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The Impact of the Euro on Equity Markets

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Abstract

This paper investigates whether comovements between euro area equity returns at national and industry level changed after the introduction of the euro. By adopting a regression quantile-based methodology, we find that after 1999 the degree of comovements among euro area national equity markets was augmented. By explicitly controlling for the impact of global factors, we show that this result cannot be explained by recent worldwide trends. A more refined analysis based on an industry breakdown suggests that the increase in national index comovements is mainly driven by financial, industrial, and consumer services sectors.

I. Introduction

The convergence of nominal interest rates, inflation rates, and fiscal budget deficit and debt to gross domestic product (GDP) ratios fostered by the Maastricht Treaty paved the way to a common monetary policy within the euro zone economies. This convergence process culminated in the launch of the euro in January 1999. The introduction of the single currency has generated a large debate among researchers, policy makers, and market participants about the impact of the euro on European financial markets.[1]

A number of contributions have attempted to quantify this impact (see, e.g., Baele, Ferrando, Hördahl, Krylova, and Monnet (2004), Eiling, Gérard, and de Roon (2005), Hardouvelis, Malliaropulos, and Priestley (2006), (2007), and European Central Bank (ECB) (2008)). A common finding is that euro area money markets have become fully integrated, as shown by the convergence of overnight interest rates. Government bond markets are also characterized by a high degree

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[1]Strictly speaking, we cannot distinguish between the impact of the introduction of the single currency from the lagged effects of the structural reforms that have led to the common monetary policy in the euro zone.

of integration, exhibiting a pronounced yield convergence. As for equity markets, the impact of the euro is harder to assess, as equity returns are not directly comparable. In principle, firms' cash flows will be more exposed to common factors, as exchange rates cannot cushion any longer adverse shocks, business cycles have become more synchronized, and regulatory harmonization has steadily progressed. Ceteris paribus this should imply an increase in comovements of equity returns.

By analyzing return dynamics, this paper investigates whether there is evidence against this hypothesis. Using the regression quantile-based methodology developed by Cappiello, Gérard, and Manganelli (2005), we document an increase in comovements between the euro area equity returns both at national and industry levels.

A large body of literature has developed over the years to measure the codependence among financial asset returns (see, e.g., the surveys of Pericoli and Sbracia (2003) and Dungey, Fry, González-Hermosillo, and Martin (2005)). In essence, one can distinguish between two different approaches: modeling first and/or second moments of returns (see, e.g., King, Sentana, and Wadhwani (1994), Forbes and Rigobon (2002), Ciccarelli and Rebucci (2007), and Eiling and Gérard (2007)), and estimating the probability of coexceedance (see, e.g., Longin and Solnik (2001), Bae, Karolyi, and Stulz (2003), and Hartmann, Straetmans, and de Vries (2004)). Each of these methodologies suffers from several drawbacks: Generalized autoregressive conditional heteroskedastic (GARCH)-type approaches and dynamic correlation-based models assume that realizations in the upper and lower tails of the distribution are generated by the same process. Probability models generally analyze only single points of the support of the distribution and adopt a two-step estimation procedure, often without correcting the standard errors.

Our methodology offers a novel approach to study comovements and possesses a number of advantages. First, it is robust to departure from normality and the well-known heteroskedasticity problem that plagues naïve correlation measures (see, e.g., Forbes and Rigobon (2002)). Second, it permits testing for asymmetries in comovement in the positive and negative parts of the distribution. Third, it is suited to analyzing changes in correlations over the long run. Finally, being based on quantiles, it provides estimates of comovements robust to outliers, as opposed to conventional, average-based measures (Kim and White (2004), White, Kim, and Manganelli (2010)).

This paper estimates the probability of comovements between equity markets before and after the introduction of the euro. We find that after 1999 the degree of comovement among euro area economies has increased. Interestingly, comovements increase significantly also for country pairs involving the U.K., Denmark, and Sweden, which are members of the European Union (EU) but have not joined the euro area. This may be due to the strong economic ties of these economies with the euro area. Alternatively, this finding can indicate that global factors, rather than a common currency, may be responsible for the observed increase in comovements.

To distinguish between these alternative hypotheses, we introduce a variable that controls for the impact of global factors. This permits us to assess to what extent changes in comovements are driven by worldwide trends in addition to euro-specific factors. Our findings show that the increase in the degree of comovement is robust to the introduction of controls for global trends.

We also analyze to what extent comovements are driven by specific industries' dynamics (see, e.g., Carrieri, Errunza, and Sarkissian (2004), Sontchik (2004), Bekaert, Hodrick, and Zhang (2005), Eiling et al. (2005), and Cappiello, Lo Duca, and Maddaloni (2008)). Lack of changes in comovements at the national level may mask offsetting changes in comovements at the sectoral level. This occurs, for instance, when within the same national indices some industries exhibit a relatively high (and others a relatively low) level of correlation. It is also possible that greater comovements are due to the increasing importance of sectors more sensitive to common shocks (see, e.g., Griffin and Karolyi (1998), Brooks and Del Negro (2006)). We address these issues by reestimating the model with a sectoral breakdown. After controlling for global factors, we document that comovements increased in coincidence of the introduction of the euro in the financial, industrials, and consumer services sectors, while they remained largely unchanged in the health care and consumer goods industries.

The remainder of the paper is structured as follows. Section II explains the links between financial integration and comovements. Section III describes the econometric methodology. Section IV discusses the data. Section V describes a Monte Carlo simulation. Section VI presents the results, and Section VII concludes. The technical details about the econometrics underlying the paper are reported in the Appendixes.

II. Asset Return Correlation and Financial Integration

As is well recognized in financial economics, accurate measures of comovements are important for portfolio allocation, risk management, and assessment about financial contagion. Estimates of comovements are also becoming increasingly important to evaluate the degree of financial integration. Previous research has proposed at least two approaches to measuring time-varying market integration. One strand of the literature exploits the implication of asset pricing models: Markets are said to be integrated when only common risk factors are priced and (partially) segmented when local risk factors also determine equilibrium returns (see, e.g., Stulz (1981), Adler and Dumas (1983), Errunza and Losq (1985), and Flood and Rose (2005)). A second group of studies relate market and economic integration to a strengthening of the financial and real linkages between economies (see, inter alia, Dumas, Harvey, and Ruiz (2003)). Typically, studies in the first group are highly parameterized and require sophisticated asset pricing tests (examples are given by Bekaert and Harvey (1995), (1997), Rockinger and Urga (2001), Gérard, Thanyalakpark, and Batten (2003), Carrieri et al. (2004), Hardouvelis et al. (2006), (2007), and Cappiello et al. (2008)). Estimates of the second group, instead, are usually conducted by investigating changes in comovements across countries between selected financial asset returns (see, e.g., Dumas et al. (2003), Aydemir (2005)). A possible problem inherent in these two approaches is that the choice of the asset pricing or more generally the economic model may affect the final results.

In two related papers, Cappiello, Gérard, Kadareja, and Manganelli (2006) and Eiling and Gérard (2007) show that measures of comovements are linked to indicators of financial integration. The two studies measure financial integration exploiting the implications derived from a factor model: As long as the share of national firms' returns volatility is increasingly explained by common rather than local factors, the degree of integration augments. Advantageously, both approaches do not require the specification of common (and local) factors and, importantly, address the issue of time-varying volatility.

As shown by Cappiello et al. (2006), there is a relationship between integration and standard correlation measures. The relationship is derived from a model for returns that distinguishes between common and idiosyncratic factors. Progress in integration is associated with an increase in the proportion of returns' variance explained by the common factors vis-à-vis country-specific factors.

This reflects the intuition that, as a country moves from being closed to an open status, the impact of foreign factors on domestic firms' cash flows increases. Hence the removal of trade barriers and the elimination of exchange rate risk within a region should be accompanied by an increase in comovements of firms' returns. In short, increased comovements in financial asset returns are consistent with greater integration and economic interdependence.

To formalize this intuition, we model returns in a national market as follows:

$$r_{i,t} \quad = \quad \omega_{ij,t} G_{ij,t} + e_{i,t}, \qquad \forall i \text{ and } j, \tag{1}$$

where $r_{i,t}$ is the return on market i, $\omega_{ij,t}$ the exposure at time t of market i to the common factor $G_{ij,t}$, and $e_{i,t}$ the idiosyncratic risk of market i assumed to be orthogonal to the common factor and to asset j idiosyncratic risk.[2] The sufficient set of statistics for the factor model (1) can be summarized as follows: $\mathrm{E}(G_{ij,t})=0$ $\forall t$, $\mathrm{E}(G^2_{ij,t})=\sigma^2_{G_{ij,t}}$, $\mathrm{E}(e_{i,t})=\mathrm{E}(e_{j,t})=0$ $\forall t$, $\mathrm{E}(e^2_{i,t})=\sigma^2_{e_{i,t}}$, $\mathrm{E}(e^2_{j,t})=\sigma^2_{e_{j,t}}$, $\mathrm{E}(e_{i,t}, e_{i,s})=0$ $\forall t \neq s$, $\mathrm{E}(e_{i,t}, e_{j,s}) = 0$ $\forall i \neq j$ and $\forall t$ and s, $\mathrm{E}(e_{i,t}, G_{ij,t}) = \mathrm{E}(e_{j,t}, G_{ij,t}) = 0$ $\forall t$.

It is possible, in principle, to explain the idiosyncratic risk in terms of local factors (i.e., $e_{i,t} = \sum_{k=1}^{K} \gamma_{k,t} F_{k,t} + \varepsilon_{i,t}$). From an asset pricing perspective, we can say that markets are perfectly integrated if only the common factor is priced (i.e., $\omega_{ij,t} \neq 0$ and $\gamma_{k,t} = 0$ for all k). On the other hand, markets would be perfectly segmented if $\omega_{ij,t} = 0$.

The variance of country i's returns can be decomposed as $\sigma^2_{r_{i,t}} = \omega^2_{ij,t}\sigma^2_{G_{ij,t}} + \sigma^2_{e_{i,t}}$. The share of volatility explained by the common factor is $\phi_{ij,t} \equiv (\omega_{ij,t}\sigma_{G_{ij,t}})/\sigma_{r_{i,t}}$. Consistently with this discussion, we adopt the following measure of integration between markets i and j:

$$\Phi_{ij,t} \quad \equiv \quad \phi_{ij,t}\phi_{ji,t}. \tag{2}$$

If markets are perfectly segmented, the volatility explained by the common factor is equal to 0 and therefore $\Phi_{ij,t} = 0$ (because $\phi_{ij,t} = 0$ and/or $\phi_{ji,t} = 0$). On the other hand, if markets are perfectly integrated, most of the source of variation

[2] $G_{ij,t}$ includes all the common components specific to markets i and j. Notice that different market pairs may have distinct common factors.

will come from the common factor, implying a strictly positive $\Phi_{ij,t}$.[3] In general, for a given level of idiosyncratic volatility, higher values of $\Phi_{ij,t}$ imply a higher degree of integration.

As suggested by Cappiello et al. (2006), it is straightforward to show that the measure of integration (2) coincides with the linear correlation coefficient:

$$\rho_{ij,t} = \frac{\sigma_{r_i r_j,t}}{\sigma_{r_{i,t}}\sigma_{r_{j,t}}} = \Phi_{ij,t}, \qquad \forall i,j \text{ and } i \neq j, \tag{3}$$

where $\sigma_{r_i r_j,t} = \omega_{ij,t}\omega_{ji,t}\sigma^2_{G_{ij,t}}$. If market i and j become more integrated, the correlation between returns on an asset in market i and j will increase.

III. The Empirical Methodology

To assess whether the degree of integration between two markets varies after the introduction of the euro, it is necessary to test for changes in correlations. These tests need to account, inter alia, for time variation in the moments of the returns distribution and departure from normality. Since changes in volatilities before and after the introduction of the euro could result in an estimation bias, a simple comparison between correlations over the two periods could lead to a spurious outcome. To solve this issue, we use a modeling strategy based on the "comovement box" of Cappiello et al. (2005). The approach is robust to heteroskedasticity, is semiparametric, does not need any assumption on the distribution of returns, and provides a direct test for changes in correlation before and after the introduction of the euro. Moreover, this methodology permits us to control for (global) factors that may, in fact, bear the ultimate responsibility for comovements between assets.

GARCH-type models could constitute an alternative empirical methodology to the comovement box. GARCH processes are also robust to volatility changes. However, differently from the comovement box approach, they are fully parametric and estimate correlation at a relatively high frequency.

Eiling and Gérard (2007) propose a nonparametric measure of instantaneous correlation based on cross-sectional dispersion and realized variance. The resulting time series of correlations are then treated as observable, which permits us to test for trends and structural breaks. This approach rests on the assumption that all the countries in a region have the same factor exposure and that the idiosyncratic country risk is diversified in the (equal-weighted) regional portfolios, which requires a large cross section. Our framework, instead, does not rely on the assumption of equal factor exposure and allows us to analyze comovements between any country pairs (as opposed to regions). Furthermore, being based on quantiles, our measures of comovements are robust to possible outliers, which are typical of financial time series.

