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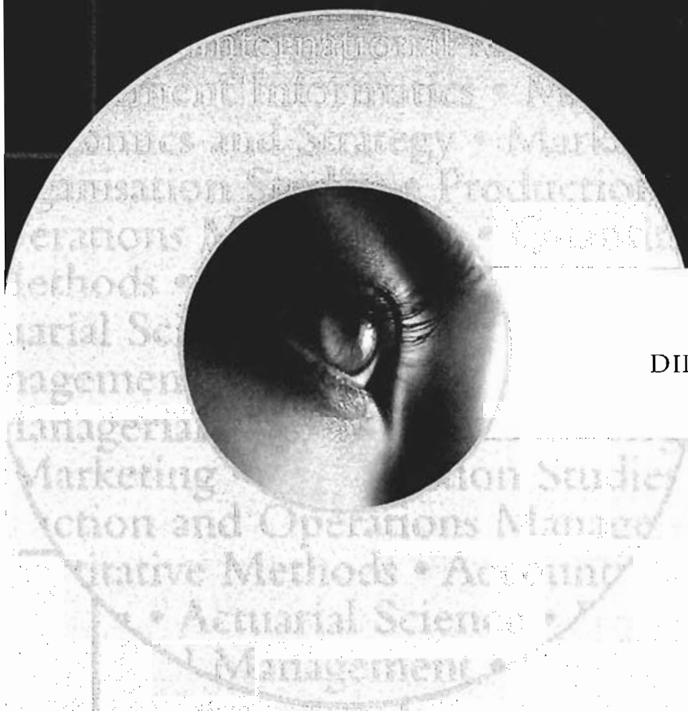

RESEARCH REPORT

INTERDEPENDENCIES BETWEEN EUROPEAN, U.S. AND JAPANESE STOCK MARKETS: DID THE EURO PROMOTE FURTHER INTEGRATION?

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Interdependencies between European, U.S. and Japanese Stock Markets: Did the Euro Promote Further Integration?

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Abstract:

This paper investigates the co-movement of stock markets in some major economic regions. Specifically, we examine the long-run interdependencies and short-term dynamics between the European market on the one hand and the U.S. and Japanese stock markets on the other hand. The results show that strong interactions exist between these markets. A shock originating in one market induces a persistent effect of the same direction in the other market on the same day. This effect generally tapers off on the second day. We further demonstrate that the interrelationships of the European market are stronger with the U.S. than with Japan. Interestingly, these interdependencies became stronger after January 1, 1999, which suggests that the introduction of the Euro has reduced international diversification benefits.

JEL: G11, G12, G15

Keywords: stock market integration, VAR modeling, persistence

I. Introduction

The European economic and monetary harmonization has fundamentally changed the relative behavior of European capital markets. Ever since the establishment of the EMU in 1992 and the introduction of the Euro as the common European currency in 1999, people have borne in mind the evolution towards a "United States of Europe" (Beckers, 1999).[1] On the academic level, a lot of studies recently have investigated the co-movement and relative behavior of various European stock markets. Friedman and Shachmurove (1997), for instance, conclude from a VAR model of daily returns that most European stock markets are closely connected. Similarly, Beckers (1999) notes that the stocks of nine core EMU member countries behave more and more similar, especially in the sectors finance and energy. From a spectral analysis of three European stock market indices, Asimakopoulos *et al.* (2000) point out strong similarities among the London FTSE100, Frankfurt DAX30 and Paris CAC40.

While the stock market integration of EMU member countries has been studied extensively and verified either directly or indirectly, the literature has paid little attention to the relation between the stock market returns of this "United States of Europe" as a whole and other major economic regions. This paper tries to fill this void by considering the EMU as a single market and exploring its stock market interdependencies with the U.S. and Japan. Specifically, we want to find out whether any synchronization or linkage exists between the European and these other stock markets. For this purpose, we build a Vector Autoregressive (VAR) framework to test cointegration, estimate Granger causality and calculate impulse response measures. In doing so, we are able to explore both the long-run interdependencies and short-term dynamics between the stock markets in Europe on the one hand and the U.S. and Japan on the other hand. Our data set covers the period between January 1, 1992 and December 31, 2002. In our estimations, we recognize one important date: January 1st of 1999, when the currencies of the EMU member countries became irrevocably fixed and the Euro became a valid

[1] The Economic and Monetary Union (EMU) includes twelve European member countries: Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, the Netherlands, Portugal and Spain.

transferable currency. This allows us to explore whether the formal introduction of the Euro has spurred further worldwide stock market integration.

The results show that the European, U.S. and Japanese stock markets bear no long-run equilibrium relationship. However, there is a strong short-term interaction mechanism between the EMU stock market index and the indices of the U.S. and Japan. Specifically, we find that a shock originating in one market generally induces a persistent effect of the same direction in the other market on the same day. When the effect lasts until the second day, it can be attributed to differences in time zone, as in Eun and Shim (1989). We further demonstrate that the interrelationships of the European market are stronger with the U.S. than with Japan. Finally, we document that the interaction mechanisms became stronger after the formal introduction of the Euro on January 1, 1999, which suggests that the Euro has reduced international diversification benefits. While the European and Japanese markets also became more closely connected from 1999 onwards, we point out that especially the European and U.S. markets became more integrated ever since the Euro was introduced.

The remainder of the paper is organized as follows. Section II briefly reviews the literature. Section III describes the data set. The testing procedures and empirical results are discussed in section IV. Finally, section V concludes the paper.

II. Literature review

For economists who need to understand the impact of worldwide share price movements on investment and consumption decisions, for capital market theorists who are interested in the segmentation of international equity markets and for individual investors who wish to diversify their portfolios, recognizing the relations among international stock markets is quite important (e.g., Panton *et al.*, 1976). Not surprisingly, research on stock market interrelationships dates already from the 1970s. This research has provided several arguments on why different stock markets may exhibit interdependencies, either in the short term or in the long run. These explanations can be categorized as economic interdependencies, market contagion and the free flow of capital and are discussed hereafter. We argue that the formal introduction of the Euro likely has further stimulated the free flow of capital. To motivate our own methodological choice, we also briefly elaborate on the various methodologies that have been used to empirically examine stock market interdependencies.

According to Choudhry (1996), strong economic ties and policy coordination among countries/regions can indirectly link their stock prices over time. Also, to the extent that one country's economy can influence that of others, for instance through imports and exports, expectations about economic developments may be somewhat similar across countries. Then, different stock markets may respond in the same way to economic shocks, inducing co-movement of stock prices (Madura, 1992).

Next, technological advances in connecting computers and instantaneous communication have contributed to the worldwide integration of stock markets by linking financial centers. As a result, shocks are being transmitted from one market to another, which is referred to as market contagion (Smith *et al.*, 1993). Also, multinational companies that simultaneously list on multiple exchanges can contribute to the worldwide co-movement of stock prices, as news about these firms will be reflected in several markets within a short period of time. As a result, stock markets of different sizes, structures and geographic locations can exhibit a high degree of co-movement after a shock in one market.

Finally, deregulation and liberalization of financial markets, innovations in financial products and services, and developments in telecommunications technology all have facilitated the free flow of capital. As a result, domestic investors can easily diversify their portfolios by investing internationally, which likely engenders worldwide stock prices changes in times of domestic fluctuation. By reducing the currency risk, the introduction of the Euro has made it easier for investors to diversify internationally (e.g., Beckers, 1999).

