**Linear and non-linear dynamics between exchange rates and stock markets returns: An application to the financial crises of Europe and Asia in the nineties**


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**Abstract**

The recent crises of the nineties have made it clear that the links between exchange rates and stock market prices are relevant factors in the transmission of the crises. Using daily exchange rates and stock index prices of the last decade (1990-1999) the interactions between the stock market and exchange rates returns of twenty-three countries of two different geographical areas (Asia and Europe) are analysed. Our results suggest that: (i) short term relationships seem to be more relevant than long term ones, (ii) it is more relevant the presence of linear and nonlinear causality in the Asian countries, and (iii) the periods of crisis affect asymmetrically the relationship between exchange rates and stock market prices.

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## 1. Introduction

The intense globalisation process of the financial and banking system as well as the increasing degree of economic and financial integration of countries have led to the arising of *global financial crises*. This term refers to crises, which in their beginning are limited to a specific geographical area, but that in a short time influence to other financial markets around the world. The financial system has experienced several financial crises during the nineties: the European crisis of 1992, the Asian crisis of 1997 and the Russian and Latin American crises of 1998. From the developed country point of view, perhaps the most severe crises were the first two ones and they are the objective of this paper.

The chronology of the European and Asian crises is well known and only their main aspects will be reviewed. In the case of the European crisis, the causes can be summarised in two. The first one was the uncertainty in the viability of the European Monetary Union (EMU), after the Danish negative answer to the *Maastrich Treaty* (June 2nd, 1992), while the second one was the asynchrony of the cycles of the European economies which made it difficult to coordinate the national economic policies. Both factors should be placed in a context characterised by free mobility of capitals in the exchange rate mechanism of the European Monetary System (EMS), a system of fixed exchange rates but anchored to the Deutsche Mark.

The origin of the crisis is generally attributed to the need of the German government to employ an expansive fiscal policy to ease unification together with a restrictive monetary policy with the objective of containing inflation. This led to an increase of the German interest rates and a strong appreciation of the Deutsche Mark, which was also favoured by the weakness of the US dollar. The rest of the countries were then forced to raise their interest rates, trying to keep the parity of their currencies within the bands, while the situation of their economies (in recession or in a decelerating process) required low interest rates. From this moment, on the EMS experienced important currency *shocks* that led to abandon the exchange rates discipline of the Italian Lira and British Pound, the devaluation of the Portuguese Escudo and the Spanish Peseta (two times), as well as to abandon the fixing of the Scandinavian currencies against the Deutsche Mark. In this context, the respective stock markets experienced a clear disparity.

On the other side, the Asian crisis (which lasted from July to October 1997) had its origin on speculative movements against the Thai *baht*. It is worth noting that, in the beginning, this crisis did not provoke important effects on the stock markets of the area, but produced a chain of devaluations and substitution processes of the exchange rates systems. The Asian crisis was composed of five sub-crises: first in Thailand, then in Philippines, Indonesia and Malaysia, and finally in South Korea and Hong Kong. It is after the crash of the Hong-Kong's stock market, when the crisis affected the stock markets around the world, which in general experienced important falls.

It is important to point out that the countries most severely affected by the crisis (South Korea, Thailand, Malaysia and Indonesia) had experienced significant growth rates along the last decades, their investment rates were over 30% of the GNP and these economies had also significant saving rates and public superavits. Then, which were the causes of the crises?. Basically, the excessive flows of incoming capitals of these economies, in the form of portfolio capitals or loans of the international banking system, provoked a strong appreciation in the real exchange rate ( especially in the case of Thailand and Malaysia) and finally led to an exponential increase of the current deficit of these economies. This process of excessive appreciation, high current deficit and high prices of financial and real-state assets led to the financial agents to lose their confidence on the *Thai baht*, which provoked an important exit of capitals from Thailand.

On July 2nd, 1997 after spending 8,700 millions of US dollars to hold their currency, the Thai Central Bank released their exchange rate and, at the end of the same year, the *Thai baht* experienced a loss in its value of about 93%, and the SET lost a 34% (in US dollars), compared to June 1997. The problem with the *baht* put in serious doubt the viability of the exchange rate systems of the other countries, and the problems extended quickly to Indonesia, Malaysia and Philippines, making clear the structural deficiencies of these economies. The measures adopted to reduce liquidity in Indonesia were not enough to stop the increasing pressures on the exchange rate markets, and authorities allowed the floating of the *rupiah* by the middle of August. The situation became worse in the following two months and the secondary effects propagated to other countries such as Hong Kong and Japan. The strong pressures on the Hong Kong dollar in October caused a significant increase on interest rates and led to the plummeted of the stock market. This led to a "*domino effect*" on the majority of world stock markets as well as speculative attacks to the currencies of developing countries. In Korea, the pressures on the *Won* intensified along October, after the attack of the Hong Kong dollar, and the stock market experienced a crash as a consequence of the lack of confidence in the future of the economy and the difficulties to renew the external debt.

As we can see, the two mentioned crises share similar origins (both had their origin on problems related to the stability of the respective exchange rates) and effects (the currencies crises were rapidly transmitted to the stock markets). This motivates our study. We will try to see if there is empirical evidence of a relation between exchange rates and stock market prices and, if so, we will try to determine the direction and nature of the relationship. We are also interested in analysing whether markets exhibit a different behaviour in the period of crises and if, from the empirical point of view, there are significant differences between both crises.

Our study may be considered an extension of previous analyses in two senses. First, opposite to several other studies, causality and cointegration between exchange rates and stock market prices are analysed in linear and nonlinear contexts. This is justified by the fact that there is considerable evidence in favour of the hypothesis that univariate financial time series may contain significant nonlinear components that could be transmitted from one market to another (e.g. Hsieh, 1991). Also, considering just linear cointegration or linear causality would rule out the possibility that short and long term relationships between the exchange and stock markets might vary over time, which seems quite implausible in the changing financial system.

Secondly, the analyses are usually conducted using a limited number of series and during short periods. Here we will expand in time and space previous analyses by examining twenty-three markets in two different monetary areas (Europe and Asia) along the period 1990-1999. This will also allow us to find out the influence of the exchange rates crises on the bivariate dynamical relationship between exchange rates and stock market prices.

The paper is structured as follows: in the second section we present the theoretical models proposed in the literature to analyse the bivariate relationships between exchange rates and stock market prices. The econometric methodology is described in section three, where we briefly describe the test routinely employed in the literature as well as some other ones not so common, which will be also employed in this paper. The database used and the results obtained will be described in section four. The main conclusions close the paper.

## 2. Theoretical setup

As we have mentioned, exchange rate markets and stock markets seem to be clearly linked. From a microeconomic point of view, the fluctuations of the exchange rates clearly affect the value of a portfolio composed by national and multinational firms: if the real exchange rate increases, the profits of the firms decrease and, as a consequence, the price of their stock also decreases. Under this point of view, the relationship can be expressed as (Jorion, 1990):

$$R_{it} = B_{0i} + B_{1i} R_{st} + \varepsilon_{it} \tag{1}$$

where $R_{it}$ is the return of firm $i$ and $R_{st}$ is the rate of variation of the exchange rate at time $t$[1]. Obviously, the effect of the variations of the exchange rate on the value of the stock is probably quite different in the case of domestic firms than for multinational firms. So, it is essential to determine the degree of exposition to exchange rates and the implication of the firm in international markets. We can express this by:

$$B_{1i} = a_0 + a_1 F_i + \mu_i \tag{2}$$

where $F_i$ is the ratio of the sales to foreign countries on the total of sales of firm $i$ and $\mu_i$ is a random perturbation. In this sense, the existing relationship between exchange rates and stock market prices can be analysed in a broader sense by:

$$R_{it} = B_{0i} + B_{1i} R_{st} + B_{2i} R_{mt} + \varepsilon_{it} \tag{3}$$

where $R_{mt}$ is the return of the domestic stock market at time $t$ and $\varepsilon_{it}$ is a random perturbation.

[1] This model is used in Abdalla and Murinde (1997, p.26) to study the linear relationships between exchange rates and stock market returns in emerging markets for the period 1985-1994 at a monthly frequency.

From a macroeconomic point of view, it is postulated that there exists a negative relationship between the strength of the local currency and the evolution of the stock market index of the country, so that:

$$R_{st} = \alpha + \beta DRS_t + \gamma Di_t + \varepsilon_t \qquad (4)$$

where $R_{st}$ is the variation of the real exchange rate, $DRS_t$ is the differential (domestic minus foreign) of the stock market return, $Di_t$ is the variation of the interest rate differential and $\varepsilon_t$ is a random perturbation. From the perspective of portfolio management, a reduction in the prices of the assets reduce the purchasing power of the domestic investors, which causes a reduction in the demand of money and a decrease of interest rates. This reduction in interest rates provokes the exit of capital which, *ceteris paribus*, will led to a depreciation of the currency. Note that, from this perspective, the stock market prices lead the process.

Under another perspective, which is known as the *traditional hypothesis*, exchange rates are the leaders of the process. An appreciation of the exchange rate would reduce the competitiveness of the exporters, which will have a negative impact on stock markets. On the other side, an appreciation of the exchange rates reduces costs and has a positive impact on the stock market prices (Granger *et al.* 2000). In a macroeconomic context, the existing relationship between exchange rates and stock market prices would be fully captured only taking into account other variables such as public debt of the respective countries as well as their current account results.