[3]We assume that the factor loading coefficients of the common factor are positive. An analogous but opposite conclusion would hold if $\text{sign}(\beta_{ij,t}) \neq \text{sign}(\beta_{ji,t})$.

A. The Comovement Box

Let $\{r_{i,t}\}_{t=1}^{T}$ and $\{r_{j,t}\}_{t=1}^{T}$ denote the time series returns of two different markets. Let $q_{\theta,t}^{i}$ be the time t θ-quantile of the conditional distribution of $r_{i,t}$. Analogously, for $r_{j,t}$, we define $q_{\theta,t}^{j}$.

Denote the conditional cumulative joint distribution of the two return series by $F_t(r_i, r_j)$. Define $F_t^{-}(r_i|r_j) \equiv \Pr(r_{i,t} \leq r_i \mid r_{j,t} \leq r_j)$ and $F_t^{+}(r_i|r_j) \equiv \Pr(r_{i,t} \geq r_i \mid r_{j,t} \geq r_j)$. Our basic tool of analysis is the following conditional probability:

$$(4)\quad p_t(\theta) \quad \equiv \quad \begin{cases} F_t^{-}\left(q_{\theta,t}^{i}|q_{\theta,t}^{j}\right) \equiv \Pr\left(r_{i,t} \leq q_{\theta,t}^{i} \mid r_{j,t} \leq q_{\theta,t}^{j}\right), & \text{if } \theta \leq 0.5 \\ F_t^{+}\left(q_{\theta,t}^{i}|q_{\theta,t}^{j}\right) \equiv \Pr\left(r_{i,t} \geq q_{\theta,t}^{i} \mid r_{j,t} \geq q_{\theta,t}^{j}\right), & \text{if } \theta > 0.5 \end{cases}.$$

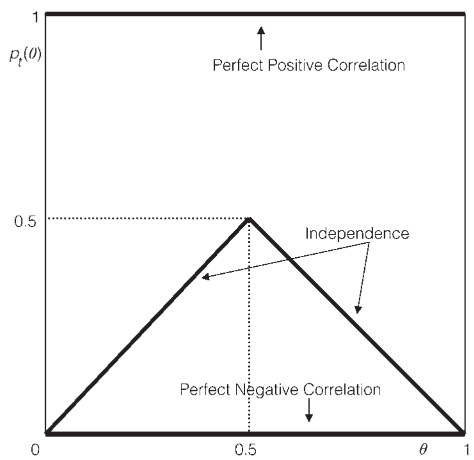
This conditional probability represents an effective way to summarizes the characteristics of $F_t(r_i, r_j)$. For each quantile θ, $p_t(\theta)$ measures the probability that, at time t, the return on market i will fall below (or above) its θ-quantile, conditional on the same event occurring in market j.

The characteristics of $p_t(\theta)$ can be conveniently analyzed in what we call the comovement box (see Figure 1). The comovement box is a square with unit side, where $p_t(\theta)$ is plotted against θ. The shape of $p_t(\theta)$ will generally depend on the characteristics of the joint distribution of the time-series returns $r_{i,t}$ and $r_{j,t}$, and therefore for generic distributions it can be derived only by numerical simulation. There are, however, three important special cases that do not require any simulation: i) perfect positive correlation, ii) independence, and iii) perfect negative correlation. If two markets are independent, which implies $\rho_{ij,t} = 0\ \forall t$, $p_t(\theta)$ will be piecewise linear, with slope equal to 1, for $\theta \in (0, 0.5)$, and slope

FIGURE 1

The Comovement Box

Figure 1 plots the probability that an asset return $r_{i,t}$ falls below (above) its θ-quantile conditional on another asset return $r_{j,t}$ being below (above) its θ-quantile, for $\theta < 0.5$ ($\theta \geq 0.5$). The cases of perfect positive correlation, independence, and perfect negative correlation are represented.



equal to -1, for $\theta \in (0.5, 1)$. When there is perfect positive correlation between $r_{i,t}$ and $r_{j,t}$ (i.e., $\rho_{ij,t} = 1\ \forall t$), $p_t(\theta)$ is a flat line that takes on unit value. Under this scenario, the two markets essentially reduce to one. The polar case occurs for perfect negative correlation (i.e., $\rho_{ij,t} = -1\ \forall t$). In this case $p_t(\theta)$ is always equal to 0: When the realization of $r_{j,t}$ is in the lower tail of its distribution, the realization of $r_{i,t}$ is always in the upper tail of its own distribution and conversely (for a more analytical description of the model, see the Appendix).

This discussion suggests that the shape of $p_t(\theta)$ provides key insights about the dependence between two asset returns $r_{i,t}$ and $r_{j,t}$. In general, the higher $p_t(\theta)$, the higher the codependence between the two time-series returns.

While $p_t(\theta)$ can be used to measure the dependence between different markets, the interest of the researcher often lies in testing whether this dependence has changed over time. Market integration is an important case in point. If increased integration can be associated to stronger comovements between markets, one can test for changes in integration by testing if the conditional probability of comovements between two markets increases after institutional changes fostering greater openness and integration.

The framework of the comovement box can be used to formalize this intuition. Let $p^B(\theta) \equiv B^{-1}\sum_{t<\tau} p_t(\theta)$ and $p^A(\theta) \equiv A^{-1}\sum_{t\geq\tau} p_t(\theta)$, where B and A denote the number of observations before and after a certain threshold date τ, respectively. We adopt the following working definition of increased integration:

Definition 1. Integration increases if $\delta(0,1) = \int_0^1 [p^A(\theta) - p^B(\theta)]d\theta > 0$.

Here, $\delta(0,1)$ measures the area between the average conditional probabilities $p^A(\theta)$ and $p^B(\theta)$.

Constructing the comovement box and testing for differences in the probability of comovements requires several steps. First, we estimate the univariate quantiles associated to the return series of interest, using the conditional autoregressive value at risk (CAViaR) model by Engle and Manganelli (2004). Second, we construct, for each series and for each quantile, indicator variables that are equal to 1 if the observed return is lower than this quantile and 0 otherwise. Finally, we regress the θ-quantile indicator variable of returns on market i on the θ-quantile indicator variable of returns on market j, interacted with time dummy variables that identify periods of greater integration. These regression coefficients will provide a direct estimate of the conditional probability of comovements and of their changes across regimes.

The average probability of comovement can be estimated by running the following regression:

$$I_t^i(\hat{\beta}_{\theta,i}) \cdot I_t^j(\hat{\beta}_{\theta,j}) = \alpha_{0,\theta} + \alpha_{1,\theta} S_{1,t} + \eta_t, \tag{5}$$

where, for each θ-quantile (with $\theta \in (0,1)$), $I_t^i(\hat{\beta}_{\theta,i}) \equiv I\left(r_{i,t} \leq q_t^i(\hat{\beta}_{\theta,i})\right)$ denotes an indicator function that takes on the value 1 if the expression in parenthesis is true and 0 otherwise, $q_t^i(\hat{\beta}_{\theta,i})$ represents the estimated quantiles, $\hat{\beta}_{\theta,i}$ is a p-vector of parameters to be estimated, and $S_{1,t}$ is the dummy variable for the test period A, $t > \tau$.[4]

[4]The "hat" denotes estimated coefficients.

Cappiello et al. (2005) show that the ordinary least squares (OLS) estimators of regression (5) are asymptotically consistent estimators of the average probability of comovement in the two periods and provide estimators for their standard errors:

$$\begin{aligned} \widehat{\alpha}_{0,\theta} &\xrightarrow{p} \mathrm{E}[p_t(\theta)|\text{ period } B] &\equiv& \; p^B(\theta), \\ \widehat{\alpha}_{0,\theta} + \widehat{\alpha}_{1,\theta} &\xrightarrow{p} \mathrm{E}[p_t(\theta)|\text{ period } A] &\equiv& \; p^A(\theta). \end{aligned} \tag{6}$$

Here, $\widehat{\alpha}_{0,\theta}$ is the parameter associated with the constant and, as such, it converges to the average probabilities in the benchmark period B. Similarly, since $\widehat{\alpha}_{1,\theta}$ is the coefficient of $S_{1,t}$, the sum of $\widehat{\alpha}_{0,\theta} + \widehat{\alpha}_{1,\theta}$ converges to the average probability of comovement in the test period A. Testing for an increase in the probability of comovement across two periods is equivalent to testing for the null that $\widehat{\alpha}_{1,\theta}$ is equal to 0. Indeed, it is only when $\widehat{\alpha}_{1,\theta} = 0$ that the two probabilities coincide. If $\widehat{\alpha}_{1,\theta} > 0$, the conditional probability during the test period will be higher than the probability during the benchmark period.

Rigorous joint tests for integration that follow from Definition 1 can be constructed as follows:

$$\begin{aligned} \widehat{\delta}\left(\underline{\theta}, \overline{\theta}\right) &= (\#\theta)^{-1} \sum_{\theta \in [\underline{\theta}, \bar{\theta}]} [\widehat{p}^A(\theta) - \widehat{p}^B(\theta)] \\ &\equiv (\#\theta)^{-1} \sum_{\theta \in [\underline{\theta}, \bar{\theta}]} (\widehat{\alpha}_{0,\theta} + \widehat{\alpha}_{1,\theta}) - \widehat{\alpha}_{0,\theta} \\ &= (\#\theta)^{-1} \sum_{\theta \in [\underline{\theta}, \bar{\theta}]} \widehat{\alpha}_{1,\theta}, \end{aligned} \tag{7}$$

where $\#\theta$ denotes the number of addends in the sum (see the Appendix for how to obtain the asymptotic distribution of this statistic).

We estimate the time-varying quantiles of the returns, $r_{i,t}$, using the following CAViaR specification:

$$\begin{aligned} q_t^i(\beta_{\theta,i}) \;=\; & \beta_{0,\theta,i} + \beta_{1,\theta,i} S_{1,t} + \beta_{2,\theta,i} r_{i,t-1} + \beta_{3,\theta,i} q_{t-1}^i(\beta_{\theta,i}) \\ & - \beta_{2,\theta,i}\beta_{3,\theta,i} r_{i,t-2} + \beta_{4,\theta,i}\left| r_{i,t-1} \right|, \end{aligned} \tag{8}$$

where $\beta_{\theta,i} \equiv [\beta_{0,\theta,i}, \beta_{1,\theta,i}, \ldots, \beta_{4,\theta,i}]'$.

This parameterization is robust to the presence of autocorrelation in our sample returns. We add the dummy variable, $S_{1,t}$, to the CAViaR specification to ensure that we have exactly the same proportion of quantile exceptions in both subperiods. This will guarantee that $\Pr(r_{i,t} \leq q_t^i(\beta_{\theta,i}^0) | r_{j,t} \leq q_t^j(\beta_{\theta,j}^0)) = \Pr(r_{j,t} \leq q_t^j(\beta_{\theta,j}^0) | r_{i,t} \leq q_t^i(\beta_{\theta,i}^0))$ will be satisfied.[5] For each market we estimate model (8) for 19 quantile probabilities ranging from 5% to 95%.

[5]Asymptotically, correct specification would imply the same number of exceedances in both periods. However, in finite samples this need not to be the case. Failure to account for this fact would affect the estimation of the conditional probabilities.

B. Regional versus Global Factors

In the factor model described by equation (1), returns on a national market are a function of common and, possibly, country-specific factors. In principle the common factor can be divided into two distinct components: i) a regional and ii) a world factor. This decomposition permits us to evaluate an increase in comovements that, on the one hand, is due to the introduction of the euro, and, on the other hand, is driven by global factors. Under this assumption, equation (1) can be rewritten as

$$r_{i,t} = \omega^R_{ij,t} G^R_{ij,t} + \omega^W_{i,t} G^W_t + \zeta_{i,t}, \quad \forall i \text{ and } j, \tag{9}$$

where $\omega^R_{ij,t}$ and $\omega^W_{i,t}$ represent the exposure at time t to the regional and world factors $G^R_{ij,t}$ and G^W_t, respectively, and $\zeta_{i,t}$ the idiosyncratic risk, which is assumed to be orthogonal to both $G^R_{ij,t}$ and G^W_t, as well as to any other asset j idiosyncratic risk. The sufficient set of statistics for the two-factor model (9) can be summarized as follows: $\mathrm{E}(G^R_{ij,t}) = 0\ \forall t$, $\mathrm{E}[(G^R_{ij,t})^2] = \sigma^2_{G^R_{ij,t}}$, $\mathrm{E}(G^W_t) = 0\ \forall t$, $\mathrm{E}[(G^W_t)^2] = \sigma^2_{G^W_t}$, $\mathrm{E}(\zeta_{i,t}) = \mathrm{E}(\zeta_{j,t}) = 0\ \forall t$, $\mathrm{E}(\zeta^2_{i,t}) = \sigma^2_{\zeta_{i,t}}$, $\mathrm{E}(\zeta^2_{j,t}) = \sigma^2_{\zeta_{j,t}}$, $\mathrm{E}(\zeta_{i,t}, \zeta_{i,s}) = 0$ $\forall t \neq s$, $\mathrm{E}(\zeta_{i,t}, \zeta_{j,s}) = 0\ \forall i \neq j$ and $\forall t$ and s, $\mathrm{E}(\zeta_{i,t}, G^R_{ij,t}) = \mathrm{E}(\zeta_{j,t}, G^R_{ij,t}) = 0$ as well as $\mathrm{E}(\zeta_{i,t}, G^W_t) = \mathrm{E}(\zeta_{j,t}, G^W_t) = 0\ \forall t$, and finally $\mathrm{E}(G^R_{ij,t}, G^W_t) = 0$.