The literature has adopted many methodologies to explore potential interdependencies between stock markets. Traditionally, the existence and strength of these relationships was examined by means of correlation coefficients and simple regression analysis. The correlation method was motivated by portfolio theory, showing that the benefits of diversification depend on the degree of correlation between asset returns. Later on, cross-market correlation coefficients were commonly used for studying stock market interrelationships. Using this method, Grubel and Fadner (1971) show that domestic and foreign stocks became more closely related over time, especially in export-oriented industries. More recently, Forbes and Rigobon (2002) derive heteroskedasticity-adjusted correlation coefficients to test for stock market contagion during the 1987 stock market crash, the 1994 Mexican peso devaluation and the 1997 East Asian crisis. As they find no significant increase in cross-market correlations after the studied crises, they conclude that stock market interrelationships during these periods are due to these markets' inherent interdependencies rather than to market contagion. Simple regression analysis is another straightforward method to test for interrelationships among markets. By setting up a regression model for monthly stock market returns in the U.S., U.K., Germany and Japan, Agmon (1972) finds support for the integrated market hypothesis.

While the traditional approach is quite easy to implement, it fails to capture the dynamics of stock market interdependencies. The correlation method only reveals the existence and strength of a relation, but leaves the direction of such a relation unanswered. Simple regression analysis only describes a static relation at one point in time and needs a priori assumptions about its nature. By contrast, the time series approach – which includes methods from simple ARIMA analysis to long-run modeling – provides researchers with more powerful tools to determine the long-run interdependencies and short-term dynamics among stock markets. Using ARIMA analysis, Schollhammer and Sand (1985), for instance, find significant lead-lag relationships among share prices in the U.S., U.K., Germany, Switzerland and the Netherlands. Hamao *et al.* (1990) estimate the variance-covariance transmission mechanisms between countries around the 1987 U.S. stock market crash in an ARCH framework and find significant price-volatility spillovers from New York to London and Tokyo, and from London to Tokyo. Eun and Shim (1989) implement a VAR analysis to show that shocks in the

U.S. are rapidly transmitted to other markets. Choudhry (1996) examines long-run interrelationships among European markets during the 1920s and 1930s by means of cointegration tests. Friedman and Shachmurove (1997) estimate a VAR model, conduct Granger causality tests and calculate variance decomposition and impulse response measures of daily returns in eight major EMU countries, showing that the stock markets of most countries are highly related.

This paper also adopts a time series approach, but extends prior research by examining both the long-run interdependencies and short-term dynamics to obtain a more complete picture of the relative behavior of the European, U.S. and Japanese stock markets. Also, given that the formal introduction of the Euro likely has increased the co-movement of share prices, we split up the sampling period into two subperiods to trace potential changes.

III. Data

The data for this study are collected from Datastream and comprise time series of the daily close values of the stock market indices for the European market (i.e., the aggregate index for the EMU member countries as compiled by Datastream), the U.S. and Japan.[2] According to Eun and Shim (1989), monthly or weekly data may obscure interactions between stock markets that last for only a few days. Our study, as a result, does not suffer from this problem. Also, all indices are expressed in local currencies, which allows us to abstract from exchange rate changes. The local currency of the EMU market in the first subperiod is referred to as the "synthetic Euro" in the Datastream database. Finally, we take the natural logarithm of all data. Daily returns then are computed as the first differences of these transformed series. This is convenient given that percentage growth rates of economic variables,

[2] The market indices as compiled by Datastream include the most important countries in the considered market, based on their market capitalization. To compute these indices, weightings are determined by the market capitalization of each constituent country and thus do not represent the relative sizes of the economies.

like stock prices, are more constant over time than absolute growth rates. Using this logarithm transformation also allows us to interpret the parameter estimates as elasticities.

The data collection starts on January 1, 1992, shortly after the Maastricht Treaty was agreed on, and ends on December 31, 2002. This period is further divided into two subperiods: 1) from January 1, 1992 to December 31, 1998; and 2) from January 1, 1999 to December 31, 2002. This split-up of the sampling period should allow us to examine whether or not the formal introduction of the Euro has contributed to further stock market integration.

IV. Empirical analysis and results

The empirical analysis proceeds in the following steps. First, all variables are pre-tested for their order of integration using both Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests. After these unit root tests, we conduct a Johansen cointegration test based on the Vector Autoregressive (VAR) framework to examine whether a long-run equilibrium relationship exists between the EMU and the other markets and to identify this relationship, if any. As stock market indices are integrated of order one and as we find no cointegration, we subsequently build Vector Autoregressive (VAR) models in first differences. These models then are used to investigate Granger causality. Finally, we implement an impulse-response analysis to determine the size and timing of any short-term dynamics.

1. Unit root testing

Before proceeding with building models, we test the stationarity of the stock market indices by means of the widely used Augmented Dicker-Fuller and Phillips-Perron tests. The latter is a generalization of the ADF procedure that allows for milder assumptions regarding the error distribution. Since the true data generating process is unknown, several concerns arise before this test can be conducted. First, we need to determine the optimal number of lags. Including too many lags will reduce the power of the test. Conversely, omitting significant lags will lead to serial correlation in the residuals and hence bias the test results. In order to decide on the appropriate number of lags, we follow the iterative top-down approach of Enders (1995). We start with a lag-length of 60 days, which should be sufficient for daily data (e.g., Friedman and Shachmurove, 1997).

The second concern is on the appropriateness of including an intercept and/or time trend in the regression model, which is not trivial either. If we inappropriately omit the intercept or time trend, the power of the test can go to zero. Conversely, when redundant regressors are added, we may fail to reject the null hypothesis of a unit root. Following Enders (1995), we start with the least restrictive specification that includes both an intercept and a time trend. If the null hypothesis of a unit root is rejected, there is no need to proceed and we can safely conclude that there is no unit root in the data. By contrast, if the null hypothesis is not rejected, it is necessary to determine whether too many deterministic regressors were included in the previous step.

The preliminary results indicate that the null hypothesis of a unit root cannot be rejected for all markets. Since the means of all series are non-zero, we keep the intercept and then test for the significance of the time trend under the null of a unit root. If the time trend is not significantly different from zero, we subsequently estimate a restricted model without time trend. Table 1 reports our final specification and test results, showing that no time trend is needed for the entire sampling period whereas it is necessary to include a time trend in the subperiods, except for Japan in the first subperiod. Finally, to judge on the appropriateness of the specifications, we calculate the Ljung-Box Q model diagnostic for each lag. These statistics show that the residuals are white noise under the selected specifications.

Overall, the ADF and PP test results show that the hypothesis of a unit root in the stock market data cannot be rejected during the entire sampling period or in the subperiods. In other words, the stock market indices follow a random walk. This conclusion is consistent with weak-form market efficiency (e.g., Fama, 1970; 1991). Finally, when testing for a unit root in the first difference of each series, it turns out that in none of the markets and periods, such a unit root exists.[3] The latter findings thus confirm that all stock market index series are integrated of order one.

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Insert Table 1
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2. Cointegration testing

Having verified that all series are integrated of order one, we now test whether a cointegrating relation exists between stock market indices such that a stationary combination can arise out of non-stationary variables. Since the focus of our study is on the relative behavior of the European market vis-à-vis the other markets, we test the bivariate relationships between the EMU index and the U.S., respectively Japanese stock market index during the entire sampling period as well as during both subperiods.

According to Stock and Watson (1988), cointegrated variables share common stochastic trends and cointegrating vectors purge the trend from the linear combination of these variables. By investigating the potentially common stochastic trends underlying the European and the other markets, we are able to determine whether there exists a long-run equilibrium relationship between these markets. It is only by identifying the long-run relationship – if one indeed exists – that the short-term dynamics can be described more accurately. The reason is that the short-term dynamic path of the studied variables has to bear some connection with their deviation from the equilibrium relationship.

