The effect of the EMU on the short and long-run stock market dynamics: New evidence on financial integration

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Abstract

This paper deals with the time evolution of stock market integration around the introduction of the euro. In particular we test whether the degree of integration between the main countries of the eurozone has increased after the European monetary union. The contribution of the paper to the extant literature is twofold: a) first, we take into account the potential long-run equilibrium relationship between stock indexes allowing for structural changes in the cointegration space that might capture the effect due to the introduction of the euro and b)we formally test the existence of a greater financial integration after the European monetary union across the main member countries and between these members and UK. Empirical evidence reveal the existence of long-run equilibrium relationships between European stock markets even before the introduction of the euro. Our empirical findings suggest that financial integration is not the direct consequence of the removal of exchange rate risk due to currency unification. It rather arises as a results of macroeconomic convergence. This aspect is corroborated by the nature of the principal component structure of estimated conditional correlations.

Classification J.E.L.: C32, E44, G15.

Key words: cointegration, dynamic financial integration, stock markets, European Monetary Union.

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1 Introduction

The increasing globalization of the world economy has lead to a significant spread in the scope of financial transactions among countries. Economic integration not only might influence the degree of capital market integration across countries over time but also contribute to international financial stability. The academic literature on comovements among international stock markets generally find that globalization is attached to increasing financial integration. For example, Chelley-Steeley (2005), which analyze some eastern European countries, show that the degree of segmentation experienced by these markets has declined significantly over the period 1994-1999; Ayuso and Blanco (2000) provide empirical evidence for the European countries of a significant increase during the nineties not only in the weight of foreign assets in agents' portfolio but also in the correlation between stock indexes, while Fujii (2005) reports that linkages between the Asia and Latin America geographic areas are strengthened around the time of major financial crisis during the nineties, i.e. the 1994-95 Mexican and the 1997-98 Asian crisis. Also, Berben and Jansen (2005) find that correlations among the German, Uk, and US stock markets have more than doubled between 1980 and 2000.

The study of the nature of financial integration is important because of the existence of real and financial effects. Financial integration tends not only to increase international correlations in both consumption and GDP fluctuations (see Imbs, 2006), but also affect the relationship between output growth and volatility (see Kose et al., 2006, which provide empirical evidence supporting that even the negative relationship between volatility and growth has survived into the 1990s, financial integration tends to weaken such relationship). Also, time varying-market integration in the world market is an important determinant of the expected returns and the cost of capital in emerging markets, basically due to the diminishing diversification benefits by trading in developed stock markets. Consistent with this idea, de Jong and de Roon (2005) which analyze 30 emerging markets from Asia, the Far East, Europe and the Mideast and Africa find that a decrease in segmentation leads to lower expected returns even when usually it is accompained by an increase in beta.

One of the most important event concerning financial economic integration is the recent past experience of introducing the euro, which has undoubtedly influence the nature of the correlations between the stock markets of the European monetary union member countries (see, for example, Kim et al., 2005, 2006). Given that capital flows across international financial markets follow return differentials, the existence of a common monetary policy between the member countries of the monetary union should leads to a greater homogeneous investment opportunities across these economies, and then reducing the potential benefits of international diversification.

This paper deals with dynamic stock market integration encouraged by the

European monetary union for the main member countries (Germany, France, Italy and Spain) and the UK. While the recent papers of Kim et al. (2005, 2006) focus on the dynamic stock market integration between either a specific country and the rest of the Europe or different financial assets (stocks and bonds) this paper deals with the dynamic interactions inside the monetary union. Though there is well documented that economic integration tends to decline the benefits of international diversification, that is, trends in stock prices become more similar, the potential existence of cointegration relationship was not taken into account to estimate time-varying correlations. This paper contributes to the literature by analyzing short and long-run financial integration. First we account for potential economic long-run equilibrium relationships between stock market indices to estimate time-varying correlations of stock returns. And secondly, we test whether the distribution of correlations across countries has changed over time using non-parametric techniques. Empirical findings can be summarized as follows: a) the introduction of the euro has undoubtedly stabilized the process of financial integration across the main member countries of the EMU; b) the existence of long-run equilibrium relationships between European stock markets even before the introduction of the euro. Cointegration implies that stock index can not deviate from each other in the long-run, revealing a strong degree of financial integration; c) however, the exogeneity of the UK stock index in the cointegration space show that long-run financial integration just appears among the eurozone member countries.

The rest of the paper is organized as follows. Section II discusses the data used. Section III describes the cointegration analysis. Section IV present the bivariate GARCH methodology to estimate time-varying correlations and section V reports empirical results. Finally, section VI summarizes and provides concluding remarks.

2 Data

In this section we discuss the data used and some preliminary statistical properties of stock index returns. The data set comprises daily closing stock market indices from the main eurozone members (Germany, France, Spain and Italy), as well as the UK, which is the nearest non-eurozone country with the highest industrialization[1]. The sample period analyzed is from 21th April 1993, up to 30th December 2004. Daily closing stock indices were collected from

[1]Despite the existence of "home bias", that is, the fact that investors across the world have small proportions of their assets allocated to foreign markets, Portes and Rey (2005) find that the most important determinant of global equity transactions between two countries is geographical proximity.

Bloomberg[2]. For each index, we compute the return as the log first differences between closing prices in trading day *t*-1 and *t*. From a perspective of policy maker concerned with financial integration, stock prices should closely reflect price discovery as it takes place in real financial markets. This way, we do not adjust stock returns of exchange rate fluctuations and dividend payments.

Table 1 reports some descriptive statistics. The sample mean negligible, as expected from a systematically long and short trading strategy on consecutive trading days. On an annual basis, the average return becomes about 18% for Germany, 13% for both France and the UK and 20% for both Spain and Italy. Considering the standard deviation as a rough measure of volatility for the overall sample the highest volatility is for the Italian stock market while the lowest one corresponds to the UK stock market. Moreover, stock returns distributions have excess of kurtosis and are slightly skewed. Both characteristics are generally associated with conditional heteroskedasticity. To asses the existence of ARCH effects in stock returns we perform the Engle's lagrange multiplier test. Empirical values of the test systematically reject the null, pointing towards a parametrization for the second order moments of stock market returns using a particular specification of the GARCH family of models.