The estimates obtained in empirical studies do not agree in the sign of the relationship. For example, Smith (1992) and Solnik (1987), consider that this relation is positive while in other papers it is negative or mixed, depending on the market (Granger *et al.* 2000). Nevertheless, there is some consensus that there exists a relationship between exchange rates and stock market returns. For example, Roll (1979) establishes that this relationship can be analysed in the context of the Law of One Price. The prices of the assets in the stock market, due to the quick reaction of these markets, are good indicators of the real economic activity and, as a consequence, could be used in exchange rate models. Giovannini and Jorion (1987) also find empirical regularities between exchange rates and stock markets in the USA and show that *ex ante* returns of the markets tend to move together in the long run. Chiang (1991) shows that excess returns of exchange rates are correlated with the relative risk of stock markets, while Bekaert and Hodrick (1992) show that dividends have predictive power on the excess returns in the exchange rate markets, and that the risk premium has predictive power on stock market returns.

Heston and Rouwenhorst (1994) analyse the causes of the difference in asset prices between 1978 and 1992 in twelve markets and show that only part of these differences can be explained by exchange rates variations. Nevertheless, they consider that the estimations are sufficiently clear to demonstrate the existence of a relationship between exchange rates and stock market prices. Dumas and Solnik (1995) show that risk premia are a fundamental component for returns of international assets. In this sense, Frankel (1996) also obtains results which sustain the hypothesis that exchange rates fluctuations have real effects on stock markets, even when these changes are the result of exogenous reasons of a particular exchange rate regime. Canzoneri and Dellas (1996) calibrate a simple model of general equilibrium, which shows that the change from a particular exchange regime to another one provokes effects on the prices of assets proportionally larger than the ones caused by real interest rates. Copeland and Copeland (1998) conclude that exchange rates are an explaining factor, significant and independent, of stock market returns, while Chelley-Steeley *et al* (1998) conclude that the reduction or disappearance of exchange rates controls cause an increase in the interdependence of stock markets. Ong and Izan (1999) use the Law of One Price and find that the returns of the domestic stock market is equal to the sum of the variation in the exchange rates plus the returns of the foreign index, so that the real return is the same after correcting for exchange rate differences and risk. What all these studies manifest is that there exist clear relationships between exchange rates and stock market prices. These relationships can be, though, of short or long run nature.

Short run relationships have been traditionally explored by the use of Granger causality test (see below), while the analysis of long run relationships if done in the context of cointegration. The concept of cointegration implies the existence of a relationship between the variables so that eventual divergences between them would tend to disappear due to forces in the market (Granger, 1986). The study of causality and cointegration relationships is relevant, since they affect market mechanisms such as price discovery, volatility transmission, and arbitrage or market efficiency. For example, some authors have suggested that causality and cointegration reduces, but not eliminates, the benefits obtained by international diversification, while other authors have suggested that they are inconsistent with market efficiency.

Our analyses will adopt a linear as well as a nonlinear perspective. There is now considerable evidence in favour of nonlinearity in high frequency financial time series (e.g. Hsieh, 1991) but, surprisingly, the vast majority of the studies analyse nonlinearity in an univariate context. It should be noted, though, that under a theoretical point of view, the existence of bivariate (or multivariate) nonlinear relationships between exchange rates and stock market returns is certainly plausible. For example, if agents and price mechanisms of stock and exchange rate markets are not the same, the apparition of new information could be processed asymmetrically in both markets, which would imply a nonlinear relationship between both markets.

In this sense, it is interesting to note that the target zone model of Krugman (1991) establishes the existence of nonlinear relationships between fundamental macroeconomic variables and the exchange rates. There is a number of studies that have tested the model, finding evidence against nonlinearity (De Jong, 1994; Meese and Rose, 1990) or in favour of it (Ma and Kanas, 2000). If we consider the evidence found in the mentioned studies (Roll, 1979; Ong and Izan, 1999), which establish that stock market prices are good indicators of real economic activity and, as a consequence, could be employed in exchange rates models, we could conclude the plausible existence of nonlinear relationships between both markets. Checking for nonlinear relations will be, therefore, a fundamental objective of our paper, since this could motivate the extension of models summarised in equations (1)-(4) to the nonlinear case.

## 3. Econometric methodology

Our research will focus on three aspects. First, we will conduct univariate test to assess the nonstationarity of the series employed. Then, we will test the existence of long run linear and nonlinear relationships by means of cointegration analyses. Finally, we will check for short run linear and nonlinear relationships.

The study of stationarity is fundamental in our context due to the well-known problems of spurious regression. Instead of using the ADF test (Dickey and Fuller 1979) we will employ the Phillips-Perron test (PP, hereafter, Phillips and Perron, 1988), which is robust to some forms of heteroskedasticity, quite common in financial data. Since it is known that both tests tend to accept the null hypothesis too frequently, we will also employ the KPSS test (Kwiatkowski *et al.* 1992) for which stationarity around a level of a trend is the null hypothesis. The ADF and PP tests are well known, and we will not describe them here, instead we will briefly describe the less known KPSS test.

Let $X_t$, $t=1,2,....T$, be the time series under analysis. Assume that $X_t$ can be decomposed as the sum of a deterministic trend, $t$, a random walk, $r_t = r_{t-1}+u_t$ and a stationary error, $\varepsilon_t$:

$$X_t = \xi\, t + r_t + \varepsilon_t \tag{5}$$

where $u_t \sim iid\,(0, \sigma_u^2)$. The null hypothesis is $H_0$: $\sigma_u^2 = 0$ while the alternative is $H_A$: $\sigma_u^2 > 0$. Under the null hypothesis, $X_t$ is stationary around a trend and, in the case that $\xi = 0$, $X_t$ is stationary around a level ($r_0$). KPSS test shows that under the null hypothesis, the statistic:

$$S_T = \frac{1}{S_{l_T}^2}\frac{1}{T^2}\sum_{t=1}^{T}\left(\sum_{j=1}^{t}\varepsilon_j\right)^2 \tag{6}$$

where

$$S_{l_T}^2 = \frac{1}{T}\sum_{t=1}^{T}\varepsilon_t^2 + \frac{2}{T}\sum_{t=1}^{l_T}\left(1-\frac{t}{l_t+1}\right)\sum_{j=t+1}^{T}\varepsilon_j\varepsilon_{j-t} \tag{7}$$

and $l_T=o(T^{1/2})$, converges to a brownian bridge of first or second order, depending on whether we regress $X_t$ on a trend or not (Kwiatkowski *et al.,* 1992, also provide the critical values of the test, which are used here). The KPSS test is employed looking for robustness: since standard statistical testing is biased to accepting the null hypothesis if no strong evidence against it is found, with the only use of ADF and PP tests one would find too frequently that univariate time series may contain a unit root, when it may not be the case.

If there exists a stable long run relationship between two time series, one could find that the residuals of the regression between them are stationary, even though none of them is stationary. This result would mean that there exists a common temporal evolution, so that the differences between the series would not tend to increase or decrease (that is, a linear combination would be stationary), in this case we say that the series are *cointegrated* (Granger, 1981). Cointegration tests are designed to verify if several time series follow common trends. Engle and Granger (1987) first proposed a simple two-step test to verify cointegration and their approach was further refined by Johansen (1988), Johansen and Juselius (1990) and Shin (1994). We will employ Johansen and Shin's approach. It is important to note that the Johansen and Juselius test takes no cointegration as the null hypothesis, while in the Shin's test it is the existence of linear cointegration. Again, we will follow both approaches to assess the results.

The method proposed by Johansen & Juselius is carried out in a multivariate maximum likelihood framework. It employs the likelihood ratio to determine not only the existence of a cointegrating vector but also its rank. This method is then useful not only to test, but also to estimate, cointegration relations of VAR models:

$$X_t = \sum_{i=1}^{n} \Pi_i X_{t-i} + \mu + \varepsilon_t \tag{8}$$

which, expressed as an Error Correcting Model (ECM), gives:

$$\Delta X_t = \sum_{i=1}^{n-1} \Gamma_i \Delta X_{t-i} + \Gamma_n X_{t-n} + \mu + \varepsilon_t \tag{9}$$

where $\varepsilon_t$ is a random k-dimensional vector with mean zero and non-singular covariance matrix Σ, and $\mu_i$ is a vector of constant terms. $\Gamma_i = - I + \Pi_1 + \ldots\ldots\ldots + \Pi_i$, $i=1,\ldots\ldots\ldots,n$ . It is considered that the roots of the characteristic implicit polynomial are placed outside the unit circle. The interesting situations arise when rank $(\Gamma_n) = r < k$; in this case, there are $k$-$r$ unit roots in the system as well as $r$ cointegration relationships and $\Gamma_n$ can be expressed as $\alpha\beta'$, where both $\alpha$ and $\beta$ are $(k \times r)$ matrices of whole rank. The $r$ first rows of $\beta'$ are the $r$ cointegration vectors, while the elements of $\alpha$ are the weights of the cointegrating vectors in the equations. With the condition that none of the elements of $X_t$ is integrated of order higher than one, the maximum likelihood estimation of the base of the cointegration space is given by the empirical canonical variables of $Xt_{-n}$ with respect to $\Delta X_t$, corrected by the short run dynamics and by the deterministic components. The number of cointegration relationships is the same as the number of significant canonical correlation and their significance can be tested by a sequence of likelihood ratios whose distribution asymptotically converges to a brownian bridge.

Additionally, we employ Shin's test (Shin, 1994) to detect the existence of cointegration. This test is a multivariate version of the KPSS test, but now the residuals, $\varepsilon_t$, are obtained by regressing $Y_t$ on $X_t$. In this case, the null hypothesis is the existence of linear cointegration while the alternative is no cointegration. The motivation to employ Shin's test is similar to the use of the KPSS: since the null hypothesis of the Johansen test is no cointegration, one would accept the null hypothesis except if the evidence is definitive, that is, one would lead to find no cointegration too frequently.