Following the reasoning of Section II, we can define the share of volatility explained by the regional and global factors as

$$\phi^R_{ij,t} \equiv \frac{\omega^R_{ij,t}\sigma_{G^R_{ij,t}}}{\sigma_{r_{i,t}}} \tag{10}$$

and

$$\phi^W_{i,t} \equiv \frac{\omega^W_{i,t}\sigma_{G^W_t}}{\sigma_{r_{i,t}}}. \tag{11}$$

In this case integration between markets i and j explained by regional factors is measured by

$$\Phi^R_{ij,t} \equiv \phi^R_{ij,t}\phi^R_{ji,t}, \tag{12}$$

and, analogously, the share of integration due to the global factor is given by

$$\Phi^W_{ij,t} \equiv \phi^W_{i,t}\phi^W_{j,t}. \tag{13}$$

The linear correlation measure is now equal to the sum of equations (12) and (13):

$$\rho_{ij,t} = \Phi^R_{ij,t} + \Phi^W_{ij,t}. \tag{14}$$

In the next section we describe how we take into account global factors in the context of the comovement box methodology.

C. The Comovement Box with a Global Factor

The comovement box methodology discussed in Section III.A can include, in addition to the temporal dummy variable $S_{1,t}$, other dummy variables. While the coefficient associated with the temporal dummy variable indicates whether comovements between two asset returns change after a certain time, other dummy variables may accommodate the impact on codependences due to other factors. Following the framework proposed by Cappiello et al. (2006), we introduce a new dummy variable, $S_{2,t}$, which controls for global factors that may also be responsible for changes in integration. We take as a control variable the correlation between average returns on the equities' market pair under study and on a world equity market index excluding the euro area.[6] We compute correlations as an exponentially weighted moving average (EWMA) with decay coefficient equal to 0.94. Next we construct $S_{2,t}$ so that it takes on the value of 1 when the underlying correlation variable is larger than a certain threshold ρ^* and 0 otherwise $\left(\text{i.e., } S_{2,t} \equiv I\left(\rho_t^{\text{EWMA}} > \rho^*\right)\right)$. Here, ρ^* is chosen so that the two dummy variables $S_{1,t}$ and $S_{2,t}$ have the same number of ones.[7] In this way we can control how much of the change in correlation after the introduction of the single currency is due to the global correlation factor.

When the $S_{2,t}$ dummy variable is introduced, equation (5) reads as follows:

$$I_t^i(\hat{\beta}_{\theta,i}) \cdot I_t^j(\hat{\beta}_{\theta,j}) \quad = \quad \alpha_{0,\theta} + \alpha_{1,\theta} S_{1,t} + \alpha_{2,\theta} S_{2,t} + \upsilon_t. \tag{15}$$

In line with equation (15), four possible cases arise: i) the comovements over the benchmark period when the global factor correlation is low, $p^{\text{BL}}(\theta)$; ii) the comovements over the test period when the global factor correlation is low, $p^{\text{AL}}(\theta)$; iii) the comovements over the benchmark period when the global factor correlation is high, $p^{\text{BH}}(\theta)$; and iv) the comovements over the test period when the global factor correlation is high, $p^{\text{AH}}(\theta)$. It can be shown that OLS estimators of equation (15) enjoy the following asymptotic properties:

$$\begin{aligned}
\widehat{a}_{0,\theta} &\xrightarrow{p} \text{E}[p_t(\theta)|\text{ period } B \text{ and low global correlation}] &\equiv\quad& p^{\text{BL}}(\theta),\\
\widehat{a}_{0,\theta} + \widehat{a}_{1,\theta} &\xrightarrow{p} \text{E}[p_t(\theta)|\text{ period } A \text{ and low global correlation}] &\equiv\quad& p^{\text{AL}}(\theta),\\
\widehat{a}_{0,\theta} + \widehat{a}_{2,\theta} &\xrightarrow{p} \text{E}[p_t(\theta)|\text{ period } B \text{ and high global correlation}] &\equiv\quad& p^{\text{BH}}(\theta),\\
\widehat{a}_{0,\theta} + \widehat{a}_{1,\theta} + \widehat{a}_{2,\theta} &\xrightarrow{p} \text{E}[p_t(\theta)|\text{ period } A \text{ and high global correlation}] &\equiv\quad& p^{\text{AH}}(\theta).
\end{aligned}$$

Therefore $\widehat{a}_{1,\theta}$ measures the changes in equity market comovements after the introduction of the euro, after controlling for global factors. Standard errors for the estimated parameters can be computed as suggested by Cappiello et al. (2005). Similarly to the case when the dummy variable $S_{2,t}$ was not included, we are interested in testing whether $\widehat{a}_{1,\theta}$ is significantly different from 0. When this

[6]Since our interest lies in the evolution over time of correlations, we use simple averages of the assets' returns, which will next provide the time series to calculate EWMA correlations.

[7]If the number of times $S_{2,t}$ is equal to 1 were quite limited (and significantly smaller than the number of times $S_{1,t}$ is equal to 1), the control dummy variable would not possess sufficient explanatory power.

occurs, integration between returns on assets' market pairs can be attributed also to region-specific factors. Tests for region-specific integration are constructed in line with equation (7):

$$\begin{aligned}(16)\qquad \widehat{\xi}\left(\underline{\theta},\overline{\theta}\right) &= (\#\theta)^{-1}\sum_{\theta\in\left[\underline{\theta},\bar{\theta}\right]}\left[\widehat{p}^{\mathrm{AL}}(\theta)-\widehat{p}^{\mathrm{BL}}(\theta)\right]\\ &= (\#\theta)^{-1}\sum_{\theta\in\left[\underline{\theta},\bar{\theta}\right]}\left[\widehat{p}^{\mathrm{AH}}(\theta)-\widehat{p}^{\mathrm{BH}}(\theta)\right]\\ &= (\#\theta)^{-1}\sum_{\theta\in\left[\underline{\theta},\bar{\theta}\right]}\widehat{\alpha}_{1,\theta}.\end{aligned}$$

By the same token, it is possible to compute joint tests for the control variable:

$$\begin{aligned}(17)\qquad \widehat{\psi}\left(\underline{\theta},\overline{\theta}\right) &= (\#\theta)^{-1}\sum_{\theta\in\left[\underline{\theta},\bar{\theta}\right]}\left[\widehat{p}^{\mathrm{BH}}(\theta)-\widehat{p}^{\mathrm{BL}}(\theta)\right]\\ &= (\#\theta)^{-1}\sum_{\theta\in\left[\underline{\theta},\bar{\theta}\right]}\left[\widehat{p}^{\mathrm{AH}}(\theta)-\widehat{p}^{\mathrm{AL}}(\theta)\right]\\ &= (\#\theta)^{-1}\sum_{\theta\in\left[\underline{\theta},\bar{\theta}\right]}\widehat{\alpha}_{2,\theta},\end{aligned}$$

where $\#\theta$ denotes the number of addends in the sum.

Returns' conditional quantiles are estimated employing a CAViaR specification similar to that of equation (8), but with the inclusion of the new dummy variable $S_{2,t}$:

$$\begin{aligned}(18)\qquad q_t^i(\beta_{\theta,i}) &= \beta_{0,\theta,i}+\beta_{1,\theta,i}S_{1,t}+\beta_{2,\theta,i}S_{2,t}+\beta_{3,\theta,i}r_{i,t-1}+\beta_{4,\theta,i}q_{t-1}^i(\beta_{\theta,i})\\ &\quad -\beta_{3,\theta,i}\beta_{4,\theta,i}r_{i,t-2}+\beta_{5,\theta,i}\left|r_{i,t-1}\right|,\end{aligned}$$

where $\beta_{\theta,i}\equiv[\beta_{0,\theta,i},\beta_{1,\theta,i},\ldots,\beta_{5,\theta,i}]'$.

IV. Data

We analyze returns on equity markets for country and sector indices. Country indices include: i) euro area countries Germany, France, Italy, the Netherlands, Spain, Austria, Belgium, Finland, Greece, Ireland, and Portugal; and ii) EU non-euro area economies Denmark, Sweden, and the U.K. The sample covers the period from March 5, 1987 to January 13, 2008.[8] Japan and the U.S. are also used in the analysis to compute the global factor indicator.

[8]Notice that the sample starts at later dates for some national and sector equity indices. In particular, observations for Finland and Portugal national equity indices commence on March 31, 1988 and January 4, 1990, respectively. Observations for: i) the Finnish, Greek, and Portuguese industrial sectors start on March 31, 1988; January 7, 1988; and January 4, 1990, respectively; ii) the Finnish, Greek, and Portuguese financial sectors start on March 31, 1988; January 4, 1990; and January 4, 1990, respectively; iii) the Finnish, Greek, Portuguese, and Swedish health sectors start on March 31, 1988; January 4, 1990; April 21, 1988; and July 18, 1991, respectively; iv) the Austrian, Belgian, Finnish, Greek, and Portuguese consumer goods sectors start on October 1, 1992; May 5, 1997; July 13, 1995; April 21, 1988; and January 4, 1990, respectively; v) the Austrian, Finnish, Greek, and Portuguese consumer services sectors start on June 16, 1988; March 31, 1988; July 14, 1994; and January 4, 1990, respectively. Data for the Danish consumer services sector are not available.

For each country, we analyze five sectors: financial, industrial (which we further divide into the subsectors construction and materials, and industrials goods and services), consumer goods (which we further divide into the subsectors automobile, food and beverages, and personal and household goods), consumer services, and health care.

We use Thomson Datastream indices at weekly frequency. Equity indices are market-value-weighted and include dividends. The use of weekly data reduces the asynchronicity effects due to different opening hours, national holidays, and administrative closures. Equity returns are continuously compounded.

Table 1 reports data summary statistics. As expected, country and sector equity index returns tend to be negatively skewed and leptokurtic. Nonnormality is confirmed by the Jarque-Bera (1987) test statistics. It is also worth noticing that for some countries the number of companies entering certain sectors is quite low (see the last column of Table 1).

V. Monte Carlo Simulation

Before discussing our empirical results, we study the finite sample properties of the comovement box methodology and the power of the associated tests. To this end, we perform a Monte Carlo experiment. We estimate the following model for French and German equity returns:

$$\boldsymbol{r}_t = \boldsymbol{\gamma}_0 + \boldsymbol{\gamma}_1 \boldsymbol{r}_{t-1} + \boldsymbol{\varepsilon}_t, \quad \boldsymbol{\varepsilon}_t \sim N(\mathbf{0}, \Sigma_t), \tag{19}$$

where $\boldsymbol{r}_t = [r_{i,t}, r_{j,t}]'$ is a 2×1 vector of equity returns and Σ_t the associated covariance matrix. Here, Σ_t, in turn, is modeled as a bivariate GARCH process:

$$\begin{aligned} \Sigma_t &\equiv \begin{bmatrix} \sigma_{i,t}^2 & \sigma_{ij,t} \\ \sigma_{ji,t} & \sigma_{j,t}^2 \end{bmatrix}, \\ \sigma_{i,t} &= \delta_{i,0} + \delta_{i,1} S_{1,t} + \delta_{i,2} |r_{i,t-1}| + \delta_{i,3} \sigma_{i,t-1}, \\ \sigma_{j,t} &= \delta_{j,0} + \delta_{j,1} S_{1,t} + \delta_{j,2} |r_{j,t-1}| + \delta_{j,3} \sigma_{j,t-1}, \\ \sigma_{ij,t} &= \delta_{ij,0} + \delta_{ij,1} r_{i,t-1} r_{j,t-1} + \delta_{ij,2} \sigma_{ij,t-1}. \end{aligned} \tag{20}$$

We estimate the return equations and the associated second moments via maximum likelihood. Differently from standard bivariate GARCH processes, we model the evolution of standard deviations instead of the evolution of the variances. It is easy to check that such a data generating process (DGP) would generate the CAViaR model described in equation (8).

