# The Euro and European Financial Market Integration

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**Abstract**

We use a time-varying copula model to investigate the impact of the introduction of the Euro on the dependence between seventeen European stock markets during the period 1994-2003. The model is implemented with a GJR-GARCH-t model for the marginal distributions and the Gaussian copula for the joint distribution, which allows capturing time-varying, non-linear relationships. The results show that, within the euro area, market dependence increased after the introduction of the common currency only for large equity markets, such as in France, Germany, Italy, the Netherlands and Spain, while transaction costs remain important barriers to investment in and thus integration of smaller markets. Structural break tests indicate that the increase in financial market integration started around the beginning of 1998 when euro membership was determined and the relevant information was announced. We also estimate time-varying dependence measures for non-euro European countries with the euro-zone equity market. The UK and Sweden, but not other countries outside the euro area, are found to exhibit an increase in equity market co-movement, which is consistent with the interpretation that these countries may be expected to join the euro in the future.

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## Abstract

We use a time-varying copula model to investigate the impact of the introduction of the Euro on the dependence between seventeen European stock markets during the period 1994-2003. The model is implemented with a GJR-GARCH-t model for the marginal distributions and the Gaussian copula for the joint distribution, which allows capturing time-varying, non-linear relationships. The results show that, within the euro area, market dependence increased after the introduction of the common currency only for large equity markets, such as in France, Germany, Italy, the Netherlands and Spain, while transaction costs remain important barriers to investment in and thus integration of smaller markets. Structural break tests indicate that the increase in financial market integration started around the beginning of 1998 when euro membership was determined and the relevant information was announced. We also estimate time-varying dependence measures for non-euro European countries with the euro-zone equity market. The UK and Sweden, but not other countries outside the euro area, are found to exhibit an increase in equity market co-movement, which is consistent with the interpretation that these countries may be expected to join the euro in the future.

*"I believe that the key question for us – public authorities as well as market participants – is how we can contribute to the further integration of financial markets in Europe.[...] The potential gains from monetary union will only be fully realised if remaining barriers to integration of European financial markets are effectively removed. There is considerable evidence that wholesale markets are now much more integrated than before. But integration in securities markets needs to proceed further. Without an integrated European securities market the outcome of the entire process of financial market integration risks falling short of expectations."*

*Keynote speech by ECB President Jean-Claude Trichet at Deutsche Börse's New Year's Reception 2004, Frankfurt am Main, January 26, 2004.*

# 1 Introduction

The introduction of the Euro has been one of the most important events for global financial markets in the last decade. An immediate consequence of the adoption of the common currency was an integration of the euro-zone money and bond markets (Adjaouté and Danthine, 2003; Hartmann et al., 2003). Increasing integration of the equity markets within the euro-zone is likely to be another consequence of the elimination of exchange rate risk across countries within the euro area as a result of the adoption of a single currency. Detken and Hartmann (2000, 2002) and Perée and Steinherr (2001) show that the euro has become one of the three major currencies in the world after its introduction, taking its place alongside the U.S. dollar and the Japanese yen. Consequently, the impact of the introduction of the euro on the integration of equity markets within Europe is an important issue with significant implications for asset management, risk management and international asset pricing.

To assess this impact of the euro, this paper provides a comprehensive analysis of financial market integration between 17 European countries during the period 1994-2003. While previous work has studied market integration based on international capital mobility (Lemmen and Eijffinger, 1998; Frankel, 1992; Frankel and MacArthur, 1988; Feldstein and Horioka, 1980), asset pricing models (Hardouvelis et al., 2001; Ferson and Harvey, 1991; Dumas and Solnik, 1995; Bekaert and Harvey, 1995), price and volatility spillovers (Koutmos and Booth, 1995; Richards, 1995; Kasa, 1992) or the development of correlation coefficients over time (Cappiello et al., 2003), we directly investigate the dependence or co-movement of stock market indices across countries using a new econometric methodology. In particular, financial market integration is assessed by estimating time-varying copula dependence models following the methodology of Patton (2005). The paper contributes to the literature by proposing a more direct and general copula model for modeling time-varying dependence between the prices of financial assets. Specifically, the model uses a GJR-GARCH-t model for the marginal distributions and the Gaussian copula for the joint distribution. The dependence parameter in the copula function is modeled as a time-varying process conditional on currently available information, allowing for time-varying, non-linear relationships. The proposed methodology can be extended to a multivariable model, which is useful for portfolio and risk management.

We successfully apply this model to the investigation of the impact of the introduction of the euro on the integration of European financial markets by assessing the dependence between stock markets in different countries. Since many papers, such as Bracker and Koch (1999), Longin and Solnik (1995), Wahab and Lashgari (1993), Madura and Soenen (1992), Fischer and Palasvirta (1990), Maldonado and Saunders (1981), Makridakis and Wheelwright (1974), demonstrate the instability of co-movements between financial asset prices, the measurement of dependence and its variation over time are important, yet difficult issues. Nevertheless, the use of a time-varying copula model allows us to investigate whether the equity markets in the euro area have experienced a structural increase in their level of dependence.

We find an increase in equity market dependence in the euro area after the introduction of the common currency, but only for relatively large markets, i.e. in France, Germany, Italy, the Netherlands and Spain. The increase in equity market dependence starts around the beginning of 1998, when euro membership was determined and the relevant information was released. We suggest that this increase in dependence reflects a higher degree of integration between European financial markets, although even without foreign exchange rate risk several remaining capital market imperfections, such as regulation, taxes, and transaction costs still prevent full integration of European equity markets. In particular, higher transaction costs and lower market liquidity are the main reasons that make smaller markets less attractive to institutional investors and thus represent important barriers to investment in and thus integration of these markets. For non-euro European countries, we find a rise in the dependence of the British and Swedish equity markets with the aggregate euro-zone stock market, which is consistent with the interpretation that these countries may be expected to join the euro in the future.

The remainder of this paper is organized as follows. Section 2 discusses financial market integration in general and in the context of the euro in particular, and it develops the hypotheses about the impact of the euro on financial market integration. Section 3 presents time-varying copula methodology in general, while Section 4 explains the implementation of the models used to test the hypotheses. The data used for the empirical analysis is presented in Section 5. Section 6 presents the empirical analysis and discusses the results. Finally, conclusions are stated in Section 7.

## 2 European Financial Market Integration

The integration of financial markets has long been an issue of interest to financial economists in academia and investment practice alike, as it has implications for the identification of opportunities for and barriers to international portfolio investment with important implications for portfolio allocation and asset pricing (Bartram and Dufey, 2001). In Europe, the harmonization of regulations and social welfare systems, most recently with the focus on pension arrangements, has been promoted as an important vehicle to reduce market frictions and barriers to cross-border mobility of all factors of production, i.e. capital and labor. In this context, the introduction of the euro has been a milestone step, triggering heated and in part controversial debate of whether the launch of the common currency represents a sensible tool to force more integration in Europe, or whether, indeed, it would require a higher degree of harmonization prior to the event in order to ensure its success. In fact, the global economic downturn that coincided with the introduction of the euro has emphasized the existing differences across European countries, and the lack of policy responses has contributed to slow economic growth in major economies (such as Germany and France) and Europe as a whole, culminating in recent violations of the Growth and Stability Pact by several countries.

In theory, if financial markets are not integrated, entailing differential investment and consumption opportunity sets across countries, investment barriers will affect investors' portfolio choices and companies' financing decisions. If purchasing power parity does not hold, exchange rates affect the cost of consumption across countries, and, thus, exchange rate risk influences the price of assets to foreign investors. International asset pricing models recognize these effects by including exchange rate risk as a systematic risk factor (e.g. Adler and Dumas, 1983; Stulz, 1981; Solnik, 1974) and can, thus, be used to empirically investigate the issue of financial market integration (Dumas and Solnik, 1995). In the same vein, the effect of the Economic and Monetary Union (EMU) on European stock market integration can be examined with a weighted average asset pricing model that includes the covariance between stock returns and exchange rate returns, suggesting that the forward interest differential between a country and Germany has played an important role for the degree of integration (Hardouvelis et al., 2001).

As the introduction of the euro means the elimination of exchange rate risk within the euro area, it has further reduced the remaining differences of investment and consumption opportunities across the member countries of the euro. As a result, there should be less regional preferences or discrimination between different national markets by investors given the risk and return characteristics of assets. Likewise, the absence of exchange rate risk allows corporations to raise funds across countries with fewer constraints and costs. In addition, the prices of assets in European markets are more determined by common factors due to the reduction of exchange rate risk.

As the degree of economic integration between countries can be measured by the extent of co-movement of their equity markets, we conjecture that the degree of dependence between the equity markets of the countries in the euro area has increased after the launch of the common currency. Since expectations about euro membership were already formed before its determination, it is likely that an increase in the dependence between euro country equity markets can be observed in the years prior to January 1, 1999, if capital markets reflect all available information efficiently. To illustrate, Danthine et al. (2001) document that there was already a consensus about euro membership among financial and economic forecasters in January 1998, and

Fratzscher (2002) suggests that European equity markets have become more integrated even since 1996.

In addition to foreign exchange rate risk, other barriers to international portfolio investment (including taxes on foreign security holdings and ownership restrictions) are crucial factors that prevent market integration. Consequently, in partially integrated economies, investors' portfolios may be biased towards home assets because the benefits of international diversification are not large enough to offset its costs (Cooper and Kaplanis, 2000; Eun and Janakiramanan, 1986; Errunza and Losq, 1985; Black, 1974; Stulz, 1981). Even without exchange rate risk, however, many differences between national markets for labor and capital in the euro area currently remain, based on regulation, language, familiarity, transaction costs, etc. Still, the launch of the common European currency was clearly associated with reduced exchange rate volatility and convergence of interest rates, lower cost of cross-country transactions, improved liquidity, breadth and depth of European capital markets, which have been noted as important drivers of integration in the euro area (Fratzscher, 2002; Danthine et al., 2001). Thus, the introduction of the euro may have increased European financial market integration, but not led to fully integrated markets.

As a result, the lack of integration may have lost some, but not all of its power as an explanation for the observed home bias in European financial markets. Consistent with stronger integration of financial markets in Europe, institutional investors increasingly organize their investment activities along industry sectors rather than countries, suggesting that the latter play a decreasing role in the investment decision (Holder et al., 2001; Tsatsaronis, 2001). At the same time, the composition of equity portfolios held by households in major European countries reveals that a strong home bias prevails in the equity market investments of retail investors, which could reflect a lack of financial market integration (Guiso et al., 2003). Nevertheless, as one of the most important obstacles for investment and financing across the countries participating in the euro has been eliminated, investors' investment decisions will be determined by other market characteristics such as size, liquidity and regulation. In fact, data on pension funds document that countries with large equity market capitalizations, such as Germany, France and Italy, exhibited particularly large capital inflows after the introduction of the euro (Adjaouté and Danthine, 2003).

There remain significant differences in transaction costs across European equity markets that suggest differential barriers to investment and thus integration of markets even within the euro area. In particular, estimates by Elkins McSherry indicate that after the introduction of the euro trading costs (represented by average market impact and total costs, in basis points) in larger European equity markets like Germany (7, 31), France (5, 29), Italy (11, 35), the Netherlands (6, 29) and Spain (13, 41) are still significantly lower than in smaller euro area markets such as Luxembourg (61, 82), Austria (17, 47), Portugal (12, 46), and Greece (11, 68). As a result, we hypothesize a stronger increase of dependence between countries with large market capitalization, which may proxy for the remaining disparities between national markets in the euro zone.

For non-euro European countries, especially the UK, Sweden and Denmark, which require a referendum for joining the euro, it is interesting to investigate whether market participants believe that these countries are likely to adopt the euro or not. If market participants expect that they will join the common European currency in the future, we conjecture that one should observe an increase in their market dependence with the euro-zone equity market. Although increasing dependence is not a sufficient criterion to conclude that these countries will definitely join the euro, it does reveal information about the expectations of market participants.