The analysis of the short term relationships between exchange rates and stock returns will be done, first, under a static point of view (using correlation matrices) and will be later extended to a dynamical perspective. Since correlation does not reveal which is the dynamical relationship among the variables, and which market is the leader, we will employ Granger causality tests. To do this, we build vector autoregressive models:

$$\begin{aligned} \Delta y_{1t} &= \alpha_0 + \sum_{i=1}^{n_1} \alpha_{1i} \Delta y_{1t-i} + \sum_{i=1}^{n_2} \alpha_{2i} \Delta y_{2t-i} + \varepsilon_{1t} \\ \Delta y_{2t} &= \beta_0 + \sum_{i=1}^{n_3} \beta_{1i} \Delta y_{2t-i} + \sum_{i=1}^{n_4} \beta_{2i} \Delta y_{1t-i} + \varepsilon_{2t} \end{aligned} \tag{10}$$

where $\Delta y_{1t}$ and $\Delta y_{2t}$ are the exchange rates and stock market returns, respectively, and $\varepsilon_{1t}$ and $\varepsilon_{1t}$ are random independent variables. The Granger test is established as an F-test on the parameters of the model, the rejection of the null hypothesis $H_0$: $\alpha_{21} = \alpha_{22} = ..... = \alpha_{2s} = 0$ implies that exchange rates returns do not Granger-cause stock market returns while the rejection of the null hypothesis $H_0$: $\beta_{21} = \beta_{22} = ..... = \beta_{2s} = 0$ suggest that stock market returns do not Granger-cause exchange rates returns.

As mentioned above, most of the studies that analyse the temporal relations between cash and futures markets employ *linear* techniques, looking for *linear* relations between the variables. Consequently, failing to reject the null hypothesis of independence can only be interpreted as evidence against *linear* relations, and not the absence of other types of dependence. Here we analyse causality and cointegration between exchange rates and stock market prices in a nonlinear context; this extension is of interest at least for two reasons: the first one is that it seems quite plausible that univariate financial time series may contain significant nonlinear components (e.g. Hsieh, 1991) that could be transmitted from one market to another one; the second reason is that the consideration of linear cointegration or linear causality rules out the possibility that the equilibrium relationships may vary over time. In the context of causality, a nonparametric test recently proposed (Baek and Brock, 1992) (the Baek and Brock test, hereafter) can be useful to reveal nonlinear dependencies between two time series, undetectable by traditional linear causality tests (see Brock, 1991 for an illustrative example).

We will provide now a brief introduction of the Baek and Brock test. In what follows, let us assume that $\{X_t\}$, $\{Y_t\}$, $t=1,2,...,n$, are two strictly stationary and weakly dependent time series. We say that $Y_t$ *does not strictly Granger-cause* $X_t$ if the probability distribution of $X_t$ conditioned on information set $I_t$ is independent of $Y_t$, that is if $F(X_t/I_t) = F(X_t/I_t\text{-}Y_t)$ (Granger, 1969).

Obviously, if $Y$ does not strictly Granger-cause $X$, lagged values of $Y$ do not provide further information to predict $X$. Standard Granger bivariate causality tests are usually implemented by estimating a vector autorregression model (VAR) and testing (by a standard $\chi^2$ or F-test) if the estimated coefficients are jointly significantly different from zero. As we mentioned, this procedure will fail to reveal many kinds of nonlinear dependence (such as dependence on conditional moments of order higher than two), so that it cannot be properly considered a test for independence.

Alternatively, Baek and Brock (1992) have proposed an extension of the BDS test (Brock *et al.* 1987) to the multivariate case. Let us note by $\boldsymbol{X}_t^m$ the m-history, $\boldsymbol{X}_t^m = (X_t, X_{t+1},...,X_{t+m-1})$, $m=1,2,...$, $t=1,2,....$; we say that $Y_t$ *does not strictly cause* $X_t$ if present and past values of $Y_t$ do not help to predict values of $X_t$, that is, for given values of $m$, $L_x$, $L_y \geq 1$ and $\varepsilon > 0$ if:

$$Pr(/\boldsymbol{X}_t^m - \boldsymbol{X}_s^m/ < \varepsilon \;/\; /\boldsymbol{X}_{t-L_x}^{L_x} - \boldsymbol{X}_{s-L_x}^{L_x}/ < \varepsilon, /\boldsymbol{Y}_{t-L_y}^{L_y} - \boldsymbol{Y}_{s-L_y}^{L_y}/ < \varepsilon) =$$
$$= Pr(/\boldsymbol{X}_t^m - \boldsymbol{X}_s^m/ < \varepsilon \;/\; /\boldsymbol{X}_{t-L_x}^{L_x} - \boldsymbol{X}_{s-L_x}^{L_x}/ < \varepsilon) \qquad (11)$$

being $/\cdot/$ the supreme norm.

The left hand side of equation (11) is the conditional probability that two arbitrary *m*-histories of $\{X_t\}$ are within a distance smaller than ε, given that the corresponding $L_x$-histories of $\{X_t\}$ and the $L_y$-histories of $\{Y_t\}$ are within a distance less than ε. The right hand side of the equation is the conditional probability that two arbitrary *m*-histories of $\{X_t\}$ are within a distance smaller of ε of each other, given that the corresponding $L_x$-histories of $\{X_t\}$ are at a distance smaller than $\varepsilon$.

Since $Pr(A/B) = Pr(A \cap B) / Pr(B)$ then,

$$Pr(/\boldsymbol{X}_t^m - \boldsymbol{X}_s^m/<\varepsilon / /\boldsymbol{X}_{t-L_x}^{L_x} - \boldsymbol{X}_{s-L_x}^{L_x}/<\varepsilon, /\boldsymbol{Y}_{t-L_y}^{L_y} - \boldsymbol{Y}_{s-L_y}^{L_y}/<\varepsilon) =$$

$$= Pr(/\boldsymbol{X}_t^m - \boldsymbol{X}_s^m/<\varepsilon, /\boldsymbol{X}_{t-L_x}^{L_x} - \boldsymbol{X}_{s-L_x}^{L_x}/<\varepsilon, /\boldsymbol{Y}_{t-L_y}^{L_y} - \boldsymbol{Y}_{s-L_y}^{L_y}/<\varepsilon) /$$

$$/ Pr(/\boldsymbol{X}_{t-L_x}^{L_x} - \boldsymbol{X}_{s-L_x}^{L_x}/<\varepsilon, /\boldsymbol{Y}_{t-L_y}^{L_y} - \boldsymbol{Y}_{s-L_y}^{L_y}/<\varepsilon),$$

and, also, since

$$Pr(/\boldsymbol{X}_t^m - \boldsymbol{X}_s^m/ < \varepsilon, /\boldsymbol{X}_{t-L_x}^{L_x} - \boldsymbol{X}_{s-L_x}^{L_x}/< \varepsilon) = Pr(/\boldsymbol{X}_{t-L_x}^{m+L_x} - \boldsymbol{X}_{s-L_x}^{m+L_x}/ < \varepsilon),$$

then equation (11) can be re-expressed as

$$\frac{C_1(m+L_x,L_y,\varepsilon)}{C_2(L_x,L_y,\varepsilon)} = \frac{C_3(m+L_x,\varepsilon)}{C_4(L_x,\varepsilon)} \tag{12}$$

where

$$C_1(m+L_x,L_y,\varepsilon) = Pr(/\boldsymbol{X}_{t-L_x}^{m+L_x} - \boldsymbol{X}_{s-L_x}^{m+L_x}/ < \varepsilon,\ /\boldsymbol{Y}_{t-L_y}^{L_y} - \boldsymbol{Y}_{s-L_y}^{L_y}/< \varepsilon)$$

$$C_2(L_x,L_y,\varepsilon) = Pr(/\boldsymbol{X}_{t-L_x}^{L_x} - \boldsymbol{X}_{s-L_x}^{L_x}/ <\varepsilon,\ /\boldsymbol{Y}_{t-L_y}^{L_y} - \boldsymbol{Y}_{s-L_y}^{L_y}/<\varepsilon)$$

$$C_3(m+L_x,\varepsilon) = Pr(/\boldsymbol{X}_{t-L_x}^{m+L_x} - \boldsymbol{X}_{s-L_x}^{m+L_x}/ < \varepsilon)$$

$$C_4(L_x,\varepsilon) = Pr(/\boldsymbol{X}_{t-L_x}^{L_x} - \boldsymbol{X}_{s-L_x}^{L_x}/ < \varepsilon)$$

An useful estimator of the probabilities involved in equation (2) (of $C_3$ and $C_4$, strictly speaking) is the correlation integral

$$C(m,\varepsilon) = \frac{2}{n_m(n_m-1)} \sum\sum_{t<s} I(x_t^m, x_s^m, \varepsilon)$$

where $n_m = n-m+1$, $\varepsilon \in \mathrm{R}^+$ and $I\_(x_t^h, x_s^h)=1$ iff $|x_t^h - x_s^h| < \varepsilon$, with $/\cdot/$ being the supreme norm.

Under the assumptions that $\{X_t\}$ and $\{Y_t\}$ are strictly stationary, weakly dependent and satisfy certain mixing conditions if $\{Y_t\}$ does not strictly cause $\{X_t\}$, Hiemstra and Jones (1994) proved that

$$\sqrt{n}\left(\frac{C_1(m+L_x,L_y,\varepsilon)}{C_2(L_x,L_y,\varepsilon)} - \frac{C_3(m+L_x,\varepsilon)}{C_4(L_x,\varepsilon)}\right) \sim N\left(0,\sigma^2\left(m,L_x,L_y,\varepsilon\right)\right)$$

where $\sigma^2(m,L_x,L_y,\varepsilon)$ can be consistently estimated.