3 Identification of the long-run structure

The empirical analysis of cointegration is based on the VAR(3) model with a constant restricted to lie in the cointegration space:

$$\Delta x_t = \Gamma \Delta x_{t-1} + \alpha\beta' x_{t-1} + \alpha\delta_0 Ds_t + \Phi_1 Dp_t + \mu_0 + \varepsilon_t \tag{1}$$

where $\mu_0 = \alpha\beta_0 + \alpha_\perp \gamma_0$, so that β_0 is an intercept in the cointegration relations and γ_0 is equal to zero. The coefficient δ_0 stands for a mean shift in $\beta' x_t$ as a result of a mean shift in the variables that do not cancel in the cointegrated relations. This mean shift is captured by a dummy which is zero until 1999 and one otherwise. This dummy is aimed to capture the introduction of the euro and the implementation of the common monetary policy. Dp_t stand for a permanent impulse dummy and corresponds to the first difference of Ds_t.

The baseline model has been carefully checked for signs of misspecification using a variety of diagnostic tests. According to these, the model seems to describe the data reasonably well. There is no deviations from the basic assumptions of residual independence, although some degree of heteroscedasticity and non-normality were detected. According to Gonzalo (1994), even under this

[2]The authors wish to thanks Emilio Palomar for providing us with the data used.

conditions, the estimation of the cointegration space by the maximum likelihood procedure proposed by Johansen (1988) is efficient. The conditional variance will be subject to further modelling by using a GARCH approach.

The choice of the cointegration rank is made based on the Bartlett corrected trace test, the roots of the characteristic polynomial and the t-statistic of the adjustment coefficients. All this information is shown in table 2. The 5% critical values for the Trace test have been simulated to account for the shift dummy restricted to the cointegration space. According to the Trace test in table 2, we might accept $r = 2$, although $r = 3$ is also a plausible result. The roots of the characteristic polynomial of the VAR might also provide useful information about the rank. Thus, if a nonstationary vector is wrongly included in the cointegration space, the largest unrestricted characteristic root will be close to the unit circle. In our case, the difference between roots for different choice of r is fairly small, so that they are not very informative. This fact can be due to the existence of small I(2) components in the data. The middle panel in table 2 presents the t-values for the adjustment coefficients. Only FTSE-100 is significantly adjusting to the third cointegration relation, which suggests that the choice $r = 2$ would imply no important lost of information.

Finally, as an additional check for the choice of the rank, table 3 present the time series properties of long-run exclusion. This test might help to decide whether a variable improves the specification of the cointegration space. The choice of $r = 3$ when comparing with $r = 2$ alters the statistical properties of the model, under the former choice, the CAC index is long-run excludable, although borderline. Hence, this index would not add significant information to the long-run analysis. However, three is no clear reason why this should be the case so that the choice $r = 3$ might be inappropriate. Moreover, the presence of multicollinearity between variables can lead to acceptance of long-run exclusion even though the variable is significant in the long-run relations. Given this possibility, we decide to keep the CAC index in the model and set $r = 2$.

Table 3 also reports the test statistics for stationarity and weak exogeneity. These tests are also presented for $r = 2$ and the two plausibles alternatives $r = 1$ and $r = 3$. Neither of the variables is stationary. Being this result independent of the choice of r. Regarding the weak exogeneity tests, for $r = 2$ and $r = 3$, it seems that both the CAC and FTSE-100 indices are weakly exogenous. If this were the case, shocks to both indices would drive the system in the sense that the other indices would adjust to CAC and FTSE-100 but the latter would not adjust to the remaining variables. This results will be subject to further formal test.

The long-run identified structure is accepted with a p-value of 0.39. The first cointegration vector corresponds to:

$$ecm1_t = CAC40_t - \underset{[-5.02]}{0.71} DAX_t - \underset{[-7.97]}{0.91} MIBTEL_t + \underset{[5.20]}{0.80} FTSE100_t \quad (2)$$

whereas the second one is given by:

$$ecm2_t = DAX_t + \underset{[3.85]}{0.65} IBEX35_t - \underset{[-8.76]}{2.01} FTSE100_t - \underset{[-2.93]}{0.10} DS_{euro} + \underset{[5.14]}{3.74} \quad (3)$$

Both cointegration vectors are plotted in figure 1. According to these graphics, the above relations are clearly stationary.

The first cointegration vector links the whole set of analyzed stock indices with the only exception of the IBEX-35. This means that the CAC-40, DAX, MIBTEL and FTSE-100 indices can not move independently from each other, at least in the long-run. Thus, for example, if either DAX or MIBTEL goes up, CAC-40 tends also to increase, whereas CAC-40 returns are inverse related to the one of FTSE-100. This results implies therefore an important degree of market integration, furthermore, the existence of a long-run equilibrium across these set of indices imply some degree of market predictability, since deviation away from equilibrium are expected to be corrected.

The second cointegration vector can be interpreted as a long-run relation for the Spanish index, IBEX-35. Is is important to notice that meanwhile the long-run relationships linking the CAC-40, DAX, MIBTEL and FTSE-100 indices are not significatively affected by the introduction of the EMU, this is not the case for the IBEX-35. In order to find an equilibrium relationship linking the IBEX-35 and (some of) the other indices, we need to include the dummy DS_{euro} to capture the introduction of the euro, otherwise, cointegration can not be found. This result is interesting in the extent that points out that market integration was already achieved before the EMU creation for some, but not for all, indices. The shift dummy in the equilibrium mean for the second cointegration vector suggest that the level of the returns for the IBEX-35 increased after the introduction of the EMU.

Parameter constancy of the cointegration space is checked using the recursive test procedures in Hansen and Johansen (1999). According to these tests the cointegration space seems reasonably stable.

Next we check for the hypothesis that some of our indices might be weakly exogenous. This hypothesis is jointly tested with the identified long-run structure. The joint hypothesis is only accepted for FTSE-100, a non EMU member, with a p-value of 0.28. This result suggest a lesser degree of market integration for the British stock exchange. In addition, the fact that FTSE-100 is weakly exogenous implies that this index constitutes a common stochastic trend driving the system. Thus, FTSE-100 might be capturing the effects to the EMU member stock indices coming from shocks originated outside the euro area. The other indices in the system are not weakly exogenous so that they are adjusting to the cointegration relations. The degree of adjustment is captured by the loadings to the cointegration vectors. According to our results, the CAC-40 index is adjusting to both cointegration realtionships. The DAX and IBEX-35 adjust to the second whereas MIBTEL adjusts to the first. It appears that the dummy variable significatively affects the dynamics for the CAC-40, DAX and IBEX-35 indices, stressing the importance of this shock to understand the degree of market integration in the EMU zone.