# 3 Time-varying Copula Dependence Theory

## 3.1 Conditional Copulas

We employ single-parameter conditional copulas to represent the dependence between two index returns, conditional upon the historical information provided by previous pairs of index returns. The parameter of the conditional copula, like the marginal densities of the separate index returns, depends upon the conditioning information. The general theory of copulas is covered in the books by Joe (1997) and Nelsen (1999) and finance applications are emphasized by Cherubini, Luciano and Vecchiato (2004). Important conditional theory has been developed and applied to financial market data by Patton (2004, 2005).

Let $X_t$ and $Y_t$ be random variables that represent two returns for period $t$ and let their conditional cumulative distribution functions (c.d.f.s) be $F_t(x_t|\Phi_{t-1})$ and $G_t(y_t|\Phi_{t-1})$ respectively, with $\Phi_{t-1}$ denoting all previous returns, i.e. $\{x_{t-i}, y_{t-i}, i > 0\}$. The conditional copula function, here denoted $C_t(u_t, v_t|\Phi_{t-1})$, is then defined by the time-varying bivariate c.d.f. of the random variables $U_t = F_t(X_t|\Phi_{t-1})$ and $V_t = G_t(Y_t|\Phi_{t-1})$, whose marginal distributions are uniform on the interval from zero to one. The conditional bivariate c.d.f. of $X_t$ and $Y_t$ is then

$$H_t(x_t, y_t|\Phi_{t-1}) = C_t(F_t(x_t|\Phi_{t-1}), G_t(y_t|\Phi_{t-1}) \mid \Phi_{t-1}).$$

Assuming that the c.d.f.s are differentiable as often as necessary, the conditional copula density function is

$$c_t(u_t, v_t \mid \Phi_{t-1}) = \frac{\partial^2 C_t(u_t, v_t \mid \Phi_{t-1})}{\partial u_t \partial v_t}.$$

Also, the bivariate conditional density function of $X_t$ and $Y_t$ is given by the product of the copula density and the two marginal conditional densities, respectively denoted by $f_t(x_t \mid \Phi_{t-1})$ and $g_t(y_t \mid \Phi_{t-1})$:

$$\begin{aligned} h_t(x_t, y_t \mid \Phi_{t-1}) &= \frac{\partial^2 H_t(x_t, y_t \mid \Phi_{t-1})}{\partial x_t \partial y_t} \\ &= c_t(F_t(x_t \mid \Phi_{t-1}), G_t(y_t \mid \Phi_{t-1}) \mid \Phi_{t-1}) \times f_t(x_t \mid \Phi_{t-1}) \times g_t(y_t \mid \Phi_{t-1}). \end{aligned} \tag{1}$$

### 3.2 Estimation of Parameters

The bivariate dynamics of the returns $X_t$ and $Y_t$ are determined by the three functions $f_t(x_t \mid \Phi_{t-1})$, $g_t(y_t \mid \Phi_{t-1})$ and $c_t(u_t, v_t \mid \Phi_{t-1})$. We specify appropriate parametric functions in Section 4.

Parameter estimation is straightforward when separate parameters are used in the functions $f_t, g_t$ and $c_t$, which we denote respectively by the vectors $\theta_x, \theta_y$ and $\theta_c$. The contribution to the log-likelihood of all the data made by the two observations at time *t* is then

$$\log h_t(x_t, y_t \big| \Phi_{t-1}, \theta) = \log c_t(u_t, v_t \big| \Phi_{t-1}, \theta_c) + \log f_t(x_t \big| \Phi_{t-1}, \theta_x) + \log g_t(y_t \big| \Phi_{t-1}, \theta_y), \tag{2}$$

with $\hat{\theta} = [\hat{\theta}_x; \hat{\theta}_y; \hat{\theta}_c]$. Summing these contributions across a set of times gives the log-likelihood of an observed time series of *n* pairs of returns $\{x_t, y_\text{t}, 1 \le t \le n\}$, which can be stated as

$$L_{x,y}(\theta) = L_{u,v}(\theta_c) + L_x(\theta_x) + L_y(\theta_y) \tag{3}$$

with $L_k$ denoting the sum of the log-likelihood function values across observations of the variable(s) *k*.

While it would be optimal to maximize $L_{x,y}(\theta)$, simultaneously for all the parameters, this is difficult to achieve in practice because the dimensions of the problem can be large. Drawing on the two-stage maximum likelihood framework of Newey and McFadden (1994) and White (1994), Patton (2004) proposes a two-stage estimation procedure that is appropriate for large samples when the dependence vector $\theta_c$ does not have any impact upon the marginal distributions.

In the first stage, the parameters of the marginal distributions parameters are estimated from univariate time series as:

$$\begin{aligned}\hat{\theta}_x &\equiv \arg\max \sum_{t=1}^{n} \log f_t(x_t \vert \Phi_{t-1}, \theta_x), \\ \hat{\theta}_y &\equiv \arg\max \sum_{t=1}^{n} \log g_t(y_t \vert \Phi_{t-1}, \theta_y).\end{aligned} \tag{4}$$

The second stage then estimates the dependence parameter(s) as:

$$\hat{\theta}_c \equiv \arg\max \sum_{t=1}^{n} \log c_t(u_t, v_t \vert \Phi_{t-1}, \theta_c). \tag{5}$$

Patton (2004) shows that the two-stage ML estimates $\hat{\theta} = [\hat{\theta}_x; \hat{\theta}_y; \hat{\theta}_c]$ are asymptotically as efficient as one-stage ML estimates.

The variance-covariance matrix of $\hat{\theta}$ has to be obtained from numerical derivatives. We have only been able to obtain satisfactory first derivatives, from which the fully efficient two-stage estimator $(n\hat{B})^{-1}$ of the variance-covariance matrix can be obtained from

$$\hat{B} = n^{-1} \sum_{t=1}^{n} \hat{s}_t \hat{s}_t{}'$$

where the score vector is $\hat{s}_t = \partial \log h_t / \partial\theta$ evaluated at $\theta = \hat{\theta}$.

## 4 Empirical Methodology

### 4.1 Models for Marginal Distributions

It is well-known that the conditional densities of equity index returns are leptokurtic and have variances that are asymmetric functions of previous returns (Nelson, 1991; Engle and Ng, 1993). Consequently, we choose to obtain our marginal distributions by fitting the GJR-GARCH(1,1) model (Glosten, Jagannathan and Runkle, 1993) with conditional Student t-distributions.

Let $R_{i,t}$ and $h_{i,t}$ respectively denote the return from equity index $i$ and its conditional variance for period $t$. The ARCH model for the returns from index $i$ is defined by:

$$\begin{aligned} R_{i,t} &= \mu_i + \varepsilon_{i,t}, \\ h_{i,t} &= \omega_i + \beta_i h_{i,t-1} + \alpha_{i,1}\varepsilon_{i,t-1}^2 + \alpha_{i,2} s_{i,t-1}\varepsilon_{i,t-1}^2, \\ \varepsilon_{i,t} \mid \Phi_{t-1} &\sim t_{\nu_i}(0, h_{i,t}), \end{aligned} \tag{6}$$

with $s_{i,t-1} = 1$ when $\varepsilon_{i,t-1}$ is negative and otherwise $s_{i,t-1} = 0$.

All of the parameters, including the degrees of freedom $\nu_i$, are estimated separately for each equity index. In the first stage of parameter estimation, the following log-likelihood function is maximized for each of two time series of index returns ($i = x, y$):

$$\log L_i = \sum_{t=1}^{n} [-\tfrac{1}{2}\log(h_{i,t}) + \log(\frac{\Gamma((\nu_i+1)/2)}{\Gamma(\nu_i/2)\sqrt{\pi(\nu_i-2)}}) - \frac{\nu_i+1}{2}\log(1+\frac{z_{i,t}^2}{\nu_i-2})]$$

where $\Gamma(.)$ is the gamma function and $z_{i,t} = (R_{i,t} - \mu_i)\big/\sqrt{h_{i,t}}$ is the standardized value of the index return $R_{i,t}$.

## 4.2 Models for Bivariate Distributions

The estimated marginal ARCH c.d.f.s provide numerical values of $u_t = F_t(R_{x,t} \mid \Phi_{t-1})$ and $v_t = G_t(R_{y,t} \mid \Phi_{t-1})$. These values are used to estimate a time-varying copula dependence parameter $\rho_t$ that is a conditional quantity determined by $\Phi_{t-1}$ and the parameter vector $\theta_c$. We first describe the conditional copula density function $c_t(u_t, v_t \mid \rho_t)$, that only depends on the single parameter $\rho_t$.

The cited textbooks describe a variety of copula density functions that have different mathematical properties. The selection of a particular copula is an important step in empirical applications. Malevergne and Sornette (2003) demonstrate that returns from most pairs of major stock indices are compatible with the Gaussian copula. Accordingly, we employ the conditional Gaussian copula.

The conditional Gaussian copula density function is the density of $(u_t, v_t)$ when a pair of variables $(x_t, y_t)$ have a bivariate Gaussian distribution with correlation $\rho_t$ between $x_t$ and $y_t$. With $\psi(.)$ representing the c.d.f. of the standard normal distribution, $a_t = \psi^{-1}(u_t)$ and $b_t = \psi^{-1}(v_t)$, the copula density can be written as:

$$c_t\,(u_t, v_t \mid \rho_t) = \frac{1}{\sqrt{1-\rho_t^2}} \exp\left\{ -\frac{1}{2(1-\rho_t^2)}[a_t^2 + b_t^2 - 2\rho_t a_t b_t] + \frac{1}{2}[a_t^2 + b_t^2] \right\}. \tag{7}$$

### 4.3 The Specification for the Dependence Parameter

Conditional copulas typically contain a time-varying dependence parameter, such as $\rho_t$ in the equation above. A few studies have already investigated how to model this time-varying process, including Rodriguez (2003), Jondeau and Rockinger (2005) and Patton (2005). Based on the observation that high correlation is associated with high volatility, Rodriguez (2003) uses a mixed copula. He lets the weights and the marginal distributions follow two-state switching processes. Jondeau and Rockinger (2005) assume that the dependence is either a function of conditional on its historical values or a deterministic function of time. Patton (2005) proposes that the current dependence is explained by the previous dependence and the historical average difference of cumulative probabilities for the two assets. A common issue in these studies is an arbitrary choice of the number of regimes or lagged periods.

We follow Patton (2005) and suppose that $\rho_t$ depends on the previous dependence $\rho_{t-1}$, to capture persistence, and historical absolute differences, $| u_{t-i} - v_{t-i} |, i > 0$, to capture variation in the dependence process. We differ by using all historical information about the absolute differences, rather than arbitrarily truncating the historical information. Also, instead of a logistic transformation function, we use a constraint in the estimation procedure to keep the dependence process within $\pm 1$. The use of a logistic transformation function would unhelpfully restrict the volatility of the dependence term when it is near its limiting values.

We estimate the following dependence process:

$$\rho_t = \kappa + \beta_1 \rho_{t-1} + \delta b_t , \tag{8}$$

with

$$b_t = \beta_2 b_{t-1} + (1 - \beta_2) \left| u_{t-1} - v_{t-1} \right| , \tag{9}$$

being the exponentially weighted moving average of all historical absolute differences, that gives higher weight to more recent observations. The intuition for the use of $| u_{t-1} - v_{t-1} |$ is that the smaller (larger) the difference between the realized cumulative probabilities, the higher (lower) is the dependence. Therefore, we expect $\delta$ to be negative, $\beta_1$ to be positive, and $\beta_2$ to be within [0, 1].

The two equations above can be simplified to one, by using the lag operator $L$ to obtain:

$$(1-\beta_1 L)(1-\beta_2 L)\rho_t = (1-\beta_2)(\kappa + \delta |u_{t-1} - v_{t-1}|) .$$

We estimate this as

$$(1-\beta_1 L)(1-\beta_2 L)\rho_t = \omega + \gamma |u_{t-1} - v_{t-1}| . \tag{10}$$

Extra assumptions are required to call this an AR(2) model, namely that a linear function of the previous absolute difference, $| u_{t-1} - v_{t-1} |$, provides an i.i.d. innovation term.