Given the estimates of the bivariate GARCH, we generate two vectors of simulated data using expressions (19) and (20). The dimension of these vectors is the same as our sample data. Next, we estimate the comovement box for these two vectors. We repeat this procedure 100 times, which results in 100 probabilities of comovements before and after the introduction of the euro. Then we compute the averages of these two groups of probabilities and obtain the comovements over the benchmark and test periods. The Monte Carlo 95% confidence bands are computed as two times the standard deviations of the comovement box estimates. The results are reported in Figure 2, where we also plot the probability

TABLE 1

Summary Statistics

Table 1 reports summary statistics relative to weekly returns on national and sectoral equity market indices. The national equity indices refer to Germany (DE), France (FR), Italy (IT), the Netherlands (NL), Spain (ES), Austria (AT), Belgium (BE), Finland (FI), Greece (GR), Ireland (IE), Portugal (PT), Denmark (DK), Sweden (SE), the United Kingdom (UK), the United States (US), and Japan (JP). For Germany, France, Italy, the Netherlands, Spain, the U.K., the U.S., and Japan, 10 sectoral indices are considered: Industrial, Construction and Materials, Industrial Goods and Services, Financial, Consumer Goods, Automobile, Food and Beverages, Personal and Household Goods, Consumer Services, and Health industries. For most of the countries and industries the data set starts on March 5, 1987 and ends on January 31, 2008. Equity indices are from Thomson Datastream. For each return series, mean and standard deviation (SD) are annualized and in percentage. "Max" and "Min" represent the weekly maximum and minimum returns and are in percentages. "Skew" and "Kurt" stand for skewness and kurtosis, respectively, while "#Comp" denotes the number of companies included in the index. The Jarque-Bera (J-B) (1987) test for normality combines excess skewness and kurtosis and is asymptotically distributed as χ^2_m with $m = 2$ degrees of freedom.

Country	Mean	Max	Min	SD	Skew	Kurt	J-B	#Comp
Panel A. Total Indices								
DE	8.33	12.40	–15.54	18.08	–0.81	7.26	935	250
FR	9.69	13.14	–10.59	17.99	–0.38	5.59	328	250
IT	6.45	11.06	–14.87	20.13	–0.37	5.14	231	160
NL	10.40	10.77	–15.69	16.52	–0.98	7.99	1,293	130
ES	11.40	9.91	–27.11	19.45	–1.35	14.40	6,185	120
AT	10.83	16.33	–17.66	17.67	–0.57	10.75	2,759	50
BE	9.67	11.65	–18.93	16.01	–0.90	10.43	2,632	90
FI	12.89	15.13	–23.64	28.58	–0.62	7.01	749	50
GR	16.57	20.25	–17.35	27.73	0.38	6.51	499	50
IE	11.18	9.37	–24.42	20.02	–1.21	11.23	3,311	50
PT	7.86	11.51	–17.24	16.14	–0.59	9.39	1,640	50
DK	12.87	9.41	–10.60	16.30	–0.35	5.24	246	50
SE	11.91	21.63	–19.64	23.44	–0.42	7.34	877	70
UK	9.48	9.44	–21.62	15.28	–1.29	14.56	6,320	549
US	10.28	9.23	–17.27	15.44	–0.90	8.53	1,525	996
JP	–0.29	13.86	–12.34	19.59	–0.16	4.91	169	999
Panel B. Industrial Sector Indices								
DE	9.22	12.09	–15.63	20.50	–0.74	6.33	596	74
FR	9.76	19.31	–14.35	21.36	–0.16	6.33	504	57
IT	0.51	16.68	–17.73	25.61	–0.21	5.14	214	35
NL	8.99	22.24	–21.68	32.98	–0.41	5.78	377	40
ES	6.62	26.36	–31.35	22.64	–0.92	17.78	9,984	26
UK	8.62	10.42	–25.56	20.19	–1.57	13.91	5,800	117
US	11.43	10.50	–20.55	18.70	–1.06	10.86	2,980	168
JP	3.27	12.35	–13.96	21.94	–0.31	5.00	197	230
Panel C. Industrial Subsector Indices—Construction and Materials								
DE	6.75	12.21	–18.76	22.74	–0.48	6.36	547	11
FR	10.13	12.33	–12.08	22.06	–0.24	4.72	142	12
IT	5.48	14.88	–15.10	24.90	–0.18	5.31	245	16
NL	15.64	14.60	–13.49	20.67	–0.46	6.08	464	9
ES	12.69	16.91	–30.93	24.28	–1.02	11.51	3,450	12
UK	9.52	12.09	–21.33	19.83	–0.56	7.96	1,161	12
US	10.89	25.30	–25.73	23.42	–0.18	11.62	3,347	19
JP	–1.71	18.37	–15.39	22.99	0.17	6.49	552	37
Panel D. Industrial Subsector Indices—Industrials Goods and Services								
DE	9.28	10.76	–15.57	20.07	–0.74	6.25	572	63
FR	5.40	15.47	–14.07	20.86	–0.53	6.11	487	45
IT	1.99	13.15	–17.48	23.42	–0.27	5.30	250	19
NL	9.24	22.17	–20.77	27.99	–0.45	6.58	613	31
ES	16.08	9.67	–15.16	18.02	–0.54	7.02	779	14
UK	7.52	9.87	–24.09	17.15	–1.54	15.59	7,559	105
US	10.45	10.86	–21.00	18.13	–1.16	11.87	3,782	149
JP	1.83	12.25	–14.55	20.05	–0.34	5.17	232	193
Panel E. Financial Sector Indices								
DE	6.11	15.87	–17.80	21.19	–0.59	7.31	898	57
FR	9.16	18.71	–15.62	21.27	–0.16	7.45	893	54
IT	6.03	11.25	–19.41	21.15	–0.34	6.08	447	50
NL	9.63	15.14	–15.43	19.28	–0.65	8.71	1,541	32
ES	10.10	13.19	–22.40	21.93	–0.74	9.41	1,944	31
UK	11.56	13.31	–20.29	19.67	–0.58	7.93	1,155	196
US	11.57	14.94	–18.12	19.12	–0.24	8.09	1,174	197
JP	–3.85	18.62	–18.80	27.03	0.07	5.45	270	185

(continued on next page)

TABLE 1 (continued)

Summary Statistics

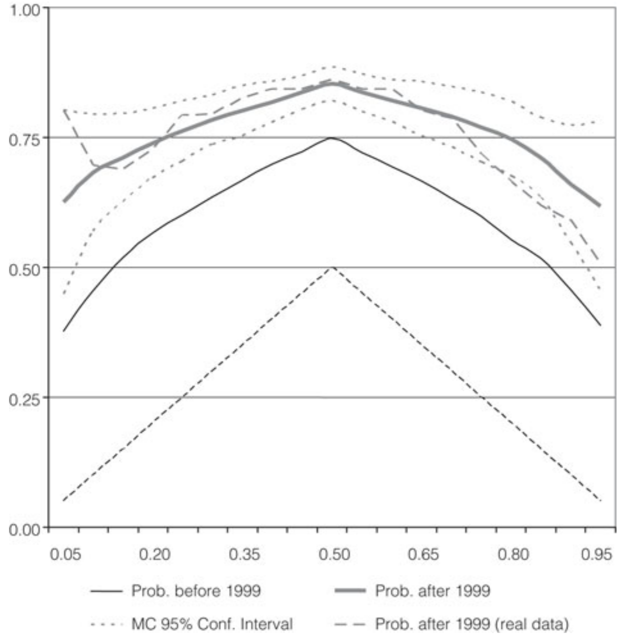
Country	Mean	Max	Min	SD	Skew	Kurt	J-B	#Comp
Panel F. Consumer Goods Sector Indices								
DE	7.66	15.86	–23.17	23.26	–0.76	9.05	1,748	32
FR	7.39	19.29	–18.81	22.98	–0.52	7.62	1,010	35
IT	2.72	12.77	–20.40	24.86	–0.52	5.92	432	22
NL	11.31	24.12	–19.05	24.13	0.10	9.27	1,772	13
ES	1.95	25.58	–24.45	34.38	0.74	9.14	1,791	16
UK	8.24	14.36	–27.23	23.25	–0.87	10.64	2,763	37
US	7.11	12.03	–22.07	19.34	–0.89	10.49	2,665	88
JP	5.25	10.34	–13.04	20.70	–0.24	4.93	177	176
Panel G. Consumer Goods Subsector Indices—Automobile								
DE	7.93	17.42	–23.25	24.52	–0.62	8.32	1,341	13
FR	7.36	17.27	–16.17	26.05	–0.40	5.53	315	7
IT	0.30	14.53	–21.16	28.66	–0.42	5.20	250	8
NL	2.64	15.98	–30.63	24.38	–1.47	16.64	8,769	0
ES	2.41	44.13	–24.67	40.48	1.45	12.29	4,262	1
UK	10.40	14.51	–25.55	26.03	–0.72	7.40	964	1
US	5.45	12.65	–22.70	22.45	–0.72	7.59	1,042	12
JP	6.38	12.63	–12.98	22.09	–0.10	4.77	142	54
Panel H. Consumer Goods Subsector Indices—Food and Beverages								
DE	8.39	10.13	–21.28	16.40	–0.79	12.70	4,347	7
FR	10.51	9.64	–10.07	16.84	–0.26	4.36	95	15
IT	5.47	11.15	–33.59	27.56	–1.23	13.04	2,778	3
NL	11.14	11.21	–12.65	17.35	–0.38	5.49	303	8
ES	6.46	14.58	–25.98	20.60	–0.74	11.74	3,538	9
UK	9.72	9.28	–20.68	16.15	–0.92	11.10	3,107	18
US	10.98	10.32	–11.04	15.51	–0.42	5.92	416	35
JP	–0.76	12.32	–15.13	16.92	–0.31	7.21	814	47
Panel I. Consumer Goods Subsector Indices—Personal and Household Goods								
DE	9.10	16.29	–12.72	17.92	–0.36	7.39	888	12
FR	11.41	18.43	–13.97	21.00	–0.18	6.43	535	13
IT	9.00	14.52	–27.95	24.32	–0.93	10.08	2,410	11
NL	10.63	20.57	–19.30	24.21	–0.04	7.71	998	5
ES	11.35	31.46	–18.66	27.84	0.40	9.73	2,064	6
UK	13.45	18.57	–24.87	18.95	–0.68	15.38	6,981	18
US	12.24	7.97	–20.22	16.92	–1.42	12.98	4,848	41
JP	2.94	14.40	–14.51	19.84	–0.31	6.22	484	75
Panel J. Consumer Services Sector Indices								
DE	6.43	17.04	–20.12	20.22	–0.52	7.63	1,023	21
FR	6.62	13.51	–19.19	20.61	–0.51	7.77	1,082	37
IT	3.97	17.30	–18.24	21.16	0.03	7.41	883	15
NL	11.89	13.89	–18.48	18.75	–0.69	8.98	1,714	17
ES	10.36	13.02	–16.48	20.97	–0.81	7.61	1,086	13
UK	7.12	10.55	–21.96	16.75	–1.23	12.85	4,682	89
US	8.50	11.54	–20.54	18.69	–0.83	9.08	1,789	139
JP	–1.01	12.41	–13.92	18.34	–0.28	5.30	252	149
Panel K. Health Sector Indices								
DE	10.78	12.82	–14.15	15.44	–0.59	8.29	1,321	18
FR	10.71	9.13	–9.76	18.73	–0.30	4.17	77	16
IT	5.30	15.10	–12.78	23.63	–0.02	4.29	74	3
NL	6.60	11.23	–18.49	19.43	–0.78	7.49	1,015	4
ES	11.57	22.12	–16.44	22.18	0.20	8.66	1,447	6
UK	9.13	11.40	–22.18	17.35	–0.64	10.70	2,742	23
US	11.53	7.34	–14.26	15.35	–0.61	6.39	582	88
JP	1.16	12.35	–13.69	16.10	–0.21	6.87	682	45

of comovements over the test period computed using observed (as opposed to simulated) data. We notice that with the chosen sample size, the methodology is powerful enough to detect statistically significant changes in comovements between the test and benchmark periods.

FIGURE 2

Monte Carlo Simulation

Figure 2 shows the average estimate of comovements and associated standard errors resulting from 100 replications of the Monte Carlo exercise described in Section IV.



VI. Structural Changes in Comovements

In this section we investigate whether comovements in equity returns have changed with the introduction of the euro. To this end we construct the time dummy variable by splitting the sample at January 1, 1999 to compare probabilities of comovement before and after the introduction of the single currency. First, we evaluate whether the comovements in national equity indices change after the introduction of the euro. Second, we introduce proxies for global factors to control that changes in comovements are not driven by worldwide trends in addition to euro-specific factors. Third, to understand the determinants of comovements between national indices, we reestimate the model with a sectoral breakdown.