4 Modelling time-varying correlations

The objective of the paper is to analyze the dynamic evolution of financial stock market integration. According to recent papers of Chelley-Steeley (2005), Fujii (2005) and Kim et al., (2005) we use the conditional correlation between stock index markets as a proxy of dynamic financial integration. It is overwhelming empirical evidence not only that distributions of daily stock index market returns are skewed with thick tails, but also that to model conditional variances of stock index returns and their respective conditional covariances GARCH models is an adequate econometric tool. Following the pathbreaking idea of Engle (1982), numerous parametric speficiations have been proposed in the literature to distinguish between the conditional and unconditional second order moments. In particular, we use the dynamic conditional correlation model proposed in Engle (2002)[3], which, as the best of our knolowgede, has not been yet used in the literature to deal with the analysis of financial stock market integration. This specification is a generalization of the constant conditional correlation multivariate GARCH proposed in Bollerslev (1990), but has the flexibility of univariate GARCH models coupled with parsimonious parametric models for the correlations. To represent the dynamics of the conditional means of the stock market returns for each pair of the coutries considered, we posit the following error correction model:

$$
\begin{aligned}
r_{i,t} &= \alpha_{1i}\, ecm1_{t-1} + \alpha_{2i} ecm2_{t-1} + \gamma_i r_{i,t-1} + \delta_i r_{j,t-1} + \varepsilon_{i,t} \\
r_{j,t} &= \alpha_{1j}\, ecm1_{t-1} + \alpha_{2j} ecm2_{t-1} + \gamma_j r_{i,t-1} + \delta_j r_{j,t-1} + \varepsilon_{j,t}
\end{aligned}
$$

[3]As pointed by Engle, Monte Carlo experiment reveals not only the bivariate version of the DDC-MV-GARCH model provides a very good approximation to a variety of time-varying correlation processes, but also the comparison of this model with simple multivariate GARCH shows that this model is often the most accurate.

$$\varepsilon_t^{'} = (\varepsilon_{i,t}, \varepsilon_{j,t}) \sim N(0, H_t), \quad H_t = D_t R_t D_t, \;\; D_t = diag\left(\sqrt{h_{11,t}}, \sqrt{h_{22,t}}\right)$$

that is an AR(1) process with short-run adjustments that allows for spillover effects between both countries, while the dynamics of the elemnets concerning H_t matrix is as follows:

$$h_{ii,t} = \omega_{ii} + \sum_{r=1}^{q} \alpha_{ir} \varepsilon_{i,t-r}^2 + \sum_{k=1}^{p} \beta_{ik} \varepsilon_{i,t-r}^2$$

$$R_t = (Q_t^*)^{-1} Q_t (Q_t^*)^{-1}, \;\; Q_t = diag\left(\sqrt{q_{11,t}}, \sqrt{q_{22,t}}\right)$$

$$Q_t = [1 - \alpha(1) - \beta(1)] \bar{Q} + \sum_{i=1}^{q^*} \alpha_i L^i \eta_{t-1} \eta_{t-1}^{'} + \sum_{j=1}^{p^*} \beta_j L^j Q_{t-1}$$

where $\eta_t^{'} = \varepsilon_t^{'} D_t^{-1}$ represents the vector of standardized residuals of the two univariate GARCH models. Finally,

$$\bar{Q} = T^{-1} \sum_{t=1}^{T} \varepsilon_t \varepsilon_t^{'}$$

is the matrix of standardized inconditional covariances where T denotes the sample size.

The model is fitted via maximum likelihood estimation. The log likelihood function for this estimator can be expressed as [4]:

[4]While a several other conditional densities might be a preferable alternative in theory, like the generalized error distribution proposed in Nelson (1991), all useful specifications must necessarily restrict the dimensionality of the parameter space to be tractable. For this reason, and taking into account that conditional normality is able to capture many stylzed facts for stock returns reasonably well (see Franses and van Dijk, 2000, chp.4), we use the conditional Gaussian distribution.

$$
\begin{aligned}
L &= -\frac{1}{2}\sum_{t=1}^{T}\left(n\log(2\pi)+\log|H_t|+r_t^{'}H_t^{-1}r_t\right)= \\
&= -\frac{1}{2}\sum_{t=1}^{T}\left(n\log(2\pi)+2\log|D_t|+r_t^{'}D_t^{-1}D_t^{-1}r_t-\varepsilon_t\varepsilon_t^{'}+\log|R_t|+\varepsilon_t^{'}R_t^{-1}\varepsilon_t\right)
\end{aligned}
$$

As stated above, the dynamic conditional correlations are estimated based on bivariates DCC-MV-GARCH models so that our five dimensional VEC model must consequently reduced. Such a dimensional reduction presents two drawbacks, first, we impose zero restrictions to the short-run adjustment coefficients in the VECM (i.e., the lagged differences) for the variables not involved in the bivariate GARCH model, second, if cointegration is not robust to GARCH effects, the loadings to the cointegration vectors (α coefficients) might present different values among the bivariate GARCH models and might also differ from those estimated in the five dimensional VECM. The models are estimated using the Broyden, Fletcher, Goldfarb and Shanno (BFGS) algorithm and the more parsimonious model that leads to uncorrelated standardized and squared standardized residuals was finally chosen

5 Empirical results

The bivariate error correction DCC-MV-GARCH model was found to be a valid parameterization of conditional first and second order moments of stock index returns. For all pair of of stock index returns considered, the Ljung-Box Q-statistic, reported in Table 5, does not systematically reject the null hypothesis of absence of autocorelation in both the standardized and squared standardize residuals.

Figures 2 to 4 depict the estimated conditional correlations from these bivariate error correction DCC-MV-models. In all cases, it can be observed since 1996-97 a significant change in the dynamics of integration between each pair of countries including the UK. A clear and positive trend in the conditional correlations takes place, revealing that financial integration significantly increases. Table 6 reports the average conditional correlation corresponding to the partitioned samples taking into the introduction of the euro. Though the financial integration has increased between any pair of countries considered, the lower rise concerns with the financial integration with the UK, the country that behaves as exogenous in the cointegration structure. Inside the monetary union the higher increase correspond to Italy.