The copula parameter vector is $\theta_c = (\beta_1, \beta_2, \omega, \gamma)'$, which is estimated in the second stage by maximizing the sum of terms $\log c_t(u_t, v_t | \Phi_{t-1}, \theta_c)$. It is not possible to uniquely identify the two $\beta$-estimates with the two $\beta$-parameters in equations (8) and (9). Consequently, we simply apply the constraints $0 \le \beta_2 \le \beta_1 \le 1$. Also, as equation (3) does not guarantee $|\rho_t| < 1$, we set the maximum and the minimum of $\rho_t$ in the estimation software as 0.9999 and –0.9999, respectively. However, the upper bound is rarely touched in the empirical implementation and the lower bound is never required.

## 5 Data and Summary Statistics

The empirical investigation is conducted for twelve euro-zone countries (France, Germany, Italy, the Netherlands, Spain, Finland, Belgium, Greece, Ireland, Portugal, Austria and Luxembourg) and five non-euro European countries (U.K., Switzerland, Sweden, Denmark and Norway). For each country, we obtain ten years of daily values of the stock market index from Datastream. The sample period is from January 1, 1994 to October 31, 2003 and excludes holidays. We also use a euro-zone stock market index from Datastream for the tests of the dependence between the euro-zone stock market and the equity market in the non-euro countries. All the indices are denominated in U.S. dollars, but we also study results for local currency returns in order to investigate the effect of different numeraires.

For every euro-zone country, we calculate a modified euro-zone stock market index by excluding the equities of that country from the euro-zone index. This is done in order to avoid mechanical relationships created by overlaps between the country indices and the euro-zone regional index. The definition of the modified euro-zone index $MPI_{i,t}$ for country *i* in period *t* is given by

$$MPI_{i,t} = MPI_{i,t-1} \frac{\sum_{j \neq i} MV_{j,t} \cdot PI_{j,t}}{\sum_{j \neq i} MV_{j,t} \cdot PI_{j,t-1}}$$

where *MV* is the market value of stocks in the country and *PI* is the country price index expressed in dollars.

There are three main reasons for using Datastream indices. Firstly, compared with other popular indices, they offer broader coverage of the markets in terms of market capitalization (at least 75%-80% for each market). Secondly, they are compiled according to the same criteria and thus are homogeneous for comparisons across markets. Moreover, the indices can be denominated in a common currency, i.e. they have the same numeraire, the impact of which on market dependence will be explored in this paper.

In order to avoid interpreting global trends as regional trends, we also investigate the time-varying dependence of European equity markets with a U.S. stock market index. As shown in Martens and Poon (2001), it is essential to have time-synchronized prices when studying equity market co-movements. Therefore, we use values of the S&P500 index at 16:00 London time recorded by Datastream to represent the U.S. stock market index.[1]

All the returns for the indices used in this study are calculated as:

$$R_{i,t} = 100\log(P_{i,t} / P_{i,t-1}) .$$

The summary statistics of these index returns are shown in Table 1. As suggested by previous research, most of the series of returns are negatively skewed, leptokurtic and do not have a high first-lag autocorrelation coefficient (independent of the currency denomination). Nevertheless, there are minor differences in skewness and kurtosis between the returns in U.S. dollars and in local currency, which may imply that the numeraire could matter in the analysis of inter-market dependence.

[1] The S&P500 is the only time-synchronized U.S. index available.

## 6 Empirical Results

### 6.1 The Euro-zone Equity Markets

Since the convergence of interest rates in the euro area has been an immediate consequence of the introduction of the euro, we are not going to document this phenomenon again. Instead, we investigate another possible consequence of the single currency, an increase in the co-movement between equity markets.

Table 2 shows the estimates of the copula dependence model for twelve euro-zone countries. The time-varying dependence model is estimated for each country index and the euro-zone stock market index excluding the examined country. For the purpose of comparison, we also include each country's dependence with the synchronized S&P500 index. All the indices are converted to the same numeraire, namely U.S. dollars. Across all countries and indices, $\beta_1$ is always larger than 0.9 and even as high as 0.99 in some cases, which indicates high dependence persistence. The other autoregressive parameter, $\beta_2$, is much smaller than $\beta_1$ and it is rarely significantly different from zero. As expected, the parameter $\gamma$ is always negative; it is also highly significant, indicating that the latest absolute difference of returns is consistently a relevant measure when modeling market dependence. Overall, the copula log-likelihood function of specifications with the euro-zone regional index is higher than that with S&P500 index.

Figure 1 shows the time-varying conditional dependence, $\rho_t$, for the parameter estimates listed in Table 2. Overall, the level of dependence within the euro-zone market is higher than the association of the euro national markets with the U.S. market. The dependence of the indices of France, Germany, Italy, the Netherlands and Spain with the euro-zone regional index exhibits an increase during our sample period, while the dependence for Finland, Belgium, Greece and Portugal does not display a regime shift, and that for Ireland, Austria and Luxembourg has actually decreased. Interestingly, some countries, especially Finland, have experienced a higher integration with the U.S. market. As shown in Figure 2, among the countries that show a rise in their dependence with the euro-zone regional market, the differences between their dependence with the euro-zone regional index and their dependence with the S&P500 index exhibit a regime shift around the middle of the sample period in France, Italy and Spain.

To test whether there are regime changes in the process generating the conditional correlations, that are statistically significant, and to determine the timing of any such regime shifts, we evaluate five ways to add one dummy variable into the conditional dependence process (equation (10)). The dependence equation now becomes:

$$(1-\beta_1 L)(1-\beta_2 L)\rho_t = \omega + \gamma |u_{t-1} - v_{t-1}|) + \lambda D_{j,t}. \tag{11}$$

Specifically, the dummy variables $D_{1,t}, D_{2,t}, D_{3,t}, D_{4,t}, D_{5,t}$ are equal to 0 before the first day of 1996, 1997, 1998, 1999 and 2000, respectively, otherwise they are equal to 1. *T*-tests and likelihood-ratio tests are employed to assess the significance of these dummy variables. For the sake of completeness and comparison, we include the remaining countries that do not exhibit an obvious dependence change in Figure 1 in these tests as well.

The results are shown in Table 3. All of the countries with an obvious dependence change in Figure 1 show a statistically significant increase in their dependence with the euro-zone regional index. For France, Germany and Spain, the most likely timing for this increase is around January 1998, because the models that contain $D_{3,t}$ have the highest increases in the copula likelihood function; these increases are all significant at the 1% level, both for t and likelihood-ratio tests. Although the highest values of the likelihood function for Italy and the Netherlands are obtained by models that contain $D_{2,t}$ (1997) and $D_{5,t}$ (2000), respectively, the difference of these likelihood function values from those of models with $D_{3,t}$ (1998) are small, 0.8 and 2.1 respectively. For the remaining countries, in contrast, there is no significant dependence increase around 1998. The Austrian stock market index even shows a significant decrease at the 5% level in 1998.

Therefore, it is reasonable to conclude that France, Germany, Italy, the Netherlands and Spain have experienced an increase in their dependence with the equity markets of other euro-zone countries, which started in late 1997 or early 1998 when the membership of the EMU was determined and the relevant information was announced. The incremental impact of the dummy variables on the unconditional dependence, $E[\rho_t | D_{j,t} = 1] - E[\rho_t | D_{j,t} = 0]$, also confirms this conclusion.

To verify that this phenomenon is unique for the euro area, we also implement tests that include the same dummy variables in the dependence models for all euro-zone stock market indices with the S&P500 index. The results, shown in Table 4, indicate that although the dependence for some indices increases during our sample period, the timing is not consistent across countries and does not match the timing of the introduction of the euro. For example, the most likely years for the increased dependence of the S&P500 index with the stock indices of France, Germany and Italy are 2000, 1996 and 1997, respectively, while there is no significant change for Spain. The Finnish stock market index shows a highly significant increase in its dependence with the S&P500 index during the second half of the 1990s, which may be due to the fact that communication companies dominate the Finnish market capitalization and this industry is strongly linked to the U.S. market.

These results largely confirm the hypothesis that only some euro-zone countries, France, Germany, Italy, the Netherlands and Spain, experienced a rise in their dependence with the other euro-zone countries. Although some of these countries also exhibit an increasing co-movement with the U.S. market, for most countries the relative degree of the increase is higher for the dependence with the other euro-area countries.

Nonetheless, according to Figure 1, there is no obvious evidence supporting the integration of the remaining euro-zone countries. We believe that other significant barriers still play a crucial role for further market integration of smaller markets. As documented earlier (Section 2), significant differences in transactions costs remain after the introduction of the euro even across euro area equity markets. The correlation coefficient between market capitalization and total transaction costs (market impact) is about –0.64 (–0.49) for the period 1998-99, which, in line with our findings, indicates that transaction costs and market liquidity are likely to remain the main concern of institutional investors regarding investment in smaller euro area markets. To this end, we have estimated a logit model where the left hand side variable indicates whether the stock market in a country shows a significant increase in dependence with the euro area market. After controlling for other country effects such as GDP per capita, legal environment and euro membership, variables proxying for transactions cost, especially market impact, show a significant negative relationship to the likelihood of market integration. Consequently, country factors may still determine the degree of regional integration (Guiso et al., 2003), as institutional investors focus on large European equity markets with low transactions cost and high liquidity.

### 6.2 Non-euro European Equity Markets

In order to investigate whether equity market dynamics say anything about beliefs that non-euro European countries will adopt the euro, we model the time-varying conditional dependence between the equity indices of these countries and the euro-zone regional stock market index. For comparison, we provide estimates for these national indices with the S&P500 index as well. All indices are denominated in U.S. dollars. As shown in Table 5, the basic properties of the estimated parameters are the same as in Table 2. Figure 3 displays the dependence processes with the euro-zone regional index and with the S&P500 index. Although there is no obvious regime change compared to euro countries, it appears that the U.K. and Sweden also experienced a slight increase in their dependence with the euro-zone market, while there is no structural change in co-movement with the U.S. market. On the other hand, Switzerland, Denmark and Norway do not exhibit a clear regime shift, neither with the euro-zone market nor with the U.S. market. Furthermore, as shown in Figure 4, the difference between the dependence with the euro-zone regional market and with the U.S. market increased in the early sample period for Sweden and in the late sample period for the U.K., respectively.

In order to test whether the changes in dependence with the euro area for the U.K. and Sweden are statistically significant and to detect the timing of these changes, we add alternative dummy variables into the dependence process as before. We also include the markets that do not exhibit an obvious dependence change in Figure 3 in this test for completeness. The results are shown in Table 6. For the U.K., the dummy variable becomes significant at the 5% level in 1999 and has the highest likelihood value in 2000, while the dummy variable for Sweden has the highest likelihood in 1996 and is significant until 1997. However, there is no significant dependence increase for the remaining non-euro countries.

For comparison, we also run the same tests for the dependence with the S&P500 index. The results are shown in Table 7. An increased dependence for both the Swedish and U.K. index is found in the early sample period. The most likely timing of the structural break is 1996, which does not match the introduction of the euro and might rather be the result of the high-tech boom or the emergent globalization of financial markets during the 1990s. Similar results are also found in Table 4 for many euro-zone countries.

The U.K. and Sweden are potential candidates for introducing the euro. Nevertheless, while we find increased dependence of their stock market indices with the euro-zone stock market index, the evidence is not sufficiently strong and thus the future development of the dependence in all financial markets still needs to be studied further before firm conclusions can be drawn. We leave these issues for future research. At present, what we can suggest is that the co-movement of the British and Swedish stock markets with the euro-zone equity market has increased in the second half of the 1990s even though they are not part of the currency union, which may reflect the expectations of market participants' about the adoption of the euro in these countries in the future.