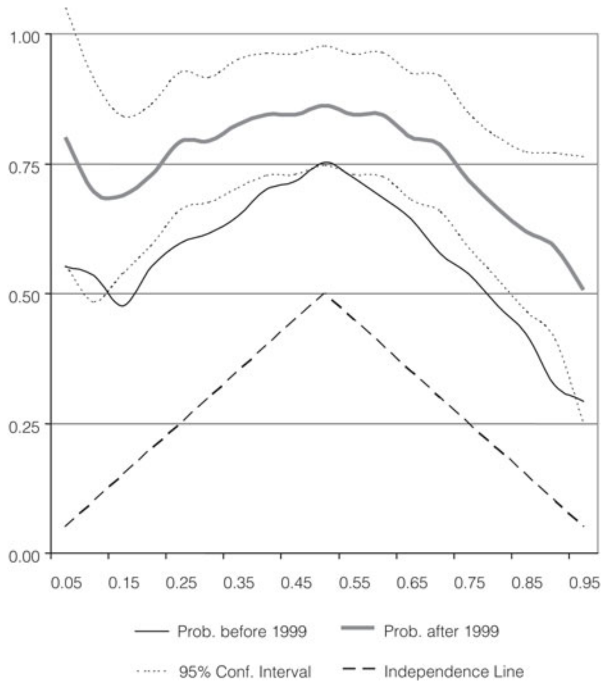
A. The Introduction of the Euro and the Comovements in National Equity Markets

We estimate the probability of comovements for the national equity indices of each possible country pair in the euro area. Since our sample includes 14 countries, we compute a total of 91 comovement boxes. In Figure 3, we report as an example the comovement box for France and Germany (together with 95% confidence bands). The chart shows that comovements between France and Germany have increased substantially with the introduction of the euro. The confidence bands indicate that the increase is also statistically significant for most of the quantiles.[9,10]

FIGURE 3

The Probability of Comovements between Returns on Equity Market Indices for France and Germany

Figure 3 plots the estimated probability of comovements between returns on French and German equity market indices over two periods. The first subsample covers the pre-monetary union period (March 1987–December 1998), while the second covers the monetary union period (January 1999–January 2008). The dashed lines denote the two standard error bounds around the estimated comovement likelihood in the monetary union period, while the thin line represents the probability of comovements in the pre-monetary union period.



Interestingly, Figure 3 shows that comovements between France and Germany are higher in the left than in the right tails of the distribution: Joint negative returns are more likely than positive ones. Such an analysis would not be possible with standard measures of correlation.

Table 2 summarizes the probability of comovements for each country pair. The upper triangular portion of the table reports the average probability of comovements across all the quantile ranges before 1999 (i.e., the average $\hat{\alpha}_{0,\theta}$ in regression (5) across all θs). The lower triangular portion of the table shows the changes in these probabilities after the introduction of the euro (statistics

[9] All the other charts are available from the authors.

[10] Since the pre- and post-1999 lines are given by the estimates of $\alpha_{0,\theta}$ and $\alpha_{0,\theta}+\alpha_{1,\theta}$ in regression (5), respectively, and the confidence bands refer to the difference between the two lines (i.e., to $\hat{\alpha}_{1,\theta}$), the standard errors associated to the estimate of $\alpha_{1,\theta}$ do not depend on the standard errors relative to $\alpha_{0,\theta}$.

significant at the 5% confidence level are reported in bold), that is, the averaged dummy variable coefficients $\hat{\alpha}_{1,\theta}$ of test (7). The average probability of comovements after 1999 can be computed by adding the probabilities of the upper and lower triangular portions of the table.

TABLE 2

The Probability of Comovements between Returns on Equity Market Indices over the Pre-Euro (upper triangular part) and the Euro Sample Periods (lower triangular part) across All Quantile Ranges

Table 2 reports the probability of comovements for each country pair. The upper triangular portion shows the average probability of comovements before the introduction of the euro. The lower triangular portion reports the changes in these probabilities after the introduction of the euro. Statistics significant at the 5% confidence level are reported in bold. Average probabilities and test statistics are estimated across all quantile ranges, for $\theta \in (0.05, 0.95)$. The first subsample covers the pre-monetary union period (March 1987–December 1998), while the second subsample covers the monetary union period (January 1999–January 2008). The equity indices refer to Germany (DE), France (FR), Italy (IT), the Netherlands (NL), Spain (ES), Austria (AT), Belgium (BE), Finland (FI), Greece (GR), Ireland (IE), Portugal (PT), Denmark (DK), Sweden (SE), and the United Kingdom (UK).

Average Probability of Comovements

	DE	FR	IT	NL	ES	AT	BE	FI	GR	IE	PT	DK	SE	UK
DE	—	0.57	0.46	0.57	0.47	0.50	0.53	0.42	0.36	0.45	0.44	0.49	0.50	0.49
FR	**0.18**	—	0.46	0.56	0.52	0.44	0.52	0.41	0.35	0.44	0.40	0.42	0.47	0.52
IT	**0.20**	**0.26**	—	0.47	0.45	0.42	0.45	0.40	0.33	0.42	0.36	0.45	0.44	0.43
NL	**0.15**	**0.17**	**0.18**	—	0.51	0.46	0.57	0.47	0.33	0.51	0.42	0.46	0.52	0.59
ES	**0.18**	**0.17**	**0.21**	**0.10**	—	0.44	0.49	0.41	0.33	0.43	0.45	0.42	0.49	0.48
AT	–0.02	0.05	**0.06**	0.02	0.03	—	0.46	0.38	0.37	0.41	0.40	0.42	0.41	0.42
BE	**0.08**	**0.11**	**0.13**	**0.08**	**0.08**	0.05	—	0.42	0.34	0.46	0.44	0.47	0.49	0.49
FI	**0.17**	**0.18**	**0.14**	**0.10**	**0.12**	0.04	**0.06**	—	0.34	0.44	0.37	0.45	0.50	0.46
GR	**0.14**	**0.13**	**0.12**	**0.15**	**0.12**	**0.07**	**0.12**	**0.06**	—	0.37	0.37	0.34	0.36	0.33
IE	**0.08**	**0.08**	**0.08**	0.01	**0.07**	0.04	0.05	0.02	**0.09**	—	0.40	0.44	0.45	0.53
PT	**0.07**	**0.11**	**0.13**	0.04	**0.07**	0.01	0.03	**0.07**	0.02	0.04	—	0.39	0.43	0.39
DK	**0.06**	**0.11**	0.04	**0.09**	**0.07**	0.01	0.05	**0.06**	**0.09**	0.02	0.04	—	0.46	0.44
SE	**0.16**	**0.20**	**0.17**	**0.11**	**0.11**	0.04	**0.06**	**0.11**	**0.13**	**0.06**	0.05	**0.11**	—	0.49
UK	**0.17**	**0.20**	**0.22**	**0.09**	**0.16**	**0.08**	**0.14**	**0.10**	**0.13**	0.00	**0.09**	**0.10**	**0.14**	—

Test for the Impact of the Euro

Table 2 offers a first set of stylized facts. First, large euro area countries exhibit higher degrees of comovement relative to the small economies already before the adoption of the euro.[11] Second, comovements increase significantly after 1999 for most of the country pairs involving at least one large economy, Austria being a noticeable exception. Third, the increase in probabilities is much higher for the large than for the small economies, despite the former having started from a higher level.

In Graphs A–D of Figure 4, we aggregate the 55 comovement boxes for euro area countries underlying Table 2. The aggregation is implemented as weighted averages of the probability of each comovement box. The weights are computed as the fraction of the average value of the country pair market capitalization relative to average value of the total euro area market capitalization. Weights are kept constant at the 2003 values. Graph A of Figure 4 shows that the overall average comovements increase after the introduction of the single currency. In line with the previous discussion, we distinguish between large and small economies

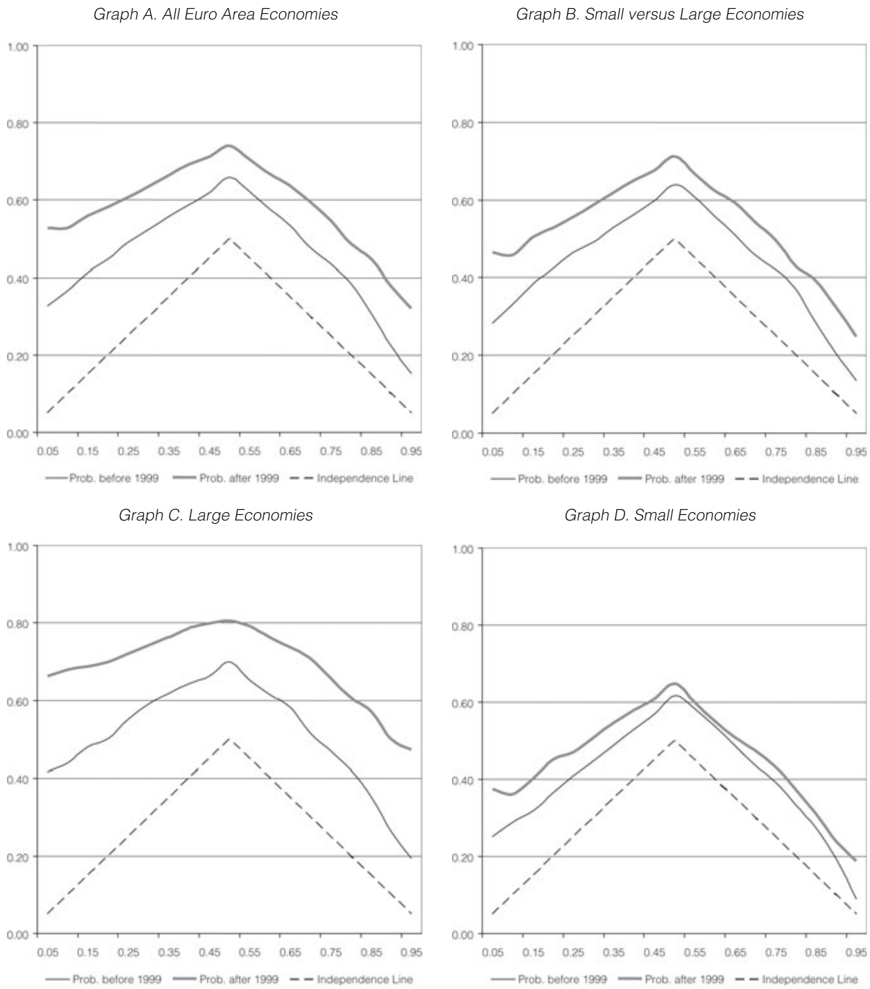
[11] The distinction between large and small economies is based on their relative market capitalization values. We consider large euro area countries Germany, France, Italy, the Netherlands, and Spain. The remaining countries of the sample form the small economies group.

(see Graphs B–D). This distinction confirms that most of the increase is driven by the large member states. Comovement changes in small economies are less pronounced.

FIGURE 4

Weighted Average Probability of Comovements between Returns on Equity Market Indices for the Euro Area Economies

Graphs A–D of Figure 4 plot weighted average estimated probability of comovements between returns on equity market indices for euro area member states over two periods. The first subsample covers the pre-monetary union period (March 1987–December 1998), while the second covers the monetary union period (January 1999–January 2008). The five largest euro area economies are France, Germany, Italy, the Netherlands, and Spain. The small economies included in the analysis are Austria, Belgium, Finland, Greece, Ireland, and Portugal. The probability of comovement of each euro area country pair is weighted by the fraction of its average market capitalization value relative to the total euro area market capitalization value at 2003.



Graphs A–D of Figure 4 also show that the asymmetric increase in comovements observed for France and Germany appears to be a stylized fact for all euro area equity markets. We further refine this analysis in Tables 3 and 4. Table 3 reports the average probability of comovements for the left and right parts of the distribution before the introduction of the euro. We notice that comovements in the left part of the distribution (reported in the upper triangular portion of the table) are substantially higher than those in the right part of the distribution (lower triangular portion of the table). Table 4 formally tests whether these differences are also statistically significant. Specifically, Table 4 reports the differences in comovements between the left and right parts of the distribution over the pre-euro and the euro sample periods. Statistics significant at the 5% level are reported in bold. The statistical significance was computed using the following tests, whose distribution can be easily derived from the joint distribution of the estimated parameters:

i) Test for asymmetries in the probability of comovements over the pre-euro sample period,

$$\hat{\xi} \equiv (10)^{-1} \sum_{\theta \in [0.05,0.5]} \widehat{\alpha}_{0,\theta} - (10)^{-1} \sum_{\theta \in [0.5,0.95]} \widehat{\alpha}_{0,\theta};$$

ii) Test for asymmetries in the probability of comovements over the euro sample period,

$$\widehat{\delta}(0.05, 0.5) - \widehat{\delta}(0.5, 0.95) = (10)^{-1} \sum_{\theta \in [0.05,0.5]} \widehat{\alpha}_{1,\theta} - (10)^{-1} \sum_{\theta \in [0.5,0.95]} \widehat{\alpha}_{1,\theta}.$$

TABLE 3

The Probability of Comovements between Returns on Equity Market Indices across the Upper (upper triangular part) and Lower (lower triangular part) Tail of the Quantile Distribution over the Pre-Euro Sample Period

Table 3 reports the probability of comovements for each country pair. The upper and lower triangular portions show the average probability of comovements before the introduction of the euro for the upper and lower parts of the distribution, respectively. The equity indices refer to Germany (DE), France (FR), Italy (IT), the Netherlands (NL), Spain (ES), Austria (AT), Belgium (BE), Finland (FI), Greece (GR), Ireland (IE), Portugal (PT), Denmark (DK), Sweden (SE), and the United Kingdom (UK).

