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 markets prices are a relevant factor in the transmission of the crises. Using daily exchange rates and stock index prices of the last decade (1990-1999) we analyze the interactions between the stock market and exchange rates returns of twenty three countries of two diferent geographical areas (Asia and Europe). Our results suggest that short term relationships seem to be more relevant than long term, that it is more relevant presence of linear and nonlinear causality in the asian countries than in the european and that the periods of crisis affect asymmetrically the relationship between exchange rates and stock market prices.*

1. Introduction

The intense process of globalization of the financial and banking system as well as the increasing degree of economic and financial integration of the countries has led to the appearance of *global financial crisis*. This term refers to situations of crises, which in their begining, are concetrated in a specific geographical area but that in short time provoke effects in other financial markets around the world. Global financial crises are very easy to remember since, during the nineties, the financial system has experimented several of them: the European crisis of 1992, the Asian crisis of 1997 and the Russian and Latin crises of

1998. Perhaps, for the developed countries, the most severe crises were first two and these will be the ones analyzed in this paper.

The chronology of the European and Asian crises is well known and we will just remember some of the main aspects. In the case of the European crisis the factors that caused can be summarized in two: the uncertainty about the viability of the European Monetary Union (EMU), after the Danish “no” to the Maastrich Treaty (june 2nd. 1992), and the asynchrony of the cycles of the european economies which made difficult the coordination of economic policies. Both factors should be framed in a context characterized by free mobility of capitals as well as by the exchange rate mechanism of the European Monetary System (EMS), that is, a system of fixed exchange rates but anchored to the Deutsche Mark.

The detonative of the crisis is generally attributed to need of the German government to employ an expansive fiscal policy to ease unification and a restrictive monetary policy with the objective of containig inflation. This led to an increase of the German interest rates and a strong apreciation of the Deutsche Mark (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 inside the bands, while the situation of their economies (in recession or in a deccelerating process) would require low interest rates. From this moment, the EMS experimented important currency *shocks* which led to the abandon of the exchange rates discipline of the Italian Lira and British Pound, the devaluation of the Portuguese Escudo and the Spanish Peseta (two times), and the abandon of the Scandinavian currencies of the fixing against the Deutsche Mark. In this context, the respective stock markets experimented a clear disparity.

On the other side, the Asian crisis (which lasted from july to october 1997) had its origins on speculative movements against the Thai *bath*. It is interesting to note that, in the begining, this crisis did not provoke important effects on the stock markets of the area but a chain of devaluations and sustitution processes of the exchange rates systems. The Asian crisis was composed of five sub-crises: first in Thailand, then in Phillipines, Indonesia and

Malaysia, and finally in South Korea and Hong Kong. It is after the crash of the Hong-Kong stock market, when the crisis affected the stock markets around the world, which generally experimented important downs.

It is important to note that the countries more severly affected by the crisis (South Korea, Thailand, Malaysia and Indonesia) had experimented significant growth rates along the last decades, their investment rates were over 30% of the GNP and these economies also had significant saving rates ahnd public superavit. Then, which were the causes of the crises?. Basically, the excessive flows of incoming capitals that experimented these economies, in form of portfolio capitals or in form of loans of the international banking system, which provoked a strong appreciation in the real exchange rate (specially 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, the high current deficit and the 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 free their exchange rate and, at the end of the same year, the *Thai baht* has experimented a lose in its value of about 93%, and the SET had lost a 34% (in US dollars), compared to june 1997. The problem with the *baht* put in serious doubt the viability of the exchange rates systems of the other countries, the problems were rapidily extended to Indonesia, Malaysia and Phillipines, making clear the structural deficiencies of these economies. The measures adopted to reduce liquidity in Indonesia were not enought to stop the increasing pressures on the exchange rates 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 secundary effects propagated to other countries such as Hong Kong and Japan. The strong pressures on the Hong Kong dollar in october provoked a significant increase on the interest rates which 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 preassures on the *Won* intensified along october, after the attack of the Hong Kong dollar, and the stock market experimented a crash as a consequence of the lack of confidence in the future of the economy and the dificulties renew the external debt.

As we see, the two mentioned crises share similar origins (both had as their detonative on problems related to the stability of the respecting exchange rates) and effects (the currencies crises were rapidily transmitted to the stock markets). This motivates our study. We will try to see if there is empirical evidence in favor of a relation between exchange rates and stock market prices and, if so, we will also try to determine the direction and nature of the relationship. Also, we are interested in analyzing whether markets exhibit a diferent behavior in the period of crises and if, from the empirical point of view, there are significant diferences between the European and Asian crises.