### 6.3 Robustness Tests

Theoretically, if markets are fully integrated, investors and corporations are indifferent to the geographical factor. When investigating the consequence of market integration by looking at market dependence, the perspective of the same investor is adopted. Therefore, all of the indices used in the empirical tests above are denominated in U.S. dollars. In order to investigate the sensitivity of the results to different base currencies, we discuss the influence of the numeraire by changing the currency of reference. To this end, we first repeat the estimations by using the euro (EUR) as the common measure to assess the dependence between the euro-zone regional index and the five major euro-zone national stock market indices that exhibit an increased dependence with the euro-zone index, and compare the fitted dependence processes with the results using the indices in U.S. dollars. As shown in Figure 5, there is little difference between these two dependence processes, since the average level, the patterns and the development over time of the correlations are very similar, which may imply that the choice of numeraire does not matter as long as the same currency is chosen for a pair of markets.

Next, we repeat the estimations by using the individual local currency for the two non-euro equity indices that exhibit an increased dependence with the euro-zone stock index, i.e. the U.K. and Sweden, but keeping the euro-zone stock index in U.S. dollars. In Figure 6, we compare the fitted dependence processes for these national indices in their local currencies and in dollars. The gap between these two processes becomes larger than that using the same currency for the examined pair of indices and the magnitude varies across countries. We suggest that this result is due to the different local currencies and the gap size may depend on the development of the exchange rate. However, for the purposes of this study, the numeraire has no effect on the conclusions.

Another potential concern is that we use price indices in our empirical implementations, rather than total return indices, and thus neglect the effect of dividends. Nonetheless, we observe that the time series of daily dividends for indices are smoothed and will not have a significant impact on our results. To validate this point, we compare the estimates of the dependence with the euro-zone stock market of returns calculated from alternatively the price indices and return indices of five non-euro stock markets. We find that for all pairs of markets, the values of marginal and copula likelihood functions are almost unchanged when we use return indices instead of price indices. All of the differences in the log-likelihood are smaller than 1. In addition, the estimated dependence processes from price indices and return indices almost overlap for all pairs of markets.

## 7 Concluding Remarks

In this paper, we propose a general time-varying copula dependence model in order to study market linkages. Subsequently, we use this model to investigate the impact of the introduction of the euro on the integration of equity markets in Europe. In particular, we investigate whether there are significant changes in the time-varying dependence structure of markets within the euro area as well as between equity markets of countries in the euro area and non-euro European countries. We find that market dependence within the euro area increased only for some countries, like France, Germany, Italy, the Netherlands and Spain, which are characterized by relatively large equity market capitalization, comprehensive regulations, high liquidity, and low transaction and information costs. When testing for alternative structural breaks in market dependence, we find that the increase in dependence started in late 1997 or early 1998 when euro membership was determined and announced. The results suggest that the introduction of the euro increased financial market integration in the euro area, but did not lead to fully integrated markets.

In contrast, most of the remaining European countries continue to lack significant integration into the euro area. Nevertheless, we do find that the dependence of the British and Swedish stock markets with the euro-zone market slightly increased. This may indicate that at least some market participants actually expected the adoption of the euro in these countries. However, we suggest further research on the development of non-euro financial markets since the existing evidence is not of sufficient strength to draw firm conclusions. Our approach can be extended to a multivariate model, which is useful for portfolio and risk management. Future research may apply this model to study changes in the dependence of other asset markets in order to provide a broader basis for conjectures about whether and when these countries may join the euro.

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# Table 1: Summary Statistics

The table shows summary statistics of the returns of the euro-zone stock market index, S&P500 index, 12 euro-zone country stock market indices and 5 non-euro European country stock market indices. All of the indices are denominated in alternatively USD or local currency. The sample period covers January 1, 1994 to October 31, 2003 and has 2319 daily observations excluding holidays. Markets are sorted by region and decreasing market capitalization.

| | Index | Currency | Mean | Std. Dev. | Skewness | Kurtosis | AR(1)* | AR(2)* |
|---|---|---|---|---|---|---|---|---|
| Euro area | Euro-zone | USD | 0.000255 | 0.0112 | -0.0913 | 5.3113 | 0.0934 | -0.0401 |
| | | EUR | 0.000249 | 0.0121 | -0.1699 | 5.2681 | 0.0501 | -0.0177 |
| | France | USD | 0.000254 | 0.0127 | -0.0246 | 4.9854 | 0.0732 | -0.0463 |
| | | EUR | 0.000234 | 0.0135 | -0.0629 | 5.1835 | 0.0490 | -0.0290 |
| | Germany | USD | 0.000157 | 0.0128 | -0.1267 | 4.9708 | 0.0620 | -0.0175 |
| | | EUR | 0.000144 | 0.0132 | -0.3158 | 5.2785 | 0.0556 | -0.0132 |
| | Italy | USD | 0.000263 | 0.0144 | -0.0533 | 4.8467 | 0.0612 | -0.0010 |
| | | EUR | 0.000252 | 0.0144 | -0.1479 | 4.9206 | 0.0370 | 0.0323 |
| | Netherlands | USD | 0.000247 | 0.0124 | -0.1251 | 8.1302 | 0.0450 | -0.0429 |
| | | EUR | 0.000237 | 0.0132 | -0.1693 | 8.1485 | 0.0227 | -0.0248 |
| | Spain | USD | 0.000347 | 0.0130 | -0.0849 | 5.0074 | 0.0772 | -0.0459 |
| | | EUR | 0.000348 | 0.0132 | -0.2301 | 5.1015 | 0.0339 | -0.0309 |
| | Finland | USD | 0.000698 | 0.0224 | -0.3690 | 9.0046 | 0.0361 | -0.0130 |
| | | EUR | 0.000645 | 0.0232 | -0.3545 | 8.8999 | 0.0241 | -0.0048 |
| | Belgium | USD | 0.000220 | 0.0108 | 0.1719 | 6.3110 | 0.1662 | -0.0079 |
| | | EUR | 0.000202 | 0.0104 | 0.2110 | 7.7974 | 0.1781 | 0.0018 |
| | Greece | USD | 0.000356 | 0.0183 | -0.0873 | 8.3628 | 0.1151 | -0.0036 |
| | | EUR | 0.000411 | 0.0174 | -0.1109 | 9.8499 | 0.1309 | -0.0012 |
| | Ireland | USD | 0.000407 | 0.0114 | -0.3315 | 6.8231 | 0.1117 | 0.0022 |
| | | EUR | 0.000387 | 0.0113 | -0.5823 | 8.7496 | 0.1124 | 0.0210 |
| | Portugal | USD | 0.000222 | 0.0109 | -0.0702 | 6.3178 | 0.1450 | 0.0164 |
| | | EUR | 0.000212 | 0.0102 | -0.5372 | 9.6835 | 0.1359 | 0.0158 |
| | Austria | USD | 0.000094 | 0.0093 | -0.1968 | 4.6855 | 0.0722 | 0.0187 |
| | | EUR | 0.000081 | 0.0080 | -0.7150 | 8.3091 | 0.0682 | 0.0068 |
| | Luxembourg | USD | 0.000181 | 0.0121 | -0.0706 | 10.2988 | 0.0755 | 0.0322 |
| | | EUR | 0.000164 | 0.0110 | -0.1806 | 15.3306 | 0.1260 | 0.0763 |
| Non-Euro Europe | UK | USD | 0.000180 | 0.0105 | -0.0557 | 5.3958 | 0.0328 | -0.0440 |
| | | GBP | 0.000120 | 0.0108 | -0.1406 | 5.5838 | 0.0217 | -0.0365 |
| | Switzerland | USD | 0.000303 | 0.0112 | -0.0850 | 5.8229 | 0.0837 | 0.0014 |
| | | SWF | 0.000257 | 0.0117 | -0.2473 | 6.4344 | 0.0684 | 0.0185 |
| | Sweden | USD | 0.000403 | 0.0165 | -0.0763 | 5.8585 | 0.0995 | -0.0267 |
| | | SEK | 0.000374 | 0.0158 | 0.0340 | 5.8404 | 0.0572 | -0.0062 |
| | Denmark | USD | 0.000417 | 0.0112 | -0.1149 | 8.2294 | 0.0370 | 0.0033 |
| | | DMK | 0.000390 | 0.0107 | -0.3742 | 11.1635 | 0.0651 | 0.0098 |
| | Norway | USD | 0.000250 | 0.0126 | -0.4716 | 7.1171 | 0.0651 | 0.0263 |
| | | NOK | 0.000224 | 0.0120 | -0.4373 | 7.1770 | 0.0565 | 0.0320 |
| United States | SP500 | USD | 0.000348 | 0.0117 | -0.1184 | 5.5530 | -0.0318 | -0.0219 |

*AR($i$) represents the $i^{th}$-lag autocorrelation coefficient of returns.

# Table 2: Estimates of Dependence Models for Euro-zone Stock Market Indices

The table shows estimates of the dependence of 12 euro-zone country stock market indices with the euro-zone stock market index and with the S&P500 index, using the following model settings. All indices are denominated in USD. Markets are sorted by region and decreasing market capitalization.

$f(x,y)=c(u,v)f(x)f(y)$ where *c(u,v)* is the Gaussian copula function defined as

$c(u,v)=\frac{1}{\sqrt{1-\rho_t^2}}e^{\{-\frac{1}{2(1-\rho_t^2)}[a^2+b^2-2\rho_t ab]+\frac{1}{2}[a^2+b^2]\}}$ with $a=\Phi^{-1}(u)$, $b=\Phi^{-1}(v)$ and $(1-\beta_1 L)(1-\beta_2 L)\rho_t=\omega+\gamma\,|u_{t-1}-v_{t-1}|$

*f(x)* and *f(y)* are modeled by the GJR-GARCH model with student *t* distribution

| Country | with | $\omega$ | $\beta_1$ | $\beta_2$ | $\gamma$ | LLF($c$) |
|---|---|---|---|---|---|---|
| France | Euro | 0.0242 | 0.9773 | 0.0000 | -0.0417 | 1458.44 |
| | | (0.0000) | (0.0000) | (0.9999) | (0.0000) | |
| | SP500 | 0.0337 | 0.9629 | 0.1759 | -0.0771 | 474.32 |
| | | (0.0001) | (0.0000) | (0.1679) | (0.0000) | |
| Germany | Euro | 0.0910 | 0.9122 | 0.0000 | -0.1463 | 1226.04 |
| | | (0.0000) | (0.0000) | (0.9999) | (0.0000) | |
| | SP500 | 0.0791 | 0.9208 | 0.0000 | -0.1741 | 321.06 |
| | | (0.0046) | (0.0000) | (0.9999) | (0.0049) | |
| Italy | Euro | 0.0640 | 0.9426 | 0.0000 | -0.1406 | 917.27 |
| | | (0.0000) | (0.0000) | (0.9999) | (0.0000) | |
| | SP500 | 0.0381 | 0.9662 | 0.0484 | -0.0988 | 274.42 |
| | | (0.0094) | (0.0000) | (0.8788) | (0.0087) | |
| Netherlands | Euro | 0.0307 | 0.9707 | 0.0112 | -0.0498 | 1439.15 |
| | | (0.0000) | (0.0000) | (0.8175) | (0.0000) | |
| | SP500 | 0.0281 | 0.9633 | 0.3186 | -0.0670 | 471.09 |
| | | (0.0000) | (0.0000) | (0.0007) | (0.0000) | |
| Spain | Euro | 0.0546 | 0.9491 | 0.0000 | -0.1015 | 1061.32 |
| | | (0.0000) | (0.0000) | (0.9999) | (0.0000) | |
| | SP500 | 0.0362 | 0.9382 | 0.3913 | -0.0747 | 322.71 |
| | | (0.0000) | (0.0000) | (0.0001) | (0.0000) | |
| Finland | Euro | 0.0519 | 0.9471 | 0.0563 | -0.1002 | 645.71 |
| | | (0.0001) | (0.0000) | (0.7900) | (0.0000) | |
| | SP500 | 0.0443 | 0.9379 | 0.3284 | -0.0969 | 431.47 |
| | | (0.0000) | (0.0000) | (0.0103) | (0.0000) | |
| Belgium | Euro | 0.0658 | 0.9258 | 0.1632 | -0.1238 | 840.79 |
| | | (0.0000) | (0.0000) | (0.1067) | (0.0000) | |
| | SP500 | 0.0531 | 0.9522 | 0.0000 | -0.1453 | 137.46 |
| | | (0.0357) | (0.0000) | (0.9999) | (0.0377) | |
| Greece | Euro | 0.0950 | 0.9013 | 0.0000 | -0.2204 | 160.07 |
| | | (0.0091) | (0.0000) | (0.9999) | (0.0100) | |
| | SP500 | 0.0653 | 0.9337 | 0.0000 | -0.1986 | 22.98 |
| | | (0.0887) | (0.0000) | (0.9999) | (0.0903) | |
| Ireland | Euro | 0.0085 | 0.9909 | 0.2420 | -0.0213 | 383.08 |
| | | (0.0096) | (0.0000) | (0.0005) | (0.0106) | |
| | SP500 | 0.0233 | 0.9570 | 0.4889 | -0.0596 | 138.34 |
| | | (0.0021) | (0.0000) | (0.0020) | (0.0021) | |
| Portugal | Euro | 0.0450 | 0.9623 | 0.0000 | -0.1078 | 488.04 |
| | | (0.0010) | (0.0000) | (0.9999) | (0.0008) | |
| | SP500 | 0.0205 | 0.9880 | 0.0000 | -0.0668 | 74.98 |
| | | (0.1216) | (0.0000) | (0.9999) | (0.1209) | |
| Austria | Euro | 0.0221 | 0.9756 | 0.2686 | -0.0574 | 407.46 |
| | | (0.0001) | (0.0000) | (0.0079) | (0.0001) | |
| | SP500 | 0.0095 | 0.9953 | 0.0513 | -0.0310 | 31.44 |
| | | (0.1653) | (0.0000) | (0.9193) | (0.1579) | |
| Luxembourg | Euro | 0.0264 | 0.9803 | 0.0015 | -0.0752 | 138.81 |
| | | (0.0003) | (0.0000) | (0.9935) | (0.0004) | |
| | SP500 | 0.0302 | 0.7421 | 0.0481 | -0.0545 | 3.47 |
| | | (0.0566) | (0.0000) | (0.0000) | (0.0544) | |

