Full sample. January 2001 - December 2012												
θ	0.003	0.004**	-0.002	0.003**	0.001	0.002	0.003**	0.003**	-0.002	0.004	0.002**	-0.008
α							-0.001**	0.009	0.28		0.003**	0.003
β	0.95**	0.68**	0.75**	0.38*	0.97**	0.90**	0.21**	0.49**	0.87**	0.76**	0.41**	0.53**
δ	-0.24	0.14	-0.17	0.29**	-0.20	-0.23	-0.17**	0.28	-0.28	0.21	0.20**	-0.23
π	.45	-.30**	-.24	.42	.56**	-.39**	-.40**	.37	-.20**	-.29**	-.30	-.38
$LB^2(10)$	7.76	9.92	5.69	7.43	8.45	8.80	7.77	5.54	5.65	4.09	9.90	8.53
K-S(D)	0.02	0.2	0.01	0.03**	0.02	0.01	0.01	0.04**	0.01	0.01	0.02	0.01
J-B	123.04	4.04	15.07*	89.57**	3.67	134.01**	5.01	16.04**	103.01**	101.51**	3.11**	91.21**
Panel B. Pre-crisis period. January 2001 - July 2007												
θ	0.008	-0.003**	0.001	0.008	0.001	-0.005	0.002**	0.002	-0.003**	-0.006	0.001	0.003**
α							0.011**	0.009**				
β	0.75**	0.88**	0.91**	0.78*	0.67**	0.65**	0.31**	0.29**	0.32**	0.36**	0.51**	0.23**
δ	0.22	0.19	0.13	0.24	0.42	0.34	0.48	0.17	0.18	0.21	0.20	0.22
π	-.31	.50	-.44**	.31**	-.76*	-.51**	-.55	.64**	-.31**	-.44**	-.55**	.31**
$LB^2(10)$	6.16	7.12	6.19	8.03	9.40	4.83	6.02	7.14	9.02	7.01	12.91	7.22
K-S(D)	0.01	0.03**	0.01	0.02	0.02	0.01	0.01	0.03**	0.01	0.01	0.02	0.01
J-B	5.04	22.030**	8.01**	24.33**	72.61**	24.20**	4.88	20.55**	33.88**	25.90**	3.61	30.20**
Panel C. Crisis period. August 2007 - March 2009												
θ	0.001	0.008**	0.031	-0.024**	0.003	0.007	0.006	0.022**	0.001**	-0.007	0.210	0.003
β	0.91**	0.87**	0.30**	0.52*	0.40**	0.37**	0.76**	0.86**	0.71**	0.67**	0.30**	0.23**
δ	0.21	0.23**	0.27	0.26**	0.19	0.21	0.20**	0.17**	0.28	0.21**	0.24	0.20**
π	-.53**	-.45**	-.37**	-.21**	.27	-.35**	-.25**	-.34**	-.23**	.20	-.25**	-.41**
$LB^2(10)$	5.01	6.02	12.09	10.42	13.00	9.83	9.23	13.51	10.62	12.00	17.91	9.52
K-S(D)	0.01	0.1	0.01	0.03**	0.01	0.01	0.01	0.04**	0.01	0.01	0.02	0.01
J-B	18.07**	19.01**	2.04	5.03	3.61	59.19**	80.77**	4.00	80.99**	12.55**	51.62**	4.26
Panel D. Post-crisis period. April 2009 - December 2012												
θ	0.005	0.471	0.021	-0.010**	0.007	0.018	-0.014	0.323	0.017	-0.401	0.031**	0.023
β	0.35**	0.48**	0.72**	0.80**	0.57**	0.68**	0.41**	0.30**	0.67**	0.80**	0.21**	0.33**
δ	0.20	0.16**	0.57	0.39	0.29	0.14	0.58	0.37	0.22	0.20	0.27	0.23
π	-.51**	.50**	-.44	-.31**	-.76	-.51**	.55	-.64**	-.31**	.44	-.55**	-.31**
$LB^2(10)$	8.70	6.90	5.09	7.44	8.05	8.01	9.19	11.50	10.15	13.10	16.80	16.50
K-S(D)	0.02	0.01	0.01	0.02	0.01	0.01	0.02	0.01	0.03**	0.04**	0.02	0.02
J-B	34.04*	20.03**	5.07	5.99**	3.60	90.29**	104.5**	1.04	71.04**	10.50**	54.30**	55.27**

Note: S represents stock price equations (3) & (5) and E represents exchange rate equations (4) & (6). θ & λ measure return spillover and error correction term, respectively. Persistence of volatility, volatility spillover from one market to another, and asymmetric spillover effect are measured by β, δ and π, respectively. The model is specified (lag length = 2) based on the Akaike Information Criterion (AIC). $LB^2(10)$ are the Ljung-Box statistics for squared standardised residuals distributed as a chi-square with 10 degrees of freedom. K-S (D) and J-B are the Kolmogorov-Smirnov and Jarque-Bera tests of null hypothesis that the residuals are normally distributed. ** denotes the level of significance at 5%.