Interestingly enough the volatility of time-varying correlations clearly reduces once the common currency is achieved. This pattern is similar to those reported by Kim et. al. (2005) between the same EMU countries and the eurozone. To statistically asses the increase of financial integration we compute the Kolmogorov-Smirnov test for the null hypothesis of both (pre-euro and after the euro) correlation distributions has the same cumulative probability distribution. This test has the advantage of making no assumption about the distribution of data. Empirical results are reported in Table7. In all cases the null hypothesis is systematically rejected, suggesting that the introduction of the euro has led to a regime shift in the process of financial integration. Table 7 also report the confidence interval for the conditional correlation after January 1, 1999[5] using the Tchebycheff inequality. The interval is carried out with at most 10% significance level. Comparing information in Table 6 and Table 7, we reject the hypoythesis of equal average correlation. Only for two cases (Germany-Uk and Spain-UK) the confidence interval contains the average value corresponding to the pre-euro subsample. However this average value is basically the border of the confidence interval, so the acceptance of equal average correlation before and after the introduction of the euro is not statistically clear.

Concerning the error correction mechanism, in our estimated five dimensional VECM, only five out of 48 short-run adjustment coefficients are statistically significative, so that we impose at most one zero restriction in the conditional mean of the bivariate GARCH models. In our case, the imposition of zero restrictions does not seem to be of big concern. Table 4 presents the loadings to the cointegration relationships for the ten bivariate GARCH models (one model for each pair of variables involved in the VECM). The estimates have been obtained taking the cointegration vectors as fixed during the estimation of the GARCH model. Trying to estimate the cointegration vectors directly as part of the maximum likelihood function for the GARCH doesn't work very well since modest changes to the cointegrating vector force rather large changes to other components. As can be seen in table 4, France, Germany and Spain are error correcting to the second cointegration relationship, whereas Italy corrects to the first one. The United Kingdom is weakly exogenous. These results are robust for all the ten bivariate GARCH models[6] and replicate the results obtained for the five dimensional VECM.

The above results suggest that the introduction of the euro being an encouragement factor of financial integration rather than the original cause. To reinforce the previous argument, next we analyze the principal component structure of the estimated time varying correlations. If the introduction of the euro had led to a strengthening of financial integration a powerful linear dimensionality reduction of the correlation data set can be achieved with the

[5] It is more appropriate the use of standard deviation as a measure of variability for time varying conditional correlation after the introduction of the euro because in this subsample the dynamic of correlations is mean stationary

[6] The only exception is Spain in model 6, which also corrects to the first cointegration relationship.

principal component technique. However the nature of the principal component structure should be similar if financial integration triggered after reaching the common currency.

5.1 Principal components structure of conditional correlations

The objective of principal component analysis is to obtain a few uncorrelated variables (principal components) in terms of linear combinations of original variables such in order to maximize the variance of these components. This objective becomes especially interesting in multivariate analysis when using a large number of interrelated variables. In our case, a reduction in dimensionality may be useful to provide additional insights on financial integration. In this section we analyze the principal component structure of the estimated conditional correlations. If financial integration has strengthened after the introduction of the euro, either the same number of components factors should explain a greater percentage of the variance or a lower number of factors should explain at least the same percentage of the variance.

The model can be stated as follows:

$\rho = Lf + Du$

where ρ is a 10×1 vevtor of standardized conditional correlations, L is a $10 \times r$ (< 10) matrix of loadings (weights of each variable in components), $f^{'} = (f_1, f_2, ..., f_m)$ is the matrix corresponding the principal components and D is a diagonal matrix of loadings corresponding to the specific factors $u^{'} = (u_1, u_2, ..., u_m)$. Let us denote the covariance matrix of correlations as S. This matrix can be transformed into a diagonal matrix using the orthonormal transformation, that is:

$\Lambda = \Gamma^{'} L L^{'} \Gamma + \Gamma^{'} E \Gamma$

where Γ is the orthonormal matrix whose columns are the eigenvectors of matrix S, $E = D^{'}D$ is the residual matrix and $tr\,(\Lambda) = tr\,(S)$. This way, for a given number of p componets, the matrix of loadings can be estimated trough the p eigenvectors which correspond to the p largest eigenvalues of matrix S using the following expression:

$\hat{L} = \Gamma \Lambda^{\frac{1}{2}}$

Table 8 reports the principal component structure for the estimated conditional correlations from the full sample, and from the both subsamples corresponding before and after January 1, 1999. The qualitative nature of the solution is similar in the full sample and after the introduction of the euro; in both cases, using as stopping rule the average root (eigenvalues greater than one), only one component arises. However, in the pre-euro sample one additional factor appears to be considered. Interestingly enough, the first factor has greater explaining ability in the post-euro sample, revealing that it is more representative of the time evolution of conditional correlations after the introduction of the euro. These aspects corroborate that though financial integration takes place before the monetary union, when the common currency is implemented financial integration becomes stronger.

6 Summary and Conclusions

This paper investigates the dynamic nature of financial integration for the main members countries of the euro (Germany, France, Spain and Italy) and the UK. The innovation of the paper relies on the joint analysis of short and long-run financial integration. Using time varying correlation as a proxy of the degree of dynamic financial integration, a bivariate error correction DCC-MV-GARCh model is used to estimate conditional correlations between stock index returns along the sample period covering 1993-2004. Our empirical findings show the existence of long-run equilibrium relationship between stock market indices, but stock market indices are cointegrated once we allow a structural change in the long run equilibrium relationship corresponding to the introduction of the euro. This results implies an strong degree of financial market integration because the existence of a long-run equilibrium across stock indices imply some degree of market predictability in the sense that deviations away from equilibrium are expected to be corrected in the short-run. This can be interpreted as a reduction of arbitrage opportunities between stock markets after the removal of exchange rate risk. However, the potential benefits of international diversification from European portfolios have reduced.

In general, the introduction of the euro has increased the degree of financial integration relate to each other. Concerning the EMU countries the higher increase correspond to Italy. As expected, financial integration of the EMU countries with the UK has experimented a lower rise. In sum, our results suggest that a new age of financial integration was due to fiscal and monetary policies implemented in the european countries to achieve the macroeconomic convergence conditions. The introduction of the EMU, rather than to be the origin of financial integration, it appears to be a factor which consolidate this dynamic process.

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Table 1: Descriptive statistics

	France	Germany	Spain	Italy	UK
Mean(in percentage)	0.0328	0.0448	0.0503	0.0506	0.0325
Standard Deviation	1.5134	1.6703	1.5833	1.6984	1.2081
Skewness	0.0266	-0.0071	0.2499	0.2235	-0.1158
Excess of Kurtosis	1.8811	1.9863	4.5277	4.3647	2.0299
ARCH(5)	44.8184	59.3966	14.1212	43.6717	62.3152
	[0.00]	[0.00]	[0.00]	[0.00]	[0.00]

ARCH(5) denotes Engle's Lagrange Multiplier test for ARCH effects considering five lags. P-values are in brackets.

