STAR-GARCH Models for Stock Market Interactions in the Pacific Basin Region, Japan and US

Giorgio Busetti and Matteo Manera

NOTA DI LAVORO 43.2003

APRIL 2003

ETA – Economic Theory and Applications

Giorgio Busetti, *Quantitative Methods, Monte Paschi Alternative Investment, Milano, Italy*

Matteo Manera, *Department of Statistics, University of Milano-Bicocca, Italy and Fondazione Eni Enrico Mattei, Milano, Italy*

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Summary

We investigate the financial interactions between countries in the Pacific Basin region (Korea, Singapore, Malaysia, Hong Kong and Taiwan), Japan and US. The originality of the paper is the use of STAR-GARCH models, instead of standard correlation-cointegration techniques. For each country in the Pacific Basin region, we find statistically adequate STAR-GARCH models for the series of stock market daily returns, using Nikkei225 and S&P500 as alternative threshold variables. We provide evidence for the leading role of Japan in the period 1988-1990 (pre-Japanese crisis years), whereas our results suggest that the Pacific Basin region countries are more closely linked with the US during the period 1995-1999 (post- Japanese crisis years).

Keywords: STAR-GARCH models, stock market integration, Pacific-Basin capital markets, outliers

JEL: C22, C51, C52, F36

Paper presented at the International Conference on Financial Engineering and Statistical Finance, held at the Graduate School of International Corporate Strategy (ICS), Hitotsubashi University, Tokyo, 17 March 2003. A special thanks to Umberto Cherubini, Toshiki Honda, Michael McAleer, Ryozo Miura, and Kazuhiko Ohashi for their many insightful suggestions. Most computations have been based on modifications of the Gauss programs whichcomplement the book by Franses and van Dijk (2000) and can be downloaded from the following URL: http://www.few.eur.nl/few/people/franses and http://www.few.eur.nl/few/people/djvandijk. This study does not necessarily reflect the views of Monte Paschi Alternative Investment.

Address for correspondence:

Matteo Manera
Department of Statistics
University of Milano-Bicocca
Via Bicocca degli Arcimboldi, 8
Building U7
20126 Milano
Italy
Phone: +39-02-64487319
Fax: +39-02-6473312
E-mail: Matteo.Manera@unimib.it

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1. Introduction

This study analyses the stock market relationships between some Eastern Asian countries using STAR-GARCH models. The idea of investigating the interactions among countries of the so-called Pacific Basin region (namely Korea, Singapore, Malaysia, Hong Kong and Taiwan) is not new in the empirical literature. Several studies (see, among others, Phylatkis, 1999) have supported both the existence of strong interrelations among those countries and the presence of a significant degree of dependence of the whole area on Japan and US. If, on the one hand, it is undisputable that the relationship with Japan is based on geographical as well as historical reasons, on the other hand it is possible to interpret the link with US as an element in favour of the thesis that there exists a global process in developed countries leading to fully integrated economies and financial markets. Moreover, it is important to emphasise that a relevant subset of those studies show that Eastern Asian financial markets are more tightly linked to Japan than to US. Using this idea as the starting point, we have looked for confirmation of the leadership exerted by the Japanese economy between the end of the 80s and the early 90s and, subsequently, we have tested the hypothesis that the Japanese stock market crisis of 1990-1991 has weakened this role. Until the early 90s, many macroeconomic and financial indicators suggest that Japan has economically dominated the Pacific Basin region. Starting from the deregulation of Hong Kong in 1973 to that of Korea in 1988, there has been a constant increase in the percentage of net capital flows originated in Japan and directed to the Pacific area. Within the same period, the Japanese currency has been widely used in Asian markets, so that some countries, among which is Malaysia, decide to re-denominate part of their debt in terms of the yen. Since mid 1990, not only has the economic expansion of Japan in the Asian area recorded a stop, but also a severe crisis has started. The behaviour of the Nikkei225 index, as shown in Graph 1, is extremely informative on that aspect: in less than ten months, from January to October 1990, the stock index has fallen to its 1986 values.

This study aims at testing the existence of a change of economic leadership in the Pacific Basin region from Japan to US, and the centrality of US as the most important financial market in this area.

The originality of this study rests in the particular use of a class of econometric tools, the so-called STAR and STAR-GARCH models, which have been introduced originally to deal with volatility in financial data, but they are also very useful in developing a concept of relation-integration among markets which is different from the one implied by cointegration or correlation. The fundamental hypothesis of our study is that the influence of the stock market of a leading country on other financial markets does not need to be linear, as correlation and cointegration are implicitly assuming. As Table 1 points out, using the simple concept of correlation of daily returns to describe integration among different stock markets can be unfruitful.

On the contrary, a STAR model offers a different view of the relationships among financial markets, which is based upon an appropriately defined threshold variable. The threshold, through the transition function, the probability that the dependent variable is in one of two, or more, states of the world. For instance, consider the behaviour of an agent who operates on a daily basis in market A. If the agent selects as the threshold variable between two regimes the past returns from the relavant stock index for market B, this means that he acts according to two views of the world represented by the two states of the STAR model and that he chooses to give a larger weight to one or the other regime depending on the levels of past returns from the investments on market B. The degree of influence of market B on market A crucially depends on the choice of the threshold variable. If two or more threshold variables are significant, we select, according to some prespecified criterion, the variable which leads to the most statistically accurate STAR model.

In order to capture non-linearities with STAR models, sufficiently large samples of data are needed. For this reason we have considered, for each selected country, daily returns on the corresponding stock price indexes during the sub-samples 1988-1990 (pre-crisis period) and 1995-1999 (post-crisis period).

2. STAR models

The idea of using non-linear models to study the behaviour of economic variables is very popular in applied econometrics, and several test statistics have been developed to empirically verify the existence of non-linear processes in many financial time series.

According to the so-called regime-switching models, the time series evolution of many economic variables is characterized by the presence of different states of the world (see, among others, Priestley, 1980, 1988). First and second moments of many time series variables depend upon the regimes and the modalities of transition from a particular state to another. Models where this transition is regulated by an observed variable are termed threshold autoregressive (TAR) or, if the transition from one regime to the other is not abrupt, smooth transition autoregressive (STAR).

2.1. Representation of STAR models

TAR and STAR (see Franses, Teräsvirta and van Dijk, 2002, for a exhaustive survey) are simple autoregressive models whose coefficients depend on a threshold variabile q. The simplest case is given by a TAR with two regimes and an autoregressive part of order one:

$$y_t = \begin{cases} f_{0,1} + f_{1,1} \cdot y_{t-1} + e_t & \text{if} \quad q_{t-1} \leq c \\ f_{0,2} + f_{1,2} \cdot y_{t-1} + e_t & \text{if} \quad q_{t-1} > c \end{cases} ,$$

where e_t is a white noise error term with $E(e_t\,|W_{t-1})=0$ and $E(e^2_t|W_{t-1})=s^2$. The value c, or threshold value, regulates the transition between the two states of the world. If the threshold variable coincides with the lagged dependent variable, the model's name modifies to SETAR (self-exciting TAR).

The major limitation of the TAR model is that the transition between one regime to the other is a sudden jump. This simplistic view has been made more realistic by increasing the number of regimes, or, alternatively, by proposing the STAR model (see Teräsvirta, 1994, among others). STAR has in its simplest version the following specification:

$$y_t = (f_{0,1} + f_{1,1} \cdot y_{t-1}) \cdot (1 - G(q_{t-1}; g; c)) + (f_{0,2} + f_{1,2} \cdot y_{t-1}) \cdot (G(q_{t-1}; g; c) + e_t \,, \qquad (1)$$

where *G(.)* is a probability function which takes values between 0 and 1, and can be interpreted as a weight between the two regimes. *G(.)* indefinitely increases the number of possible combinations between the states of the world, smoothing the transition between one regime and the other. It is obvious that STAR models introduce non-linearities in the conditional mean.