Our study extends previous analyses in two senses. First, on the contrary of several other studies, we analyze causality and cointegration between exchange rates and stock market prices in a linear as well as a nonlinear context. This is justified by the fact that that there is now considerable evidence favouring the hipothesis that univariate financial time series may contain significant nonlinear components that could be transmited 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 of time. Here we will expand in time and space previous analyses by analyzing twenty three markets in two diferent monetary areas (Europe and Asia) along the period 1990-1999 which also permit us to we analyze the incidence of the exchange rates crises on the bivariate dynamical relationship between exchange rates and stock market prices.

The remaining of the paper is structured as follows: in the second section we present the theoretical models proposed in the literature to analyze the bivariate relationships between exchange rates and stokmarket prices. The econometric methodology is described in section three, whwre we briefly describe the test routinely employed in the literature as well as some others not so common, which will be also employed in this paper. The database used and the results obtained will be covered section four. Conclusions and references close the paper.

2. Theoretical setup

As we have previously mentioned, from our brief review of the European and Asian crises, exchange rates 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, also descends the price of its stock. Under this point of view, the relation (Jorion, 1990) can be expressed as:

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

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 plausibly quite diferent 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 implicattion 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 μ_i is a random perturbation. In this sense, the existing relationship between exchange rates and stock market prices can be analyzed in a broader sense by:

[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.

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

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

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

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

where D_{st} is the variation of the real exchange rate, DRS_t is the diferential (domestic minus foreign) of the stock market return, Di_t is the variation of the interest rate diferential and ε_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 descense 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 led the process.

Under another perspective, which is known as the *traditional hypothesis*, exchange rates liderate the process. An appreciation of the exchange rate would reduce the competitivity of the exporters, which will have a negative impact on stock markets. On the other side, an apreciation of the exchange rates reduces costs and has a possitive impact on the exchange rates (Granger *et al*, 1998). In a macroeconomic context, the existing relationship between exchange rates and stock market prices would be fully capturated 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 on the sign of the realtionship. For example, Smith (1992) and Solnik (1987), consider that this relation is possitive (in contrast with economic theory) while in others it is negative or mixed, depending on the market (Granger *et al,* 1998). Nevertheless there is some consensus that there exist a relationship between exchange rates and stock market returns. For example, Roll (1979) stablishes that this relationship can be analyzed in the contex of the Law of One Price. The price of the assets in the stock market, due to the rapid reaction of these markets, sare 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 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 rates markets and that the risk premium has predictive power on stock market returns.

Heston and Rouwenhorst (1994) analyze the causes of the diference in assets prices between 1978 and 1992 in twelve markets and show that only part of these diferences can be explained by exchange rates variations. Nevertheless, they consider that the estimations are sufficiently important 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 that shows that the change from a particular exchange regime to another provokes effects on the prices of assets proportionaly bigger than the ones caused by real interest rates. Copeland and Copeland (1998) conclude that exchange rates are an explicative factor, significant and independent, of stock market returns, while Chelley-Steeley *et al* (1998) conclude that the reduction or desaparition of exchange rates controls provoke 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 diferences 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 the affect to market mechanisms such as price discovery, volatility transmission, 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 others have suggested that they are inconsistent with market efficiency).

Our analyses will adopt a linear as well as a nonlinear perpective. There is now considerable evidence in favor of nonlinearity in high frequency financial time series (e.g. Hsieh, 1991) but, surprisingly, the vast majority of the studies analyze 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, assuming that agents and price mechanisms of stock and exchange rates markets are not the same, the aparition 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) stablishes the existence of nonlinear relationships between fundamental macroeconomic variables and the exchange rartes. There is a number of studies that have test the model, finding evidece against nonlinearity (De Jong, 1994; Meese y Rose, 1990) or in favor of it Ma and Kanas (2000). If we consider the evidence found in the mentioned studies (Roll, 1979; Ong and Izan, 1999), which stablish 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 summarized in equations (1)-(4) to the nonlinear case.