The numbers in brackets ( ) are P values and 0.0000 means that the value is less than 0.00005.
LLF(c) is the maximum of the copula component of the log-likelihood function.

# Table 3: Tests of Dependence Change between Euro-zone National Stock Market Indices and Euro-zone Stock Market Index

The table shows estimates of the dependence of 12 euro-zone country stock indices with the euro-zone regional stock index. All indices are denominated in USD. The model settings, except the process of dependence variable defined as below, are the same as those in Table 2. Markets are sorted by region and decreasing market capitalization.

$(1-\beta_1 L)(1-\beta_2 L)\rho_t = \omega + \gamma \mid u_{t-1} - v_{t-1} \mid + \lambda Dummy_t$

$Dummy_t = D_1 = 1$ when $t >= 1/1/1996$, otherwise $D_1 = 0$. $Dummy_t = D_2 = 1$ when $t >= 1/1/1997$, otherwise $D_2 = 0$.
$Dummy_t = D_3 = 1$ when $t >= 1/1/1998$, otherwise $D_3 = 0$. $Dummy_t = D_4 = 1$ when $t >= 1/1/1999$, otherwise $D_4 = 0$.
$Dummy_t = D_5 = 1$ when $t >= 1/1/2000$, otherwise $D_5 = 0$.
*f(x)* and *f(y)* are modeled by the GJR-GARCH model with student t distribution

| Country | *Dummy* | $D_1$ | $D_2$ | $D_3$ | $D_4$ | $D_5$ |
|---|---|---|---|---|---|---|
| France | $\lambda$ | 0.0020 | 0.0021 | *0.0025* | 0.0018 | 0.0020 |
| | | (0.0026) | (0.0007) | *(0.0001)* | (0.0014) | (0.0014) |
| | $\Delta E(\rho)$ | 0.0728 | 0.0749 | *0.0873* | 0.0625 | 0.0694 |
| | $\Delta LLF(c)$ | 5.44*** | 7.43*** | *11.71**** | 7.03*** | 7.93*** |
| Germany | $\lambda$ | 0.0017 | 0.0055 | *0.0167* | 0.0108 | 0.0072 |
| | | (0.3059) | (0.0194) | *(0.0007)* | (0.0039) | (0.0080) |
| | $\Delta E(\rho)$ | 0.0181 | 0.0469 | *0.0914* | 0.0651 | 0.0484 |
| | $\Delta LLF(c)$ | 0.50 | 3.36*** | *11.48**** | 6.91*** | 5.62*** |
| Italy | $\lambda$ | 0.0129 | *0.0164* | *0.0141* | 0.0059 | 0.0046 |
| | | (0.0003) | *(0.0003)* | *(0.0004)* | (0.0089) | (0.0145) |
| | $\Delta E(\rho)$ | 0.2179 | *0.2582* | *0.2265* | 0.0772 | 0.0608 |
| | $\Delta LLF(c)$ | 13.59*** | *24.12**** | *23.32**** | 5.06*** | 4.53*** |
| Netherlands | $\lambda$ | -0.0000 | 0.0004 | *0.0009* | 0.0007 | *0.0011* |
| | | (0.9828) | (0.3313) | *(0.0612)* | (0.0782) | *(0.0179)* |
| | $\Delta E(\rho)$ | -0.0003 | 0.0136 | *0.0293* | 0.0223 | *0.0330* |
| | $\Delta LLF(c)$ | 0.02 | 0.60 | *2.50*** | 2.35** | *4.60**** |
| Spain | $\lambda$ | 0.0047 | 0.0062 | *0.0095* | 0.0066 | 0.0096 |
| | | (0.0041) | (0.0017) | *(0.0027)* | (0.0030) | (0.0032) |
| | $\Delta E(\rho)$ | 0.0671 | 0.0854 | *0.0907* | 0.0626 | 0.0786 |
| | $\Delta LLF(c)$ | 4.74*** | 9.54*** | *11.88**** | 7.98*** | 10.88*** |
| Finland | $\lambda$ | 0.0022 | 0.0022 | 0.0000 | -0.0006 | -0.0004 |
| | | (0.2642) | (0.2412) | (0.9623) | (0.4610) | (0.6311) |
| | $\Delta E(\rho)$ | 0.0416 | 0.0376 | 0.0009 | -0.0123 | -0.0080 |
| | $\Delta LLF(c)$ | 0.87 | 1.04 | 0.02 | 0.29 | 0.13 |
| Belgium | $\lambda$ | -0.0006 | -0.0001 | 0.0011 | 0.0009 | *0.0018* |
| | | (0.5409) | (0.9452) | (0.1490) | (0.2201) | *(0.0312)* |
| | $\Delta E(\rho)$ | -0.0078 | -0.0008 | 0.0148 | 0.0127 | *0.0247* |
| | $\Delta LLF(c)$ | 0.35 | 0.24 | 0.85 | 0.73 | *2.04*** |
| Greece | $\lambda$ | 0.0001 | 0.0011 | 0.0057 | 0.0016 | 0.0040 |
| | | (0.9773) | (0.7791) | (0.1996) | (0.6282) | (0.2813) |
| | $\Delta E(\rho)$ | 0.0012 | 0.0114 | 0.0523 | 0.0162 | 0.0392 |
| | $\Delta LLF(c)$ | 0.00 | 0.03 | 0.93 | 0.10 | 0.59 |
| Ireland | $\lambda$ | -0.0001 | -0.0007 | -0.0003 | -0.0004 | 0.0001 |
| | | (0.8344) | (0.4342) | (0.4420) | (0.3703) | (0.7265) |
| | $\Delta E(\rho)$ | -0.0140 | -0.0519 | -0.0346 | -0.0379 | 0.0112 |
| | $\Delta LLF(c)$ | 0.50 | 0.74 | 0.53 | 0.54 | 0.03 |
| Portugal | $\lambda$ | 0.0001 | 0.0009 | 0.0006 | 0.0001 | 0.0004 |
| | | (0.9270) | (0.4503) | (0.5345) | (0.9372) | (0.6109) |
| | $\Delta E(\rho)$ | 0.0026 | 0.0224 | 0.0165 | 0.0016 | 0.0110 |
| | $\Delta LLF(c)$ | 0.00 | 0.29 | 0.21 | 0.00 | 0.14 |
| Austria | $\lambda$ | -0.0000 | -0.0006 | *-0.0012* | -0.0019 | -0.0019 |
| | | (0.8381) | (0.2533) | *(0.0406)* | (0.0919) | (0.1232) |
| | $\Delta E(\rho)$ | -0.0056 | -0.0319 | *-0.0690* | -0.0923 | -0.0858 |
| | $\Delta LLF(c)$ | 0.12 | 0.79 | *3.47**** | 3.41*** | 2.65** |
| Luxembourg | $\lambda$ | *-0.0039* | -0.0013 | -0.0008 | -0.0004 | -0.0006 |
| | | *(0.0616)* | (0.4434) | (0.4211) | (0.5675) | (0.4454) |
| | $\Delta E(\rho)$ | *-0.1910* | -0.0668 | -0.0497 | -0.0278 | -0.0398 |
| | $\Delta LLF(c)$ | *5.30**** | 0.91 | 0.76 | 0.38 | 0.61 |

The numbers in brackets ( ) are P values and 0.0000 means that the value is less than 0.00005.
$\Delta E(\rho)$: $E(\rho_t \mid D_t=1) - E(\rho_t \mid D_t=0)$. $\Delta LLF(c)$: Copula LLF(with $D_t$) - Copula LLF(without $D_t$).
** : Significance at 5% level for the likelihood ratio test. ***: Significance at 1% level for the likelihood ratio test.

# Table 4: Tests of Dependence Change between Euro-zone National Stock Market Indices and S&P500 Index

The table shows estimates of the dependence of 12 euro-zone country stock indices with S&P500 index. All indices are denominated in USD. The model settings, except the process of dependence variable defined as below, are the same as those in Table 2. Markets are sorted by region and decreasing market capitalization.

$$(1-\beta_1 L)(1-\beta_2 L)\rho_t = \omega + \gamma \,|\, u_{t-1} - v_{t-1} \,|\, + \lambda Dummy_t$$

$Dummy_t = D_1 = 1$ when $t$ >= 1/1/1996, otherwise $D_1 = 0$. $Dummy_t = D_2 = 1$ when $t$ >= 1/1/1997, otherwise $D_2 = 0$.
$Dummy_t = D_3 = 1$ when $t$ >= 1/1/1998, otherwise $D_3 = 0$. $Dummy_t = D_4 = 1$ when $t$ >= 1/1/1999, otherwise $D_4 = 0$.
$Dummy_t = D_5 = 1$ when $t$ >= 1/1/2000, otherwise $D_5 = 0$.
*f(x)* and *f(y)* are modeled by the GJR-GARCH model with student t distribution