Based on this, Hiemstra and Jones (1994) propose to employ the residuals from estimated VAR models to check for nonlinear dependence between $\{X_t\}$ and $\{Y_t\}$. A problem related with this procedure is that the asymptotic distribution of the statistic when it is applied to VAR residuals is not known. Nevertheless, results in Baek and Brock (1992), and Monte Carlo evidence presented in Hiemstra and Jones (1993) suggest that the statistic is robust against nuisance parameter problems and that the test has an adequate size and power for moderate sample sizes.

In addition, in the context of cointegration, and as an alternative to *linear* cointegration, several authors have proposed nonlinear extensions (e.g. Balke and Fomby, 1997; Aparicio and Escribano, 1998; Bierens, 2000; Corradi *et al.* 2000). Bierens (2000) has suggested the term nonlinear cotrending to analyse the situation where series have common nonlinear deterministic trends. Specifically, suppose that $z_t = g(t) + u_t$, where $g(t) = \beta_0 + \beta_1 t + f(t)$, where $z_t$ is a k-variant time series process, $u_t$ is a k-variant stationary process with mean zero and $f(t)$ is a deterministic k-variant nonlinear trend. Nonlinear cotrending is the phenomenon that there exists a non-zero vector θ such that $\theta^T f(t) = 0$. The idea of Bierens is to test the number of co-trending vectors based on the generalised eigenvalues of two stochastic matrices.
Specifically, let

$$\hat{M}_1 = (1/n)\sum_{t=1}^{n}\hat{F}(t/n)\hat{F}(t/n)^T \tag{13}$$

where

$$\hat{F}(x) = (1/n)\sum_{t=1}^{[nx]}(z_t - \hat{\beta}_0 - \hat{\beta}_1 t) \tag{14}$$

if $x\in[1/n,1]$ and $F(x)=0$ if $x\in[0,1/n)$, and where $[nx]$ denotes the integer part of $nx$ and $\beta_0$ and $\beta_1$ are the OLS in the above regression for time $t = 1, 2, ..., n$. And let

$$\hat{M}_2 = (1/n)\sum_{j=m}^{n}\left((1/m)\sum_{j=0}^{m-1}(z_{t-j} - \beta_0 - \beta_1(t-j))\right)\left((1/m)\sum_{j=0}^{m-1}(z_{t-j} - \beta_0 - \beta_1(t-j))\right)^T \tag{15}$$

Bierens suggests a test for the null hypothesis that the space of all cotrending vectors $\theta$ has dimension 1, against the alternative that the dimension is zero based on the minimum solution $\lambda_1$ of the generalised eigenvalue problem det($M_1$-$\lambda M_2$)=0. Bierens has also computed the quantiles of the converging distribution (for both detrended and non-detrended data) so that the null hypothesis can be tested at the usual significance levels.

## 4. Database and results

The stock market database employed consists on daily prices from January 1st, 1990 to August 3rd, 1999 obtained from Morgan Stanley Capital International indexes (MSCI) and expressed in local currency. We analyse twenty-three stock markets corresponding to two different geographical areas, Asia (Hong Kong, Indonesia, Japan, Korea, Malaysia, Singapore, Taiwan and Thailand) and Europe (Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Norway, Portugal, Spain, Sweden, Switzerland and United Kingdom).

We also use the corresponding exchange rates against the US dollar, calculated as the mean of the best bid price and best ask price of a limited number of financial institutions negotiated directly among them or through brokers (23 hours, GMT)[2]. The currencies employed are the following: for Asia the *Dollar* (Hong Kong), *Rupiah* (Indonesia), *Yen* (Japan), *Won* (South Korea), *Ringgit* (Malaysia), *Dollar* (Singapore), *Dollar* (Taiwan), *Baht* (Thailand), and for Europe the *Tschelling* (Austria), *Franc* (Belgium), *Krone* (Denmark), *Mark* (Finland), *Franc* (France), *Mark* (Germany), *Dracma* (Greece), *Pound* (Irland), *Lira* (Italy), *Krone* (Norway), *Escudo* (Portugal), *Peseta* (Spain), *Krone* (Sweden), *Franc* (Switzerland) and *Sterling Pound* (United Kingdom). Since some of the level series of exchange rates suffered strong devaluations (significantly Indonesia and Korea), we corrected the return series by eliminating returns exceeding 15%.

First, we will try to determine whether the level series contain a unit root, since this will be a necessary condition to our subsequent analyses. In Table 1 we show the results obtained with the PP test. As we can see, there is clear evidence of the existence of a unit root in the level series, while the returns show stationary. Note that Japan is an exception, the null hypothesis of a unit root can be rejected for both levels and returns. The results obtained reversing the null hypothesis and using the KPSS test coincide, as we can see in Table 2, since there is a clear rejection of stationarity around a level or a trend. Note that, in this case, Japan has an identical behaviour to the other markets.

### 4.1 Linear relationships

For the cointegration analyses between exchange rates and stock market indexes, we employ Johansen (1988) and Shin's test (1994). The results obtained with Johansen's test (Table 3) show that, for the whole period, in virtually all the cases the null hypothesis of no cointegration cannot be rejected. Exceptions are Hong Kong and Japan, where the null hypothesis is rejected at the 5% level. To check for possible changes along the period of study we conduct the analyses for the periods that include the crises (January 1st, 1990 to June 30th, 1993 for European markets and July 2nd, 1997 to August 3rd, 1999 for Asian countries). We find that, again, no European markets show evidence of cointegration between exchange rates and stock market prices, while for the Asian ones there is some evidence for Malaysia and Thailand. Finally, during the periods that do not include the crises July 1st, 1997 to August, 3rd, 1999 for Europe and January 1st, 1990 to July 1st, 1997 for Asia) we find only strong evidence for Japan. In the same line, Shin's test (Table 4) allows to reject the null hypothesis of linear cointegration in all cases. Note that the null hypothesis is rejected in *all* cases and for *all* periods. These results reveal some weakness of the cointegration test for moderate sample sizes, but they also allow to affirm that there is strong evidence against long run relationships between exchange rates and stock markets returns in both crisis and non-crisis periods.

The absence of long run relationships does not rule out short term relationships. For this reason, we will conduct the following analyses. First, we calculated the correlation matrices for both exchange rates and stock market returns[3]. We found significant correlations in the exchange rates of the same geographical area but the correlations were small among the countries of two different geographical areas (Japan and Singapore were exceptions since they also exhibited high correlations with the European markets). Interestingly, Korea showed negative correlation coefficients against all European markets. For the stock market returns, the correlations are significantly higher, the European

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[2] For an extended exposition see Martens and Kofman (1998, p.349).
[3] To save space, we do not include them here; they can be obtained from the authors upon request.

markets seemed to be more integrated than the Asian ones and, again, Japan and Singapore showed higher correlations with the European markets.

Comparing the degree of correlation in the exchange rates before and after the Asian crisis we found that it increases among European countries (an exception is the United Kingdom) while for the Asian countries it happens exactly the opposite. Also, it is interesting to note that the correlation among European and Asian countries reduces during the crises. Regarding the stock market indexes, we find that the correlation increases in the period of the Asian crisis and among European and Asian countries. Note that these facts reveal that the Asian crises produced a segmentation of the exchange rate markets, but an integration of the stock markets of the two-different areas, and also produced integration in the exchange rates as well as stock markets of the countries in the same geographical area.

During the European crisis, we find an increase in the correlation of exchange rates among the European countries as well as with the Asian ones, but the correlation among the Asian countries do not increase. Regarding stock markets, we find stronger correlations among the European countries and Japan, Malaysia, Singapore, Thailand and Taiwan, but not among the European and Asian countries. Note that these facts reveal that the European crisis produced an integration of the exchange rate markets as well as an integration of the stock markets of the two different areas, and also produced a segmentation in the stock markets of the countries in the same geographical area as well as a segmentation in the exchange rate markets among the Asian countries.

The study of short term relationships is completed with the application of the Granger causality tests. Table 5 (second and third columns) shows the results of testing the null hypothesis that exchange rates do not Granger cause stock market returns, as well as the results for testing the null hypothesis that stock market returns do not Granger cause exchange rates returns (columns fourth and fifth). For the European countries, exchange rates seem to cause stock market returns for the cases of Belgium, Denmark, Germany, Ireland, Portugal, Sweden, and United Kingdom while for Italy, Norway, Austria, and Denmark we find evidence in the opposite direction. For the Asian countries, we can verify the existence of bivariate causality in some cases (Korea, Indonesia, Malaysia, Singapore and Thailand). For Hong Kong we find one-way causality (exchange rates returns causes stock markets returns). Overall, we find more evidence in favour of the hypothesis that exchange rates return Granger-cause stock market returns than the opposite. In addition, we find no substantial differences between the two geographical areas.

To see which are the effects of the crises, we re-calculate the statistics for the subperiods described above. The results in favour of exchange rates and stock market causality (Table 6) show that it is not more evident in the Asian case than in the European one, and that the influx goes mainly from the exchange rate market to the stock market. It should be also noted that the evidence in favour of causality in either of the directions is more relevant during the periods of crisis both in Europe (January 1st, 1990 - June 30th, 1993) and Asia (July 2nd, 1997 - August 3rd, 1999).

As a conclusion, we can say that, from a linear point of view, the European and Asian crisis had similar effects on the relationship between exchange rates and stock markets of the countries in the geographical area affected by the crisis, increasing the degree of influence of exchange rate markets on the stock markets.

### 4.2. Nonlinear relationships

To check for nonlinear long run relationships, we apply the Bierens' test (Bierens, 2000)[4]. About the European countries and for the whole period (Table 7), we do not find evidence in favour of nonlinear cointegration for any of the countries. Interestingly, we find significant evidence for the period of crisis (only in the cases of Portugal and Switzerland we reject the null hypothesis) but these relations seem to disappear in the post-crisis period (only for Austria, Ireland, Italy and Sweden we fail to reject the null hypothesis). These results may be a consequence of the increasing instability in the relationship between exchange rates and stock markets in the most recent period when the economic policies of the European countries (especially those belonging to the EMU) have imposed a strict control to some macroeconomic magnitudes (exchange rates) trying to reach a higher degree of convergence. This fact could have led to significant differences since, regarding the stock markets and on the contrary to exchange rates, have experimented a clear revaluation (in the presence, though, of a high volatility).