3. Econometric methodology

Basically, our research will analyze 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. Finaly, 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 (1988) (PP, hereafter), which is robust to some forms of heteroskedasticity, quite common in financial data. Since, it is known that both test tend to accept the null too frequently, we will also employ the KPSS test (Kwiatkowski *et al,* 1992) for wich stationarity around a level of a trend is the null. The ADF and PP tests are well known, and we will not describe them here, instead we will briefly describe the more recent KPSS test.

Let X_t, $t=1,2,....T$, be the time series under analysis. Assume that ot 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, ε_t:

$$X_t = \xi\, t + r_t + \varepsilon_t \qquad (5)$$

where $u_t \sim iid\ (0, \sigma_u^2)$. The null hipothesis is $H_0 : \sigma_u^2 = 0$ while the alternative is $H_A : \sigma_u^2 > 0$. Under the null, X_t is stationary around a trend and, in the case that $\xi = 0$, X_t is stationary around a level (r_0). KPSS show that under the null, 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 whether we regress X_t on a trend or not (KPSS also provide in the critical values of the test, which are used here). The KPSS test is here employed looking for robustness: since standard statistical testing is biased through accepting the null if no strong evidence against it is found, whith 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 diferences 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 test 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 while it the Shin’s test it is the existence of linear cointegration. Againg, we will follow both approaches to asses the results.

The method proposed by Johansen and Juselius is done in a multivariate maximum likelihood framework. It employs the likelihhod 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:

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

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

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

where ε_t is a random k-dimensional vector with mean zero and non singular covariance matrix Σ and μ_{i} is a vector of constant terms. Γ_i = - I +Π_1 + ...+ Π_i, i=1,.......,n . It is considered that the roots of the characteristic implicit polynomio are outside the unit circle. The interesting situations appear 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 Γ_n can be expressed as $\alpha\beta'$, where both α and β are $(k \times r)$ matrices of whole rank. The r first rows of β' are the r cointegration vectors, while the elements of α are the weightings of the cointegrating vectors in the equations. With the condition that none of the elements of X_t is integrated or order higher than one, the maximum likelihhod estimation of the base of the cointegration space is given by the empirical canonical variables of X_{t-n} with respect to ΔX_t, corrected by the short run dynamics and by the determinstic components. The number of cointegration relationships is the same as the number of significant canonical correlations and their significance can be tested by a sequence of likelihood ratios whose distribution asimptotically converges to a brownian bridge.

Additionally, we employ Shin's test (Shin, 1994) to detect the existence of cointegration. This test is a mutivariate version of the KPSS test, but now the residuals, ε_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 of the Johansen test is no cointegration, one would accept the null 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 correaltion matrices) and will be later extended to a dynamical perspective. Since correlation does not reveal which is the dynamical relationship between the variables and which market is the leader, we will employ Granger causality tests. To do this, we build vector autoregressive models:

$$\Delta y_{1t} = \alpha_0 + \sum_{i=1}^{n_1} \alpha_{1i} \Delta y_{1t-i} + \sum_{i=1}^{n_2} \alpha_{2i} \Delta y_{2t-i} + \varepsilon_{1t}$$

$$\Delta y_{2t} = \beta_0 + \sum_{i=1}^{n_3} \beta_{1i} \Delta y_{2t-i} + \sum_{i=1}^{n_4} \beta_{2i} \Delta y_{1t-i} + \varepsilon_{2t} \qquad (10)$$

where Δy_{1t} and Δy_{2t} are the exchange rates and stock market returns, respectively, and ε_{1t} and ε_{1t} are random independent variables. The Granger test is stablished as an F-test on the parameters of the model, the rejection of the null hipothesis 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 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 analyzing the temporal relations between cash and futures markets employ *linear* techniques, looking for *linear* relations between the variables. Consequently, failing to reject the null of independence can only be interpreted as evidence against *linear* relations, and not the absence of other types of dependence. Here we analyze 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 now it seems quite plausible that univariate financial time series may contain significant nonlinear components (e.g. Hsieh, 1991) that could be transmited from one market to another, the second one 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 dependences 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 no provide further information to predict X. Standard Granger bivariate causality tests are usually implemented by estimating a vector autorregression model (VAR) and testing (by an standard χ^2 or F-test) if the estimated coefficients are jointly signifficantly 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 that 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:

$$Pr(\| \boldsymbol{X}_t^m - \boldsymbol{X}_s^m \| < \varepsilon \;/\; \| \boldsymbol{X}_{t-Lx}^{Lx} - \boldsymbol{X}_{s-Lx}^{Lx} \| < \varepsilon, \| \boldsymbol{Y}_{t-Ly}^{Ly} - \boldsymbol{Y}_{s-Ly}^{Ly} \| < \varepsilon) =$$
$$= Pr(\| \boldsymbol{X}_t^m - \boldsymbol{X}_s^m \| < \varepsilon \;/\; \| \boldsymbol{X}_{t-Lx}^{Lx} - \boldsymbol{X}_{s-Lx}^{Lx} \| < \varepsilon)$$

(11)

being $\| \cdot \|$ the suprem norm.