Panel D shows that there is volatility spillover between the markets only in Ireland. The asymmetric volatility spillover runs from stock prices to exchange rates and the sign is positive, which means that 'good news' has a greater impact on volatility than 'bad news'. This may be because 'good news' on stock prices may have a greater impact on demand for local currency as foreign investors want to increase their holding of rising stock. On the other hand, 'bad news' may induce a less pronounced tendency to sell declining stock denominated in local currency (as this would entail acceptance of a loss).

In comparing the findings between developed and emerging economies, the volatility spillover from stock prices to exchange rates confirms that stock prices is the lead variable in the developed economies. The findings are mixed in the case of the emerging economies. There are instances of a unidirectional volatility spillover from exchange rates to stock prices in Brazil and Turkey for the full sample period, and a bi-directional volatility spillover in Brazil during the financial crisis. The bi-directional spillover can be explained by a process whereby an increase in exchange rates causes an increase in output, export and ultimately stock prices (particularly in export dominated countries). The demand for local currencies increases as a result of the increased stock prices through the wealth effect which, in turn, causes a consequent appreciation of the local currency.

Our findings are consistent with those by Kanas (2000), and Yang and Doong (2004), who also found a volatility spillover from stock prices to exchange rates for developed countries. However, contrary to the findings of Raghavan and Dark (2008), our study has not produced evidence in support of a volatility spillover in the opposite direction. Our findings also corroborate those by Wu (2005) and Mishra *et al.* (2007), who found a bi-directional volatility spillover between stock prices and exchange rates in emerging countries. Furthermore, as per the results obtained by Apte (2001) and Walid *et al.* (2011), we find a unidirectional volatility spillover from exchange rates to stock prices in emerging economies.

In order to assess the robustness of our results, we undertake diagnostic checks on the EGARCH model (see Table 6). The model appears adequately to describe the volatility spillover effect between stock prices and exchange rates. The fact that the Ljung-Box statistics are not significant indicates that there are no residual dependencies, linear or non-linear. The statistics of the Kolmogorov-Smirnov and Jarque-Bera tests of the null hypothesis of residual normality provide mixed results, though this is not unusual when working with financial series and normality is not a crucial issue here due to quite a large sample size.

To gauge the sensitivity of our results we also re-run our cointegration, causality and EGARCH estimations using Nominal Effective Exchange Rates (NEER) instead of nominal bilateral exchange rates. The results (not reported to conserve space but available from the authors upon request) are broadly in line with those reported in the Tables above, thus corroborating our earlier findings.

6. Conclusions

In this paper, we have re-examined empirically the volatility spillover effects between exchange rates and stock prices in selected developed and emerging countries across pre-financial-crisis, crisis and post-crisis periods using an EGARCH model.

Our results indicate that there is a unidirectional volatility spillover effect running from stock prices to exchange rates in the developed countries. The direction of the volatility spillover between the two markets is opposite in the emerging countries. However, there is evidence of a bi-directional volatility spillover between markets in Brazil.

Significantly, we also find that there are asymmetric volatility spillover effects between exchange rates and stock prices in both developed and emerging countries, particularly during the financial crisis period.

Whilst acknowledging that additional variables we did not control for (such as interest rates and inflation) may also play a role on the volatility spillover effects between exchange rates and stock prices (which constitutes, in itself, a profitable extension for future research), two important implications flow from our findings.

First, evidence that stock and foreign exchange markets are interrelated, in both developed and emerging countries, implies that lagged information from one market can be used to forecast changes in the other. This also signifies that markets are 'informationally' inefficient, with one market having significant predictive power on the other. Second, the finding of the volatility spillover effect between stock prices and exchange rates in all countries except Spain has important implications for portfolio managers and investors, suggesting that they should not include both assets in the same basket if aiming to diversify risk in their asset portfolio.

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Endnotes

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2. For example, in July 2005, after more than a decade of strictly pegging the renminbi to the US dollar, the People's Bank of China announced a revaluation of the currency and a reform of the exchange rate regime that entailed taking a 'reference bas-

ket' of currencies when choosing its target for the renminbi. India too is well known for having managed pragmatically many episodes of volatility of exchange rate by officially pegging the rupee to the US dollar.

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