The most widely used specifications for the function *G(.)* are exponential (E-STAR) and logistic (L-STAR). The *G(.)* function for the L-STAR model is:

$$G(q_t, g, c) = (1 + \exp(-g \prod_{i=1}^{n} (q_t - c_i)))^{-1} \qquad (2)$$

with $n=1$ in a two-regime model.

The value of c has to lie between the maximum and minimum value of q_t, and the smoothness of the model depends on the parameter g. The behaviour of the L-STAR model is asymmetric. In the limiting case where $g \rightarrow \infty$, *G(.)* becomes an indicator function $I[q_t > c]$ with $I[A] = 1$ if *A* is true, and $I[A] = 0$ otherwise. On the contrary, if $g \rightarrow 0$, the model is linear with a constant logistic function whose value is 0.5. Classical, non-financial applications of the logistic function involve modelling asymmetries (i.e. recessions and expansions) in economic cycles (see, for example, Teräsvirta, Tjøstheim and Granger, 1994).

2.2. Hypothesis testing in STAR models

The procedure suggested by Teräsvirta (1994) for a correct specification of a STAR model involves first the definition of an appropriate AR(p) model for the states of the world, then testing for non-linearity.

As far as the specification of the autoregressive order p is concerned, the approach followed in the STAR context is similar to the one of any standard AR model (i.e. partial autocorrelation function of the series, AIC).

On the contrary, there is no standard procedure to identify the threshold variable. The approaches that are more frequently adopted in the empirical literature range from the simple use of some economic intuition to an ex-post informal test based on AIC or BIC, or the choice of the threshold variable that gives rise to the smallest p-value associated with a particular test for non-linearity.

The general structure of any test of non-linearity is to compare the fit of a STAR model with that of a linear model. That is, given a L-STAR model with two regimes

$$y_t = (f_1 \cdot x_t) \cdot (1 - G(q_{t-1}; g; c)) + (f_2 \cdot x_t) \cdot (G(q_{t-1}; g; c) + e_t \qquad (3)$$

where $x_t = (1, y_{t-1}, \ldots, y_{t-p})$, $f_i = (f_{i,o}, f_{i,1}, \ldots, f_{i,p})$, i=1,2, linearity implies H_0: $f_1 = f_2$.

Under H_0, unfortunately, parameters g and c are unidentified nuisance parameters. Eitrheim and Teräsvirta (1996), Luukkonen, Saikkonnen and Teräsvirta (1998), and Escribano and Jordà (1999) suggest to approximate the *G(.)* function with a Taylor series. In this way the identification problem is solved and the null hypothesis of linearity can be tested with standard LM-type statistics. Specifically, equation (3) can be re-expressed as:

$$y_t = (f_1 \cdot x_t) + (f_2 - f_1) \cdot x_t \cdot (G(q_{t-1}; g; c) + e_t \, .$$

Provided that *G(.)* is given by (2), and it is differentiable, it can be approximated around g=0, thus obtaining the auxiliary regression:

$$y_t = b_0^{'} \cdot x_t + b_1{}' \cdot x_t \cdot s_t + e_t ,$$

where $b_i = (b_{i,0}, b_{i,1},\ldots,b_{i,p})$, $e_t = e_t + (f_2 - f_1)' x_t \cdot R_1(q_t ; g ; c)$, with R indicating the error in the Taylor approximation. In this context, under the null hypothesis g=0 we have $b_0 \neq 0$ and b_1=0, where the b's are functions of the parameters of the STAR model. Thus the null hypothesis becomes b_1=0. In the applied literature this test is referred to as the LM_1-test, which has a c^2 asymptotic distribution with $p+1$ degrees of freedom. If one is interested in testing the null hypothesis of linearity against a L-STAR model where $q_t = y_{t-i}$, $0 < i < p+1$, the term $b_{1,o}\, q_t$ must be omitted in order to avoid perfect multicollinearity. It can also be shown that if the two states of the world differ in the value of the constant only, an alternative test should be used. In this case, $G(.)$ has to be approximated up to the third degree, yielding the auxiliary regression:

$$y_t = b_0^{'} \cdot x_t + b_1{}' \cdot x_t \cdot q_t + b_2{}' x_t q_t^{\,2} + b_3{}' x_t q_t^3 + e_t . \qquad (4)$$

The null hypothesis becomes H_0:b_1=b_2=b_3=0. The name for this test is LM_3, and it has an asymptotic c^2 distribution with $3(p+1)$ degrees of freedom. Both test can be calculated using the asymptotic c^2 version as well as the small-sample F counterpart.

2.3. Estimation and diagnostic tests in STAR models

STAR models are typically estimated with Non-linear Least Squares (NLS). NLS is equivalent to Maximum Likelihood or Quasi-Maximum Likelihood, according to whether normality of e_t is assumed or not.

If we define $q=(f_1', f_2', g, c)$, then the NLS estimator is

$$\hat{q} = \arg\min_{f} Q_T(q) = \arg\min_{f} \sum_{t=1}^{T} (y_t - F(x_t ; q))^2 ,$$

where $F(.)$ is the STAR model (4).

Given the regularity condition $E(e_t^2) < \infty$ (see Ling and McAleer, 1999), NLS is consistent and asymptotic normal. It is well-known that the choice of appropriate initial conditions play a crucial role in any estimator which uses numerical methods such as Newton-Raphson or Gauss-Newton. The procedure proposed by Leybourne, Newbold and Vougas (1998) notes that the STAR model is linear for given values of g and c. Conditional to g and c, it is possible to estimate f with OLS and obtain $\hat{f} = (\hat{f}'_1, \hat{f}'_2)$, i.e.:

$$\hat{f}(g,c) = \left(\sum_{t=1}^{T} x_t(g,c) x_t(g,c)' \right)^{-1} \left(\sum_{t=1}^{T} x_t(g,c) y_t \right),$$

where $x_t(g,c) = (x_t'(1 - G(s_t; g,c)), x_t' G(s_t; g,c))'$.

In this way the minimand function Q_T has a reduced dimension, that is:

$$Q_T(g,c) = \sum_{t=1}^{T} (y_t - f(g,c)' x_t(g,c))^2 ,$$

which can now be minimized with respect to c and g.

From the empirical viewpoint, it is worth noticing that many studies (see, for instance, Bates and Watts, 1988), point out that estimating g is far from being an easy task. This is mainly due to the fact that a large number of observations in the neighborhood of c is required in order to obtain a reliable estimate of g.

In general even reasonable estimates of g come with very high standard errors and t-statistics which apparently do not reject the null hypothesis of linearity. In this context, however, the t-test is not reliable, for its distribution is not standard. In any case, this problem is mitigated by considering that large variations in g do not have a significant impact on the transition function. The estimation of c is subject to similar problems (see Hansen, 1997), even if a value of c which is not statistically different from zero does not affect the overall validity of the model.

The asymptotic variance/covariance matrix C of $\hat{q}$ can be consistently estimated using the robust estimator $C = A(q_0)^{-1} B(q_0) A(q_0)^{-1}$, where $A(q_0) = \lim_{n\to\infty} E\left[J^2 F_n(q) / JqJq' \rfloor_{q_0} \right]$ and $B(q_0) = \lim_{n\to\infty} E[nJF_n(q)/Jq \cdot JF_n(q)/Jq' \rfloor_{q_0}]$, i.e. *A(.)* is the limit of the Hessian of the objective function and *B(.)* is the limit of the cross-product of the score function.