														$\theta \in (0.05, 0.50)$
	DE	FR	IT	NL	ES	AT	BE	FI	GR	IE	PT	DK	SE	UK
DE	—	0.61	0.53	0.61	0.54	0.55	0.57	0.45	0.38	0.50	0.48	0.52	0.54	0.55
FR	0.54	—	0.52	0.59	0.57	0.48	0.56	0.44	0.35	0.48	0.46	0.45	0.50	0.56
IT	0.42	0.42	—	0.52	0.49	0.47	0.50	0.43	0.36	0.46	0.39	0.50	0.49	0.47
NL	0.54	0.53	0.44	—	0.56	0.52	0.59	0.50	0.36	0.51	0.49	0.49	0.57	0.63
ES	0.42	0.49	0.43	0.48	—	0.47	0.54	0.44	0.36	0.46	0.49	0.46	0.54	0.51
AT	0.46	0.42	0.39	0.43	0.42	—	0.50	0.43	0.41	0.44	0.47	0.46	0.47	0.47
BE	0.51	0.50	0.42	0.56	0.47	0.44	—	0.44	0.36	0.48	0.46	0.47	0.52	0.51
FI	0.41	0.40	0.39	0.46	0.41	0.35	0.43	—	0.39	0.43	0.39	0.45	0.50	0.47
GR	0.36	0.37	0.33	0.32	0.33	0.36	0.35	0.32	—	0.40	0.39	0.35	0.37	0.35
IE	0.42	0.42	0.39	0.52	0.42	0.39	0.45	0.47	0.37	—	0.42	0.46	0.47	0.56
PT	0.41	0.38	0.36	0.37	0.42	0.36	0.43	0.37	0.36	0.40	—	0.39	0.46	0.44
DK	0.47	0.42	0.42	0.45	0.41	0.39	0.47	0.46	0.35	0.43	0.42	—	0.51	0.46
SE	0.48	0.46	0.40	0.49	0.46	0.37	0.47	0.52	0.37	0.45	0.43	0.42	—	0.50
UK	0.46	0.50	0.42	0.56	0.47	0.40	0.48	0.46	0.33	0.53	0.37	0.43	0.50	—
	$\theta \in (0.50, 0.95)$													

The results highlight that the asymmetry between left and right parts of the distribution was already present before 1999 and was not further increased by the

TABLE 4

Tests for Asymmetries in the Probability of Comovements between Returns on Equity Market Indices over the Pre-Euro (upper triangular part) and the Euro Sample Periods (lower triangular part) across All Quantile Ranges

Table 4 reports tests for asymmetries in the probability of comovements for each country pair. The upper and lower triangular portions show whether there is more probability mass in the left or right parts of the distribution before and after the introduction of the euro, respectively. Statistics significant at the 5% level are reported in bold. The test for asymmetries in the probability of comovements over the pre-euro sample period is

$$\hat{\xi} \equiv (10)^{-1} \sum_{\theta \in [0.05,0.5]} \widehat{\alpha}_{0,\theta} - (10)^{-1} \sum_{\theta \in [0.5,0.95]} \widehat{\alpha}_{0,\theta}.$$

The test for asymmetries in the probability of comovements over the euro sample period is

$$\widehat{\delta}(0.05, 0.5) - \widehat{\delta}(0.5, 0.95) = (10)^{-1} \sum_{\theta \in [0.05,0.5]} \widehat{\alpha}_{1,\theta} - (10)^{-1} \sum_{\theta \in [0.5,0.95]} \widehat{\alpha}_{1,\theta}.$$

The equity indices refer to Germany (DE), France (FR), Italy (IT), the Netherlands (NL), Spain (ES), Austria (AT), Belgium (BE), Finland (FI), Greece (GR), Ireland (IE), Portugal (PT), Denmark (DK), Sweden (SE), and the United Kingdom (UK).

Pre-Euro Sample Period

	DE	FR	IT	NL	ES	AT	BE	FI	GR	IE	PT	DK	SE	UK
DE	—	**0.08**	**0.12**	**0.08**	**0.13**	**0.11**	**0.07**	0.04	0.02	**0.08**	**0.08**	0.05	**0.07**	**0.10**
FR	0.00	—	**0.11**	**0.07**	**0.09**	**0.06**	**0.06**	0.04	–0.02	**0.07**	**0.09**	0.03	0.05	**0.07**
IT	–0.03	–0.06	—	**0.08**	**0.07**	**0.10**	**0.10**	0.04	0.04	**0.07**	0.04	0.09	**0.09**	0.05
NL	0.03	0.05	0.00	—	**0.09**	**0.10**	0.03	0.04	0.04	–0.01	**0.13**	0.04	**0.08**	**0.09**
ES	–0.06	–0.03	–0.02	0.01	—	**0.06**	**0.08**	0.03	0.04	0.04	**0.07**	**0.06**	**0.08**	0.05
AT	0.03	0.05	0.00	0.01	–0.01	—	**0.07**	**0.09**	0.06	**0.06**	**0.12**	**0.08**	**0.11**	**0.07**
BE	0.06	0.05	–0.03	0.05	–0.03	0.04	—	0.01	0.01	0.03	0.04	0.00	0.05	0.03
FI	0.04	0.01	–0.03	0.01	–0.01	–0.05	0.04	—	**0.08**	–0.05	0.03	–0.01	–0.02	0.02
GR	0.03	0.08	0.02	0.04	–0.01	–0.01	0.08	–0.06	—	0.04	0.03	0.00	0.00	0.02
IE	0.01	0.03	0.03	0.13	0.04	0.01	0.07	0.08	0.05	—	0.01	0.03	0.03	0.04
PT	–0.03	–0.02	0.04	0.00	0.02	–0.01	0.08	–0.01	0.05	0.08	—	–0.03	0.03	**0.07**
DK	0.07	0.05	0.00	0.05	0.01	0.05	0.09	0.06	0.11	0.03	**0.14**	—	**0.10**	0.03
SE	0.03	–0.02	–0.03	–0.03	0.00	0.01	0.01	0.05	0.05	0.07	0.06	0.01	—	0.00
UK	0.00	–0.02	–0.01	0.05	0.01	0.08	0.14	0.05	0.06	0.10	0.02	0.06	0.07	—

Euro Sample Period

introduction of the single currency. Previous studies also document the presence of asymmetric correlations in equity markets. For instance, Ang and Chen (2002) and Hong, Tu, and Zhou (2007) find that comovements between selected portfolios of equities and the whole aggregate stock market are larger in down than in up markets.

B. Controlling for Global Factors

It is interesting to notice from Table 2 that comovements increase significantly also for country pairs involving the U.K., Denmark, and Sweden, which are members of the EU but have not joined the euro area. This may suggest that, after the introduction of the euro, the degree of integration among the financial markets of these countries increased as well. This is plausible, considering the strong economic ties of these economies with the euro area. Alternatively, this finding can indicate that the increase in comovements between euro area economies coincides with the augmented strength in global factors rather than the introduction of the common currency.

To distinguish between the introduction of the euro and possibly enhanced global financial trends, we control whether our results are robust to the inclusion of a factor proxying worldwide comovements. The control is implemented following the procedure described in Section III.C, estimating equation (15).[12] To the extent that high global correlations reflect markets' reactions to worldwide shocks, country pair correlations are likely to be affected as well. If the effect of these global shocks is not taken into account, the estimated comovements after the introduction of the euro would be biased. The implication is that one could erroneously associate the increase in comovements to the introduction of the euro, when in fact it is driven by global factors. For example, the burst of the dotcom bubble or recent geopolitical risks occurred after the introduction of the euro. If the uncertainty generated by these episodes increased correlations worldwide, neglecting these global correlation patterns would result in higher but spurious changes in comovements after the introduction of the single currency.

Results are reported in Table 5. Panel A reports the average probability of comovements across all the quantile ranges before 1999 (the average $\hat{\alpha}_{0,\theta}$ in regression (15) across all θs). Panel B shows the average dummy coefficients $\hat{\alpha}_{1,\theta}$ (upper portion of the table) and $\hat{\alpha}_{2,\theta}$ (lower portion) of tests (16) and (17), respectively. As discussed in Section III.C, these tests can be interpreted as the increase in the probability of comovements due to the euro and global dummy variables, respectively.

Not surprisingly, the inclusion of the global factor generally reduces the magnitude of the increase in country pair comovements occurring after the introduction of the euro. However, it is worth noticing that the euro dummy variable remains strongly significant for the large country pairs, while it becomes insignificant or marginally significant for small economies. The euro dummy variable continues to be significant also for the non-euro area countries vis-à-vis the large euro area economies. This suggests that the economic linkages among these nations have been strengthened by the creation of the single currency area. As for the global factor, it generally has positive coefficients, although only in a few cases is it statistically significant.

To sum up, there is substantial empirical evidence that after the introduction of the single currency the degree of comovement among euro area countries increased beyond what can be accounted for by global trends. The increase appears to be particularly pronounced and statistically significant, especially for large economies. In the next section we analyze to what extent these developments are driven by specific sectoral dynamics.

C. A Sectoral Decomposition

National aggregates may hide interesting developments occurring at a more disaggregate level. For instance, recent studies have shown that more tradable sectors are more sensitive to common shocks than less tradable industries. Brooks and Del Negro (2006) find that a firm increasing its international sales by 10% raises the exposure of its stock returns to global shocks by 2%. In a similar vein, Griffin and Karolyi (1998) find that global industry effects are more relevant than

[12]The global index used to compute the EWMA correlations is constructed as a weighted average of the Japanese, U.K., and U.S. equity indices. Weights are based on averages of market capitalization values over the period under consideration.

TABLE 5

The Probability of Comovements between Returns on Equity Market Indices over the Pre-Euro (upper triangular part) and the Euro Sample Periods (lower triangular part) across Different Quantile Ranges—Do Global Factors Play a Role?

Table 5 reports the probability of comovements for each country pair. Panel A shows the average probability of comovements before the introduction of the euro. Panel B reports the changes in these probabilities due to the introduction (upper triangular portion) of the euro and the global factor (lower triangular portion), respectively. Statistics significant at the 5% confidence level are reported in bold. Average probabilities and test statistics are estimated across all quantile ranges, for $\theta \in (0.05, 0.95)$. The first subsample covers the pre-monetary union period (March 1987–December 1998), while the second subsample covers the monetary union period (January 1999–January 2008). The equity indices refer to Germany (DE), France (FR), Italy (IT), the Netherlands (NL), Spain (ES), Austria (AT), Belgium (BE), Finland (FI), Greece (GR), Ireland (IE), Portugal (PT), Denmark (DK), Sweden (SE), and the United Kingdom (UK).