The two most widely used cointegration tests are the Engle-Granger test and the Johansen test. The Engle-Granger test is based on a single-equation regression of the variables that are potentially cointegrated. Then, the stationarity of the residuals from this regression model is used to decide on the equilibrium relation. This methodology can be implemented rather easily using OLS regression analysis. However, the Engle-Granger methodology has two important defects (Enders, 1995). First is the normalization problem. The test result is subject to the choice of dependent variable

[3] These results are not reported, but are available from the authors upon request.

and regressors in the regression model. Typically, different orderings of variables are used to ensure the validity of the results. Asymptotic theory suggests that the test for a unit root in the residuals from the different models will lead to the same conclusion as the sample size becomes infinitely large. However, in practice, the sample size generally is not large enough to satisfy this asymptotic condition. Second is the lack-of-power problem due to its reliance on a two-step estimator. The first step is to generate the error series, which in a second step are used to estimate a regression of the form $\Delta e_t = a_1e_{t-1} + a_2e_{t-2} + \ldots$ The coefficients a_i thus are obtained by estimating a regression that uses the residuals from another regression model. Therefore, any errors made in Step 1 are carried into Step 2.

By focusing on the relation between the rank of a matrix and its characteristic roots, the Johansen test solves the normalization and lack-of-power problems associated with the Engle-Granger test. Moreover, the Johansen test is able to provide estimates of all cointegrating vectors in the multivariate case and offers a framework for testing restrictions on the parameters of the implied long-run relation.[4] This study therefore will use the Johansen methodology to test for cointegration. The Johansen test is based on a Vector Autoregressive framework and has the following specification:

$$X_t = A_0 + \pi X_{t-1} + \varepsilon_t \tag{1}$$

where X_t is a (nx1) vector containing the n variables of interest

A_0 is a (nx1) vector of constants (allowing for a linear time trend in the data)

π is a (nxn) matrix of parameters

ε_t is a (nx1) vector of error terms (white noise)

The methodology centers on estimating the matrix π in unrestricted form and testing whether the restrictions, as reflected in the reduced rank of π, can be rejected. The number of cointegrating

[4] However, the Johansen test suffers from severe interpretation problems when more than one cointegrating relation is found. Given that the focus of our study in on the bivariate relation between two data series, this problem is unlikely.

vectors depends on the rank of π, which in turn is determined by its number of non-zero characteristic roots. For this purpose, we focus on the λ trace statistic, which tests the null hypothesis that the number of cointegrating vectors is less than or equal to r against a general alternative. Again, we need to think about the optimal lag-length and including additional regressors before the test can be conducted.

We follow the procedure suggested by Enders (1995) to first estimate a VAR model using the undifferenced data and then implement a lag-length test.[5] Also, we have to decide on including an intercept in the cointegrating vector and allowing for a linear time trend in the data. Given that all series have a non-zero mean, we include an intercept. As far as the trend term is concerned, we take into account the results from the unit root tests in Table 1 and visually inspect the data plots in Figure 1. The unit root test results and the data plots indicate that there is no time trend during the entire sampling period even though the trend terms are significantly different from zero during the two subperiods. During the first subperiod, the trend is increasing in the European and U.S. markets whereas the trend is declining in all markets during the second subperiod. So, we allow for a linear trend in the data during both subperiods, but assume no such trend over the entire sampling period.

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Insert Figure 1

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Table 2 now reports the specification and test results from the Johansen cointegration test. The results indicate that during each of the studied periods, no long-run equilibrium relationship exists

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[5] The lag-length test starts with the longest possible length to estimate the VAR model, yielding a variance/covariance matrix of residuals. Thereafter, the model is re-estimated using a shorter length, which yields another variance/covariance matrix. A Likelihood Ratio test then is used to determine whether the restriction is binding; if not, the shorter length is chosen. Given that our sample is large enough to satisfy the asymptotic distribution of the Likelihood Ratio test, we use the LR statistic for the purpose of model selection. Two other frequently used model selection criteria – the multivariate generalizations of AIC and SBC – favor more parsimonious models but the residuals from these models are not white noise.

between the European market on the one hand and the U.S. and Japanese markets on the other hand. This conclusion even holds at the 10% significance level. So, we can safely conclude that the European and the U.S., respectively Japanese markets do not share a common stochastic trend; in other words, these markets evolve independently from one another in the long run.

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Insert Table 2
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3. Model specification, Granger causality and impulse response functions

3.1. Model specification

In this section, we further investigate the relationship between the studied stock market indices. Since the time path of one stock market index may affect or be affected by the time path of another index, we cannot say anything about exogeneity. Then, VAR modeling is an appropriate choice as it enables us to treat each index series symmetrically. Since we want to describe the behavior of the European market relative to the others, we build bivariate models. Given that stock market indices are integrated of order one but share no common stochastic trend, we build VAR models in first differences.[6] Each VAR model in first differences then takes the following form:

$$\Delta X_t = A_0 + A_1 \Delta X_{t-1} + A_2 \Delta X_{t-2} + \ldots + A_p \Delta X_{t-p} + \varepsilon_t \qquad (2)$$

where ΔX_t is a (2x1) vector containing log(stock market index) in first differences

A_0 is a (2x1) vector of constants

A_i is a (2x2) matrix of parameters

ε_t is a (2x1) vector of error terms

[6] In case of cointegration, an error correction term has to be incorporated in the model; this special form of VAR modeling is referred to as Vector Error Correction (VEC) modeling.

An important concern regarding model specification again is the selection of the appropriate lag-length. In order to preserve the system's symmetry and make OLS-estimates consistent and asymptotically efficient, we use the same lag-length in both equations of the VAR model. Then, we apply the lag-length test suggested by Enders (1995) and use the LR statistic as model selection criterion. As the means of the series in first differences are not significantly different from zero, we include no intercept in the models. Finally, to obtain a more parsimonious specification and more accurate estimates, we implement a lag-exclusion test to remove redundant regressors from the system.[7] Table 3 reports the lag-determination test results, i.e. the lag-length and the significant lags kept in the VAR model after performing lag-exclusion tests. The Ljung Box-Q statistics suggest that our final specification is appropriate for each of the VAR models.

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Insert Table 3

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A serious disadvantage of VAR modeling is that it is hard to make sense out of the multitude of parameter estimates.[8] Rather than analyzing the parameter estimates directly, we use these estimates to calculate some other measures, such as Granger causality and impulse response functions, which are discussed in the following two sections.

3.2. Granger causality testing

Granger causality is an econometric relationship, testing whether the information contained in one variable helps to explain the other. Given that the variables are not cointegrated, it is sufficient to

[7] Lag-exclusion test: for each lag, the Wald statistic for the joint significance of all endogenous variables at that lag is calculated for each equation separately and jointly. In order to save on degrees of freedom and get more accurate estimates, we use this statistic to exclude the non-significant lags.

[8] The parameter estimates corresponding to each of the lags kept in the model are not reported but can be obtained from the authors upon request.

perform a standard Granger causality test on the VAR model in first differences to examine possible short-term market linkages. An important advantage of the test is that its results are unaffected by the ordering of the VAR system. Granger (1969) offers four definitions of causality, which in this context comprise: 1) unidirectional causality from the European market to the other market; 2) unidirectional causality from the other market to the European market; 3) feedback causality between the two considered markets; and 4) independence between these markets. The most direct way to determine Granger causality is to perform a standard F-test on the coefficients of the earlier estimated VAR models (see section 3.1.).

Table 4 summarizes the Granger causality test results. The null hypothesis of no Granger causality is strongly rejected in all cases, which indicates a strong feedback causality between the European and the other markets during the entire sampling period as well as during both subperiods (definition 3). These results, however, do not imply that we can realize abnormal returns in one market based on historical information from another market and, thus, that markets are inefficient. Rather, the feedback mechanism points to strong interdependencies between two markets in the sense that information in one market is (quickly) incorporated into the information set of the other market. If this indeed happens fast, one cannot earn abnormal returns in one market by using past information from other markets. To determine the exact speed of adjustment, we need to calculate impulse response functions.