The most common test for residual autocorrelation in STAR models is based on the auxiliary regression:

$$\hat{e}_t = a + \vec{b} \cdot \hat{\hat{z}}_t + \sum_{i=1}^{q} r_i \cdot \hat{e}_{t-i} ,$$

where $\hat{e}_t$ are the residuals under the null of independence of the errors e_t and $\hat{\hat{z}}_t = JF(x_t; \hat{q}) / Jq$, with *F(.)* being a twice-differentiable function. The analogy with Breusch and Pagan (1980) is clear once we consider that in a linear context the partial derivatives of *F(.)* with respect to the parameters correspond to the regressors x_t. As usual, the null hypothesis is H_0: $r_i = 0$, and the test in its LM form has an asymptotic c^2 distribution with q degrees of freedom.

The idea which is behind a test for remaining non-linearity is to detect if a STAR model has captured all the non-linearity which is in the series by evaluating the statistical adequacy of a STAR model with an additional regime (see Eitrheim and Teräsvirta, 1996).

2.4. Evaluating the forecasting performance of STAR models

A standard STAR model with $q_t = y_{t-1}$ can be represented as $y_t = F(x_t, q) + e_t$, where $F(.)$ is defined as in (4), with $x_t = (1, y_{t-1}, \ldots, y_{t-p})'$. The optimal forecast of y_{t+h} made at time t is $\hat{y}_{t+h|t} = E[y_{t+h} \mid \Omega_t]$, where $e_{t+h|t} = y_{t+h} - y_{t+h|t}$ is the forecast error. The 1-step ahead forecast is given by $\hat{y}_{t+1|t} = E[y_{t+1} \mid \Omega_t] = F(x_{t+1}; q)$, with $E[e_{t+1}|W_t]=0$. In this study we concentrate on static forecasts only. The indicators which are most commonly used to evaluate the forecasting performance of a STAR model are the mean squared error (MSE) and the mean absolute error (MAE):

$$\text{MSE} = \frac{1}{m} \sum_{j=0}^{m-1} (\hat{y}_{T+h+j|T+j} - y_{T+h+j})^2$$

and

$$\text{MAE} = \frac{1}{m} \sum_{j=0}^{m-1} | \hat{y}_{T+h+j|T+j} - y_{T+h+j} | .$$

3. GARCH models

ARCH models (Autoregressive Conditional Heteroskedasticity) have been introduced for the first time by Engle (1982) in order to model two phenomena which are typical of many financial time series, namely non-constant conditional variances and volatility clustering.

The simples formulation of a ARCH(r) model is:

$$e_t = h_t h_t^{1/2} \quad \text{with} \quad h_t = a_0 + a_1 e_{t-1}^2 + .. + a_r e_{t-r}^2 ,$$

where $a_0 > 0$, $a_i \geq 0$ ($i=1,\ldots r$); h_t indicates a white noise error and h_t is the conditional variance of the process. In other terms, if F_t is the s-algebra generated by $\{h_t, h_{t-1}, \ldots\}$, then $E(e^2_t|F_t) = h_t$.

The first generalization of the ARCH model (GARCH) has been proposed by Bollerslev (1986):

$$e_t = h_t h_t^{1/2} \quad \text{with } h_t = a_0 + \sum_{i=1}^{r} a_{i1} e_{t-i}^2 + \sum_{i=1}^{s} b_i h_{t-i} \,, \tag{5}$$

where $a_0 > 0$, $a_{i1} \geq 0$ and $b_i \geq 0$ ($i=1,\ldots r \backslash s$). The main advantage of this model with respect to the ARCH specification is that the additional term h_{t-i} allows us to reduce the number of parameters in the ARCH component.

The second, more natural generalization is to use a STAR model in the conditional mean of the process and a GARCH specification for the conditional variance (STAR-GARCH models). From an estimation viewpoint, it is easy to deal with STAR-GARCH models since, according to Engle (1982), parameters for the conditional mean and the conditional variance can be estimated separetely, provided that the GARCH specification is symmetric. In this case, given the Hessian for the STAR-GARCH model

$$H_t(q) = \begin{pmatrix} \dfrac{J^2 l_t(q)}{JxJx'} & \dfrac{J^2 l_t(q)}{JxJy'} \\ \dfrac{J^2 l_t(q)}{JyJx'} & \dfrac{J^2 l_t(q)}{JyJy'} \end{pmatrix} = \begin{pmatrix} H_t^{xx}(q) & H_t^{xy}(q) \\ H_t^{xy}(q) & H_t^{yy}(q) \end{pmatrix}$$

where $q = (\vec{x}, \vec{y})$, with $\vec{x}$ indicating the vector of parameters in the conditional mean, and $\vec{y}$ the vector of parameters of the conditional variance, elements H^{xy} and H^{yx} are both zero.

4. Outliers in STAR and GARCH models

4.1. Effects of outliers on STAR models

Van Dijk, Franses and Lucas (1999) show that LM-type tests tend to reject the null of linearity too often in presence of outliers. Possible solutions to this problem are based on robust estimation techniques which involve different weighting functions. Such techniques can be briefly sketched starting from a simple AR(p) model $y_t = f' x_t + e_t$, and modifying the first-order condition $\sum_{t=1}^{T} w_r(r_t) \cdot x_t (y_t - f' x_t) = 0$, where r_t are the standardized residuals, $r_t \equiv \dfrac{(y_t - f' x_t)}{(s_e w_x(x_t))}$, with weights $w_x(.)$, $w_r(.)$ between 0 and 1. The main common characteristic shared by all weighting functions discussed in the literature is to give small weights to values of $\frac{(y_t - f' x_t)}{s_e}$ that are exceptionally large. For the test LM_3, van Dick, Franses and Lucas (1999) suggest to calculate the R^2 from regressing the weighted residuals $\hat{y}(\hat{r}_t) = \hat{w}_r(\hat{r}_t)\hat{r}_t$ on the weighted regressors $\hat{w}_x(x_t) \otimes (x_t', x_t' s_t, x_t' s_t^2, x_t' s_t^3)'$. Both weighting functions are obtained from estimating the AR(p) for y_t under H_0. This test has an asymptotic c^2 distribution with $3(p+1)$ degrees of freedom. Monte Carlo simulations point out that, in the presence of additive as well as innovation outliers, robust tests have more power, and power increases as sample size increases.

Van Dick (1999) analyzes the effects of outliers on the parameters of a STAR model simulating a two-regime logistic specification:

$$y_t = (f_{0,1} + f_{1,1} \cdot y_{t-1}) \cdot (1 - G(q_{t-1}; g; c)) + (f_{0,2} + f_{1,2} \cdot y_{t-1}) \cdot (G(q_{t-1}; g; c) + e_t .$$

Estimates of the threshold parameter c do not seem to be particularly affected, provided that the outliers are not numerous enough to justify the existence of a specific regime. This can well happen in presence of many outliers of the same sign. The same is true also for g. Conversely, parameters $f_{1,i}$, i=1,2, are generally biased, as it happens with standard AR(p) models. Here the magnitude of the bias depends on the specific regime that prevails at the time of the outliers. If, for instance, all outliers fall in just one regime, the bias affecting f_1 would be noticeable only for that regime.

Several different solutions are available in the literature to reduce the bias in estimating the autoregressive parameters of a STAR model. In this study we consider a simple symmetric trimming algorithm (STA), which is composed by the following steps: i) calculation of the standard deviation of the series over the whole sample; ii) if a single observation is larger than 4 standard deviations, it is trimmed to 4 standard deviations; iii) if an observation is between 3 and 4 standard deviations, it is trimmed to to 3 standard deviations; iv) if an observation is between 2.5 and 3 standard deviations, it is trimmed to 2.5 standard deviations; v) repetition of steps i)-iv) for each obervations in the sample. This procedure has simplicity as its main advantage, but suffers from at least two drawbacks. First, simple STA tends to reduce variability in the series. Second, if different regimes are located far from the mean of the series, simple STA could select as an outlier (and hence exclude from the sample) an observation which instead should be modelled.