Table 2: The cointegration rank

Modulus of the 6 largest characteristic roots

$r = 3$	1.00	1.00	1.00	1.00	0.98	0.21
$r = 2$	1.00	1.00	1.00	0.98	0.97	0.22
$r = 1$	1.00	1.00	0.99	0.98	0.97	0.22

Adjustment coeficientes (*t-ratios*)

	CAC-40	DAX	IBEX-35	MIBTEL	FTSE-100
$\hat{\alpha}_1$	**-2.09**	**3.47**	**3.18**	**4.71**	**2.41**
$\hat{\alpha}_2$	**-2.22**	**-2.46**	**-2.76**	-0.12	0.37
$\hat{\alpha}_3$	-0.83	-2.03	-0.31	0.58	**-2.38**

Trace test

$p - r$	r	Trace	Trace*	Cval95	Cval95*
5	0	90.25	89.63	76.81	89.07
4	1	60.14	58.30	53.94	61.84
3	2	34.24	32.97	35.07	43.52
2	3	18.36	16.97	20.16	26.38
1	4	6.80	5.68	9.14	13.00

Probability in brackets. Bold face indicates $|t{-}ratio| > 2$. $Trace^{*}$ stands for the Barlett corrections to the standard Trace test for the I1-model. $Cval95$ is the critical values corresponding to a model without dummies. $Cval95^{*}$ is the simulated critical values at 5% for the I1-model with shift dummies.

Table 3: Time series properties in the VAR

Test for variable exclusion

r	dgf	$\chi^2(r)$	CAC	DAX	IBEX-35	MIBTEL	FTSE-100
1	1	3.84	0.27 [0.59]	3.14 [0.08]	0.06 [0.78]	1.49 [0.22]	1.39 [0.24]
2	2	5.99	7.13 [0.02]	8.16 [0.01]	9.83 [0.00]	8.12 [0.01]	11.32 [0.00]
3	3	7.81	7.36 [0.06]	11.19 [0.01]	11.70 [0.00]	13.34 [0.00]	16.57 [0.00]

Test for stationarity

r	dgf	$\chi^2(r)$	CAC	DAX	IBEX-35	MIBTEL	FTSE-100
1	4	9.48	20.58 [0.00]	18.40 [0.00]	22.61 [0.00]	20.78 [0.00]	20.43 [0.00]
2	3	7.81	18.22 [0.00]	16.24 [0.00]	19.98 [0.00]	18.76 [0.00]	17.50 [0.00]
3	2	5.99	9.00 [0.01]	7.22 [0.02]	10.40 [0.00]	10.74 [0.00]	8.81 [0.01]

Test for weak exogeneity

r	dgf	$\chi^2(r)$	CAC	DAX	IBEX-35	MIBTEL	FTSE-100
1	1	3.84	2.61 [0.10]	3.16 [0.07]	3.13 [0.08]	2.78 [0.09]	0.38 [0.54]
2	2	5.99	5.26 [0.07]	6.48 [0.03]	7.83 [0.01]	12.78 [0.00]	3.37 [0.18]
3	3	7.81	7.12 [0.07]	11.63 [0.00]	10.49 [0.01]	14.47 [0.00]	6.48 [0.09]

For the test for stationarity, a restricted intercept is included in the cointegrating relation(s). Probability in brackets.

Table 4: Estimated error correction coefficients in the DCC-MV-GARCH

		α_1	α_2
Model 1	Germany	0.0010 [0.69]	**-0.0124** [0.00]
	France	-0.0022 [0.46]	**-0.0111** [0.00]
Model 2	Germany	0.0004 [0.93]	**-0.0126** [0.00]
	Italy	**0.0128** [0.00]	-0.0027 [0.50]
Model 3	Germany	0.0019 [0.70]	**-0.0085** [0.04]
	Spain	-0.0011 [0.98]	**-0.0112** [0.01]
Model 4	Germany	0.0009 [0.82]	**-0.0113** [0.00]
	United Kingdom	0.0052 [0.08]	0.0021 [0.52]
Model 5	France	-0.0071 [0.75]	**-0.0151** [0.00]
	Italy	**0.0101** [0.03]	-0.0087 [0.51]
Model 6	France	0.0005 [0.79]	**-0.0113** [0.00]
	Spain	**0.0050** [0.00]	**-0.0126** [0.00]
Model 7	France	0.0003 [0.94]	**-0.0092** [0.01]
	United Kingdom	0.0059 [0.08]	0.0021 [0.52]
Model 8	Italy	**0.0153** [0.00]	-0.0063 [0.17]
	Spain	0.0142 [0.44]	**-0.0147** [0.00]
Model 9	Italy	**0.0126** [0.00]	-0.0024 [0.59]
	United Kingdom	0.0032 [0.23]	-0.0004 [0.90]
Model 10	Spain	0.0046 [0.32]	**-0.0083** [0.03]
	United Kingdom	0.0042 [0.22]	0.0025 [0.46]

Asymptotic p-values in brackets. Significant coefficients in bold. α_1 and α_2 stand respectively for the loadings to the first and second cointegration relationships.

Table 5: Diagnosis test for the DCC-MV-GARCH model

		Standardized residuals	Squared standardized residuals
Model 1	Germany	13.17 [0.87]	15.67 [0.61]
	France	29.73 [0.07]	17.65 [0.61]
Model 2	Germany	13.42 [0.86]	13.35 [0.50]
	Italy	19.22 [0.51]	29.23 [0.08]
Model 3	Germany	14.32 [0.81]	13.71 [0.84]
	Spain	30.56 [0.06]	7.58 [0.99]
Model 4	Germany	15.52 [0.75]	12.10 [0.91]
	United Kingdom	20.45 [0.43]	13.55 [0.85]
Model 5	France	31.40 [0.05]	17.05 [0.65]
	Italy	17.36 [0.63]	20.85 [0.40]
Model 6	France	28.73 [0.09]	18.75 [0.54]
	Spain	28.84 [0.09]	10.56 [0.96]
Model 7	France	30.61 [0.06]	21.86 [0.35]
	United Kingdom	22.99 [0.29]	21.69 [0.36]
Model 8	Italy	21.19 [0.39]	22.49 [0.31]
	Spain	28.55 [0.10]	8.53 [0.98]
Model 9	Italy	17.85 [0.59]	33.60 [0.03]
	United Kingdom	21.13 [0.39]	26.64 [0.15]
Model 10	Spain	31.60 [0.05]	13.77 [0.84]
	United Kingdom	21.19 [0.38]	11.32 [0.94]