In relation to the Asian markets, and similarly to the European ones, we do not find significant evidence in favour of nonlinear cointegration for the whole period (the exception is Thailand, where we fail to reject the null hypothesis at the 5% level). We find a number of significant long run relationships during the period preceding the crisis (for half of the countries), which disappear in the crisis period (an exception is Hong Kong). This contrasts with the results obtained in the European case where the opposite behaviour is observed.

The explanation of these results is, if one avoids speculation, quite difficult. At present, there are not testable models, which support or reject the hypothesis of long run nonlinear relationships between exchange rates and stock market prices. What our results seem to suggest is that the period of crises effectively affects this relationship, but in an unpredictable manner (making it stronger or weaker, depending on the particular economy studied). A more simple explanation is that the econometric methodology is not powerful enough to discriminate among competing alternatives.

Now we will focus on the possible existence of nonlinear causal relationships. Since the modified Baek & Brock test may have power against linear alternatives, we employ the procedure suggested by Hiemstra and Jones (1993). Each one of the pairs of exchange rates and stock market returns series is filtered through a VAR system where the number of lags is determined by minimising the Akaike Information Criterion. Then we calculate the residuals and run the nonlinear causality test. Following suggestions by Hiemstra and Jones (1994) we employ embedding dimensions from 2 to 5, and $\varepsilon$ equal to the unconditional standard deviation of the corresponding series. The results for the whole period are shown in Table 8. In Europe, only for Ireland, Sweden and Norway (weak in this case) we find evidence of bivariate causality. We also find causality from exchange rates to stock markets for Greece and Germany, and the opposite for Portugal and Finland (very weak). If we now turn our attention to Asia, it is worth noting that nonlinear causality is much stronger; we find clear bivariate evidence for all the countries except Hong Kong (the stock market leads exchange rates) and Taiwan (no relationship is found). Note that in this case the differences between the two geographical areas are clear: weak evidence for Europe and strong evidence for Asia.

When we analyse the results for subperiods (Table 9, Panels A and B) we find that the evidence of nonlinear relationships is quite different for Asia and similar for Europe. During the crisis period we find seven cases where the exchange rates seem to lead stock

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[4] In the implementation of the test we use the default parameters of the *EasyReg* software of Bierens.

market prices (the exception is Taiwan and for Hong Kong and Thailand the evidence is very weak) and four links in the opposite direction (although only the case of Indonesia seem clear), while for the non crisis period we find no evidence (an exception is Japan where the evidence that exchange rates cause stock market returns is very weak). For Europe, we find evidence that exchange rates cause stock returns in five and six cases for the crisis and non-crisis periods, respectively, and three and four cases for the opposite link, respectively. It should be noted that the values of the Baek & Brock statistic differ significantly along the embedding dimension considered. This could be due to the loss of power of the test for moderate sample sizes, when the nonlinear structure is not strong enough.

Overall, our results suggest that there is some evidence of nonlinear causality between exchange rates and stock returns, and that it is clearer from exchange rates to stock returns, for the Asian markets and for the periods of crisis.

## 5. Conclusions

In this paper, we have analysed short and long term relationships between exchange rates and stock market prices of twenty-three markets corresponding to two different geographical areas. We have extended previous studies by using econometric tests that are useful in detecting nonlinear structure and by specifically accounting for the crises periods.

The existence of linear causality is clear in Asia: for Korea, Indonesia, Malaysia, Singapore and Thailand we find bidirectional relationships while for Hong Kong it is unidirectional. From the nonlinear point of view, we also found bidirectional relationships for Singapore, Malaysia, Thailand, Indonesia and Japan.

In the European case, we found less evidence in favour of causality, and it was manifested only in one direction (from exchange rates to stock markets) for Belgium, Denmark, Germany, Ireland, Portugal, Sweden and United Kingdom. Only for three countries we found evidence of nonlinear causality (Greece, Ireland and Norway). The weak evidence of causality relationships from stock market returns to exchange rates returns in Europe could be explained by the special characteristic of the monetary policy in the EU, and especially due to the implications on economic policy imposed by the European common currency.

Finally, it is interesting to note that, for Asia, short run linear and nonlinear relationships are much clearer in the period that includes both crises. This is not the case for long-run relationships since linear cointegration seem to be absent and nonlinear cointegration is more relevant in the crisis period for Europe and in the non crisis period for Asia.

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**Table 1**
**Phillips-Perron's test**

| | **Exchange rates** | | | | **Stock market prices** | | | |
|---|---|---|---|---|---|---|---|---|
| | Levels | | Returns | | Levels | | Returns | |
| | $\eta_\mu$ | $\eta_\tau$ | $\eta_\mu$ | $\eta_\tau$ | $\eta_\mu$ | $\eta\tau$ | $\eta_\mu$ | $\eta_\tau$ |
| ***Europe*** | | | | | | | | |
| *Austria* | -1.89 | -2.16 | -49.55 | -49.55 | -2.40 | -2.42 | -40.99 | -40.98 |
| *Belgium* | -1.84 | -2.11 | -13326.19 | -13362.47 | 0.74 | -1.81 | -42.74 | -42.80 |
| *Denmark* | -1.97 | -2.20 | -13466.75 | -13493.45 | 0.16 | -1.60 | -44.15 | -44.17 |
| *Finland* | -1.44 | -1.85 | -10925.34 | -10916.57 | 3.27 | 0.76 | -44.14 | -44.17 |
| *France* | -2.05 | -2.32 | -13445.74 | -13471.55 | 1.10 | -1.29 | -46.99 | -47.07 |
| *Germany* | -1.88 | -2.15 | -13150.07 | -13176.05 | 0.34 | -1.79 | -48.83 | -48.89 |
| *Greece* | -1.54 | -2.21 | -13597.66 | -13592.75 | 1.28 | -0.09 | -42.65 | -42.64 |
| *Ireland* | -1.84 | -2.73 | -13628.39 | -13642.58 | 0.48 | -1.98 | -44.61 | -44.62 |
| *Italy* | -1.07 | -2.18 | -13282.61 | -13273.09 | 0.02 | -1.97 | -44.70 | -44.72 |
| *Norway* | -1.57 | -2.47 | -13907.91 | -13917.13 | -1.21 | -1.95 | -44.71 | -44.70 |
| *Portugal* | -1.15 | -2.51 | -13518.66 | -13532.19 | -0.28 | -2.11 | -41.99 | -41.97 |
| *Spain* | -0.90 | -2.57 | -12677.26 | -12679.54 | 0.67 | -1.56 | -44.47 | -44.50 |
| *Sweden* | -1.21 | -1.99 | -12387.44 | -12379.60 | 1.68 | -1.41 | -44.04 | -44.08 |
| *Switzerland* | -2.16 | -2.12 | -11578.43 | -11605.85 | 0.42 | -2.01 | -47.28 | -47.29 |
| *U.Kingdom* | -2.28 | -2.55 | -13146.43 | -13136.60 | 0.78 | -2.11 | -46.06 | -46.08 |
| ***Asia*** | | | | | | | | |
| *HongKong* | -3.32 | -3.14 | -307841.7 | -309409.4 | -1.56 | -1.96 | -47.69 | -47.70 |
| *Indonesia* | n.a. | n.a. | n.a. | n.a. | -1.94 | -2.05 | -39.91 | -39.90 |
| *Japan* | -1.66 | -1.50 | -11452.15 | -11472.32 | -4.74 | -4.16 | -46.89 | -46.97 |
| *Korea* | n.a. | n.a. | n.a. | n.a. | -0.90 | -0.94 | -47.92 | -47.97 |
| *Malaysia* | 0.33 | -0.80 | -15167.01 | -15306.78 | -1.41 | -1.29 | -47.03 | -47.02 |
| *Singapore* | -1.81 | -0.75 | -26621.44 | -27229.40 | -1.08 | -1.41 | -41.14 | -41.13 |
| *Thailand* | -0.21 | -1.38 | -11141.57 | -11169.17 | -0.94 | -1.08 | -42.53 | -42.52 |
| *Taiwan* | -0.27 | -1.40 | -26038.28 | -26100.07 | -1.99 | -3.03 | -48.49 | -48.49 |

**Note**: Phillips-Perron (1988) test for the null hypothesis of a unit root against the alternative of stationarity against a a level ($\eta_\mu$) and a trend ($\eta_\tau$). The critical values at 1% level are -3,4360 ($\eta_\mu$) and -3,9671 ($\eta_\tau$). The number of lags employed is 8.