4.2. Effects of outliers on GARCH models

Van Dijk, Franses and Lucas (1999) show that the empirical behaviour of the tests for heteroskedasticity in presence of outliers is very similar to the performance of the tests for linearity in STAR models, that is the null hypothesis is rejected too often. Almost all these tests are based on the residuals of robustly estimated conditional means, once the outliers have been removed and an auxiliary regression of the Breusch-Pagan type is taken into account to check for the presence of linear ARCH effects.

Verhoeven and McAleer (1999), and Chan and McAleer (2002a, 2002b), among others, point out that the presence of outliers tends to increase $\hat{a}_1$ and to reduce $\hat{b}_1$ in the simple GARCH(1,1) model $h_t = a_0 + a_1 e_{t-1}^2 + b_1 h_{t-1}$. Although several procedures are available in the literature to cope with this problem (e.g. Franses and Ghijsels, 1999), the algorithm we use in this study to take into account for outliers is a modification of the simple STA described in Section 4.1, which is motivated by the following considerations: i) irrespective of the selected model, simple STA often produces estimation and forecasting results which are worse than those obtained on the original series; ii) alternative methods to simple STA, such as the Generalized M estimator, heavily depend on the choice of the weighting function (see van Dick, 1999) and are not easy to implement; iii) many estimation procedures which correct for the presence of outlier in GARCH models are based on a priori choices (see the selection of *c* in Franses and Ghijsels, 1999).

The most important features of our alternative procedure are simplicity (since it is based on the simple STA) and flexibility (the mean of a STAR model is not forced to be constant). Specifically, our modified STA algorithm is formed by the following steps: 1) estimation of a STAR model; 2) residual analysis using a Breusch-Pagan-type test for GARCH effects and estimation of a STAR-GARCH model if appropriate; 3) 1-step ahead forecasts of the conditional mean and variance; 4) application of simple STA, using the 1-step ahead forecast of the mean as the actual mean and the 1-step ahead forecast of the variance as the actual variance; 5) re-estimation of the STAR model on the series corrected for the presence of outliers; 6) overall evaluation of model improvements by comparing the MSE and MAE of the forecasts obtained at steps 3) and 5).

5. Empirical analysis

5.1. Data

The first important decision concerns the starting date of the Japanese financial crisis. On the basis of the temporal behaviour of the Nikkei225 stock price index (see Graph 1), which clearly shows a dramatic fall at the beginning of 1990, we set the end of the first period (i.e. pre-crisis) on the 6th june 1990.

The second choice is about the starting date of the pre-crisis period, which is intimately linked to the deregulation characterizing the countries under analysis. As already stated, some countries (e.g. Korea and Taiwan) have deregulated only since 1988, while others (i.e. Hong Kong and Singapore) have started to liberalize their financial markets during the 70s. In order to evaluate different models for different countries on a homogenous basis, we consider the period 5th January 1988 – 6th June 1990 as the maximum common sample.

The second sample (post-crisis) goes from 5th January 1995 to 5th November 1999. The choice of 1995 as the starting date for this period assumes that agents, at the time of implementation of any financial decision, have already incorporated the fall of the Japanese stock price index in their information sets.

We have decided to concentrate our attention on five countries (Korea, Hong Kong, Malaysia, Singapore and Taiwan), since these are the countries which are analyzed by those studies aiming at testing the existence of a strong link between the so-called Pacific Basin region and Japan.

The novelty of our approach is how the potential existence of this link is modelled and tested, that is using STAR and STAR-GARCH models. Moreover, we are also interested in investigating to what extent the financial Japanese crisis has weakened those relationships, and whether US has substituted Japan in its economic leadership in this area.

Our empirical investigation proceeds as follows. First, we check the autocorrelation properties of the original series and determine the order *p* of the AR model. Second, we test for linearity and explanatory power of alternative threshold variables, that is one-period lagged Japan Nikkei225 and US S&P500 stock price indexes. Third, we estimate a L-STAR model and test for its statistical adequacy using residual based diagnostic tests. Fourth, we test for ARCH effects in STAR residuals. Fifth, we estimate an appropriate STAR-GARCH model (if ARCH effects are present). Sixth, we correct for the presence of outliers using both STA and modified STA algorithms. Seventh, we re-estimate the STAR model and check for estimation improvements using MSE and MAE calculated on the forecasting horizon 8th November 1999 - 28th December 2001. It is worth noticing that the comparison between simple and modified STA is not feasible on the first period. In fact, if it were, the implied forecasting exercise would have been conducted on the Japanese crisis period, in violation of the ceteris paribus condition which is the maintained assumption of our comparative analysis.

5.2. Results

Korea has started the process of liberalization in 1988. The Korean stock price index over the first period (Graph 2a) shows a sharp drop in 1990, which is probably due to the Japanese financial crisis. The correlogram of the corresponding series of returns (Graph 2b) suggests an AR model with p=1. The estimation of a STAR(1) model reveals that only the Japanese Nikkei225 stock price index is the appropriate threshold variable. The STAR (1) residuals are not autocorrelated, although they exhibit significant ARCH effects, which lead to the estimation of a STAR(1)-GARCH(1) specification (see Table 3a). The adjustment parameter g has a quite high value, which is nevertheless justified in the literature. The first regime is predominant, whereas the second regime has a weight equal to one when the threshold variable takes extreme values (i.e. less than 1%, Graph 7). This is the reason why the weighting function takes the value one more frequently at the end of the sample, that is close to the Japanese crisis. Even in the absence of significant outliers, an application of the simple STA produces a strong bias in the series, since the distinction between the two states of the world disappears. In fact, STA eliminates from the sample all the extreme observations which are responsible for the identification of the second regime. Instead, the application of our modified STA does not alter substantially the estimation results. If we take into consideration the second period (Graphs 2c-2d), it is evident that the Japanese crisis affects both the levels of the index and the volatility of returns. The STAR(1)-GARCH(1) specification which uses the US S&P500 as the threshold variable is preferable on the basis of both the non-linearity test (Table 2b) and in-sample AIC. The presence of strong kurtosis in the residuals is an indicator of outliers. Both procedures, simple STA and modified STA, produce results which are very close to the starting model in terms of MSE and MAE. Table 3b reports the estimated coefficients of the STAR(1)-GARCH(1) model obtained after the application of the simple STA, since this specification is superior in terms of more accurate forecasts and reduced standard errors. With respect to the first sample, the value of g is larger, whereas the standard error for the threshold parameter c is smaller. The transition function indicates the absence of polarization towards a particular state of the world. In summary, in the first period only Nikkei225 can be used as a threshold variable, whereas in the post-crisis period the most statistically adequate threshold is US

S&P500. One possible interpretation is that, prior to 1988, the Korean financial market is heavily dependent on Japan, while, after 1995, progresses in world integration have strenghtened the relationships with both Japan and US, attributing the stock market leadership to the latter.

Hong Kong’s stock exchange is the most developed within the Pacific Basin region, since it dates back to 1973. A quick inspection to the Hong Kong stock price index during the first period (Graph 3a) reveals the presence of a strong increasing trend, with two marked corrections on july 1989 and july 1990. Obviously, daily returns are characterized by large outliers around the 300th observations (Graph 3b). The correlogram of returns is compatible with an AR(1) model, whereas the tests for non-linearity suggest that both Nikkei225 and S&P500 are statistically adequate threshold variables, with a slight preference for the former (Table 2a). Once estimated, the two STAR(1) behave very differently. In particular, the model with S&P500 as the threshold variable is characterized by very high standard errors even after the use of both simple and modified STA. In the light of these results, the selected model is STAR(1)-GARCH(1) with Nikkei225 as the threshold variable after adjusting the residuals with simple STA (Table 3a, Graph 8). The results obtained in the second period (Graphs 3c-3d) are more controversial. On the one hand, the STAR(1) model with S&P500 as the threshold after applying simple STA is preferable in terms of residual sum of squares and AIC; on the other hand, the STAR(1) model with Nikkei225 as the threshold produces lower MSE and MAE, when both simple and modified STA are used (Tables 2b, 4). Irrespective of the trimming algorithm and the threshold variables, STAR residuals are affected by ARCH effects. To summarize, the empirical results are quite clear in the first period, where the statistically adequate model has Nikkei225 as the threshold. Conversely, in the post-crisis period both threshold variables yield statistically adequate STAR models, although it is not possible to define which is the leading country for Hong Kong.