Empirical values of the Ljung-Box Q-statistics for the null hypothesis of absence of autocorrelation allowing for 20 lags. P-values in brackets

Table 6: Average estimated conditional correlations

	pre-EMU	post-EMU	Variation Rate
Correlations			
Germany-France	0.57	0.85	47.68%
Germany-Italy	0.34	0.79	133.96%
Germany-Spain	0.53	0.76	43.78%
Germany-United Kingdom	0.53	0.69	30.89%
France-Italy	0.47	0.86	82.25%
France-Spain	0.63	0.83	32.63%
France-United Kingdom	0.60	0.75	24.61%
Italy-Spain	0.47	0.81	74.54%
Italy-United Kingdom	0.38	0.70	84.26%
Spain-United Kingdom	0.54	0.66	22.82%

Table 7: Non-parametric testing from estimated conditional correlations

	KS test	Confidence Interval
Germany-France	0.82 [0.00]	[0.66, 1.00]
Germany-Italy	0.88 [0.00]	[0.60, 1.00
Germany-Spain	0.77 [0.00]	[0.56, 1.00]
Germany-United Kingdom	0.65 [0.00]	[0.52, 1.00]
France-Italy	0.84 [0.00]	[0.74, 1.00]
France-Spain	0.72 [0.00]	[0.73, 1.00]
France-United Kingdom	0.48 [0.00]	[0.61, 1.00]
Italy-Spain	0.68 [0.00]	[0.66, 1.00]
Italy-United Kingdom	0.73 [0.00]	[0.47, 1.00]
Spain-United Kingdom	0.49 [0.00]	[0.48, 0.98]

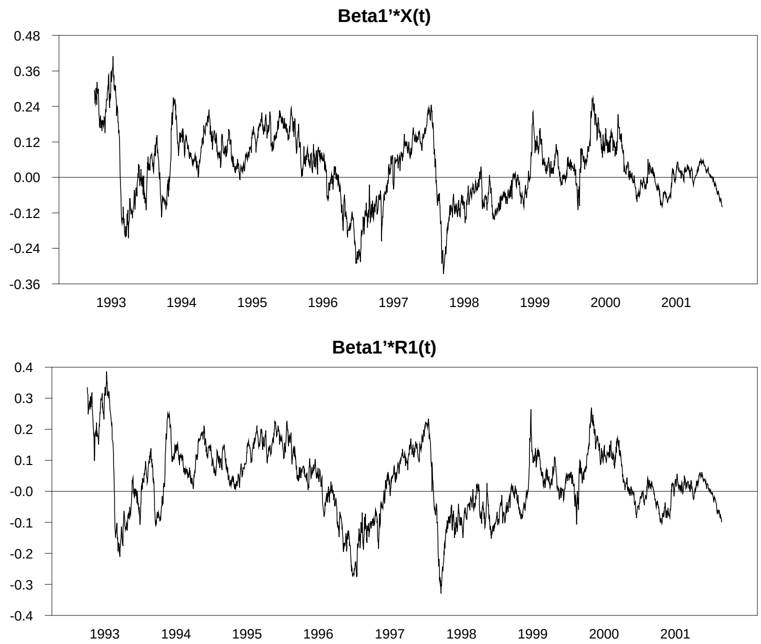
The KS column reports the Kolmogorov-Smirnov test for the null hypothesis that both correlations distributions are identical. The eigenvalues greater than one of each component in the pre-euro period are 6.60 and 1.16, while for the unique component in the post-euro period and the full sample are 7.37 and 8.13, respectively. The confidence interval, which is carried out using the Tchebycheff inequality with at least the 90% of confidence level, corresponds to the correlation after the introduction of the euro.

Table 8: Principal component structure of conditional correlations

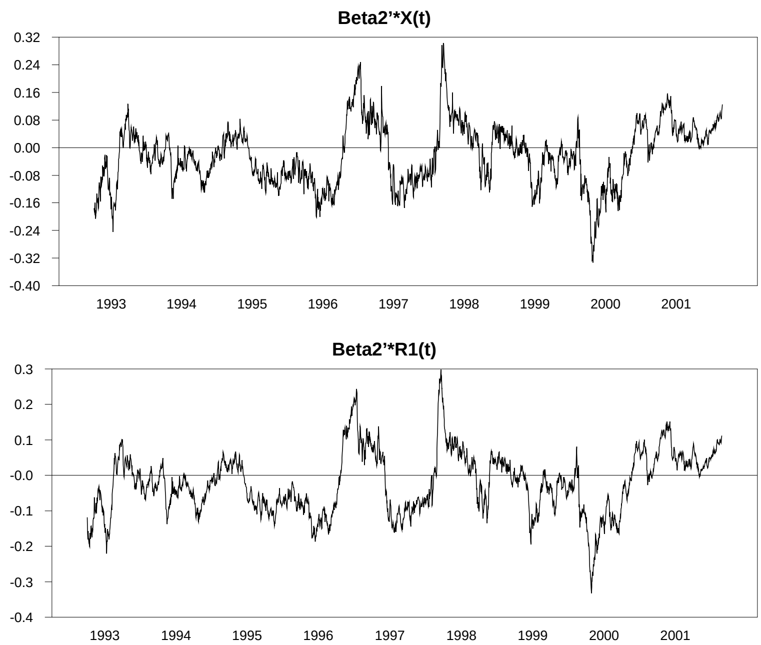
	pre-EMU		**post-EMU**	**Full sample**
Correlations	Component 1	Component 2	Component 1	Component 1
Germany-France	0.851	0.242	0.764	0.924
Germany-Italy	0.904	-0.154	0.865	0.945
Germany-Spain	0.817	0.205	0.870	0.923
Germany-United Kingdom	0.721	0.534	0.880	0.872
France-Italy	0.809	-0.444	0.880	0.920
France-Spain	0.807	-0.223	0.935	0.928
France-United Kingdom	0.791	0.385	0.862	0.849
Italy-Spain	0.751	-0.560	0.707	0.854
Italy-United Kingdom	0.884	-0.110	0.906	0.952
Spain-United Kingdom	0.908	0.163	0.891	0.835
Percentage of explained variance	66.6%	11.6%	73.7%	81.3%

The eigenvalues greater than one of each component in the pre-euro period are 6.60 and 1.16, while for the unique component in the post-euro period and the full sample are 7.37 and 8.13, respectively