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Insert Table 4
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3.3. Impulse response functions

Impulse response functions provide a more direct way for extracting information on interrelationships from the VAR system. Through a vector of moving average transformation (VMA), these functions trace out the time path of various shocks on the variables contained in the VAR system (Enders, 1995). However, computing impulse response functions is not straightforward, especially not when innovations in the VAR model are correlated. A widely used method to solve the problem of identifying impulse response coefficients is Cholesky decomposition, whereby an asymmetry in the system is assumed. In a bivariate VAR model, it is assumed that a shock originating in one variable does not have a contemporaneous effect on the other variable whereas a shock originating in the other variable is able to affect both variables simultaneously. This assumption thus imposes a causal ordering between stock markets.

According to Enders (1995), the consequences of ordering depend on the magnitude of the correlation coefficient between two shocks. If the absolute value of this correlation coefficient is larger than 0.2, then different ordering may lead to quite different results. In our case, all correlation coefficients turn out to exceed 0.2. Fortunately, Pesaran and Shin (1998) propose a generalized impulse response approach to tackle this problem.[9] This approach does not require orthogonization of shocks and is invariant to the ordering of the variables in the VAR system. Hence, our study computes impulse response functions using this generalized impulse response approach.

We investigate the impact of a one-standard deviation shock originating in one market on the returns in the other market during a window of 20 days afterwards.[10] This window should be long enough to capture the dynamics between markets, even when there is some delay in market reactions. Figures 2a to 2c display the impulse response graphs of the European and U.S. markets. Figure 2a is estimated over the entire sampling period whereas figures 2b and 2c are estimated over the subperiods 1/1/1992-12/31/1998, respectively 1/1/1999-12/31/2002. The upper graphs depict the impact of a shock on the stock market return (first difference in log stock prices) whereas the lower graphs

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[9] The basic idea of the generalized impulse response approach is to shock only one element by ε_t and to cancel out the effects of other shocks using an assumed or historically observed error distribution. See Pesaran and Shin (1998) for more technical details.

[10] In this study, we only examine the time path of a shock in one market on the returns in the other market; the impact of shocks arising simultaneously in multiple markets thus is beyond the scope of our study.

illustrate the accumulated (persistent) effects on the level of (log) stock prices. The solid line represents the time path of the effect whereas the dashed lines trace out a two-standard deviation band around this effect.

As indicated by the graphs, the impulse response measures are consistent with the Granger causality test results that show a strong feedback mechanism between the European and U.S. markets. Both during the entire sampling period as in the two subperiods, a one-standard deviation positive shock originating in one market engenders a significant increase in the returns of the other market. Also, the impact on the level of stock prices is not reversed as time goes by. For the entire sampling period, a one-standard deviation positive shock originating in the European market increases the U.S. stock market index by 0.49% on the same day whereas a one-standard deviation positive shock in the U.S. market impacts the European stock market index by 0.45%. The effect of a European shock on the U.S. stock market tapers off on the next day, while the impact of a U.S. shock on the European stock market lasts for two days. These findings likely reflect that the two markets belong to different time zones (see also Eun and Shim, 1989). The stock markets of the EMU member countries actually close shortly after the opening of markets in the U.S. Further inspection of the graphs reveals that the effect of a shock from either source lasts for quite a number of days. Since the magnitude of these effects is relatively limited and mostly insignificant, these findings cannot be interpreted as evidence against market efficiency. Also, indirect effects from other sources may be responsible for some of the observed effects. The latter issue is beyond the scope of our paper, however.

Overall, the impact of a shock in one market on the stock market returns in the other region is not reversed over time. Assuming that the whole effect is realized over the studied twenty-day window, the impact of a one-standard deviation European shock on the U.S. market converges to a permanent change of about 0.6% whereas the impact of a one-standard deviation U.S. shock on the European market converges to a permanent change of 0.89%. Similar patterns are found in both subperiods. Nevertheless, shocks prove to have a much larger impact on stock price levels during the second period. Specifically, in response to a one-standard deviation positive European shock, U.S. stock prices rise by a permanent 0.40% in the first subperiod whereas the impact amounts to 0.90% in the second subperiod. In response to a one-standard deviation positive U.S. shock, European stock prices rise by a permanent 0.67% in the first subperiod whereas the impact amounts to 1.2% in the second subperiod.

Finally, the impact of a shock in the U.S. market on the European stock market is larger than that of a European shock on the U.S. stock market. This result is consistent with earlier conclusions on the leading role of the U.S. market in the world (e.g., Eun and Shim, 1989). Moreover, the two stock markets became more closely related over time, given the significantly increased interaction effects in the second subperiod. These results suggest that the introduction of the Euro has made a significant contribution to the further integration of European and U.S. stock markets.

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Insert Figure 2
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Figures 3a to 3c now display the impulse response graphs of the European and Japanese stock markets. These two markets are shown to interact to some extent, which corroborates the results from the Granger causality tests. During the entire sampling period as well as during both subperiods, the index of one market rises significantly in response to a one-standard deviation positive shock originating in the other market. Also, the impact on stock price levels is shown to persist over time. For the entire sampling period, a one-standard deviation positive shock originating in the European market induces the Japanese stock market index to rise by 0.29% on the same day whereas a similar shock in the Japanese market impacts the European stock market index by 0.25%. Also, European shocks continue to impact Japanese stock market returns on the next day, leading to another increase by 0.31% whereas the impact of Japanese shocks in the European market generally tapers off on the second day. Again, this result can be explained by the time zone factor: the Japanese stock market closes before markets in Europe open. Nevertheless, there are some minor resiliencies in European and Japanese stock market reactions, especially during the first subperiod.

Overall, the impact of one-standard deviation European shock on the Japanese market converges to an increase of about 0.62% after twenty days while the impact of a one-standard deviation Japanese shock on the European market converges to a permanent increase of only 0.26%. These patterns hold across both subperiods: in response to a European shock, there is a long-run impact of 0.60% during the first subperiod, respectively 0.68% during the second subperiod. Conversely, in response to a Japanese shock, the permanent increase amounts to 0.21% during the first subperiod, respectively 0.33% during the second subperiod. Together, these results indicate that stock markets in Europe have a much larger impact on the Japanese market than vice versa. Also, given that the persistence measure increased during the second subperiod, we can safely conclude that the European and Japanese stock markets became more closely related after the formal introduction of the Euro in 1999. The enlarged market integration, however, is less spectacular than that between Europe and the U.S.

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Insert Figure 3
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V. Conclusions

This paper uses a time series approach to analyze the long-run interdependencies and short-term dynamics between stock markets in Europe, the U.S. and Japan. We study these relationships over the period January 1, 1992 – December 31, 2002. Also, we split up the sampling period, using the formal introduction of the Euro on January 1, 1999 as a dividing line. Unit root tests indicate that the stock market indices of the studied economic regions follow a random walk, which is consistent with weak-form market efficiency. Also, we find no evidence of cointegration between the index series. In other words, these stock market indices exhibit no long-run equilibrium relationship. Therefore, we build VAR models in first differences to derive Granger causality and calculate impulse response measures that capture the short-term dynamics between stock markets. The Granger causality test results indicate that there is a strong feedback mechanism between the European market on the one hand and the U.S. and Japanese markets on the other hand. This mechanism manifests during the entire sampling period as well as during both subperiods. Impulse response analysis based on VAR modeling corroborates the Granger causality test results and provides more details about the short-term dynamics. We find that a shock originating in one market induces an effect of the same direction in the other market, on the same day. The effect generally tapers off on the second day, although some minor resiliencies remain during a period as far as twenty days after the original shock. Also, we show that shocks in the U.S. market have a larger impact on the European market whereas shocks in the European market have more effect on the Japanese market. Given that all markets belong to different time zones, we cannot attribute differences in shock adjustment speed to stock market inefficiencies; rather, the pre-established time differences likely are responsible for the observed delays (see also Eun and Shim, 1989). Overall, shocks in the European market are quickly transmitted to the other markets whereas shocks in the U.S. and Japanese markets are also rapidly incorporated in European share prices.