Malaysia is the country in the Pacific Basin region with the strongest commercial and financial relationships with Japan. This condition is confirmed by the decision of denominating part of Malaysia’s debt in tems of the Japanese currency. During the period 1998-1991, the Malaysia stock price index has shown a continuous, upward trend up to march 1990 (Graph 4a). From april 1990 the index records a sharp drop, due to the incoming Japanese financial crisis. The corresponding series of returns shows some important outliers at the beginning and at the end of the sample, while, contrary to the previous cases, the correlogram points out the presence of second-order autocorrelation (Graph 4b). Thus, we have implemented the tests for non-linearity on the STAR(2) specification. In this case, both threshold variables are statistically adequate, but Nikkei225 seems to be preferable on standard goodness of fit considerations (Table 2a). Since the residuals are characterized by ARCH effects and outliers, we have re-estimated the model using the simple STA (Table 3a, Graph 9). The second period (Graph 4c) is characterized by a severe financial crisis in 1998, which affects the series of returns in terms of the number of outliers, excess kurtosis and increased volatility (Graph 4d). The correlogram of this series is compatible with third-order autocorrelation. Using the simple STA, the best model in terms of estimated standard errors and AIC is STAR(1)-GARCH(1,1) with S&P500 as the threshold (Table 3b). Although Malaysia has experienced the strongest economic links with Japan, over the analyzed period our empirical approach seems to suggest a progressive substitution between Japan and US as the most influential country on the Malaysian stock market.

The liberalization of Singapore dates back to 1978, which makes this country the second best developed financial market in the region, after Hong Kong. If we concentrate on the first period (Graphs 5a-5b), the Singapore stock price index shows a steady upward trend until the end of 1990, while the corresponding returns are highly volatile and affected by outliers at the beginning of the sample and around the 400^{th} observation. Both threshold variables are statistically adequate, although the lowest p-value of the test for non-linearity is recorded when Nikkei225 is the threshold. In fact, if the STAR(1)-GARCH(1,1) model with Nikkei225 is satisfactory (Table 3a, Graph 10a), the estimation of the corresponding specification after replacing the Japanese stock price index with S&P500 has been problematic, due to difficulties of convergence of the underlying numerical estimation algorithm. The use of the simple STA has contributed to solve this problem only partially. The second period (Graphs 5c-5d) shows the gradual decline of the Singapore stock market. The non-linearity test indicates Nikkei225 as the most appropriate threshold variable (Table 2b). Both thresholds produce STAR(1) model residuals which pass standard diagnostics, whereas the specification with S&P500 has lower residual sum of squares and AIC, but higher parameter standard errors. Due to the massive presence of outliers in the post-crisis period, we have re-estimated both models after applying the simple as well as the modified STA. The model with Nikkei225 increases its forecasting performance in both cases, while the model with S&P500 shows some improvement only after using the modified STA (Table 4). In any case, AIC is favourable to the latter. In the transition function computed on the selected model the second regime is over-represented, whereas the first regime indicates a limiting situation, since the threshold value is around –2% (Graph 10b). The results are in line with the original motivation of this study, that is the presence of a switch, from the first to the second period, between Japan and US in the role of financial leader within the Pacific Basin region.

As Korea and Malaysia, also Taiwan has started to liberalize its financial market since 1988 only. During the first period of analysis (Graph 6a) the stock price index has recorded a strong upward trend up to mid 1990, followed by a steep drop caused by the Japanese financial crisis. With respect to the series of returns, the Taiwan index is almost free from the presence of outliers (Graph 6b). The absence of significant threshold variables in the tests for non-linearity prevents us from estimating any appropriate STAR model (Table 2a). This picture is confirmed in the second period (Graphs 6c-6d); here the threshold variable Nikkei225 appears to be significant in a STAR(1) specification, which is by no means different from a simple AR(1) model according to the values of the estimated coefficients, residual sum of squares and AIC (Table 3b). Undoubtly, among the countries we have analyzed in this study, Taiwan is the biggest exception. If it is not difficult to relate the results obtained on the pre-crisis period to the immaturity of its stock market, the empirical findings of the second sub-sample suggest that Taiwan has built up very strong links with the Japanese financial system and that these relationships, which were unsignificant during the late 80s, have been reinforced by the Japanese crisis.

6. Conclusion

The aim of this work is to analyse the financial markets of Korea, Hong Kong, Malaysia, Singapore and Taiwan in order to empirically verify the effects of the Japanese financial crisis of 1990-1991 on the relationships between the Pacific Basin region, on the one hand, and Japan and US, on the other. The econometric investigation has been conducted with statistically adequate STAR and STAR-GARCH models, where Nikkei225 and S&P500 have played the role of threshold variables. The underlying idea is that the Japanese crisis has affected the financial leadership of this country in Eastern Asia in favour of US. The statistical significance of the variable S&P500 in appropriate STAR-GARCH specifications, together with the irrelevance of Nikkei225, is then interpreted as empirical evidence of a strong and progressive process of global integration which looses even the strongest links among the countries of a specific geographical area. At the same time, the switch between Japan and US in the financial leadership over the Pacific Basin region can be read as an indication of openess of the Asian countries towards a unique, global financial market.

The leadership of Japan within the Pacific Basin region is undisputable, at least until the early 90s. Both economic data and our empirical results on the first sub-sample agree to indicate Japan as the main financial market in the area. It is noticeable that Japan is the most important reference not only to the well-developed financial systems, such as Singapore, but also to the newly-liberalized stock markets, such as Korea. Among the five countries under scrutiny, only Taiwan is atypical, since in this case we were unable to find any statistically adequate STAR model. This result can be justified by noting that Taiwan financial market was born in 1988 only, the starting year of our empirical investigation. Possibly, economic and financial dynamics internal to Taiwan have produced greater impacts than any exogenously driven process of integration.

In the first period, it is interesting to point out that S&P500 is statistically adequate as threshold variable in only two cases, namely Singapore and Malaysia. It is relatively easy to justify the influence of the S&P500 index on Singapore stock market, which has been open to the international context since 1978, On the contrary, one possible explanation for the behaviour of the Malaysian stock market is that it has been oriented to integration since the beginning of its liberalization.

In the second period, the significance of S&P500 is evident in all models.

All the analyzed time series are characterized by volatility, excess kurtosis and outliers. Nevertheless, the STAR-GARCH models we have finally selected are statistically robust, and the trimming algorithms we have adopted have proved to be effective. When the comparison has been possible, the modified STA has shown its superiority relative to the simple STA, since the latter suffers from the tendency to bias the series with ample fluctuations in their conditional means.

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Tables

Table 1. Correlation coefficients between daily returns on selected stock market indexes and the thresholds Nikkei225 and S&P500 (1988-1990)

Threshold	Korea	Hong Kong	Malaysia	Singapore	Taiwan
Nikkei225	-0.069	0.250	0.025	-0.014	0.055
S&P500	0.010	0.072	0.502	0.484	0.079

Table 2a. Tests of STAR-type non-linearity (selected countries, 1988-1990)

Threshold	Hong Kong	Malaysia	Taiwan
Nikkei225 (*F*)	3.88 (0.01)	14.31 (0.00)	1.09 (0.35)
S&P500 (*F*)	3.73 (0.01)	6.74 (0.00)	3.26 (0.35)
Nikkei225 (c^2)	11.65 (0.01)	42.94 (0.00)	0.93 (0.43)
S&P500 (c^2)	11.18 (0.01)	20.23 (0.00)	2.78 (0.43)

Notes: The tests reported in this table are based on the LM_3 test described in equation (4); *F* = finite-sample F distribution of LM_3; c^2 = asymptotic χ^2 distribution of LM_3; p-values are reported in parentheses.