**Table 2**
**KPSS test**

| | Stock market Indexes | | Exchange Rates | |
|---|---|---|---|---|
| | $\eta_\mu$ | $\eta_\tau$ | $\eta_\mu$ | $\eta_\tau$ |
| ***Europe*** | | | | |
| *Austria* | 3.533 | 3.004 | 4.156 | 2.951 |
| *Belgium* | 21.11 | 5.525 | 3.917 | 2.845 |
| *Denmark* | 19.83 | 5.257 | 3.194 | 2.346 |
| *Finland* | 18.79 | 4.336 | 11.65 | 2.620 |
| *France* | 20.38 | 5.034 | 3.616 | 2.432 |
| *Germany* | 21.13 | 5.437 | 4.152 | 2.960 |
| *Greece* | 16.05 | 4.894 | 23.75 | 3.162 |
| *Ireland* | 22.29 | 5.354 | 9.300 | 1.012 |
| *Italy* | 19.17 | 4.329 | 21.03 | 2.606 |
| *Norway* | 16.98 | 2.873 | 10.62 | 1.508 |
| *Portugal* | 18.37 | 5.051 | 16.25 | 0.924 |
| *Spain* | 20.39 | 5.627 | 21.48 | 1.270 |
| *Sweden* | 24.16 | 5.292 | 15.40 | 1.813 |
| *Switzerland* | 24.11 | 5.186 | 2.908 | 2.759 |
| *U.Kingdom* | 24.93 | 4.891 | 7.038 | 3.162 |
| ***Asia*** | | | | |
| *Hong Kong* | 18.67 | 3.599 | 8.919 | 4.792 |
| *Indonesia* | n.a. | n.a. | n.a. | n.a. |
| *Japan* | 9.180 | 1.913 | 7.580 | 5.526 |
| *Korea* | n.a. | n.a. | n.a. | n.a. |
| *Malaysia* | 7.184 | 4.544 | 10.76 | 4.860 |
| *Singapore* | 8.316 | 4.218 | 10.53 | 5.404 |
| *Thailand* | 6.046 | 5.826 | 14.82 | 4.574 |
| *Taiwan* | 10.55 | 1.764 | 17.06 | 4.857 |

**Note**: KPSS test (Kwiatkowski *et al.* 1992) for the null hypothesis of stationarity around a level ($\eta_\mu$) or a trend ($\eta_\tau$). Critical values for $\eta_\mu$ are 0,463 (5%) and 0,347 (10%), and for $\eta_\tau$ are 0,146 (5%) and 0,119 (10%). The number of lags employed is 8.

**Table 3**
**Johansen´s test**

| | January 1st, 1990-August 3rd, 1999 | January 1st, 1990-June 30th, 1993 | July 1st, 1993-August 3rd, 1999 |
|---|---|---|---|
| ***Europe*** | | | |
| *Austria* | 12.80 | 7.520 | 11.11 |
| *Belgium* | 7.021 | 12.72 | 3.860 |
| *Denmark* | 7.147 | 13.86 | 4.620 |
| *Finland* | 13.75 | 10.61 | 7.000 |
| *France* | 8.607 | 9.330 | 4.760 |
| *Germany* | 6.944 | 14.00 | 5.950 |
| *Greece* | 5.287 | 9.230 | 5.870 |
| *Ireland* | 7.720 | 12.61 | 6.200 |
| *Italy* | 6.258 | 11.12 | 7.460 |
| *Norway* | 6.036 | 7.950 | 7.470 |
| *Portugal* | 6.993 | 10.06 | 4.640 |
| *Spain* | 4.508 | 10.76 | 5.150 |
| *Sweden* | 6.971 | 13.76 | 3.290 |
| *Switzerland* | 5.717 | 6.110 | 6.260 |
| *U.Kingdom* | 6.071 | 7.630 | 13.28 |

| ***Asia*** | January 1st, 1990-August 3rd, 1999 | January 1st, 1990-July 1st, 1997 | July 2nd, 1997-August 3rd, 1999 |
|---|---|---|---|
| *Hong Kong* | 16.97* | 11.88 | 6.670 |
| *Indonesia* | n.a. | 7.910 | n.a. |
| *Japan* | 27.10** | 24.80** | 10.32 |
| *Korea* | n.a. | 9.370 | n.a. |
| *Malaysia* | 3.650 | 8.180 | 17.50* |
| *Singapore* | 6.060 | 8.940 | 9.490 |
| *Thailand* | 3.700 | 14.67 | 35.70** |
| *Taiwan* | 8.390 | 7.490 | 11.650 |

**Note**: Johansen´s test (Johansen, 1988) for the null hypothesis of no cointegration. Critical values are 15,41 (5%) and 20,04 (1%). The number of lags is 8. One asterisk denotes a rejection at the 5% level, while two asterisks denote a rejection at the 1% level.

**Table 4**
**Shin´s test**

| | January 1st, 1990-August 3rd, 1999 | January 1st, 1990-June 30th, 1993 | July 1st, 1993-August 3rd, 1999 |
|---|---|---|---|
| ***Europe*** | | | |
| *Austria* | 4.803 | 7.636 | 0.640 |
| *Belgium* | 20.92 | 1.393 | 11.13 |
| *Denmark* | 21.14 | 4.670 | 13.79 |
| *Finland* | 11.95 | 3.467 | 12.01 |
| *France* | 20.75 | 2.626 | 11.53 |
| *Germany* | 21.62 | 3.275 | 11.19 |
| *Greece* | 6.188 | 0.776 | 2.042 |
| *Ireland* | 14.62 | 5.917 | 14.98 |
| *Italy* | 7.583 | 5.916 | 7.017 |
| *Norway* | 9.632 | 8.652 | 8.678 |
| *Portugal* | 6.368 | 7.976 | 9.690 |
| *Spain* | 7.140 | 3.440 | 7.867 |
| *Sweden* | 12.80 | 4.091 | 16.37 |
| *Switzerland* | 25.85 | 6.692 | 14.18 |
| *U.Kingdom* | 22.05 | 5.756 | 7.200 |

| ***Asia*** | January 1st, 1990-August 3rd, 1999 | January 1st, 1990-July 1st, 1997 | July 2nd, 1997-August 3rd, 1999 |
|---|---|---|---|
| *Hong Kong* | 13.27 | 12.03 | 1.566 |
| *Indonesia* | n.a. | 3.423 | n.a. |
| *Japan* | 7.036 | 1.915 | 1.45 |
| *Korea* | n.a. | 6.697 | n.a. |
| *Malaysia* | 15.53 | 8.486 | 0.815 |
| *Singapore* | 1.151 | 1.869 | 1.579 |
| *Thailand* | 8.441 | 9.097 | 2.752 |
| *Taiwan* | 2.939 | 2.749 | 1.093 |

**Note**: Shin's test (Shin, 1994) for the null hypothesis of linear cointegration. The critical values are 0,314 (5%) and 0,533 (10%). The number of lags is 8.

**Table 5**
**Granger causality tests (whole period)**

| | $H_0$: Exchange rates returns do not cause stock market returns | $H_0$: Stock market returns do not cause exchange rates returns |
|---|---|---|
| ***Europe*** | | |
| *Austria* | 1.20106 | 2.12629 |
| | (0.3026) | (0.0474) |
| *Belgium* | 2.53642 | 1.72605 |
| | (0.0189) | (0.1109) |
| *Denmark* | 3.20756 | 2.51638 |
| | (0.0039) | (0.0198) |
| *Finland* | 0.61384 | 0.67197 |
| | (0.7194) | (0.6724) |
| *France* | 0.81594 | 0.48340 |
| | (0.5574) | (0.8212) |
| *Germany* | 3.51265 | 1.21995 |
| | (0.0018) | (0.2927) |
| *Greece* | 1.23269 | 1.28541 |
| | (0.2862) | (0.2604) |
| *Ireland* | 2.56193 | 1.37056 |
| | (0.0178) | (0.2227) |
| *Italy* | 1.55522 | 2.41744 |
| | (0.1563) | (0.0248) |
| *Norway* | 1.36137 | 2.51946 |
| | (0.2265) | (0.0196) |
| *Portugal* | 2.09153 | 1.80600 |
| | (0.0512) | (0.0941) |
| *Spain* | 0.30482 | 1.21663 |
| | (0.9347) | (0.2944) |
| *Sweden* | 2.19037 | 1.31469 |
| | (0.0412) | (0.2469) |
| *Switzerland* | 1.61999 | 1.98568 |
| | (0.1375) | (0.0644) |
| *U.Kingdom* | 2.36793 | 0.25780 |
| | (0.0277) | (0.9563) |
| ***Asia*** | | |
| *Hong Kong* | 4.26599 | 0.18010 |
| | (0.0003) | (0.9823) |
| *Indonesia* | 8.56782 | 6.56206 |
| | (3.1E-09) | (6.9E-07) |
| *Japan* | 0.32131 | 1.83380 |
| | (0.9261) | (0.0888) |
| *Korea* | 7.82024 | 2.94266 |
| | (2.3E-08) | (0.00729) |
| *Malaysia* | 8.31326 | 4.10937 |
| | (6.1E-09) | (0.0025) |
| *Singapore* | 6.80350 | 4.61421 |
| | (3.6E-07) | (0.0001) |
| *Thailand* | 6.16394 | 3.68670 |
| | (2.0E-06) | (0.0012) |
| *Taiwan* | 1.11398 | 0.95333 |
| | (0.3514) | (0.4555) |

**Note**: Test for the null hypothesis that exchange rates returns do not Granger cause stock market returns and for the null hypothesis that stock market rates returns do not Granger cause exchange rates returns (6 lags). We show the value of the statistic and the significance level (in parentheses). The cases where the null hypothesis is rejected are underlined.