Panel A. The Average Probability of Comovements across the Whole Quantile Ranges

	DE	FR	IT	NL	ES	AT	BE	FI	GR	IE	PT	DK	SE	UK
DE		0.56	0.44	0.55	0.45	0.49	0.50	0.41	0.34	0.43	0.42	0.47	0.48	0.49
FR			0.45	0.55	0.52	0.43	0.51	0.41	0.33	0.43	0.39	0.42	0.46	0.52
IT				0.46	0.44	0.41	0.43	0.39	0.31	0.39	0.35	0.42	0.43	0.42
NL					0.49	0.46	0.55	0.46	0.31	0.50	0.40	0.46	0.52	0.58
ES						0.43	0.48	0.41	0.34	0.42	0.44	0.41	0.49	0.47
AT							0.44	0.36	0.35	0.39	0.38	0.42	0.40	0.42
BE								0.41	0.32	0.45	0.42	0.46	0.48	0.48
FI									0.30	0.42	0.34	0.45	0.48	0.44
GR										0.35	0.34	0.32	0.35	0.32
IE											0.39	0.43	0.45	0.52
PT												0.38	0.43	0.39
DK													0.45	0.42
SE														0.48

Panel B. Test for the Impact of the Euro and the Global Factor across the Whole Quantile Ranges

Test for the Impact of the Euro

	DE	FR	IT	NL	ES	AT	BE	FI	GR	IE	PT	DK	SE	UK
DE	—	**0.14**	**0.12**	**0.10**	**0.12**	–0.08	–0.02	**0.10**	0.06	0.02	0.02	0.02	0.06	**0.12**
FR	0.07	—	**0.20**	**0.17**	**0.12**	0.01	0.04	**0.16**	**0.07**	0.04	0.05	0.05	**0.12**	**0.19**
IT	**0.13**	0.09	—	**0.15**	**0.14**	–0.01	0.05	**0.10**	0.01	**–0.08**	0.06	–0.04	**0.11**	**0.18**
NL	0.08	0.02	0.04	—	**0.09**	0.03	**0.07**	0.09	**0.12**	0.01	0.01	**0.10**	**0.10**	**0.08**
ES	**0.09**	0.06	**0.11**	0.05	—	0.02	0.05	**0.12**	**0.11**	0.04	**0.07**	0.04	**0.12**	**0.11**
AT	0.08	0.05	**0.11**	–0.02	0.03	—	0.03	**0.09**	0.03	**0.07**	0.01	**0.10**	–0.01	0.05
BE	**0.15**	0.10	**0.13**	0.04	0.06	0.04	—	0.02	**0.08**	0.05	–0.01	0.03	0.01	–0.01
FI	0.09	0.03	0.08	0.03	0.00	0.02	0.06	—	–0.06	**–0.10**	0.06	0.06	0.03	0.04
GR	**0.12**	**0.10**	**0.17**	0.06	0.01	0.03	**0.09**	**0.19**	—	0.03	–0.03	0.01	0.06	0.09
IE	**0.08**	0.05	**0.21**	0.02	0.05	–0.02	0.01	**0.15**	**0.10**	—	0.01	0.00	**0.08**	0.02
PT	**0.09**	0.08	0.08	**0.08**	0.00	0.02	**0.09**	0.07	**0.11**	0.04	—	0.01	0.01	**0.09**
DK	0.05	0.08	**0.12**	0.00	0.04	–0.03	0.04	–0.01	**0.12**	0.04	0.05	—	0.11	0.03
SE	**0.15**	**0.10**	0.06	0.02	–0.03	**0.09**	0.07	**0.12**	**0.09**	–0.02	0.05	0.02	—	0.03
UK	0.06	0.00	0.07	0.02	0.05	0.03	**0.09**	**0.11**	0.04	**0.12**	–0.05	0.07	**0.08**	—

Test for the Impact of the Global Factor

country effects for traded than nontraded goods industries. Other factors that may have a differential impact on sectors include externalities generated by scientific discoveries (such as advances in information technology) or increased international mobility of production factors, notably financial capital and labor force. One cannot rule out the possibility that global factors affect different sectors with different intensities. Alternatively, lack of changes in comovements at the national level may reflect offsetting changes in comovements at the sectoral level.

To better understand the sources of comovements between national indices, we reestimate the model with a sectoral breakdown for the five largest euro area countries and the U.K. In Table 6 we present the probability of comovements for five industries: financial, industrial, consumer goods, consumer services, and

TABLE 6

Tests for Differences in the Probability of Comovements over the Pre-Euro and the Euro Sample Periods across All Quantile Ranges—A Sectoral Analysis

Table 6 reports the changes in the probability of comovements after the introduction of the euro (upper triangular portion) and controlling for the global factor (lower triangular portion). The analysis is carried with a sectoral breakdown. Test statistics are estimated across all quantile ranges, for $\theta \in (0.05, 0.95)$. The first subsample covers the pre-monetary union period (March 1987–December 1998), while the second subsample covers the monetary union period (January 1999–January 2008). Statistics significant at the 5% confidence level are reported in bold. The equity indices refer to Germany (DE), France (FR), Italy (IT), the Netherlands (NL), Spain (ES), and the United Kingdom (UK).

	DE	FR	IT	NL	ES	UK	DE	FR	IT	NL	ES	UK
	Industrial Sector						*Construction and Materials Sector*					
						Test for the Impact of the Euro						*Test for the Impact of the Euro*
DE	—	0.07	0.06	**0.16**	0.06	**0.14**	—	0.04	**0.11**	0.02	0.04	**0.08**
FR	0.03	—	**0.12**	**0.12**	0.06	**0.15**	**0.10**	—	**0.12**	0.03	**0.08**	0.03
IT	0.09	0.06	—	**0.11**	–0.01	**0.09**	0.05	**0.07**	—	0.02	0.02	**0.11**
NL	–0.03	0.06	**0.15**	—	**0.07**	**0.10**	0.00	–0.02	**0.08**	—	0.00	0.00
ES	0.03	0.03	**0.14**	0.02	—	0.02	**0.08**	0.05	**0.09**	**0.06**	—	0.02
UK	0.03	0.02	**0.10**	0.07	**0.08**	—	0.02	**0.07**	0.01	–0.01	**0.09**	—
	Test for the Impact of the Global Factor						*Test for the Impact of the Global Factor*					
	Industrial Goods and Services Sector						*Financial Sector*					
						Test for the Impact of the Euro						*Test for the Impact of the Euro*
DE	—	**0.09**	0.06	**0.09**	0.06	**0.14**	—	0.07	**0.13**	**0.09**	**0.14**	**0.09**
FR	0.06	—	**0.11**	**0.14**	0.05	**0.13**	**0.13**	—	**0.15**	**0.19**	**0.09**	**0.14**
IT	0.04	**0.10**	—	**0.10**	0.03	0.05	**0.09**	**0.12**	—	**0.13**	**0.11**	**0.18**
NL	0.04	0.06	**0.10**	—	**0.10**	**0.11**	0.06	0.03	0.05	—	**0.10**	**0.12**
ES	0.04	0.04	**0.14**	**0.10**	—	**0.06**	0.08	**0.12**	**0.14**	0.02	—	**0.11**
UK	0.00	0.04	**0.09**	0.06	0.02	—	0.05	0.03	0.04	–0.01	**0.07**	—
	Test for the Impact of the Global Factor						*Test for the Impact of the Global Factor*					
	Consumer Goods Sector						*Automobile Sector*					
						Test for the Impact of the Euro						*Test for the Impact of the Euro*
DE	—	0.06	0.06	0.01	–0.01	**0.09**	—	0.06	0.03	—	**–0.07**	**0.08**
FR	**0.09**	—	**0.08**	**0.06**	–0.02	0.02	**0.10**	—	0.05	—	–0.04	–0.01
IT	**0.14**	0.06	—	0.03	–0.03	**0.07**	**0.08**	0.04	—	—	–0.05	0.04
NL	0.05	0.02	**0.06**	—	–0.04	0.02	—	—	—	—	—	—
ES	**0.11**	**0.11**	**0.14**	0.04	—	0.02	**0.08**	0.02	**0.07**	—	—	–0.04
UK	0.05	**0.12**	0.05	0.02	0.02	—	0.05	**0.10**	0.02	—	**0.06**	—
	Test for the Impact of the Global Factor						*Test for the Impact of the Global Factor*					
	Food and Beverages Sector						*Personal and Household Goods Sector*					
						Test for the Impact of the Euro						*Test for the Impact of the Euro*
DE	—	–0.04	0.01	0.00	0.00	0.03	—	0.01	0.03	0.00	0.04	0.00
FR	0.03	—	0.01	**0.09**	0.02	**0.07**	**0.07**	—	**0.08**	0.04	–0.03	–0.02
IT	0.01	0.03	—	0.04	–0.01	0.01	**0.07**	0.03	—	0.03	**–0.05**	–0.03
NL	**–0.06**	**–0.06**	–0.04	—	**–0.05**	**0.07**	**0.08**	–0.02	**0.08**	—	–0.02	–0.04
ES	0.04	0.04	**0.06**	–0.02	—	–0.02	0.01	0.00	0.04	0.01	—	0.04
UK	–0.04	**–0.08**	0.00	**–0.11**	–0.04	—	0.04	0.03	**0.08**	0.01	0.06	—
	Test for the Impact of the Global Factor						*Test for the Impact of the Global Factor*					
	Consumer Services Sector						*Health Sector*					
						Test for the Impact of the Euro						*Test for the Impact of the Euro*
DE	—	**0.11**	**0.12**	0.06	**0.13**	**0.16**	—	0.01	0.03	0.00	0.03	0.04
FR	0.03	—	**0.18**	**0.13**	**0.13**	**0.19**	0.03	—	0.01	0.02	–0.02	0.04
IT	0.03	0.00	—	**0.14**	**0.16**	**0.19**	**0.11**	0.05	—	0.01	0.01	0.03
NL	**0.09**	0.00	0.06	—	**0.08**	**0.10**	–0.02	–0.02	0.04	—	–0.02	**–0.08**
ES	0.05	0.01	–0.04	0.02	—	**0.14**	–0.02	–0.04	0.05	0.03	—	–0.04
UK	–0.02	0.00	–0.01	0.01	–0.01	—	0.05	0.05	0.05	0.03	0.03	—
	Test for the Impact of the Global Factor						*Test for the Impact of the Global Factor*					

health care. For the industrial and consumer goods sectors we report a further breakdown. The industrial sector is split into "Construction and Materials" and "Industrial Goods and Services." The consumer goods sector is split into "Automobile," "Food and Beverages," and "Personal and Household Goods."

Similarly to the analysis reported in Table 3, we estimate the changes in comovements after the introduction of the euro on the different sectors, controlling for global factors. These controls are implemented following the procedure discussed in Section III.C.[13]

We notice that the changes in comovements observed at the aggregate level after the introduction of the euro continue to hold for the financial and consumer services sectors, and to a lesser extent for the industrial sector.

Changes in comovements in the financial sector after the advent of the euro are consistent with the recent findings by ECB (2008) on the progress of financial integration in Europe. The introduction of the euro has been complemented in the financial sector by an enhanced EU framework aimed at removing barriers to cross-border activities and safeguarding the stability of the single market. In particular, the financial services action plan (FSAP), launched in 1999, constituted a major overhaul of the EU legislation for financial services. While the FSAP targeted the entire financial sector, most of the initiatives related to securities markets (e.g., the Markets in Financial Instruments Directive (MiFID)). These initiatives contributed to creating an EU-wide level playing field in the financial sector, thus explaining the strong increase in comovements for the stocks in this sector after 1999.

Changes in comovements in consumer goods and health care industries after the introduction of the single currency are significantly less important. The further sectoral breakdown reveals that the comovement in the industrial sector is mostly driven by comovements in the industrial goods and services sector. As for the global factor, it has an impact in the financial and consumer goods sectors, while it appears to have almost no impact on industrial, consumer services, and health sectors.

These results suggest that looking at sectoral breakdowns uncovers interesting dynamics that could not be observed at more aggregate levels. The positive changes in comovements in the national index after the introduction of the euro appear to be mainly driven by the financial, consumer services, and industrial sectors.

VII. Conclusions

In this paper we employ a new methodology to investigate changes in comovements in European equity markets after the introduction of the euro. The methodology is based on quantile regression and evaluates comovements by estimating the probability that the returns on two different indices simultaneously exceed a given quantile. The advantage of this approach is that it is robust to the heteroskedasticity biases and departures from normality that typically plague financial data. Furthermore, it allows us to draw precise statistical inferences about the changes in comovements after the introduction of the euro. By properly addressing time-varying volatility issues, our measures of comovements permit us

[13]Global sector indices are constructed as a weighted average of the Japanese, U.K., and U.S. sector indices. Weights are based on averages of industry market capitalization values over the period under consideration.

to evaluate whether the introduction of the euro has coincided with an increased degree of financial integration.

We document an overall increase in the degree of comovement between European equity markets upon the introduction of the single currency. This increase is robust to controls accounting for changes in global correlations. A more refined analysis on sector indices confirms that after the introduction of the single currency, overall comovements among euro area economies did increase. It also reveals that most of the comovements are driven by the financial, industrial, and consumer services sectors.

Appendix A. Time-Varying Regression Quantiles

Let $q_t^i(\beta_{\theta,i})$ denote the empirical specification for the $q_{\theta,t}^i$ time-varying quantile conditional on Ω_t, where $\beta_{\theta,i}$ denotes the p-vector of parameters to be estimated. Let $\rho_\theta(\lambda) \equiv [\theta - I(\lambda \leq 0)]\,\lambda$ be a piecewise linear "check function," where $I(\cdot)$ denotes an indicator function that takes the value of 1 if the expression in parenthesis is true and 0 otherwise. The unknown parameters of the quantile specification can be consistently estimated by solving the following minimization problem (Koenker and Bassett (1978)):

$$\text{(A-1)} \qquad \min_{\beta_{\theta,i}} T^{-1} \sum_{t=1}^{T} \rho_\theta \left(r_{i,t} - q_t^i(\beta_{\theta,i}) \right),$$

where T denotes the sample size

Engle and Manganelli (2004) provide sufficient conditions for consistency and asymptotic normality results of individual quantile specifications.