Table 2b. Tests of STAR-type non-linearity (selected countries, 1995-1999)

Threshold	Korea	Hong Kong	Singapore
Nikkei225 (*F*)	1.89 (0.13)	1.12 (0.00)	8.75 (0.00)
S&P500 (*F*)	12.11 (0.00)	1.12 (0.00)	6.71 (0.00)
Nikkei225 (c^2)	5.67 (0.13)	33.69 (0.00)	26.24 (0.00)
S&P500 (c^2)	36.34 (0.00)	33.60 (0.00)	20.14(0.00)

Notes: see Table 2a.

Table 3a. Estimated STAR-GARCH models (selected countries, 1988-1990)

Specification	Korea	Hong Kong	Malaysia	Singapore
STAR(*p*) GARCH(*r*,*s*)	STAR(1) GARCH(1,1)	STAR(1) GARCH(1,1)	STAR(2) GARCH(1,1)	STAR(1) GARCH(1,1)
Threshold	Nikkei225	Nikkei225	Nikkei225	Nikkei225
Trimming	Modified STA	Simple STA	Simple STA	Simple STA
g	31.433**	16.502**	1427.965	13.080**
c	-0.016**	0.004**	-0.009	-0.009**
d_1	0.001	-	-0.004	-0.039*
f_1	-0.731**	-0.160**	-0.476**; -0.322	-0.512**
d_2	5.0E-4	-	1.0E-03	4.9E-4
f_2	0.088	0.192	0.087; 0.108	-0.173*
a_0	3.1E-5*	2.0E-5**	8.9E-6**	9.5E-6**
a_1	0.106**	0.098**	0.122**	0.064**
b_1	0.707**	0.774**	0.791**	0.810**
RSS	0.085	0.140	0.064	0.054
AIC	-8.675	-8.156	-8.929	-9.113

Notes: RSS = residual sum of squares; AIC = Akaike information criterion; * (**) = significance at 5% (1%); nE-m = $n10^{-m}$.

Table 3b. Estimated STAR-GARCH models (all countries, 1995-1999)

Specification	Korea	Hong Kong	Malaysia	Singapore	Taiwan
STAR(p) GARCH(r,s)	STAR(1) GARCH(1,1)	STAR(1) GARCH(1)	STAR(1) GARCH(1)	STAR(1) GARCH(1,1)	STAR(1) GARCH(0,0)
Threshold	S&P500	S&P500	S&P500	S&P500	Nikkei225
Trimming	Simple STA	Modified STA	Simple STA	Modified STA	-
g	410.000	153.988**	34.000**	10.005	2857.676
c	-0.018	-0.016**	-0.006**	-0.027**	0.012*
d_1	-7.4E-3	0.004**	0.006**	0.012	-0.001**
f_1	-0.329**	-0.325**	-0.476**	-0.504	0.026
d_2	1.3E-4	-	-	4.0E-4	4.0E-3*
f_2	0.133**	0.108**	0.220**	0.238**	-0.086*
a_0	2.0E-6**	3.2E-6**	1.8E-6**	2.1E-7	-
a_1	0.079**	0.101**	0.116**	0.131**	-
b_1	0.919**	0.891**	0.887**	0.865**	-
RSS	0.522	0.368	0.394	0.174	0.257
AIC	-7.545	-7.896	-7.828	-8.615	-8.255

Notes: see Table 3a.

Table 4. Comparison between MSE and MAE of alternative STAR-GARCH models (selected countries, 1995-1999)

		Hong Kong		Singapore	
Threshold	Trimming	MSE	MAE	MSE	MAE
Nikkei225	-	0.032	1.383	-	-
Nikkei225	Simple STA	0.030	1.331	-	-
Nikkei225	Modified STA	0.029	1.330	-	-
S&P500	-	-	-	0.024	1.168
S&P500	Simple STA	0.033	1.431	0.027	1.290
S&P500	Modified STA	-	-	0.023	1.143

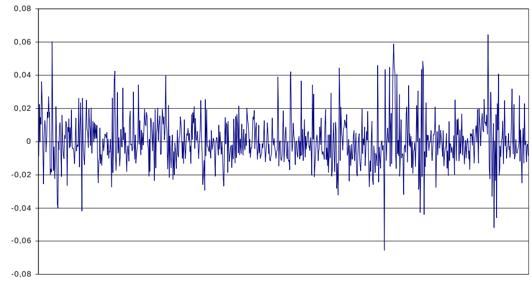
Notes: MSE = Mean Squared Error; MAE = Mean Absolute Error; MSE and MAE are calculated on the forecasting horizon 8th November 1999 – 28th December 2001.

Graphs

Graph 1. Nikkei225 stock price index (1985-1992)

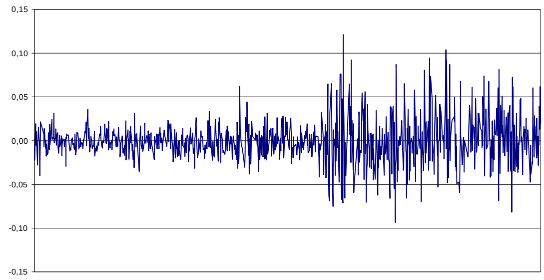
Graph 2a. Stock price index (Korea, 1988-1990)

Graph 2b. Daily returns (Korea, 01/01/88-07/06/90)



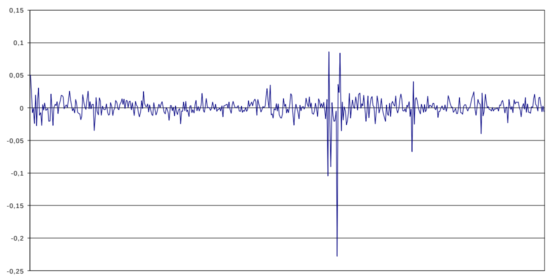
Graph 2c. Stock price index (Korea, 1995-1999)

Graph 2d. Daily returns (Korea, 06/01/95-03/11/99)



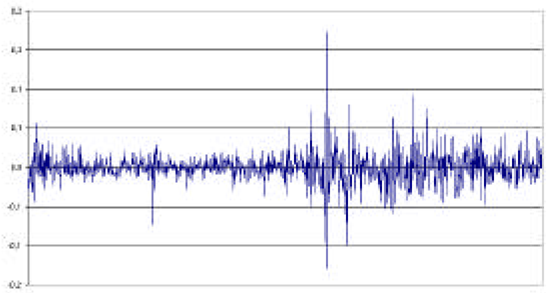
Graph 3a. Stock price index (Hong Kong, 1988-1990)

Graph 3b. Daily returns (Hong Kong, 05/01/88-07/06/90)

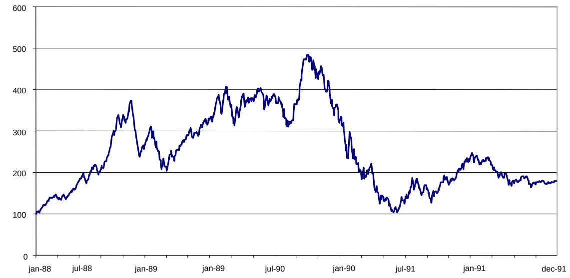


Graph 3c. Stock price index (Hong Kong, 1995-1999)