The left hand side of equation (11) is the conditional probability that two arbitrary m-histories of $\{X_t\}$ are whithin a distance smaller that ε, given that the corresponding Lx-histories of $\{X_t\}$ and the Ly-histories of $\{Y_t\}$ are whithin 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 whithin a distance smaller of ε of each other, given that the corresponding Lx-histories of $\{X_t\}$ are at a distance smaller than ε.

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

$$Pr(|\mathbf{X}_t^m - \mathbf{X}_s^m| < \varepsilon \,|\, |\mathbf{X}_{t-Lx}^{Lx} - \mathbf{X}_{s-Lx}^{Lx}| < \varepsilon, |\mathbf{Y}_{t-Ly}^{Ly} - \mathbf{Y}_{s-Ly}^{Ly}| < \varepsilon) =$$
$$= Pr(|\mathbf{X}_t^m - \mathbf{X}_s^m| < \varepsilon, |\mathbf{X}_{t-Lx}^{Lx} - \mathbf{X}_{s-Lx}^{Lx}| < \varepsilon, |\mathbf{Y}_{t-Ly}^{Ly} - \mathbf{Y}_{s-Ly}^{Ly}| < \varepsilon) /$$
$$/ Pr(|\mathbf{X}_{t-Lx}^{Lx} - \mathbf{X}_{s-Lx}^{Lx}| < \varepsilon, |\mathbf{Y}_{t-Ly}^{Ly} - \mathbf{Y}_{s-Ly}^{Ly}| < \varepsilon)$$

and also since

$$Pr(|\mathbf{X}_t^m - \mathbf{X}_s^m| < \varepsilon, |\mathbf{X}_{t-Lx}^{Lx} - \mathbf{X}_{s-Lx}^{Lx}| < \varepsilon) = Pr(|\mathbf{X}_{t-Lx}^{m+Lx} - \mathbf{X}_{s-Lx}^{m+Lx}| < \varepsilon)$$

then equation (11) can be re-expressed as

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

where

$$C1(m+Lx,Ly,\varepsilon) = Pr(|\mathbf{X}_{t-Lx}^{m+Lx} - \mathbf{X}_{s-Lx}^{m+Lx}| < \varepsilon, |\mathbf{Y}_{t-Ly}^{Ly} - \mathbf{Y}_{s-Ly}^{Ly}| < \varepsilon)$$
$$C2(Lx,Ly,\varepsilon) = Pr(|\mathbf{X}_{t-Lx}^{Lx} - \mathbf{X}_{s-Lx}^{Lx}| < \varepsilon, |\mathbf{Y}_{t-Ly}^{Ly} - \mathbf{Y}_{s-Ly}^{Ly}| < \varepsilon)$$
$$C3(m+Lx,\varepsilon) = Pr(|\mathbf{X}_{t-Lx}^{m+Lx} - \mathbf{X}_{s-Lx}^{m+Lx}| < \varepsilon)$$
$$C4(Lx,\varepsilon) = Pr(|\mathbf{X}_{t-Lx}^{Lx} - \mathbf{X}_{s-Lx}^{Lx}| < \varepsilon)$$

An useful estimator of the probabilities involved in equation (2) (of $C3$ and $C4$, 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 $n_m=n-m+1$, $\varepsilon \in R^+$ and $I\_(x_t^h,x_s^h)=1$ iff $|x_t^h-x_s^h|<\varepsilon$, with $/\cdot/$ the suprem 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

$$\sqrt{n}\left(\frac{C1(m+Lx,Ly,\varepsilon)}{C2(Lx,Ly,\varepsilon)}-\frac{C3(m+Lx,\varepsilon)}{C4(Lx,\varepsilon)}\right)\sim_a N(0,\sigma^2(m,Lx,Ly,\varepsilon))$$

where $\sigma^2(m,Lx,Ly,\varepsilon)$ 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 adecuate size and power for moderate sample sizes.