Finally, we find that the interrelationships between the European market on the one hand and the U.S. and Japanese markets on the other hand started long before the formal introduction of the Euro in 1999. Interestingly, we find that the interactions became even stronger after January 1, 1999. While European and Japanese markets also became more integrated from 1999 onwards, we especially point out the increased integration between European and U.S. markets ever since the Euro was introduced. In sum, our study confirms that the segmented market hypothesis does not hold and concludes that the benefits from diversifying investment portfolios between the European, U.S. and Japanese stock markets have decreased over time.

References

Agmon, T., 1972, The relations among equity markets: a study of share price co-movements in the United States, United Kingdom, Germany and Japan, Journal of Finance 27(4), 839-856.

Asimakopoulos, I., Goddard, J. and C. Siriopoulos, 2000, Interdependence between the US and major European equity markets: evidence from spectral analysis, Applied Financial Economics 10(1), 41-47.

Beckers, S., 1999, Investment implications of a single European capital market, Journal of Portfolio Management 25(3), 7-17.

Choudhry, T., 1996, Interdependence of stock markets: evidence from Europe during the 1920s and 1930s, Applied Financial Economics 6(3), 243-249.

Enders, W., 1995, Applied econometric time series (John Wiley & Sons, Inc.).

Engle, R. and C. Granger, 1987, Cointegration and error correction: representation, estimation, and testing, Econometrica 55, 251-276.

Eun, C.S. and S. Shim, 1989, International transmission of stock market movements, Journal of Financial & Quantitative Analysis 24(2), 241-256.

Fama, E.F., 1970, Efficient capital markets: a review of theory and empirical work, Journal of Finance 25(2), 383-417.

Fama, E.F., 1991, Efficient capital markets: II, Journal of Finance 46(5), 1575-1617.

Forbes, K.J. and R. Rigobon, 2002, No contagion, only interdependence: measuring stock market co-movements, Journal of Finance 57(5), 2223-2262.

Friedman, J. and Y. Shachmurove, 1997, Co-movements of major European community stock markets: A vector autoregression analysis, Global Finance Journal 8(2), 257-277.

Granger, C., 1969, Investigating causal relations by econometric models and cross-spectral methods, Econometrica 37, 424-436.

Grubel, H.G. and K. Fadner, 1971, The interdependence of international equity markets, Journal of Finance 26(1), 75-94.

Hamao, Y., Masulis, R.W., and V.K. Ng, 1990, Correlations in price changes and volatility across international stock markets, The Review of Financial Studies 3, 281-307.

Johansen, S., 1991, Estimation and hypothesis testing of cointegration vectors in Gaussian vector autoregressive models, Econometrica 59, 1551-1580.

Madura, J., 1992, Financial markets and institutions (West Publishing Company, New York).

Panton, D.B., Lessig, V.P., and O.M. Joy, 1976, Co-movement of international equity markets: a taxonomic approach, Journal of Financial & Quantitative Analysis 11(3), 415-432.

Perron, P., 1989, The great crash, the oil price shock, and the unit root hypothesis, Econometrica 57, 1361-1401.

Pesaran, M.H. and Y. Shin, 1998, Generalized impulse response analysis in linear multivariate models, Economics letters 58, 17-29.

Schollhammer, H. and O. Sand, 1985, The interdependence among the stock markets of major European countries and the United States: An empirical investigation of interrelationships among national stock prices movements, Management International Review 25(1), 17-26.

Smith, K.L., Brocato, J and J.E. Rogers, 1993, Regularities in the data between major equity markets: evidence from Granger causality tests, Applied Financial Economics 3(1), 55-60.

Stock, J. and M. Watson, 1988, Testing for common trends, Journal of the American Statistical Association 83, 1097-1107.

Figure 1: Data Plots

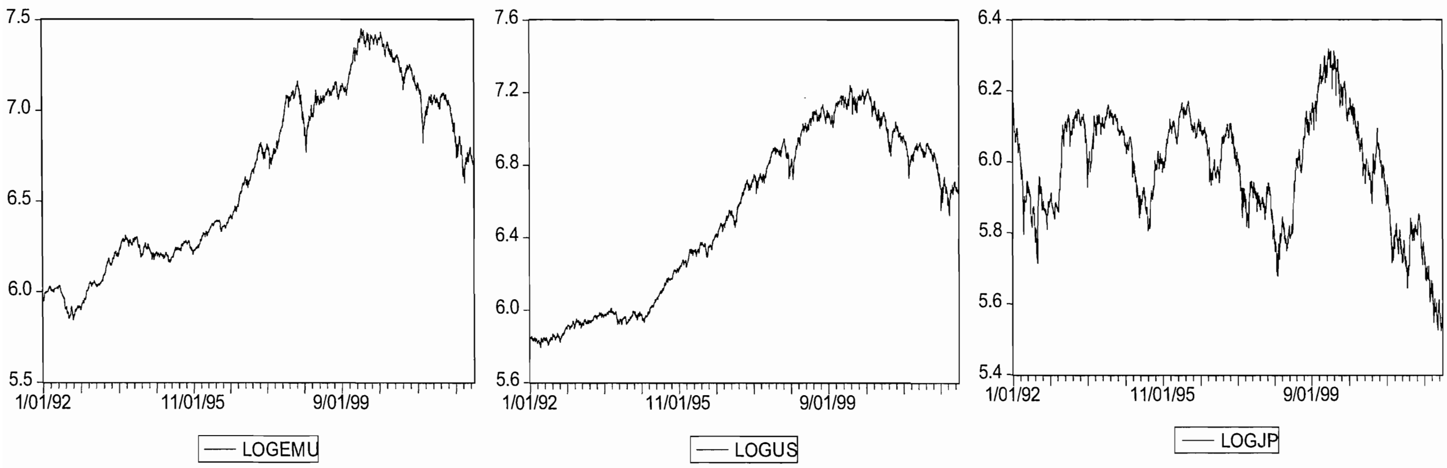
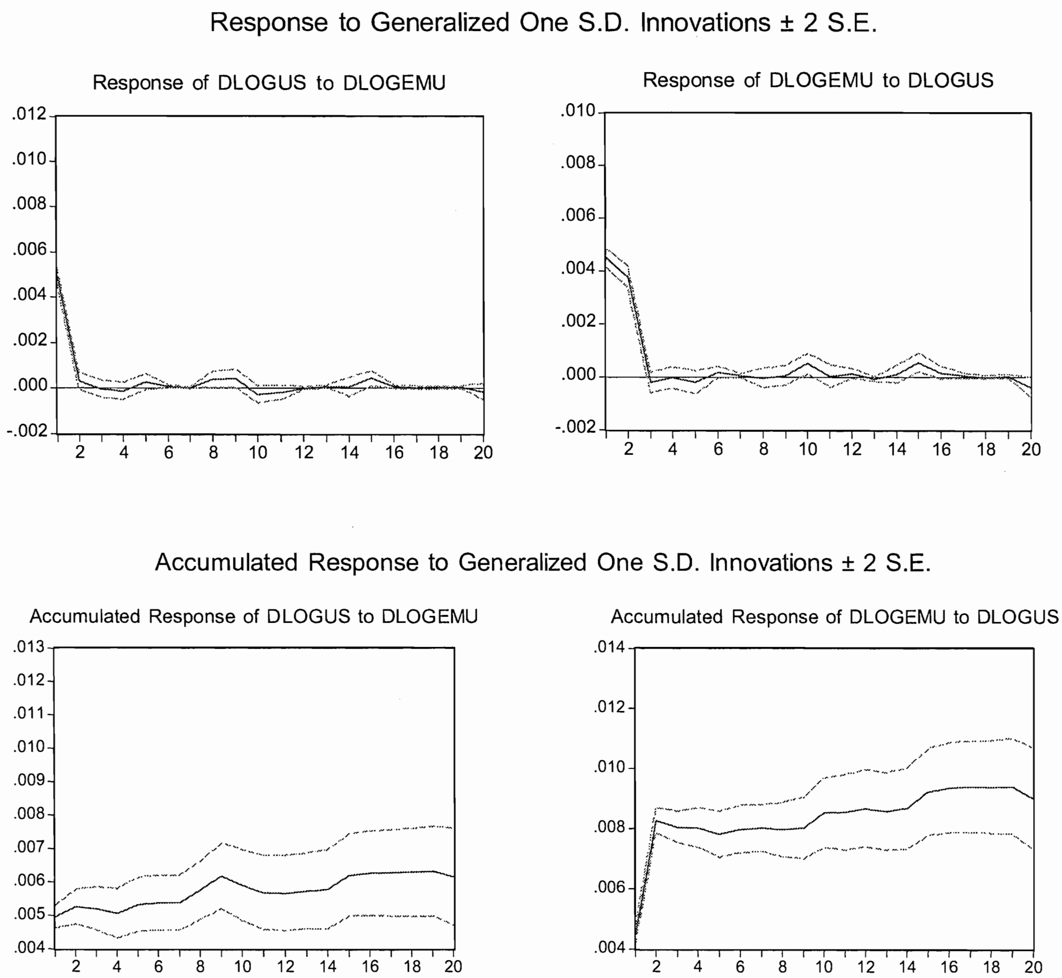