**Table 6. Granger causality test (subperiods)**

| | H$_0$: Exchange rates returns do not cause stock market returns | | H$_0$: Stock market returns do not cause exchange rates returns | |
|---|---|---|---|---|
| ***Europe*** | **January 1st, 1990-June 30th, 1993** | **July 1st, 1993-August 3rd, 1999** | **January 1st, 1990-June 30th, 1993** | **July 1st, 1993-August 3rd, 1999** |
| *Austria* | 0.92211 | 0.27749 | 0.7612 | 2.36178 |
| | (0.47808) | (0.94770) | (0.60001) | (0.02827) |
| *Belgium* | 2.99627 | 1.20208 | 1.07217 | 0.91881 |
| | (0.00665) | (0.30229) | (0.37737) | (0.48022) |
| *Denmark* | 2.03420 | 2.49101 | 2.33653 | 1.51727 |
| | (0.05871) | (0.02109) | (0.03030) | (0.16864) |
| *Finland* | 1.25782 | 0.40576 | 1.61213 | 0.78175 |
| | (0.27443) | (0.87558) | (0.14054) | (0.58421) |
| *France* | 1.37753 | 0.35502 | 0.61279 | 1.12840 |
| | (0.22065) | (0.90723) | (0.72023) | (0.34318) |
| *Germany* | 1.84740 | 3.55053 | 0.73171 | 0.89651 |
| | (0.08712) | (0.00170) | (0.62415) | (0.49648) |
| *Greece* | 0.49563 | 1.73099 | 0.49555 | 2.12930 |
| | (0.81192) | (0.11010) | (0.81198) | (0.04735) |
| *Ireland* | 1.18263 | 2.37531 | 0.63131 | 0.85706 |
| | (0.31319) | (0.02742) | (0.70530) | (0.52593) |
| *Italy* | 2.42589 | 0.55782 | 0.90339 | 2.15989 |
| | (0.02480) | (0.76412) | (0.49166) | (0.04428) |
| *Norway* | 2.03954 | 0.72043 | 3.31639 | 1.08974 |
| | (0.05804) | (0.63317) | (0.00310) | (0.36618) |
| *Portugal* | 2.30337 | 0.87432 | 0.94210 | 1.66244 |
| | (0.03262) | (0.51294) | (0.46382) | (0.12653) |
| *Spain* | 1.00154 | 0.74541 | 0.69078 | 1.68864 |
| | (0.42291) | (0.61311) | (0.65715) | (0.12001) |
| *Sweden* | 2.77526 | 2.09564 | 0.38226 | 2.45107 |
| | (0.01115) | (0.05095) | (0.89059) | (0.02310) |
| *Switzerland* | 1.27182 | 0.86259 | 1.09334 | 1.86379 |
| | (0.26765) | (0.52175) | (0.36440) | (0.08363) |
| *U.Kingdom* | 3.83920 | 0.96413 | 0.51891 | 0.68091 |
| | (0.00087) | (0.44810) | (0.79428) | (0.66512) |
| ***Asia*** | **January 1st, 1990-July 1st, 1997** | **July 2nd, 1997-August 3rd, 1999** | **January 1st, 1990-July 1st, 1997** | **July 2nd, 1997-August 3rd, 1999** |
| *Hong Kong* | 0.48579 | 6.05558 | 0.77337 | 1.46499 |
| | (0.81937) | (3.8E-06) | (0.59082) | (0.18809) |
| *Indonesia* | 1.32793 | 2.61094 | 2.39359 | 2.01837 |
| | (0.24109) | (0.01672) | (0.02623) | (0.06154) |
| *Japan* | 0.67952 | 0.57205 | 0.77076 | 2.53021 |
| | (0.66626) | (0.75272) | (0.59288) | (0.02006) |
| *Korea* | 0.55820 | 2.66112 | 0.79523 | 1.32979 |
| | (0.76384) | (0.01492) | (0.57358) | (0.24185) |
| *Malaysia* | 1.77821 | 2.49422 | 2.10587 | 0.73879 |
| | (0.09978) | (0.02175) | (0.04972) | (0.61857) |
| *Singapore* | 1.41137 | 5.37802 | 0.37694 | 2.41175 |
| | (0.20640) | (0.00002) | (0.89403) | (0.02615) |
| *Thailand* | 6.62417 | 1.38033 | 1.72652 | 1.87542 |
| | (6.0E-07) | (0.22042) | (0.11096) | (0.08306) |
| *Taiwan* | 2.37376 | 0.85326 | 1.31350 | 1.50531 |
| | (0.02743) | (0.52925) | (0.24755) | (0.17416) |

**Note**: Granger´s test for the null hypothesis that exchange rates returns do not Granger cause stock market returns and for the null hypothesis that stock market rates returns do not Granger cause exchange rates returns (6 lags). We show the value of the statistic and the significance level (in parentheses). The cases where the null hypothesis is rejected are underlined.

**Table 7**
**Bierens´s nonlinear cointegration test**

| | January 1st, 1990- August 3rd, 1999 | January 1st, 1990- June 30th, 1993 | July 1st, 1993- August 3rd, 1999 |
|---|---|---|---|
| ***Europe*** | | | |
| *Austria* | 0.34548 | 0.13205** | 0.12570** |
| *Belgium* | 0.15982 | 0.14326** | 0.32165 |
| *Denmark* | 0.23706 | 0.12757** | 0.17193 |
| *Finland* | 0.42461 | 0.08868* | 0.46432 |
| *France* | 0.37462 | 0.14499** | 0.33971 |
| *Germany* | 0.19133 | 0.13697** | 0.24307 |
| *Greece* | 0.32278 | 0.13355** | 0.37169 |
| *Ireland* | 0.19635 | 0.13947** | 0.07329* |
| *Italy* | 0.21540 | 0.11663* | 0.12320** |
| *Norway* | 0.28618 | 0.14083** | 0.33897 |
| *Portugal* | 0.15629 | 0.17632 | 0.16026 |
| *Spain* | 0.18565 | 0.13379** | 0.23875 |
| *Sweden* | 0.21719 | 0.07582* | 0.09411* |
| *Switzerland* | 0.28132 | 0.15588 | 0.18564 |
| *U.Kingdom* | 0.14308 | 0.11512* | 0.18441 |
| | **January 1st, 1990- August 3rd, 1999** | **January 1st, 1990- July 1st, 1997** | **July 2nd, 1997- August 3rd, 1999** |
| ***Asia*** | | | |
| *Hong Kong* | 0.26736 | 0.17034 | 0.1081** |
| *Indonesia* | n.a. | 0.11255* | n.a. |
| *Japan* | 0.35987 | 0.44228 | 0.19239 |
| *Korea* | n.a. | 0.48124 | n.a. |
| *Malaysia* | 0.17960 | 0.13881** | 0.26528 |
| *Singapore* | 0.17810 | 0.19189 | 0.19379 |
| *Thailand* | 0.13412** | 0.10092* | 0.26667 |
| *Taiwan* | 0.23732 | 0.11849* | 0.22020 |

**Note**: Bierens's test (Bierens, 2000) for the null hypothesis of nonlinear cotrending. Critical values 0.119 (10%) and 0.151 (5%). A failure to reject the null hypothesis is noted by *(10%) and **(5%)

**Table 8**
**Hiemstra and Jones nonlinearity test (whole period)**

| | *m=2* | *m=3* | *m=4* | *m=5* | | *m=2* | *m=3* | *m=4* | *m=5* |
|---|---|---|---|---|---|---|---|---|---|
| ***Europe*** | | | | | | | | | |
| *Ausie* | 0.028 | -0.044 | 0.064 | 1.079 | *Ausei* | 0.663 | 0.105 | 1.381 | 1.231 |
| *Belie* | 0.116 | -0.21 | -1.053 | -0.734 | *Belei* | -0.556 | -0.64 | 0.712 | 1.488 |
| *Denie* | -0.558 | 0.099 | 0.763 | 0.783 | *Denei* | 0.212 | 0.008 | 0.156 | -0.056 |
| *Finie* | 1.478 | 2.262** | 1.275 | 1.523 | *Finei* | 0.383 | 0.431 | 0.796 | 0.061 |
| *Fraie* | 0.197 | -0.687 | 0.223 | 0.25 | *Fraei* | 0.53 | 0.663 | 1.565 | 1.617 |
| *Gerie* | 0.599 | 0.036 | 0.209 | 0.763 | *Gerei* | 1.927** | 0.723 | 1.644 | 1.741** |
| *Greie* | -0.08 | -0.116 | 1.19 | 1.733** | *Greei* | 2.505** | 3.027** | 3.284** | 2.74** |
| *Irlie* | 1.942** | 2.538** | 2.889** | 2.509** | *Irlei* | 3.345** | 3.152** | 2.99** | 2.701** |
| *Itaie* | 0.793 | -1.273 | -0.247 | 0.123 | *Itaei* | 0.333 | 0.199 | 1.058 | 1.367 |
| *Norie* | 1.364 | 1.694** | 1.086 | 1.653 | *Norei* | 3.002** | 2.015** | 1.04 | 1.262 |
| *Porie* | 1.225 | 1.964** | 1.719** | 1.776** | *Porei* | 0.314 | 0.118 | 0.607 | 0.561 |
| *Spaie* | 2.714** | 1.212 | 1.662 | 1.683 | *Spaei* | 0.485 | -0.096 | 0.854 | 0.246 |
| *Sweie* | 1.601 | 2.119** | 2.269** | 2.193** | *Sweei* | 1.255 | 1.961** | 2.248** | 3.241** |
| *Swiie* | 0.881 | 0.107 | -0.353 | -0.693 | *Swiei* | 0.658 | -0.092 | 0.892 | -0.275 |
| *Ukgie* | 0.536 | 0.32 | 0.327 | 0.822 | *Ukgei* | 0.085 | -0.728 | -1.21 | -0.823 |
| ***Asia*** | | | | | | | | | |
| *Hkgie* | 1.231 | -0.441 | 0.065 | -0.101 | *Hkgei* | -1.195 | -1.868** | -2.803** | -2.278** |
| *Indie* | 5.766** | 5.519** | 5.166** | 4.303** | *Indei* | 5.711** | 6.000** | 5.71** | 4.805** |
| *Japie* | 1.966** | 2.108** | 1.693** | 1.767** | *Japei* | 2.722** | 3.239** | 2.395** | 2.011** |
| *Korie* | 4.232** | 4.834** | 3.88** | 3.18** | *Korei* | 4.518** | 5.076** | 4.162** | 3.366** |
| *Malie* | 4.257** | 3.723** | 2.789** | 1.339** | *Malei* | 5.647** | 5.028** | 4.222** | 2.907** |
| *Sinie* | 5.021** | 4.355** | 3.303** | 3.034** | *Sinei* | 4.744** | 3.547** | 4.082** | 3.306** |
| *Thaie* | 5.136** | 5.209** | 4.589** | 3.745** | *Thaei* | 5.689** | 6.137** | 5.897** | 5.63** |
| *Tawie* | 0.621 | 0.378 | 0.86 | 1.454 | *Tawei* | 0.026 | 0.191 | 0.591 | 0.505 |

**Note**: Hiemstra and Jones (1994) test for the null hypothesis of no nonlinear causality. In first column we label the cases by using three letters of the corresponding countries, the last two letters indicate whether the null hypothesis is that the exchange rates do not nonlinearly cause stock market returns (*ei*) or stock market returns does not nonlinearly cause exchange rates returns (*ie*). The rejection of the null hypothesis at the 5% level is indicated by **.