| Country | *Dummy* | $D_1$ | $D_2$ | $D_3$ | $D_4$ | $D_5$ |
|---|---|---|---|---|---|---|
| France | $\lambda$ | 0.0017 | 0.0011 | 0.0012 | 0.0012 | *0.0022* |
| | | (0.1278) | (0.2726) | (0.2197) | (0.1429) | *(0.0600)* |
| | $\Delta$E($\rho$) | 0.0611 | 0.0358 | 0.0361 | 0.0359 | *0.0593* |
| | $\Delta$LLF($c$) | 1.27 | 0.68 | 0.89 | 1.13 | *2.82*$^{**}$ |
| Germany | $\lambda$ | *0.0154* | 0.0129 | 0.0112 | 0.0068 | 0.0049 |
| | | *(0.0686)* | (0.0820) | (0.0970) | (0.1454) | (0.1535) |
| | $\Delta$E($\rho$) | *0.1710* | 0.1376 | 0.1114 | 0.0699 | 0.0557 |
| | $\Delta$LLF($c$) | *5.57*$^{***}$ | 5.25$^{***}$ | 4.53$^{***}$ | 2.18$^{**}$ | 1.68 |
| Italy | $\lambda$ | 0.0041 | *0.0151* | 0.0072 | 0.0022 | 0.0035 |
| | | (0.0147) | *(0.0895)* | (0.0865) | (0.1541) | (0.1047) |
| | $\Delta$E($\rho$) | 0.1665 | *0.2012* | 0.1174 | 0.0493 | 0.0748 |
| | $\Delta$LLF($c$) | 5.20$^{***}$ | *9.53*$^{***}$ | 4.51$^{***}$ | 1.43 | 2.76$^{**}$ |
| Netherlands | $\lambda$ | *0.0025* | 0.0015 | 0.0007 | 0.0011 | 0.0013 |
| | | *(0.0057)* | (0.0614) | (0.2311) | (0.0525) | (0.0381) |
| | $\Delta$E($\rho$) | *0.1266* | 0.0624 | 0.0300 | 0.0464 | 0.0542 |
| | $\Delta$LLF($c$) | *4.68*$^{***}$ | 1.94$^{**}$ | 0.70 | 1.90 | 2.43$^{**}$ |
| Spain | $\lambda$ | 0.0011 | 0.0012 | 0.0009 | 0.0008 | 0.0019 |
| | | (0.1277) | (0.3182) | (0.3782) | (0.3649) | (0.0548) |
| | $\Delta$E($\rho$) | 0.1276 | 0.0332 | 0.0241 | 0.0219 | 0.0507 |
| | $\Delta$LLF($c$) | 1.59 | 0.47 | 0.34 | 0.31 | 1.62 |
| Finland | $\lambda$ | 0.0080 | *0.0114* | 0.0116 | 0.0063 | 0.0063 |
| | | (0.0088) | *(0.0093)* | (0.0636) | (0.0931) | (0.1025) |
| | $\Delta$E($\rho$) | 0.1719 | *0.1819* | 0.1204 | 0.0755 | 0.0803 |
| | $\Delta$LLF($c$) | 5.65$^{***}$ | *9.14*$^{***}$ | 5.66$^{***}$ | 2.76$^{**}$ | 3.51$^{***}$ |
| Belgium | $\lambda$ | *0.0076* | 0.0041 | 0.0045 | 0.0030 | 0.0032 |
| | | *(0.1115)* | (0.1651) | (0.0939) | (0.1230) | (0.1252) |
| | $\Delta$E($\rho$) | *0.1497* | 0.0827 | 0.0879 | 0.0634 | 0.0671 |
| | $\Delta$LLF($c$) | *2.28*$^{**}$ | 0.87 | 1.75 | 0.78 | 0.90 |
| Greece | $\lambda$ | 0.0001 | *0.0417* | *0.0498* | 0.0153 | 0.0177 |
| | | (0.9793) | *(0.0360)* | *(0.0383)* | (0.0602) | (0.0673) |
| | $\Delta$E($\rho$) | 0.0015 | *0.4139* | *0.3538* | 0.1692 | 0.1815 |
| | $\Delta$LLF($c$) | 0.08 | *16.50*$^{***}$ | *16.44*$^{***}$ | 4.72$^{***}$ | 5.36$^{***}$ |
| Ireland | $\lambda$ | 0.0001 | -0.0000 | -0.0001 | -0.0018 | -0.0001 |
| | | (0.9250) | (0.9772) | (0.8928) | (0.1203) | (0.8468) |
| | $\Delta$E($\rho$) | 0.0046 | -0.0011 | -0.0046 | -0.0681 | -0.0063 |
| | $\Delta$LLF($c$) | 0.00 | 0.00 | 0.00 | 1.48 | 0.01 |
| Portugal | $\lambda$ | *0.0056* | *0.0080* | 0.0005 | 0.0002 | 0.0006 |
| | | *(0.0001)* | *(0.0043)* | (0.6272) | (0.6933) | (0.3840) |
| | $\Delta$E($\rho$) | *0.5090* | *0.4118* | 0.0371 | 0.0188 | 0.0452 |
| | $\Delta$LLF($c$) | *10.85*$^{***}$ | *10.55*$^{***}$ | 0.10 | 0.07 | 0.42 |
| Austria | $\lambda$ | *0.0008* | 0.0007 | -0.0253 | -0.0305 | -0.0202 |
| | | *(0.0008)* | (0.0053) | (0.3502) | (0.3791) | (0.3010) |
| | $\Delta$E($\rho$) | *0.2407* | 0.1675 | -0.2111 | -0.2217 | -0.2069 |
| | $\Delta$LLF($c$) | *3.31*$^{**}$ | 2.11$^{**}$ | 1.01 | 0.00 | 1.57 |
| Luxembourg | $\lambda$ | 0.0073 | 0.0089 | 0.0080 | 0.0052 | *0.0102* |
| | | (0.2728) | (0.1954) | (0.1775) | (0.2633) | *(0.0947)* |
| | $\Delta$E($\rho$) | 0.1090 | 0.1307 | 0.1210 | 0.0785 | *0.1510* |
| | $\Delta$LLF($c$) | 0.66 | 1.56 | 1.74 | 0.82 | *3.18*$^{**}$ |

The numbers in brackets ( ) are P values and 0.0000 means that the value is less than 0.00005.
$\Delta$E($\rho$): E($\rho_t$ | $D_t$=1) - E($\rho_t$ | $D_t$=0). $\Delta$LLF($c$): Copula LLF(with $D_t$) - Copula LLF(without $D_t$).
** : Significance at 5% level for the likelihood ratio test. ***: Significance at 1% level for the likelihood ratio test.

# Table 5: Estimates of Dependence Models for Non-euro European Stock Market Indices

The table shows estimates of the dependence of 5 non-euro country stock market indices with the euro-zone stock market index and with the S&P500 index, using the following model settings. All indices are denominated in USD. Markets are sorted by region and decreasing market capitalization.

$f(x,y) = c(u,v)f(x)f(y)$ where *c(u,v)* is the Gaussian copula function defined as

$c(u,v) = \frac{1}{\sqrt{1-\rho_t^2}} e^{\{-\frac{1}{2(1-\rho_t^2)}[a^2+b^2-2\rho_t ab]+\frac{1}{2}[a^2+b^2]\}}$ with $a = \Phi^{-1}(u)$, $b = \Phi^{-1}(v)$ and

$(1-\beta_1 L)(1-\beta_2 L)\rho_t = \omega + \gamma \mid u_{t-1} - v_{t-1} \mid$

*f(x)* and *f(y)* are modeled by the GJR-GARCH model with student *t* distribution

| Country | with | $\omega$ | $\beta_1$ | $\beta_2$ | $\gamma$ | LLF(*c*) |
|---|---|---|---|---|---|---|
| UK | Euro | 0.0679 | 0.9330 | 0.0494 | -0.1301 | 845.06 |
| | | (0.0000) | (0.0000) | (0.5832) | (0.0000) | |
| | SP500 | 0.0192 | 0.9776 | 0.1459 | -0.0381 | 468.35 |
| | | (0.0158) | (0.0000) | (0.2239) | (0.0083) | |
| Switzerland | Euro | 0.0970 | 0.9041 | 0.0000 | -0.1497 | 946.59 |
| | | (0.0000) | (0.0000) | (0.9999) | (0.0000) | |
| | SP500 | 0.0479 | 0.9283 | 0.2962 | -0.1061 | 234.55 |
| | | (0.0085) | (0.0000) | (0.2953) | (0.0101) | |
| Sweden | Euro | 0.0339 | 0.9701 | 0.0000 | -0.0732 | 870.09 |
| | | (0.0000) | (0.0000) | (0.9999) | (0.0000) | |
| | SP500 | 0.0698 | 0.9189 | 0.1209 | -0.1387 | 417.55 |
| | | (0.0003) | (0.0000) | (0.5748) | (0.0003) | |
| Denmark | Euro | 0.0293 | 0.9575 | 0.3422 | -0.0605 | 412.56 |
| | | (0.0005) | (0.0000) | (0.0001) | (0.0004) | |
| | SP500 | 0.0122 | 0.9566 | 0.6468 | -0.0301 | 60.75 |
| | | (0.0153) | (0.0000) | (0.0000) | (0.0259) | |
| Norway | Euro | 0.0394 | 0.9310 | 0.3725 | -0.0589 | 532.20 |
| | | (0.0016) | (0.0000) | (0.0000) | (0.0007) | |
| | SP500 | 0.0140 | 0.9604 | 0.5282 | -0.0239 | 188.30 |
| | | (0.0299) | (0.0000) | (0.0000) | (0.0652) | |

The numbers in brackets ( ) are P values and 0.0000 means that the value is less than 0.00005.

# Table 6: Tests of Dependence Change between Non-euro European Country Stock Market Indices and Euro-zone Stock Market Index

The table shows estimates of the dependence of 5 non-euro country stock market indices with the euro-zone stock market index, using the following model settings including a dummy variable. All indices are denominated in USD. The model settings, except the process of dependence variable defined as below, are the same as those in Table 5. Markets are sorted by region and decreasing market capitalization.

$$(1-\beta_1 L)(1-\beta_2 L)\rho_t = \omega + \gamma \mid u_{t-1} - v_{t-1} \mid + \lambda Dummy_t$$

*$Dummy_t$* = $D_1$ = 1 when $t$ >= 1/1/1996, otherwise $D_1$ = 0. *$Dummy_t$* = $D_2$ = 1 when $t$ >= 1/1/1997, otherwise $D_2$ = 0.
*$Dummy_t$* = $D_3$ = 1 when $t$ >= 1/1/1998, otherwise $D_3$ = 0. *$Dummy_t$* = $D_4$= 1 when $t$ >= 1/1/1999, otherwise $D_4$ = 0.
*$Dummy_t$* = $D_5$ = 1 when $t$ >= 1/1/2000, otherwise $D_5$ = 0.
*f(x)* and *f(y)* are modeled by the GJR-GARCH model with student t distribution

| Country | *Dummy* | $D_1$ | $D_2$ | $D_3$ | $D_4$ | $D_5$ |
|---|---|---|---|---|---|---|
| UK | $\lambda$ | 0.0012 | 0.0020 | 0.0032 | 0.0033 | *0.0049* |
| | | (0.3733) | (0.1764) | (0.0935) | (0.0267) | *(0.0105)* |
| | $\Delta E(\rho)$ | 0.0175 | 0.0283 | 0.0400 | 0.0436 | *0.0603* |
| | $\Delta LLF(c)$ | 0.79 | 0.47 | 2.44** | 4.24*** | *7.47**** |
| Sweden | $\lambda$ | *0.0035* | 0.0042 | 0.0005 | 0.0002 | 0.0006 |
| | | *(0.0228)* | (0.0468) | (0.5252) | (0.6836) | (0.1587) |
| | $\Delta E(\rho)$ | *0.0958* | 0.0797 | 0.0151 | 0.0054 | 0.0208 |
| | $\Delta LLF(c)$ | *5.03**** | 4.61*** | 0.16 | 0.07 | 1.00 |
| Switzerland | $\lambda$ | 0.0020 | 0.0020 | 0.0023 | 0.0003 | 0.0005 |
| | | (0.2557) | (0.2310) | (0.1332) | (0.7869) | (0.6568) |
| | $\Delta E(\rho)$ | 0.0203 | 0.0203 | 0.0228 | 0.0032 | 0.0053 |
| | $\Delta LLF(c)$ | 0.64 | 0.81 | 1.36 | 0.03 | 0.08 |
| Denmark | $\lambda$ | -0.0019 | -0.0024 | -0.0018 | *-0.0043* | -0.0027 |
| | | (0.1337) | (0.0775) | (0.0886) | *(0.0818)* | (0.1152) |
| | $\Delta E(\rho)$ | -0.0564 | -0.0659 | -0.0573 | *-0.1017* | -0.0738 |
| | $\Delta LLF(c)$ | 0.33 | 1.40 | 1.19 | *4.65**** | 1.82 |
| Norway | $\lambda$ | -0.0042 | *-0.0055* | -0.0030 | -0.0037 | -0.0024 |
| | | (0.1418) | *(0.1055)* | (0.1086) | (0.0868) | (0.3985) |
| | $\Delta E(\rho)$ | -0.0696 | *-0.0818* | -0.0563 | -0.0645 | -0.0468 |
| | $\Delta LLF(c)$ | 1.40 | *3.59*** | 1.08 | 2.09** | 0.04 |

The numbers in brackets ( ) are P values and 0.0000 means that the value is less than 0.00005.
$\Delta E(\rho)$: $E(\rho_t \mid D_t{=}1) - E(\rho_t \mid D_t{=}0)$. $\Delta LLF(c)$: Copula LLF(with $D_t$) - Copula LLF(without $D_t$).
** : Significance at 5% level for the likelihood ratio test. ***: Significance at 1% level for the likelihood ratio test.