To derive the joint distribution of the regression quantile estimators of the two time series, $r_{i,t}$ and $r_{j,t}$, define

$$\underset{p\times p}{D_\theta^i} \quad \equiv \quad \mathrm{E}\left[T^{-1} \sum_{t=1}^{T} h_{\theta,t}^i(0) \nabla q_t^i(\beta_{\theta,i}^0) \nabla' q_t^i(\beta_{\theta,i}^0) \right],$$

where $h_{\theta,t}^i(0)$ is the value at 0 of the density of $\varepsilon_{\theta,t}^i \equiv r_{i,t} - q_t^i(\beta_{\theta,i}^0)$, $\nabla q_t^i(\beta_{\theta,i}^0)$ is the gradient of the quantile function evaluated at the true parameter $\beta_{\theta,i}^0$, and

$$\underset{p\times 1}{\psi_{\theta,t}^i(\beta_{\theta,i}^0)} \quad \equiv \quad [\theta - I(r_{i,t} \leq q_t^i(\beta_{\theta,i}^0))] \nabla q_t^i(\beta_{\theta,i}^0).$$

Next, let $\beta_i \equiv [\beta_{\theta,i}]_{\theta=1}^m$ denote the pm-vector stacking the $\beta_{\theta,i}$ regression quantile parameters, $D^i \equiv \mathrm{diag}\left([D_\theta^i]_{\theta=1}^m\right)$ the $(pm \times pm)$ block diagonal matrix with the matrices D_θ^i along the main diagonal, and $\psi_t^i(\beta_i^0) \equiv [\psi_{\theta,t}^i(\beta_{\theta,i}^0)]_{\theta=1}^m$ the pm-vector stacking all the $\psi_{\theta,t}^i(\beta_{\theta,i}^0)$.

Consider analogous terms for $r_{j,t}$ and, finally, define

$$\underset{2pm\times 1}{\beta} \quad \equiv \quad [\beta_i', \beta_j']',$$

$$\text{(A-2)} \qquad \underset{2pm\times 2pm}{D} \quad \equiv \quad \mathrm{diag}([D^i, D^j]),$$

(A-3) $$\underset{2pm\times 1}{\psi_t(\beta^0)} \equiv [\psi_t^i(\beta_i^0)', \psi_t^j(\beta_j^0)']'.$$

The following corollary derives the joint asymptotic distribution of the regression quantile estimators.

Corollary 1. Under assumptions C0–C7 and AN1–AN4 in Appendix A of Cappiello et al. (2005),

$$\sqrt{T}A^{-1/2}D(\hat{\beta} - \beta^0) \quad \xrightarrow{d} \quad N(0, I),$$

where $\hat{\beta}$ is the vector containing the solutions to expression (A-1) and

$$A \equiv \mathrm{E}\left[T^{-1}\sum_{t=1}^{T}\psi_t(\beta^0)\psi_t(\beta^0)'\right].$$

Engle and Manganelli (2004) provide asymptotically consistent estimators of the variance-covariance matrix (see their Theorem 3).

Appendix B. Estimation of the Conditional Probability of Comovements

The average probability of comovements between $r_{i,t}$ and $r_{j,t}$ can be estimated by running the following regression:

(B-1) $$I_t(\hat{\beta}_{\theta,i}) \cdot I_t(\hat{\beta}_{\theta,j}) = W_t\alpha_\theta^0 + \varepsilon_t, \qquad \theta = 1, \ldots, m,$$

where $I_t(\beta_{\theta,i}) \equiv I\left(r_{i,t} \leq q_t^i(\beta_{\theta,i})\right)$, $I_t(\beta_{\theta,j})$ is defined analogously, $W_t \equiv [1,\ S_t]$, S_t is an $(s-1)$ row vector of dummy variables (possibly indicating alternative time periods identified by economic variables), and α_θ^0 a $(s, 1)$ vector of unknown coefficients.

Let $\hat{\alpha}_\theta$ be the ordinary least squares (OLS) estimator of (B-1) and denote with $\hat{\alpha}_{l,\theta}$ the $(l+1)$th element of this vector, $l = 0, 1, \ldots, s-1$. Analogously, let $S_{l,t}$ denote the lth element of S_t. Let C_l be the number of observations identified by the dummy variables $\{S_{l,t} = 1, S_t^{-l} = \mathbf{0}\}_{t=1}^T$, where S_t^{-l} represents the vector S_t without its lth element and $\mathbf{0}$ is a vector of 0s of appropriate dimension. Define also

$$\bar{F}_l^\theta \equiv C_l^{-1} \sum_{t\in\{t:S_{l,t}=1, S_t^{-l}=\mathbf{0}\}} F_t(q_t^i(\beta_{\theta,i}), q_t^j(\beta_{\theta,j})).$$[14]

The following theorem shows that $\hat{\alpha}_\theta$ is a consistent estimator of the average probability of comovements in the time periods defined by the dummy variables.

[14]We denote with C_0 the number of observations in the benchmark period. Here, $\bar{F}_0^\theta$ is correspondingly defined as the average cumulative distribution function (CDF) in the benchmark period.

Theorem 1 (Consistency). Assume that $C_l/T \overset{T\to\infty}{\longrightarrow} k_l$, where $k_l \in (0,1)$, $l = 0,\ldots,s-1$, is the asymptotic ratio between the number of observations identified by the lth dummy variable (C_l) and the total number (T) of observations. Under the same assumptions of Corollary 1,

$$\hat{\alpha}_{0,\theta} \overset{p}{\to} \operatorname{plim}(\bar{F}_0^{\theta}), \qquad \theta = 1,\ldots,m,$$

$$[\hat{\alpha}_{0,\theta} + \hat{\alpha}_{l,\theta}] \overset{p}{\to} \operatorname{plim}(\bar{F}_l^{\theta}), \qquad \theta = 1,\ldots,m, \quad \text{and} \quad l = 1,\ldots,s-1.$$

Here, $\hat{\alpha}_{0,\theta}$ is the parameter associated with the constant and, as such, it converges to the average probability of comovements in the benchmark period (i.e., the period when all other dummy variables are equal to 0s). Similarly, since $\hat{\alpha}_{l,\theta}$ for $l = 1,\ldots,s-1$ is the coefficient of the lth dummy variable, $S_{l,t}$, the sum of $\hat{\alpha}_{0,\theta} + \hat{\alpha}_{l,\theta}$ converges in probability to the average probability of comovements in the corresponding dummy period. According to this theorem, testing for a change in the conditional probability of comovements in the periods identified by the dummy variable, $S_{l,t}$, is equivalent to testing for the null that $\alpha_{l,\theta}$ is equal to 0. Indeed, it is only when $\alpha_{l,\theta} = 0$ that there is no change in the probability of comovements relative to the benchmark period. Otherwise, if $\alpha_{l,\theta} < 0$, the probability over the lth dummy period will be lower than the probability during the benchmark period, while if $\alpha_{l,\theta} > 0$, the probability will be higher.

To obtain the asymptotic distribution of this estimator, first define the following terms:

$$\underset{s\times 1}{g_t(\beta_{\theta,i},\beta_{\theta,j})} \equiv W_t' I_t(\beta_{\theta,i}) \cdot I_t(\beta_{\theta,j}) - T^{-1}(W'W)\alpha_\theta^0,$$

and

$$\underset{s\times 2pm}{G_\theta} \equiv \mathrm{E}\left\{T^{-1}\sum_{t=1}^{T} W_t' \left[\nabla_\beta' q_t^j(\beta_{\theta,j}^0)\int_{-\infty}^{0} h_{\theta,t}(\eta,0)d\eta + \nabla_\beta' q_t^i(\beta_{\theta,i}^0)\int_{-\infty}^{0} h_t(0,\upsilon)d\upsilon\right]\right\},$$

where $h_{\theta,t}(\eta,\upsilon)$ is the joint probability density function (PDF) of $(r_{i,t} - q_t^i(\beta_{\theta,i}^0), r_{j,t} - q_t^j(\beta_{\theta,j}^0))$, and ∇_β denotes the derivative with respect to the $2pm$-vector β. Next let $g_t(\beta^0) \equiv [g_t(\beta_{\theta,i}^0,\beta_{\theta,j}^0)]_{\theta=1}^m$ be the sm-vector stacking all the m possible vectors $g_t(\beta_{\theta,i},\beta_{\theta,j})$, and construct the $(sm \times T)$ matrix $R \equiv [g_t(\beta^0)]_{t=1}^T$. Define also $G \equiv [G_\theta]_{\theta=1}^m$, an $(sm \times 2pm)$ matrix stacking all the G_θ matrices; $\Psi \equiv [\psi_t(\beta^0)]_{t=1}^T$, a $(2pm \times T)$ matrix where $\psi_t(\beta^0)$ was defined in equation (A-3); and $W \equiv [W_t]_{t=1}^T$, a $(T \times s)$ matrix containing all the vectors of dummy variables from regression (5).

Finally, let $\alpha^0 \equiv [\alpha_{\theta_1}^{0\prime}, \alpha_{\theta_2}^{0\prime},\ldots,\alpha_{\theta_m}^{0\prime}]'$ be the sm-vector of true unknown parameters to be estimated in regression (5). Similarly, define $\hat{\alpha} \equiv [\hat{\alpha}_{\theta_1}', \hat{\alpha}_{\theta_2}', \ldots, \hat{\alpha}_{\theta_m}']'$.

Theorem 2 (Asymptotic Normality). Under the assumptions of Corollary 1 and AN5 in Appendix A of Cappiello et al. (2005),

$$\sqrt{T}M^{-1/2}Q\left(\hat{\alpha}-\alpha^0\right) \quad \xrightarrow{d} \quad N(0, J_{sm}),$$

where

$$\underset{sm\times sm}{M} \quad \equiv \quad \mathrm{E}[T^{-1}(R+GD^{-1}\Psi)(R+GD^{-1}\Psi)'],$$

$$\underset{sm\times sm}{Q} \quad \equiv \quad J_m \otimes (T^{-1}W'W),$$

J_k is the identity matrix of dimension k, and D is defined in equation (A-2).

Without the correction term $GD^{-1}\Psi$ in matrix M, we would get the standard OLS variance-covariance matrix. The correction is needed in order to account for the estimated regression quantile parameters that enter the OLS regression. This correction term is similar to the one derived by Engle and Manganelli (2004) for the in-sample dynamic quantile test. The main difference is related to the composition of matrix G. Since two different random variables ($r_{i,t}$ and $r_{j,t}$) enter the regression, G contains the terms $\int_{-\infty}^{0} h_{\theta,t}(\eta,0)d\eta$ and $\int_{-\infty}^{0} h_{\theta,t}(0,\upsilon)d\upsilon$, which can be interpreted as the bivariate analogue of the height of the density function of the quantile residuals evaluated at 0 that typically appears in standard errors of regression quantiles.

The variance-covariance matrix can be consistently estimated using plug-in estimators. The only nonstandard term is G, whose estimator is provided by the following theorem:

Theorem 3 (Variance-Covariance Estimation). Under the same assumptions of Theorem 2 and assumptions VC1–VC3 in Appendix A of Cappiello et al. (2005), $\hat{G}_\theta \xrightarrow{p} G_\theta$, where

$$\begin{aligned}\hat{G}_\theta \quad \equiv \quad & (2T\hat{c}_T)^{-1}\sum_{t=1}^{T}\Big\{I(|r_{j,t}-q_t^j(\hat{\beta}_{\theta,j})|<\hat{c}_T)I(r_{i,t}<q_t^i(\hat{\beta}_{\theta,i}))W_t'\nabla_\beta' q_t^j(\hat{\beta}_{\theta,j}) \\ & +I(|r_{i,t}-q_t^i(\hat{\beta}_{\theta,i})|<\hat{c}_T)I(r_{j,t}<q_t^j(\hat{\beta}_{\theta,j}))W_t'\nabla_\beta' q_t^i(\hat{\beta}_{\theta,i})\Big\}\end{aligned}$$

and $\hat{c}_T$ is defined in assumption VC1.

Appendix C. Hypothesis Testing

Using Theorems 2 and 3, a test of linear restrictions on the estimated probability of comovement can be easily constructed.

Corollary 2. Suppose that α is subject to u ($\leq sm$) linearly independent restrictions $U\alpha^0 = b$, where U is an (u, sm) matrix of rank u and b is an u-vector. Under the assumptions of Theorem 3,

$$\sqrt{T}(UQ^{-1}\hat{M}Q^{-1}U')^{-1/2}\left(U\hat{\alpha}-b\right) \quad \xrightarrow{d} \quad N(0, I_u),$$

which can be equivalently restated as a Wald test,

$$T(U\hat{\alpha} - b)'(UQ^{-1}\hat{M}Q^{-1}U')^{-1}(U\hat{\alpha} - b) \quad \stackrel{d}{\rightarrow} \quad \chi^2(u),$$

where the ˆ indicates estimated quantities.

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