Also, in the context on 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, 1999; Corradi *et al*, 2000). Bierens (1999) has sugggested the term *nonlinear cotrending* to analyze 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-variate time series process, u_t is a k-variate stationary process with mean zero and $f(t)$ is a deterministic k-variate nonliner trend. Nonlinear co-trending 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 on the basis of the generalized 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 β_0 and β_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 suggest a test for the null that the space of all cotrending vectors θ has dimension 1 against the alternative that the dimension is zero based on the minimum solution λ_1 of the generalized eigenvalue problem det(M_1-λM_2)=0, Bierens has also computed the quantiles of the converging distribution (for both detrended and non-detrended data) so that the null can be tested at the usual significance levels .

4. Database and results

The stock market database employ consists on daily prices from 1/1/1990 to 8/3/1999 obtained from Morgan Stanley Capital International indexes (MSCI) and expressed in local currency. We analyze twenty-three stock markets corresponding to two diferent 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 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 returns 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 in favor of the existence of a unit root in the level series while the returns show stationary. Note that Japan is an exception, the null of a unit root can be rejected for both levels and returns. The results obtained reversing the null and using the KPSS test coincide with preceeding, as we can see in Table 2, there is a clear rejection of stationarity around a level or a trend. Note that, in this case, Japan has a identical behavior 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 of no cointegration cannot be rejected. Exceptions are Hong Kong and Japan where the null is rejected at the 5%. To check for posible changes along the period of study we conduct the analyses for the periods that include the crises (1/1/90 to 6/30/93 for European markets and 7/2/97 to 8/3/99 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

[2] For an extended exposition see Martens and Kofman (1998, p.349).

evidence for Malaysia and Thailand. Finally, during the periods that do not include the crises (7/1/93 to 8/3/99 for Europe and 1/1/90 to 7/1/97 for Asia) we find only strong evidence for Japan. In the same line, Shin's test (Table 4) allows to reject the null of linear cointegration in all the cases. Note that the null is rejected in *all* the cases and for *all* the periods. These results reveal some weakness of the cointegration test for moderate sample sizes but they also allows 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 our 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 diferent geographical areas (Japan and Singapore were exceptions since they also exhibited high correlations with the european markets). Interestingly Korea showed negative correlation coeficients against all the European markets. For the stock market returns, the correlations are significantly higher than in the preceeding case, the European 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 happened exactly the opposite, Also, it is interesting to note that the correlation among European and Asian countries reduces during the crises. Respecting the stock market indexes, we find that the correlation increases in the period of the Asian crisis an also among European and Asian countries. Note that these facts reveal that the Asian crises produced a segmentation of the exchange rates markets but an integration of the stock markets of the two-diferent areas and also produced integration

[3] To save space, we do not include them here; they can be obtained from the authors upon request.

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. Respecting 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 rates markets as well as an integration of the stock markets of the two diferent 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 rates 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 column) shows the results of testing the null that exchange rates do not Granger cause stock market returns as well as the results for testing the null 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 oposite 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 favor of the hipothesis those exchange rates returns Granger-cause stock market returns than the opposite. Also, we find no substantial differences between the two geographical areas.

To check which are the effects of the crises, we re-calculate the statistics for the subperiods described above. The results in favor 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 exhange rate market to the stock market. It should be also noted that the evidence in favor of causality in either of the directions is more relevant during the periods of crisis both in Europe (1/1/90 - 6/30/93) and Asia (7/2/97 - 8/3/99).

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 rates markets on the stock markets.

4.2. Nonlinear relationships

To check for nonlinear long run relationships, we apply the Bierens' test (Table 7). In the implementation of the test we use the default parameters of the *EasyReg* program of Bierens. Respecting the European countries and for the whole period, we do not find evidence in favor 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) but these relations seem to dissappear in the post-crisis period (only for Austria, Ireland, Italy and Sweden we fail to reject the null). 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 (specially those belonging to the EMU) have impossed a strict control to some macroeconomic magnitudes (exchange rates) trying to reach a higher degree of convergence. This fact could have led to significant diferences since, respecting the stock markets and on the contrary to exchange rates, have experimented a clear revaluation (in the presence, though, of a high volatility).