# Table 7: Tests of Dependence Change between Non-euro European Country Stock Market Indices and S&P500 Index

The table shows estimates of the dependence of 5 major non-euro country stock market indices with the S&P500 index, using the following model settings including a dummy variable. All indices are denominated in USD. The model settings, except the process of dependence variable defined as below, are the same as those in Table 5. Markets are sorted by region and decreasing market capitalization.

$$(1-\beta_1 L)(1-\beta_2 L)\rho_t = \omega + \gamma \mid u_{t-1} - v_{t-1} \mid + \lambda Dummy_t$$

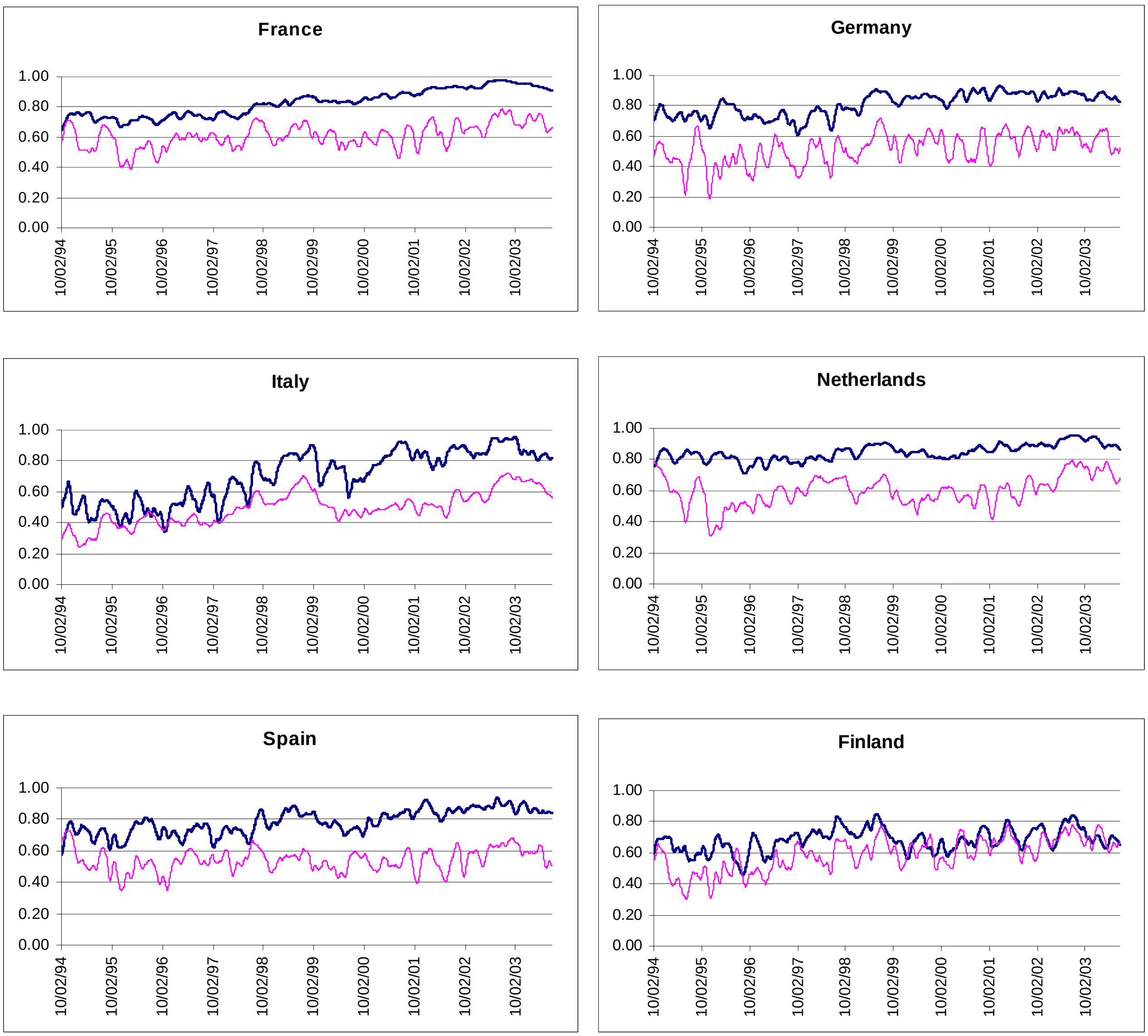
$Dummy_t = D_1 = 1$ when $t >= 1/1/1996$, otherwise $D_1 = 0$. $Dummy_t = D_2 = 1$ when $t >= 1/1/1997$, otherwise $D_2 = 0$.
$Dummy_t = D_3 = 1$ when $t >= 1/1/1998$, otherwise $D_3 = 0$. $Dummy_t = D_4 = 1$ when $t >= 1/1/1999$, otherwise $D_4 = 0$.
$Dummy_t = D_5 = 1$ when $t >= 1/1/2000$, otherwise $D_5 = 0$.
*f(x)* and *f(y)* are modeled by the GJR-GARCH model with student t distribution

| Country | *Dummy* | $D_1$ | $D_2$ | $D_3$ | $D_4$ | $D_5$ |
|---|---|---|---|---|---|---|
| UK | $\lambda$ | *0.0009* | 0.0009 | 0.0015 | 0.0011 | 0.0009 |
| | | *(0.0035)* | (0.0415) | (0.1982) | (0.2106) | (0.1931) |
| | $\Delta E(\rho)$ | *0.1307* | 0.0871 | 0.0671 | 0.0436 | 0.0496 |
| | $\Delta LLF(c)$ | *3.46*$^{***}$ | 2.89$^{**}$ | 2.91$^{**}$ | 1.80 | 1.84 |
| Sweden | $\lambda$ | *0.0072* | 0.0060 | 0.0021 | 0.0012 | 0.0036 |
| | | *(0.0390)* | (0.0419) | (0.2632) | (0.4452) | (0.1092) |
| | $\Delta E(\rho)$ | *0.0967* | 0.0765 | 0.0283 | 0.0173 | 0.0455 |
| | $\Delta LLF(c)$ | *3.15*$^{**}$ | 2.96$^{**}$ | 0.58 | 0.25 | 1.65 |
| Switzerland | $\lambda$ | *0.0083* | 0.0064 | 0.0037 | 0.0031 | 0.0023 |
| | | *(0.0905)* | (0.1101) | (0.1711) | (0.1739) | (0.2545) |
| | $\Delta E(\rho)$ | *0.1573* | 0.1180 | 0.0660 | 0.0556 | 0.0417 |
| | $\Delta LLF(c)$ | *4.00*$^{***}$ | 3.55$^{***}$ | 1.69 | 1.44 | 0.81 |
| Denmark | $\lambda$ | *0.0016* | 0.0013 | 0.0009 | -0.0004 | 0.0005 |
| | | *(0.0202)* | (0.0616) | (0.2576) | (0.5909) | (0.4704) |
| | $\Delta E(\rho)$ | *0.2240* | 0.1374 | 0.0598 | -0.0245 | 0.0328 |
| | $\Delta LLF(c)$ | *3.17*$^{**}$ | 2.23$^{**}$ | 0.71 | 0.13 | 0.23 |
| Norway | $\lambda$ | 0.0001 | 0.0008 | 0.0003 | -0.0006 | 0.0000 |
| | | (0.8452) | (0.1784) | (0.6530) | (0.4397) | (0.9482) |
| | $\Delta E(\rho)$ | 0.0083 | 0.0547 | 0.0149 | -0.0298 | 0.0021 |
| | $\Delta LLF(c)$ | 0.26 | 0.77 | 0.09 | 0.35 | 0.00 |

The numbers in brackets ( ) are P values and 0.0000 means that the value is less than 0.00005.
$\Delta E(\rho)$: $E(\rho_t \mid D_t=1) - E(\rho_t \mid D_t=0)$. $\Delta LLF(c)$: Copula LLF(with $D_t$) - Copula LLF(without $D_t$).
** : Significance at 5% level for the likelihood ratio test. ***: Significance at 1% level for the likelihood ratio test.

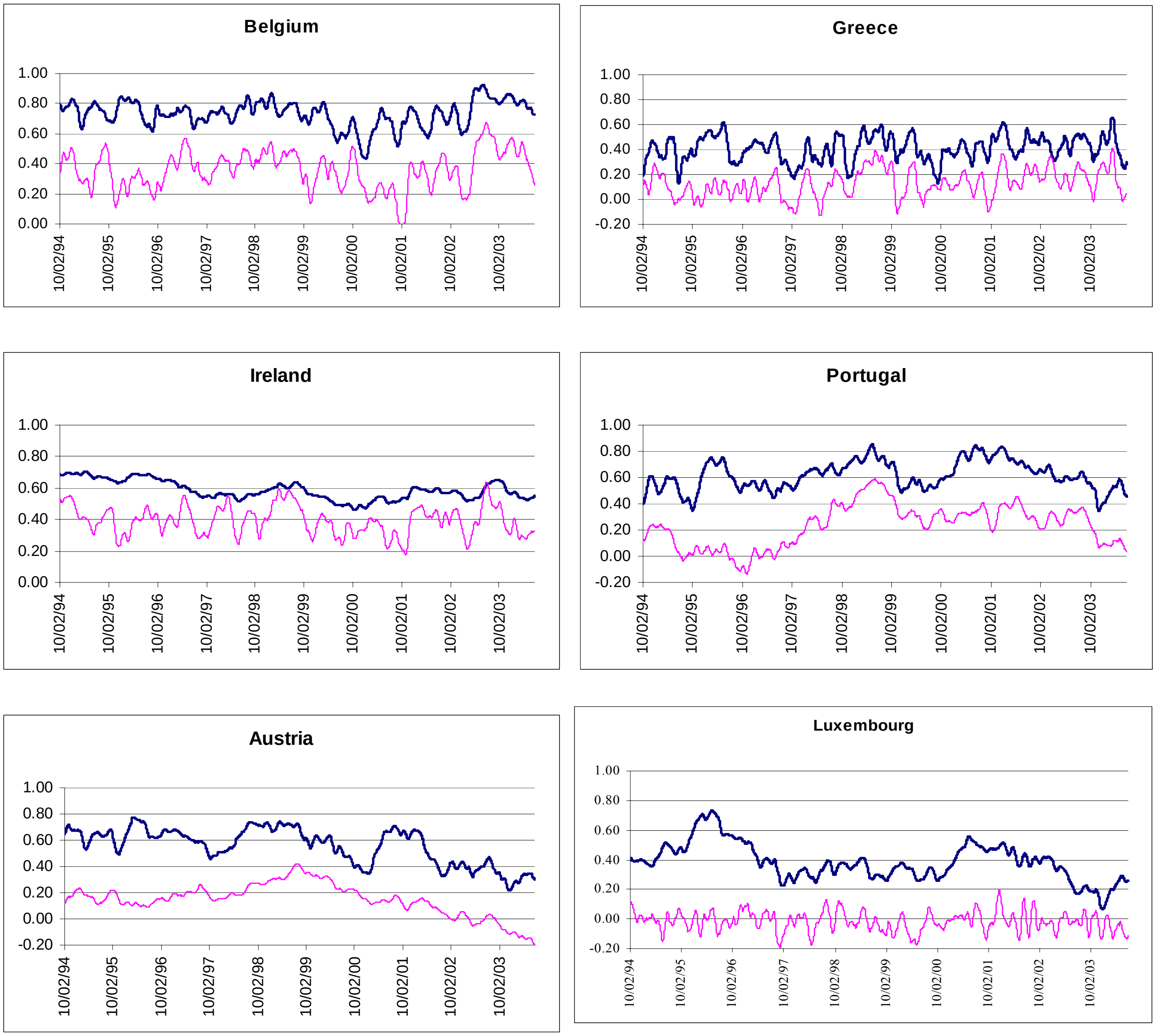
# Figure 1: Dependence of Euro-zone Country Stock Indices with Euro-zone Stock Index and with S&P500 Index

The figure shows the time-varying conditional dependence of 12 euro-zone country stock indices with the euro-zone regional stock index and with the S&P500 index. All indices are denominated in USD. The euro-zone stock index excludes the examined country. The S&P500 index is observed at 16.00 London time. The fat line shows the dependence with euro-zone stock index, the thin line shows the dependence with S&P500 index.