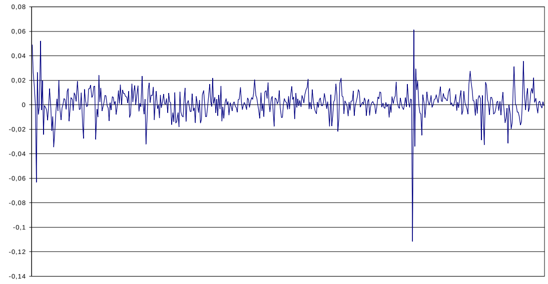
Graph 3d. Daily returns (Hong Kong, 06/01/95-03/11/99)



Graph 4a. Stock price index (Malaysia, 1988-1990)

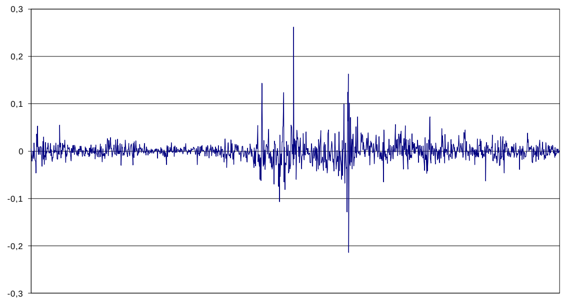


Graph 4b. Daily returns (Malaysia, 1988-1990)

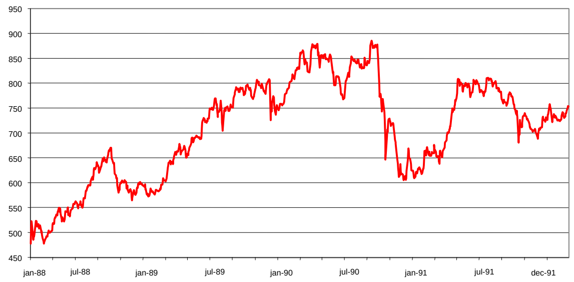


Graph 4c. Stock price index (Malaysia, 1995-1999)

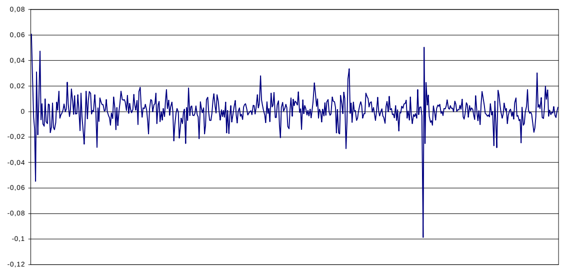
Graph 4d. Daily returns (Malaysia, 1995-1999)



Graph 5a. Stock price index (Singapore, 1988-1990)

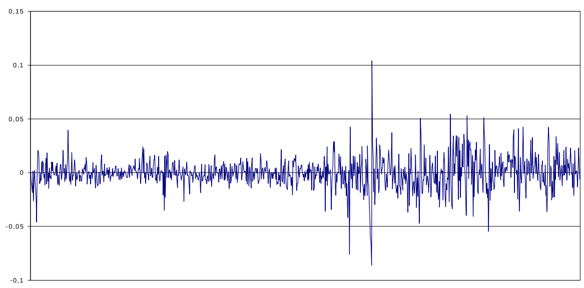


Graph 5b. Daily returns (Singapore, 05/01/88-07/06/90)



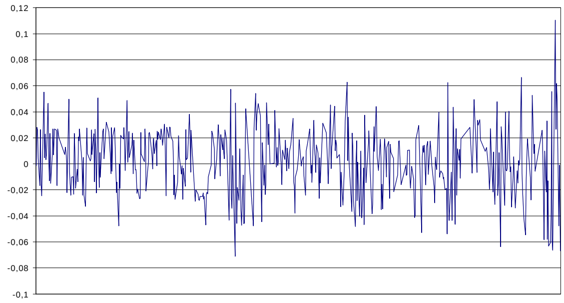
Graph 5c. Stock price index (Singapore, 1995-1999)

Graph 5d. Daily returns (Singapore, 1995-1999)



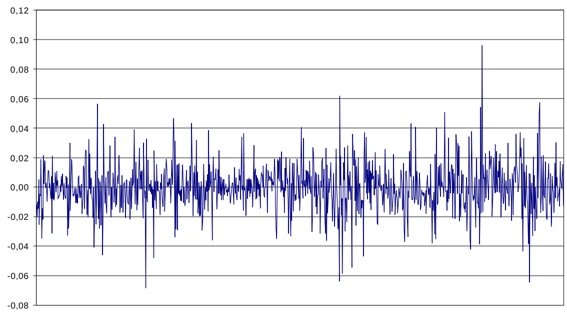
Graph 6a. Stock price index (Taiwan, 1988-1990)

Graph 6b. Daily returns (Taiwan, 1988-1990)

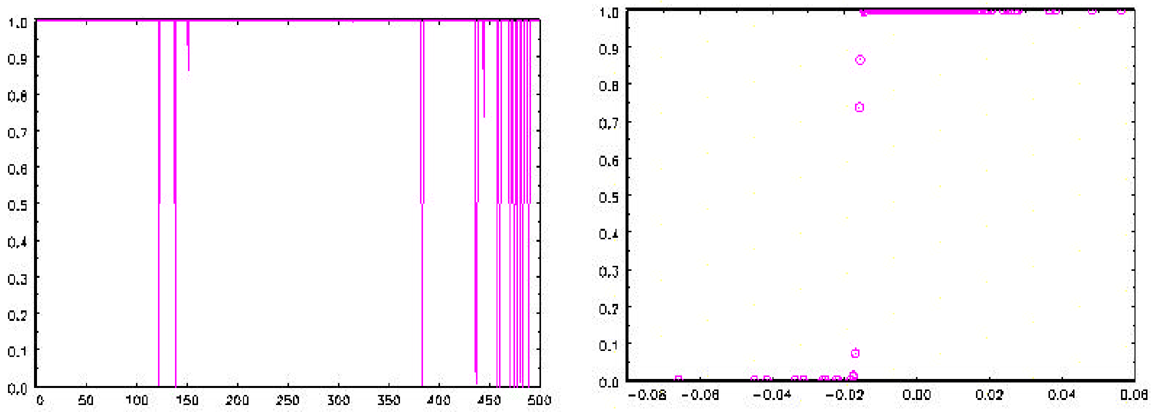


Graph 6c. Stock price index (Taiwan, 1995-1999)

Graph 6d. Daily returns (Taiwan, 1995-1999)

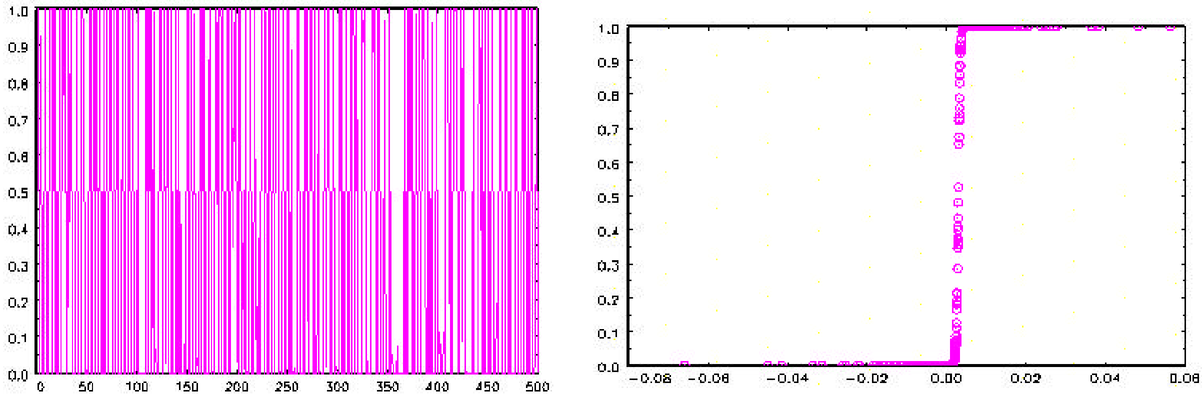


Graph 7. Transition function of selected STAR model with Nikkei225 (Korea, 1988-1990)