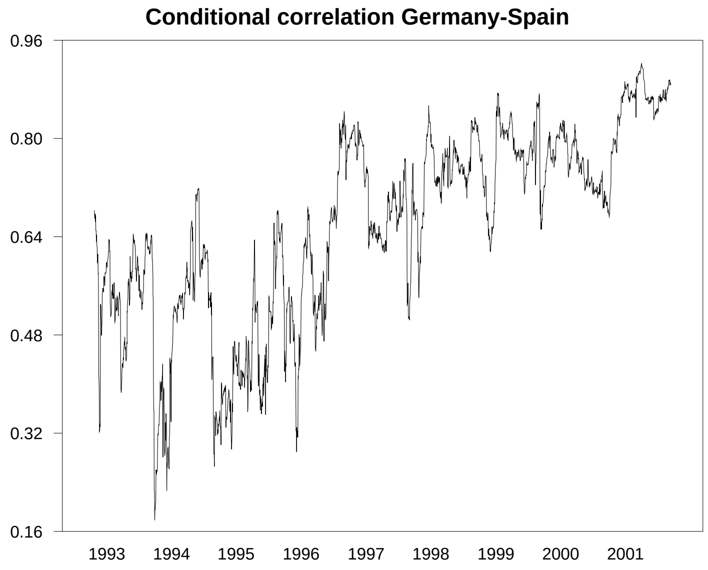
Figure 1: Cointegration vectors



(a) First vector



(b) Second vector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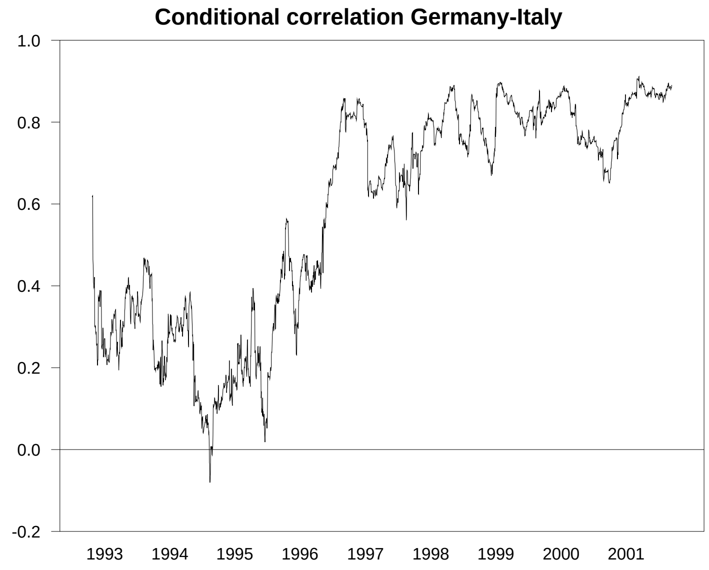
0.64
0.48
0.32
0.16
1993 1994 1995 1996 1997 1998 1999 2000 2001


Conditional correlation Germany-Italy
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0.4
0.2
0.0
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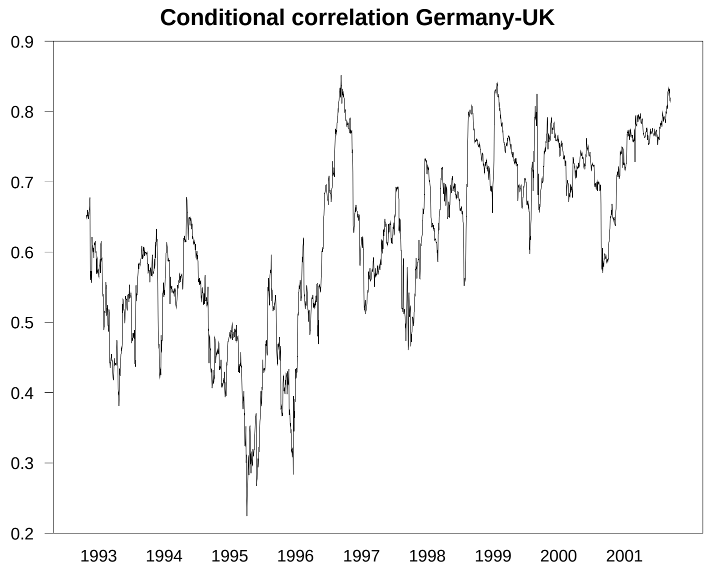
1993 1994 1995 1996 1997 1998 1999 2000 2001


Conditional correlation Germany-UK
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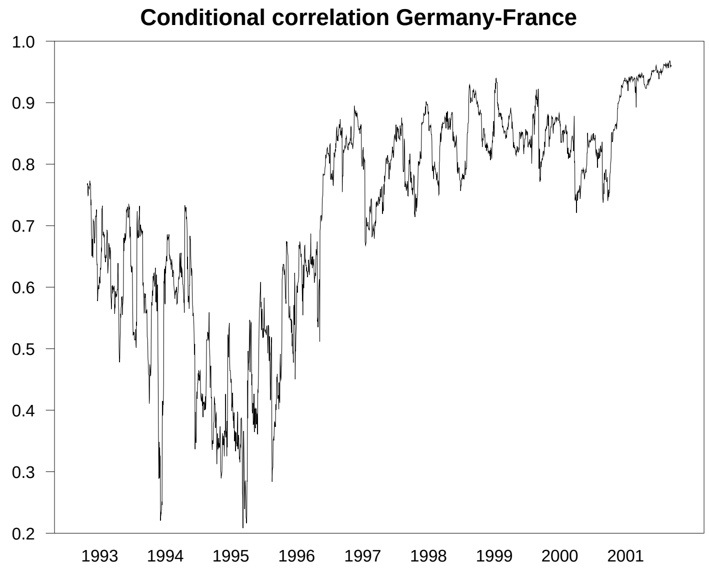
1993 1994 1995 1996 1997 1998 1999 2000 2001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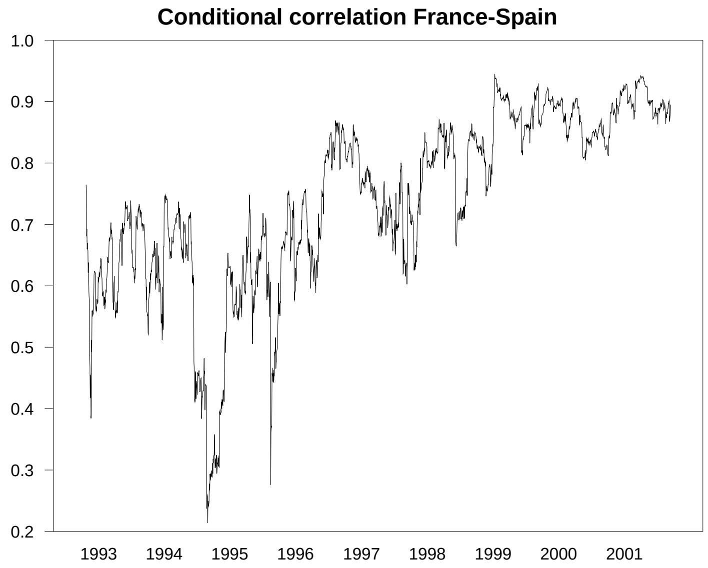
1993 1994 1995 1996 1997 1998 1999 2000 2001