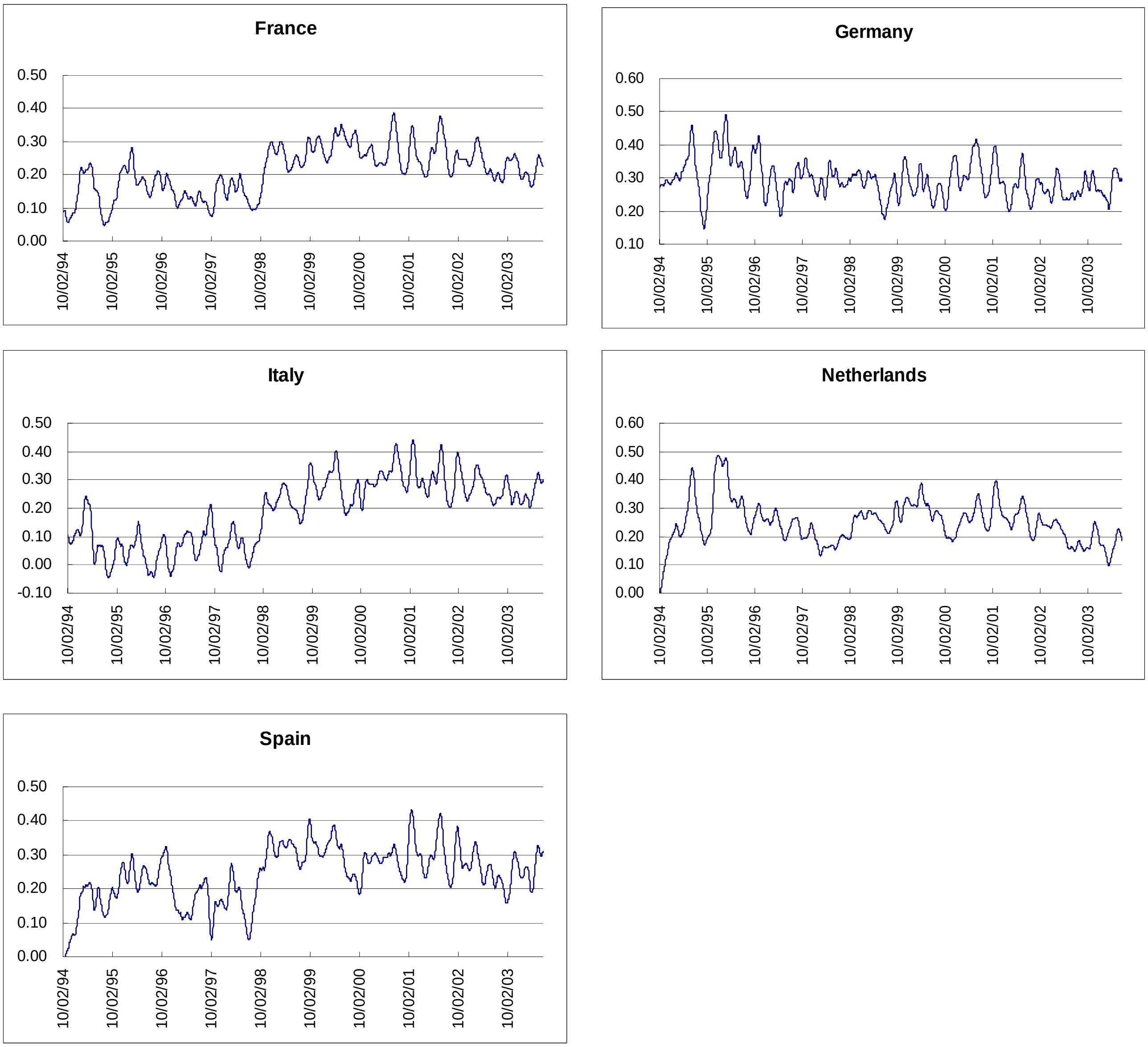

*(continued)*

Figure 1: Dependence of Euro-zone Country Stock Indices with Euro-zone Stock Index and with S&P500 Index (continued)



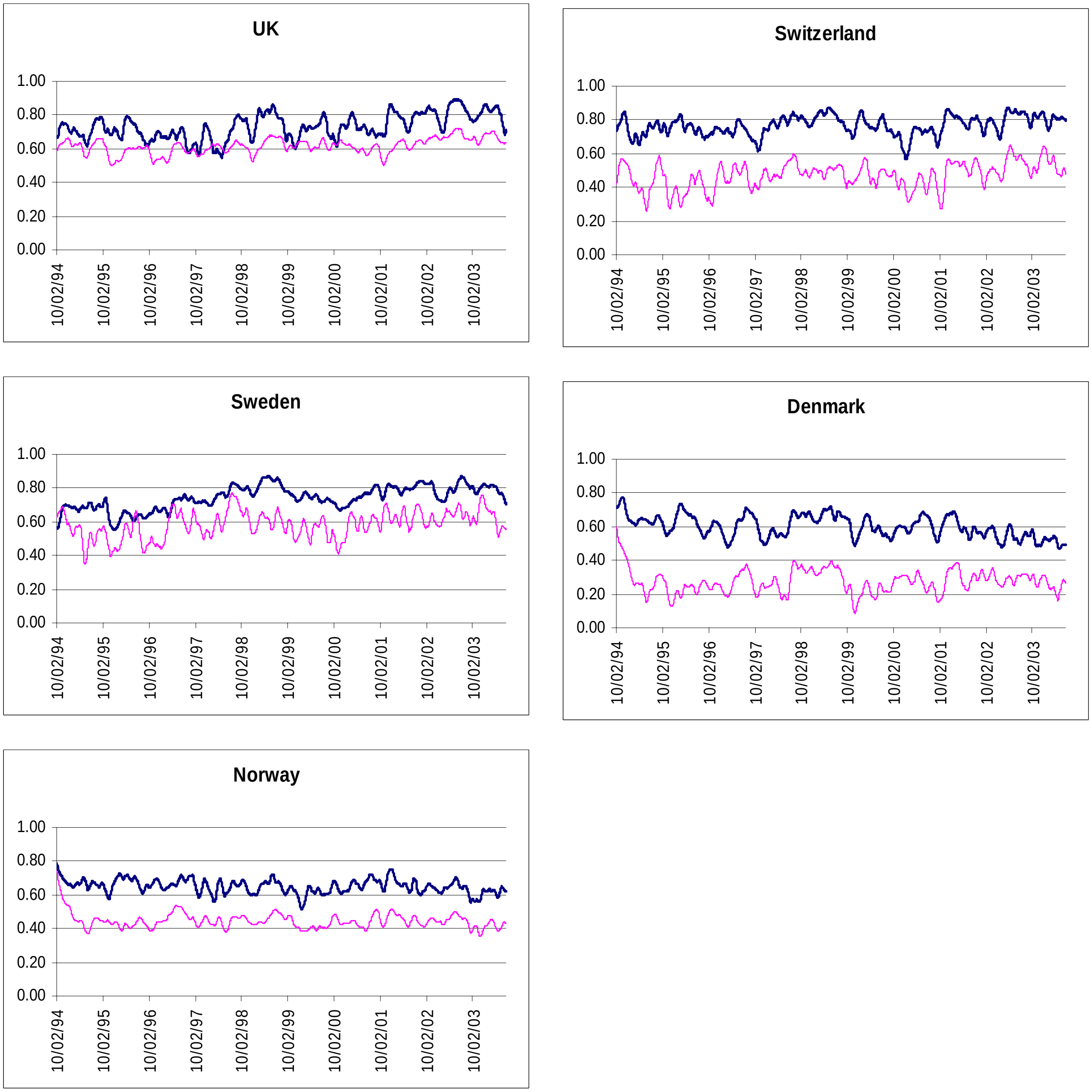
# Figure 2: Differences between Dependence of Euro-zone Country Indices with Euro-zone Regional Index and that with S&P500 Index

The figure shows the time-varying differences of the conditional dependence of 5 major euro-zone countries with euro-zone stock index and that with S&P500 index. All indices are denominated in USD. The S&P500 index is observed at 16.00 London time.



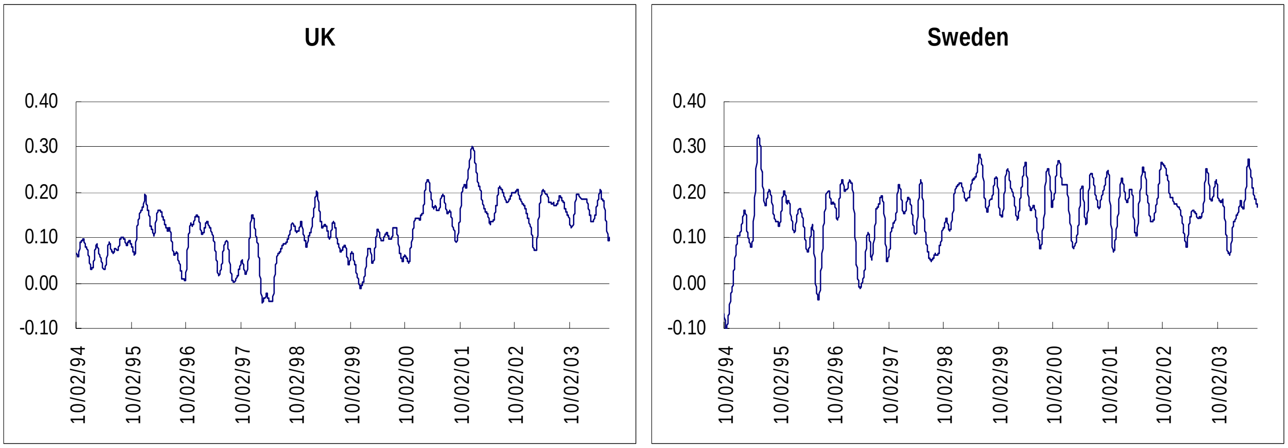
# Figure 3: Dependence of Non-euro Country Stock Indices with Euro-zone Stock Index and with S&P500 Index

The figure shows the time-varying conditional dependence of 5 non-euro country stock indices with the euro-zone stock index and the S&P500 index. All indices are denominated in USD. The S&P500 index is observed at 16.00 London time. The fat line shows the dependence with the euro-zone stock index, the thin line shows the dependence with the S&P500 index.



# Figure 4: Differences between Dependence of Non-euro European Country stock Indices with Euro-zone Regional Index and that with SP500 Index

The figure shows the time-varying differences between the conditional dependence of 2 major non-euro country stock indices with the euro-zone stock index and with the S&P500 index. All indices are denominated in USD. The S&P500 index is observed at 16.00 London time.



# Figure 5: Dependence of Euro-zone Country Stock Indices with Euro-zone Stock Index in EUR and in USD

The figure shows the time-varying conditional dependence of 5 major euro-zone country stock indices with the euro-zone stock index in EUR and in USD. The euro-zone stock index excludes the examined country. The fat line represents returns denominated in EUR, the thin line represents returns denominated in USD.

## Figure 6: Dependence of Non-euro Country Stock Indices with Euro-zone Stock Index in Different Currencies

The figure shows the time-varying conditional dependence of 2 major non-euro country indices with the euro-zone stock index in local currency and in USD. The fat line represents returns in local currency, the thin line represents returns in USD.