Respecting the Asian markets, and similarly to the European ones, we do not find significant evidence in favor of we find nonlinear cointegration for the whole period (the exception is Thailand where we fail to reject the null at the 5% level). We find a number of significant long run relationships during the period preceeding the crisis (for half of the countries), which disappear in the crisis period (a exception is Hong Kong). This contrast with the results obtained in the European case where it happened the opposite.

The explanation of these results is, if one avoids to speculate, quite difficult. At the present state of things 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 affect 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 competing alternatives.

Now we will focus our attention to the possible existence of nonlinear causal relationships. Since the modified Baek and Brock test may have power against linear alterantives, we employ the procedure suggested by Hiemtra 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 minimizing the Akaike Information Criterion. Then we calculate the residuals and run the nonlinear causality test. Following suggestions by Hiemstra and Jones (1994) we wmploy embedding dimensions from 2 to 5 and ε equal to the unconditional standard deviation of the corresponding series. The results for the whole period are shown in Table 8. For Europe only for Ireland, Sweden and Norway (weak in this case) we find evidence of bivariate causality 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 interesting to note that nonlinear causality is much stronger; we find clear bivariate evidence for all the countries excepting Hong Kong (the stock market leads exchange rates) and Taiwan (no relationship). Note that, in this case, the diferences between the two geographical are are clear: wek evidence for Europe and clear for Asia.

When we analyze the results for subperiods (Table 9, Panels A and B) we find that the evidence of nonlinear relationships is quite diferent for Asia and similar for Europe. During the crisis period we find seven cases where the exchange rates seem to lead stock market prices (the exception is Taiwan and for Hong Kong and Thailand the evidence is very weak) and four links in the opposite direction (only the case of Indonesia seem clear, though) 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 and Brock statistic differ significantly along the embedding dimension considered. This could be due to a lose of power of the test for moderate sample sizes when the nonlinear structure is not strong enough.

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

6. Conclusions

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

The existence of linear causality it is clear is Asia, for Korea, Indonesia, Malaysia, Singapore and Thailand we find bi-directional 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 find less evidence in favor of causality and it is manifested in only 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).

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

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 because of the implications on economic policy impossed by the European single currency.

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Table 1
Phillips-Perron test

	Exchange rates				**Stock market prices**			
	Levels		Returns		Levels		Returns	
	ημ	ητ	ημ	ητ	ημ	ητ	ημ	ητ
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
Norwey	-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 of a unit root against the alternative of stationarity against a a level (ημ) and a trend (ητ). The critical levels are -3,4360 (ημ) and -3,9671 (ητ). (1%). 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
Norwey	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 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 test

	90-99	**1/1/90-30/6/93**	**1/7/93-3/8/99**
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
Norwey	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	**90-99**	**1/1/90-1/7/97**	**2/7/97-3/8/99**
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 test (Johansen, 1988) for the null 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 test

	90-99	**1/1/90-30/6/93**	**1/7/93-3/8/99**
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	**90-99**	**1/1/90-1/7/97**	**2/7/97-3/8/99**
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 test (Shin, 1994) for the null 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 that exchange rates returns do not Granger cause stock market returns and for the null 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 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	**1/1/90-30/6/93**	**1/7/93-3/8/99**	**1/1/90-30/6/93**	**1/7/93-3/8/99**
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	**1/1/90-1/7/97**	**2/7/97-3/8/99**	**1/1/90-1/7/97**	**2/7/97-3/8/99**
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 test for the null that exchange rates returns do not Granger cause stock market returns and for the null 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 is rejected are underlined.

Table 7
Bierens nonlinear cointegration test

	90-99	1/1/90-30/6/93	1/7/93-3/8/99
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
	90-99	**1/1/90-1/7/997**	**2/7/97-3/8/99**
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 test (Bierens, 1999) for the null of nonlinear co-trending. Critical values 0.119 (10%) and 0.151 (5%). A failure to reject the null 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 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 is that the exchange rates do not nonlinearly cause stock market returns (*ei*) or stockimarket returns does not nonlinearly cause exchange rates returns (*ie*). The rejection of the null at the 5% level is noted 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 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 is that the exchange rates do not nonlinearly cause stock market returns (*ei*) or stocikmarket returns does not nonlinearly cause exchange rates returns (*ie*). The rejection of the null at the 5% level is noted 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 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 is that the exchange rates do not nonlinearly cause stock market returns (*ei*) or stocikmarket returns does not nonlinearly cause exchange rates returns (*ie*). The rejection of the null at the 5% level is noted by *.