Conditional correlation France-Spain
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1994
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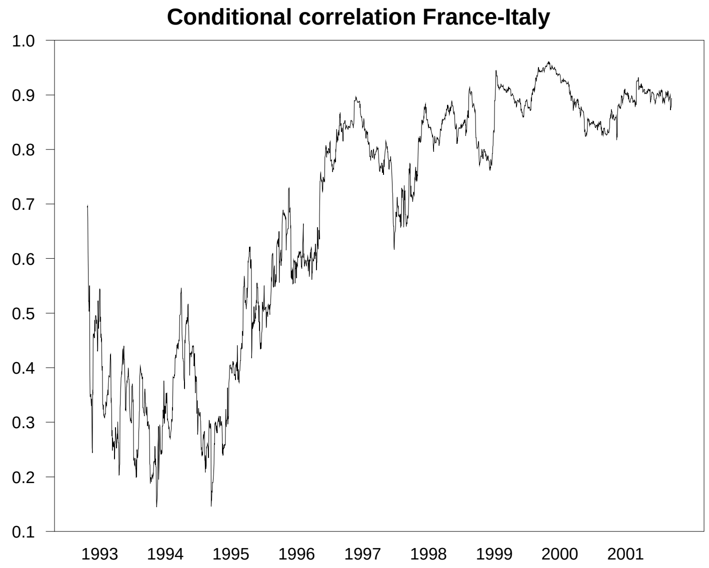
Conditional correlation France-Italy
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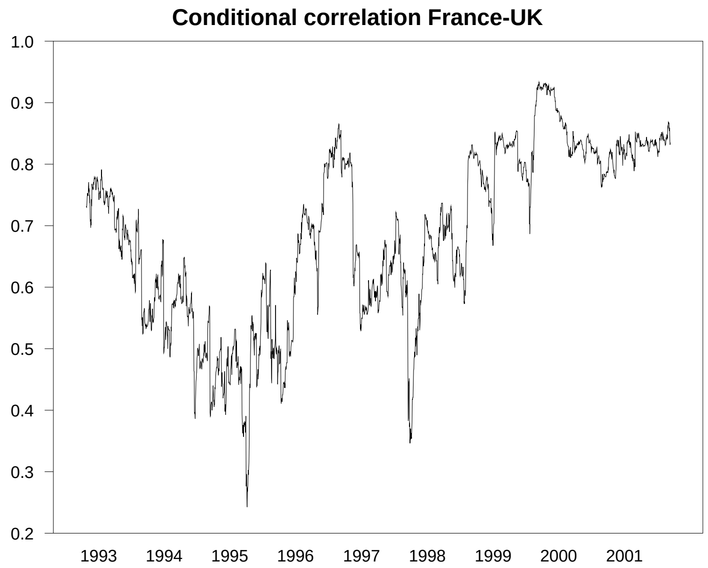
Conditional correlation France-UK
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0.8
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0.2
1993
1994
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1997
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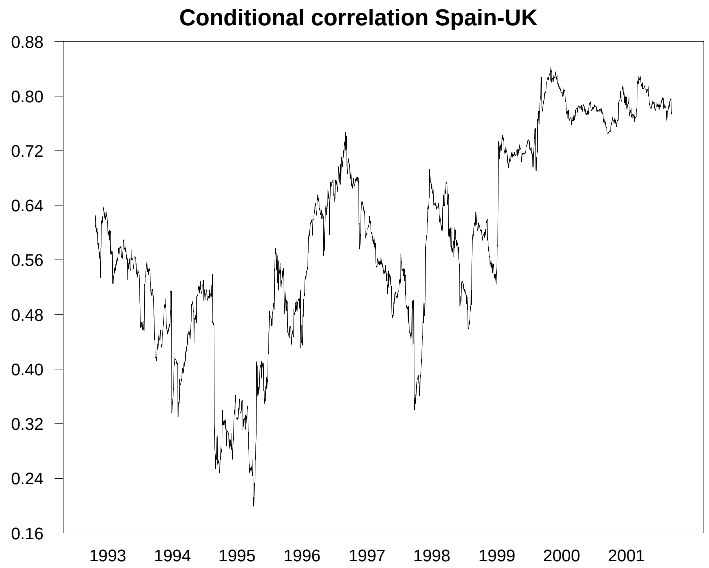
Conditional correlation Spain-UK
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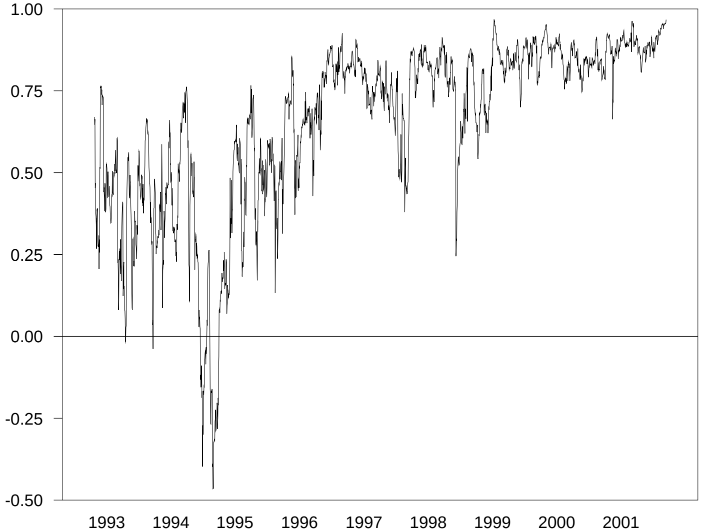
0.72
0.64
0.56
0.48
0.40
0.32
0.24
0.16
1993
1994
1995
1996
1997
1998
1999
2000
2001

Conditional correlation Italy-Spain



Conditional correlation Italy-UK