**Table 9, Panel A**
**Crisis**
**Hiemstra and Jones nonlinearity test (subperiods)**

| | *m=2* | *m=3* | *m=4* | *m=5* | | *m=2* | *m=3* | *m=4* | *m=5* |
|---|---|---|---|---|---|---|---|---|---|
| ***Europe*** | | | | | | | | | |
| *Ausie* | -2.484 | -2.052 | -1.512 | -2.071 | *Ausei* | 0.645 | 0.031 | 0.398 | 0.261 |
| *Belie* | -0.442 | 0.516 | -0.54 | -0.403 | *Belei* | 0.844 | 0.127 | 1.231 | 1.526 |
| *Denie* | -0.017 | -0.353 | 0.771 | 0.069 | *Denei* | 1.079 | 0.069 | -0.019 | 0.17 |
| *Finie* | 2.983** | 2.572** | 2.154** | 2.18** | *Finei* | 2.477** | 1.339 | 1.229 | 0.137 |
| *Fraie* | 0.175 | -1.482 | -0.261 | -0.812 | *Fraei* | 0.847 | 0.575 | 1.868 | 2.102** |
| *Gerie* | -1.235 | -1.796 | -0.961 | -0.391 | *Gerei* | 1.134 | 0.696 | 0.619 | 2.043** |
| *Greie* | -0.872 | -2.066 | -1.645 | -1.451 | *Greei* | 0.664 | -0.251 | 0.249 | -0.382 |
| *Irlie* | 0.127 | 0.703 | 1.221 | 0.634 | *Irlei* | 1.933** | 1.299 | 0.394 | 0.807 |
| *Itaie* | 2.025** | 0.412 | 0.785 | 1.556 | *Itaei* | 0.722 | 1.341 | 1.819 | 1.621 |
| *Norie* | 1.98** | 1.81 | 1.236 | 1.51 | *Norei* | 1.525 | 0.975 | -0.607 | 0.58 |
| *Porie* | -1.271 | -0.158 | 0.507 | 0.014 | *Porei* | 0.047 | -0.405 | -0.299 | -0.406 |
| *Spaie* | 1.656 | -0.036 | 0.174 | 0.069 | *Spaei* | 1.161 | 0.225 | 1.357 | 0.709 |
| *Sweie* | 0.299 | 0.303 | 0.483 | 0.092 | *Sweei* | 2.294** | 2.173** | 2.122** | 2.498** |
| *Swiie* | 0.196 | -0.048 | -0.964 | -1.735 | *Swiei* | 0.384 | 0.177 | 0.461 | 0.547 |
| *Ukgie* | 0.266 | -0.494 | -1.28 | -1.702 | *Ukgei* | 0.436 | -1.007 | -1.538 | -2.49 |
| ***Asia*** | | | | | | | | | |
| *Hkgie* | 1.816 | 0.967 | 1.193 | 0.768 | *Hkgei* | 2.206** | 0.742 | 1.371 | 2.25 |
| *Indie* | 2.159** | 1.35 | 1.031 | 2.321** | *Indei* | 1.962** | 1.342 | 1.85** | 2.51** |
| *Japie* | -0.158 | -0.702 | -0.458 | -0.297 | *Japei* | 1.697 | 2.558** | 1.863 | 2.144** |
| *Korie* | 2.205** | 1.728 | 1.175 | 1.292 | *Korei* | 1.745 | 1.836 | 1.562 | 2.345** |
| *Malie* | 0.71 | 1.34 | 1.343 | 1.32 | *Malei* | 1.069 | 1.854 | 2.006** | 2.67** |
| *Sinie* | 1.064 | 1.421 | 1.574 | 2.192** | *Sinei* | 0.768 | 1.371 | 1.413 | 2.197** |
| *Thaie* | 0.949 | 1.795 | 2.127** | 1.017 | *Thaei* | 2.144** | 1.721 | 1.484 | 1.968** |
| *Tawie* | 0.106 | 1.079 | 0.538 | -0.477 | *Tawei* | -0.421 | 0.324 | -0.04 | -0.247 |

**Note**: Hiemstra and Jones (1994) test for the null hypothesis of no nonlinear causality. In first column we label the cases by using three letters of the corresponding countries, the last two letters indicate whether the null hypothesis is that the exchange rates do not nonlinearly cause stock market returns (*ei*) or stock market returns does not nonlinearly cause exchange rates returns (*ie*). The rejection of the null hypothesis at the 5% level is indicated by **.

**Table 9, Panel B**
**No crisis**
**Hiemstra and Jones nonlinearity test (subperiods)**

| | *m=2* | *m=3* | *m=4* | *m=5* | | *m=2* | *m=3* | *m=4* | *m=5* |
|---|---|---|---|---|---|---|---|---|---|
| ***Europe*** | | | | | | | | | |
| *Ausie* | 1.362 | 0.951 | 0.434 | 1.674 | *Ausei* | -0.735 | -0.622 | 0.652 | 0.889 |
| *Belie* | 0.457 | 0.01 | -0.126 | 0.185 | *Belei* | -0.719 | 0.023 | 0.598 | 1.318 |
| *Denie* | -0.349 | 0.792 | 0.736 | 1.19 | *Denei* | 0.386 | 1.027 | 1.159 | 0.695 |
| *Finie* | 0.375 | 1.08 | 0.651 | 1.124 | *Finei* | -0.591 | 0.497 | 1.007 | 1.286 |
| *Fraie* | 0.691 | 0.948 | 1.243 | 1.191 | *Fraei* | 0.325 | 0.89 | 1.824 | 1.997* |
| *Gerie* | 1.876 | 1.33 | 1.531 | 1.956 | *Gerei* | 2.351* | 1.716 | 2.459* | 2.337* |
| *Greie* | -0.426 | 0.658 | 1.759 | 2.326* | *Greei* | 1.803 | 3.457* | 3.232* | 3.369* |
| *Irlie* | 1.667 | 1.635 | 1.217 | 1.067 | *Irlei* | 1.95* | 2.357* | 2.489* | 1.594 |
| *Itaie* | 0.982 | -0.112 | 1.041 | 1.583 | *Itaei* | 0.828 | -0.123 | 1.057 | 1.57 |
| *Norie* | -0.613 | 0.188 | -0.582 | -0.212 | *Norei* | 2.359* | 2.015* | 1.563 | 1.065 |
| *Porie* | 2.447* | 2.633* | 2.096* | 2.165* | *Porei* | 0.741 | 1.332 | 1.953 | 1.722 |
| *Spaie* | 2.697* | 2.354* | 2.107* | 2.031* | *Spaei* | 0.417 | 0.943 | 0.953 | 1.193 |
| *Sweie* | 2.096* | 2.79* | 2.216* | 2.44* | *Sweei* | -0.331 | 1.005 | 1.63 | 2.49* |
| *Swiie* | 1.648 | 0.601 | 0.119 | -0.262 | *Swiei* | 0.686 | 0.203 | 1.003 | -0.099 |
| *Ukgie* | 1.292 | 1.431 | 1.103 | 2.077* | *Ukgei* | -0.203 | -0.021 | -1.151 | -0.349 |

| | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| ***Asia*** | | | | | | | | | |
| *Hkgie* | 1.41 | 0.302 | 0.625 | 0.499 | *Hkgei* | -0.794 | -1.011 | -2.092 | -1.865 |
| *Indie* | 1.577 | 1.73 | 1.484 | 1.468 | *Indei* | 1.225 | 0.821 | -0.574 | -0.662 |
| *Japie* | 1.578 | 1.673 | 1.143 | 1.153 | *Japei* | 2.168* | 1.704 | 1.245 | 0.896 |
| *Korie* | -1.44 | -2.957 | -2.144 | -1.909 | *Korei* | -1.384 | -1.218 | 0.228 | 0.544 |
| *Malie* | -0.477 | -0.257 | -0.2 | 0.486 | *Malei* | 0.431 | 0.307 | 0.483 | 0.489 |
| *Sinie* | 1.712 | -0.148 | -0.306 | 0.385 | *Sinei* | 0.713 | -0.296 | 0.118 | 0.021 |
| *Thaie* | 1.161 | 1.613 | 0.641 | 0.316 | *Thaei* | 2.32 | 1.233 | -0.031 | -0.941 |
| *Tawie* | 0.613 | 0.62 | 1.237 | 1.322 | *Tawei* | -0.866 | -1.731 | -0.864 | -0.313 |

**Note**: Hiemstra and Jones (1994) test for the null hypothesis of no nonlinear causality. In first column we label the cases by using three letters of the corresponding countries, the last two letters indicate whether the null hypothesis is that the exchange rates do not nonlinearly cause stock market returns (*ei*) or stock market returns does not nonlinearly cause exchange rates returns (*ie*). The rejection of the null hypothesis at the 5% level is indicated by *.