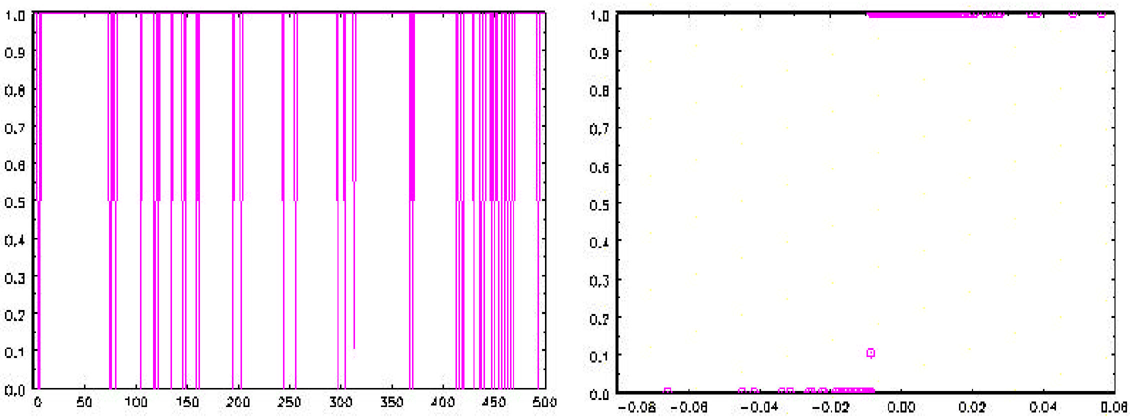
a) time b) threshold

Graph 8. Transition function of selected STAR model with Nikkei225 (Hong Kong, 1988-1990)



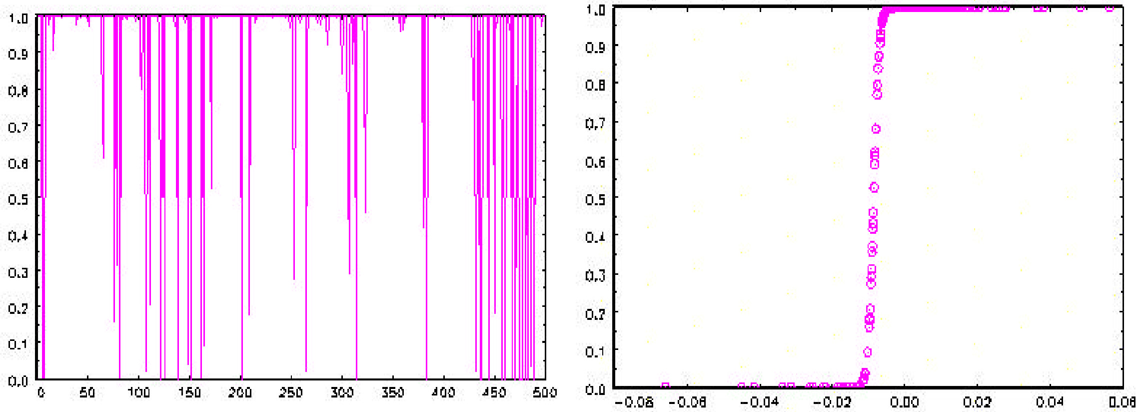
a) time b) threshold

Graph 9. Transition function of selected STAR model with Nikkei225 (Malaysia, 1988-1990)



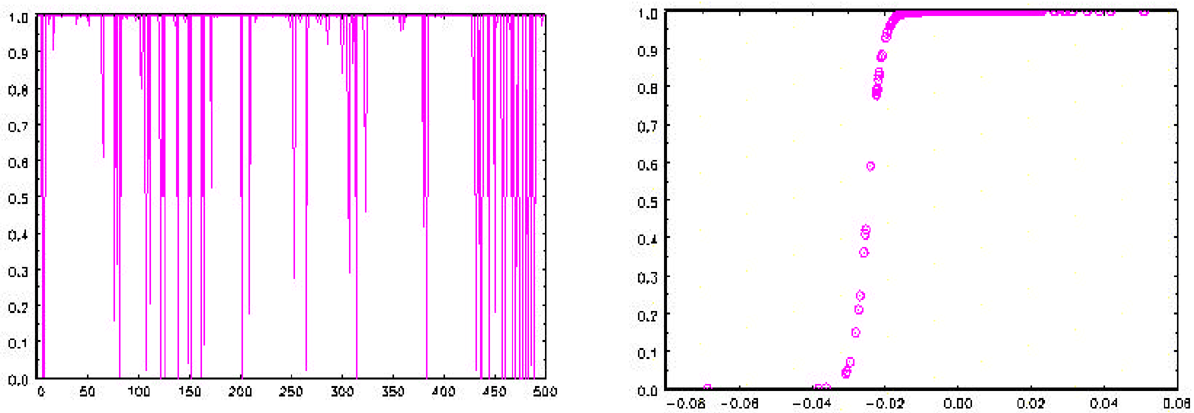
a) time b) threshold

Graph 10a. Transition function of selected STAR model with Nikkei225 (Singapore, 1988-1990)



a) time b) threshold

Graph 10b. Transition function of selected STAR model with S&P500 (Singapore, 1995-1999)



a) time b) threshold

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(l) This paper was presented at the Workshop “Growth, Environmental Policies and Sustainability” organised by the Fondazione Eni Enrico Mattei, Venice, June 1, 2001

(li) This paper was presented at the Fourth Toulouse Conference on Environment and Resource Economics on “Property Rights, Institutions and Management of Environmental and Natural Resources”, organised by Fondazione Eni Enrico Mattei, IDEI and INRA and sponsored by MATE, Toulouse, May 3-4, 2001

(lii) This paper was presented at the International Conference on “Economic Valuation of Environmental Goods”, organised by Fondazione Eni Enrico Mattei in cooperation with CORILA, Venice, May 11, 2001

(liii) This paper was circulated at the International Conference on “Climate Policy – Do We Need a New Approach?”, jointly organised by Fondazione Eni Enrico Mattei, Stanford University and Venice International University, Isola di San Servolo, Venice, September 6-8, 2001

(liv) This paper was presented at the Seventh Meeting of the Coalition Theory Network organised by the Fondazione Eni Enrico Mattei and the CORE, Université Catholique de Louvain, Venice, Italy, January 11-12, 2002

(lv) This paper was presented at the First Workshop of the Concerted Action on Tradable Emission Permits (CATEP) organised by the Fondazione Eni Enrico Mattei, Venice, Italy, December 3-4, 2001

(lvi) This paper was presented at the ESF EURESCO Conference on Environmental Policy in a Global Economy “The International Dimension of Environmental Policy”, organised with the collaboration of the Fondazione Eni Enrico Mattei , Acquafredda di Maratea, October 6-11, 2001

(lvii) This paper was presented at the First Workshop of “CFEWE – Carbon Flows between Eastern and Western Europe”, organised by the Fondazione Eni Enrico Mattei and Zentrum fur Europaische Integrationsforschung (ZEI), Milan, July 5-6, 2001

(lviii) This paper was presented at the Workshop on “Game Practice and the Environment”, jointly organised by Università del Piemonte Orientale and Fondazione Eni Enrico Mattei, Alessandria, April 12-13, 2002

(lvix) This paper was presented at the ENGIME Workshop on “Mapping Diversity”, Leuven, May 16-17, 2002

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