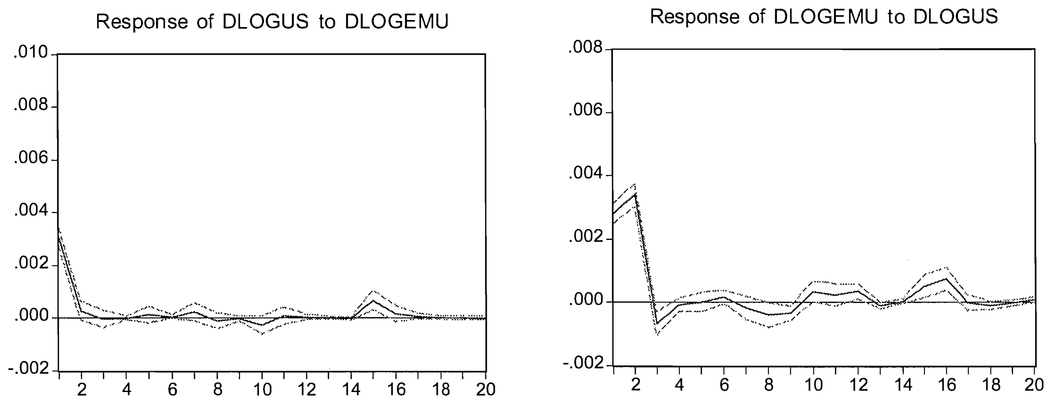
Figure 2: Impulse response graph of the European and U.S. markets

2a. Entire sample period from 1/1/1992-12/31/2002

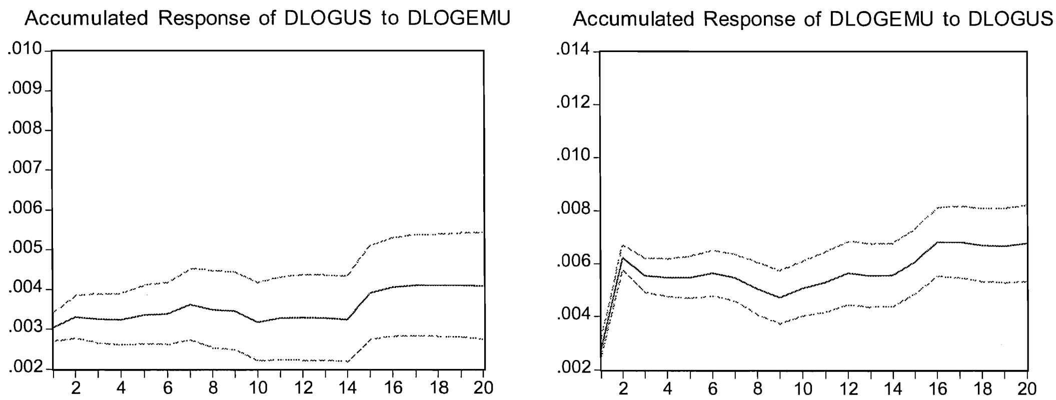


2b. The first subperiod from 1/1/1992 to 12/31/1998

Response to Generalized One S.D. Innovations ± 2 S.E.

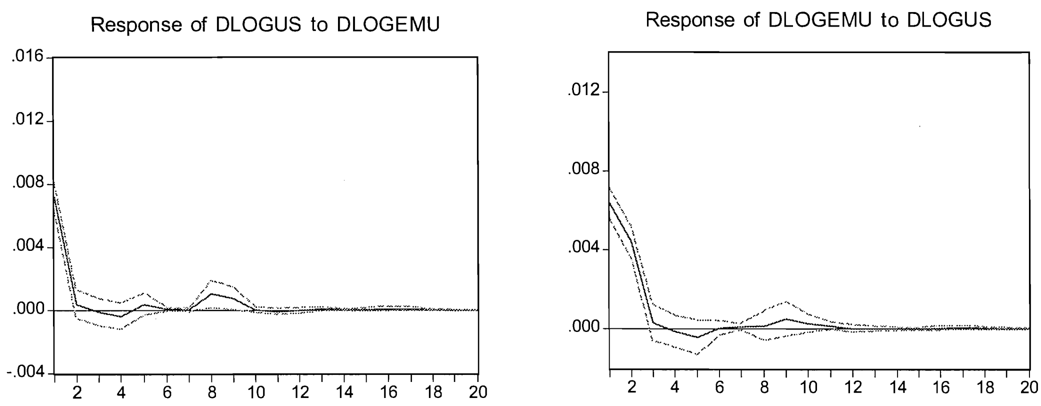


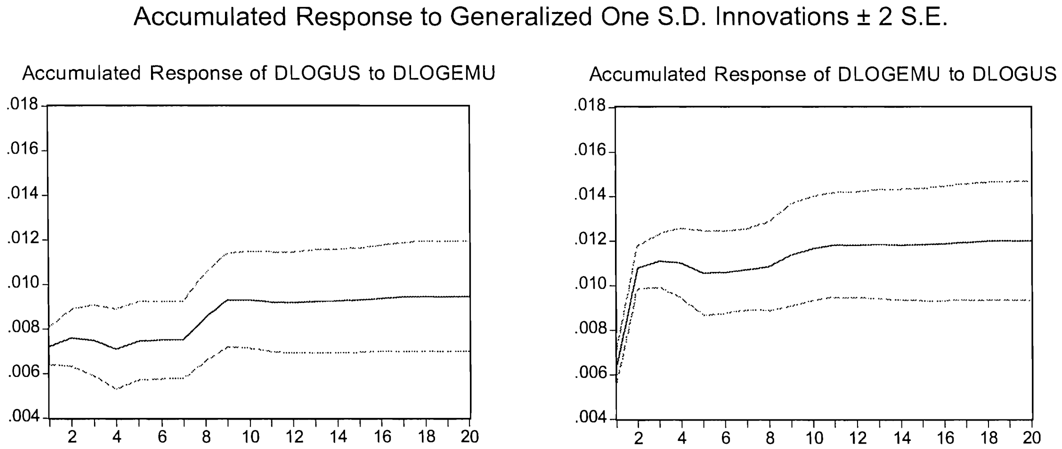
Accumulated Response to Generalized One S.D. Innovations ± 2 S.E.



2c. The second subperiod from 1/1/1999 to 12/31/2002

Response to Generalized One S.D. Innovations ± 2 S.E.

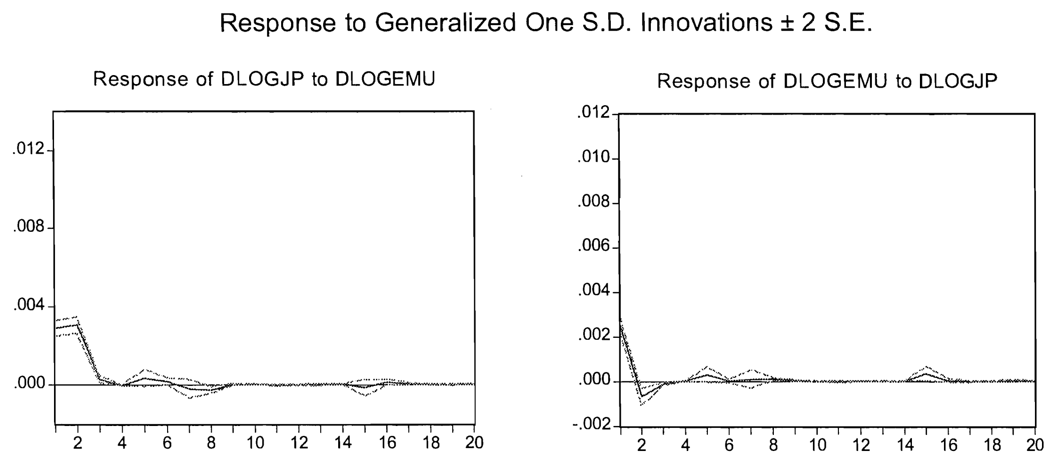




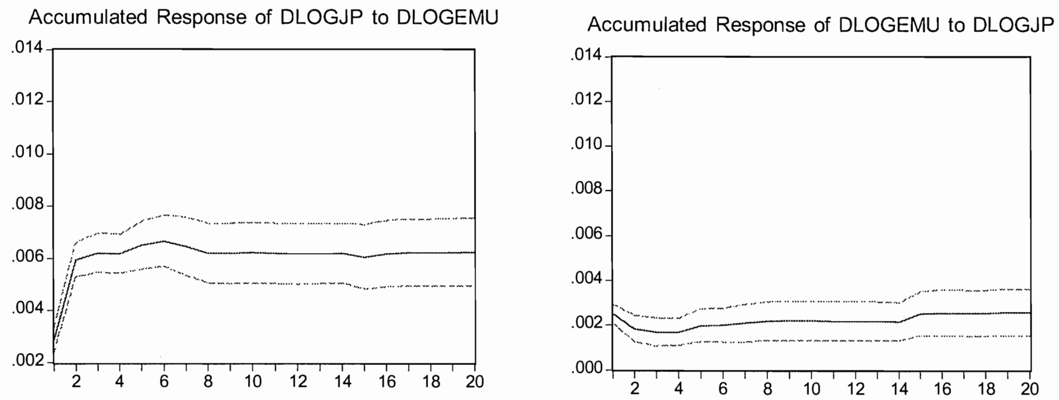
Note: "Response of DLOGUS to DLOGEMU" denotes the effect on U.S. stock market returns in response to an EMU shock and vice versa. "Accumulated response of DLOGUS to DLOGEMU" denotes the persistent effect of an EMU shock on U.S. stock market returns and vice versa.

Figure 3: Impulse response graph of the European and Japanese markets

3a. Entire sample period from 1/1/1992-12/31/2002

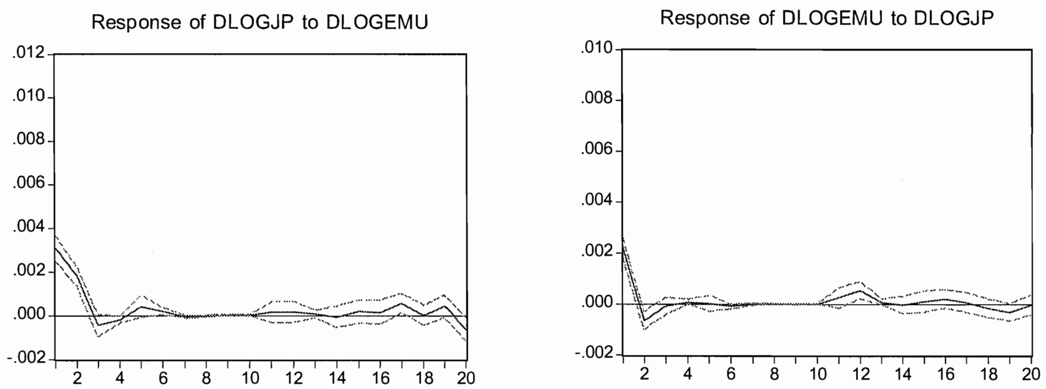


Accumulated Response to Generalized One S.D. Innovations ± 2 S.E.

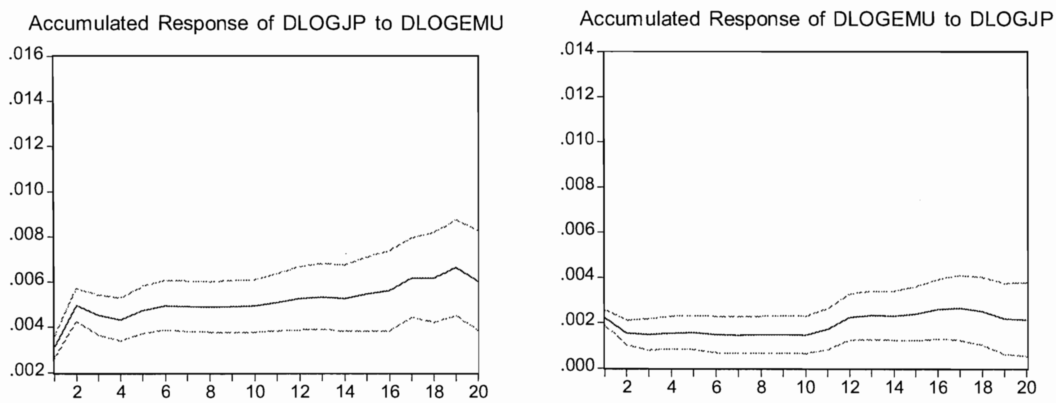


3b. The first subperiod from 1/1/1992 to 12/31/1998

Response to Generalized One S.D. Innovations ± 2 S.E.

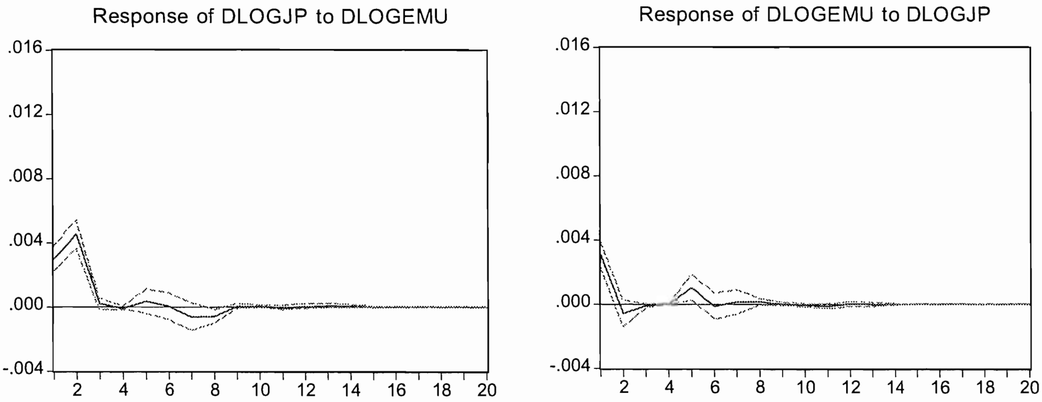


Accumulated Response to Generalized One S.D. Innovations ± 2 S.E.

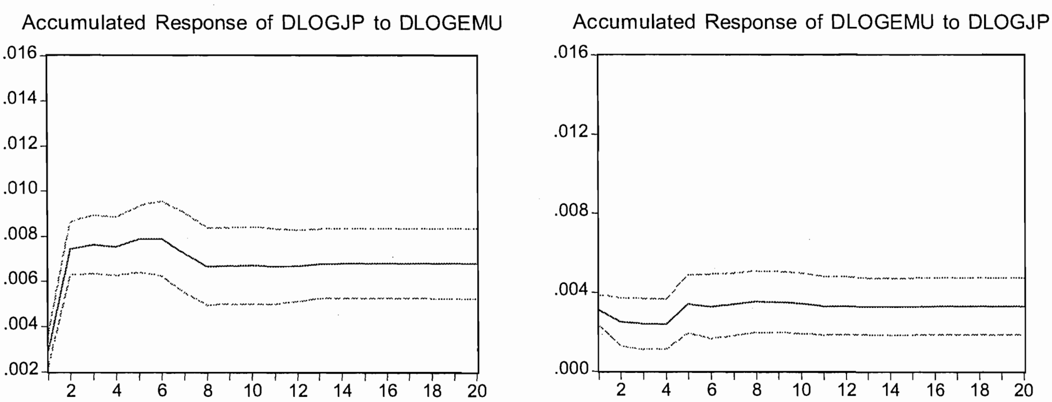


3c. The second subperiod from 1/1/1999 to 12/31/2002

Response to Generalized One S.D. Innovations ± 2 S.E.



Accumulated Response to Generalized One S.D. Innovations ± 2 S.E.



Note: "Response of DLOGJP to DLOGEMU" denotes the effect on Japanese stock market returns in response to an EMU shock and vice versa. "Accumulated response of DLOGUS to DLOGEMU" denotes the persistent effect of an EMU shock on Japanese stock market returns and vice versa.

Table 1: Unit root test results

Panel A: Entire sample period from 1/1/1992 to 12/31/2002

Market	Number of lags	Deterministic components	Test statistic	
			ADF	PP
EMU	50	Intercept only	-1.4782	-1.5238
U.S.	23	Intercept only	-1.4697	-1.4378
Japan	27	Intercept only	-1.5333	-1.3754

Panel B: The first subperiod from 1/1/1992 to 12/31/1998

Market	Number of lags	Deterministic components	Test statistic	
			ADF	PP
EMU	58	Intercept and trend	-1.9054	-1.5173
U.S.	59	Intercept and trend	-1.5518	-1.8952
Japan	35	Intercept only	-2.1544	-2.0654

Panel C: The second subperiod from 1/1/1999 to 12/31/2002

Market	Number of lags	Deterministic components	Test statistic	
			ADF	PP
EMU	50	Intercept and trend	-1.2936	-1.4407
U.S.	43	Intercept and trend	-2.1270	-2.5673
Japan	55	Intercept and trend	-2.6528	-2.7073

Note: The critical values are –2.5722 at the 1% level, –1.9406 at the 5% level and –1.6162 at the 10% level when there is no intercept nor trend term; –3.4540 at the 1% level, –2.8714 at the 5% level and –2.5720 at the 10% level when there is only an intercept; –3.9912 at the 1% level, –3.4262 at the 5% level and –3.1358 at the 10% level if there is both an intercept and a trend term. The critical values for the Phillips-Perron test are the same as those for the ADF test.

Table 2: Cointegration test results

Panel A: EMU and U.S.

	Lag-length	Trend assumption	λ_{trace} value	10% critical value	5 % critical value	1 % critical value	No. of CE(s)
Entire sample period (1/1/1992-12/31/2002)	21	No	17.816	17.85	19.96	24.60	None
First subperiod (1/1/1992-12/31/1998)	17	Yes	10.032	13.33	15.41	20.04	None
Second subperiod (1/1/1999-12/31/2002)	4	Yes	7.7115	13.33	15.41	20.04	None

Note: H_0: r = 0; H_1: r = 1

Panel B: EMU and Japan

	Lag-length	Trend assumption	λ_{trace} value	10% critical value	5 % critical value	1 % critical value	No. of CE(s)
Entire sample period (1/1/1992-12/31/2002)	15	No	9.8186	17.85	19.96	24.60	None
First subperiod (1/1/1992-12/31/1998)	20	Yes	7.8043	13.33	15.41	20.04	None
Second subperiod (1/1/1999-12/31/2002)	7	Yes	6.8432	13.33	15.41	20.04	None

Note: H_0: r = 0; H_1: r = 1

Table 3: Lag-determination test results

	Lag-determination by LR	
	EMU and U.S.	EMU and Japan
Entire sample period (1/1/1992-12/31/2002)	24 (1-4, 7-10, 13-14, 19-20, 24)	14 (1, 4, 6, 14)
First subperiod (1/1/1992-12/31/1998)	20 (1-2, 4, 6-7, 9-10, 14-15, 20)	19 (1-2, 4, 10-11, 13-19)
Second subperiod (1/1/1999-12/31/2002)	24 (1-4, 7-8, 24)	6 (1, 4-6)

Note: The figures in parentheses indicate the significant lags kept in the VAR model after performing the lag-exclusion test

Table 4: Granger causality test results

Panel A: EMU and U.S.

	Null Hypothesis	F-statistic	p-value	Conclusion
Entire sample period (1/1/1992-12/31/2002)	logEMU dnc logUS	3.067	0.000	Reject Null
	logUS dnc logEMU	35.318	0.000	Reject Null
First subperiod (1/1/1992-12/31/1998)	logEMU dnc logUS	2.110	0.020	Reject Null
	logUS dnc log EMU	41.359	0.000	Reject Null
Second subperiod (1/1/1999-12/31/2002)	logEMU dnc logUS	3.322	0.002	Reject Null
	logUS dnc logEMU	24.404	0.000	Reject Null

Panel B: EMU and Japan

	Null Hypothesis	F-statistic	p-value	Conclusion
Entire sample period (1/1/1992-12/31/2002)	logEMU dnc logJP	44.919	0.000	Reject Hull
	logJP dnc logEMU	6.689	0.000	Reject Null
First subperiod (1/1/1992-12/31/1998)	logEMU dnc logJP	4.725	0.000	Reject Null
	logJP dnc logEMU	3.061	0.000	Reject Null
Second subperiod (1/1/1999-12/31/2002)	logEMU dnc logJP	32.621	0.000	Reject Null
	logJP dnc logEMU	2.476	0.043	Reject Null

Note: "dnc" denotes "does not Granger cause"