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Have Euro Area Government Bond Spreads converged to their Common State ?

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Have euro area government bond spreads converged to their common state ?*

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Abstract

We derive a model in which a standard international capital asset pricing (ICAPM) model is nested within an ICAPM model with market imperfections. In the latter model an idiosyncratic stochastic factor affects the return of risky assets (over a risk-free rate) on top of the systematic component that is common to all countries (and that is interacted with a time-varying idiosyncratic "beta"). We introduce asymptotic convergence from the full ICAPM model with imperfections to the standard model by multiplying the idiosyncratic factor by convergence operators. The model is then estimated using the weekly 10 year government bond spreads of Belgium, France, Italy, and the Netherlands versus Germany over the period 1991-2006. We find that the idiosyncratic components have converged towards zero for all countries after the introduction of the euro implying that the efficiency of the euro area government bond markets under consideration has increased. Full convergence has not yet occurred however.

JEL Classifications: E43, G12

Keywords: Government bonds, euro area, interest rate spreads, state space methods

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1 Introduction.

The run-up to the introduction of the euro in 1999 has been characterized by a significant decrease in the interest rate spreads of bonds of euro area governments over German benchmark bonds, reflecting primarily a decrease in the exchange rate risk premium. As a result, after the start of the European Economic and Monetary Union (EMU), these interest rate spreads in principle only reflect a liquidity premium and/or a credit risk premium (see Codogno et al. 2003).

However, the covariance of the spreads of different euro area countries is higher than what can be explained by liquidity indicators or credit risk factors (i.e. local fundamentals) (see Favero et al. 2007). The common component that is present in government bond spreads reflects international risk. In the literature on (EMU) government bond spreads two approaches have been followed to take this common factor into account.

First, a number of studies investigating euro area countries include a proxy for this common factor in their analysis. These studies also allow for interaction effects of this common factor with credit risk factors and/or liquidity indicators. Codogno et al. (2003) relate US risk factors (corporate and banking risk) to country-specific default risk (measured via the government debt) for a number of countries among which are Belgium, France, Italy, and the Netherlands. They conclude that the impact of international risk factors is higher in countries with a higher government debt. Bernoth et al. (2004) examine yield spreads of EU countries versus Germany and the United States. They conclude that international risk factors captured by US corporate risk affect spreads, but evidence on the interaction of global risk with local fundamentals or liquidity indicators is mixed. Favero et al. (2007) argue that the changing risk attitude of international investors interacts with liquidity indicators.

Second, a number of studies explicitly filter the common factor out of the government bond spreads through the use of factor analysis and state space methods. Dungey et al. (2000) find a common factor in the long-term bond yield differentials of Australia, Japan, Germany, Canada, and the UK versus the US. Geyer et al. (2004) filter out a common component in the bond spreads of Austria, Belgium, Italy, and Spain versus Germany. They find a significant impact of European risk proxies on the common factor and interpret the common factor as a risk premium that reflects the risk of a failure of EMU that could lead to the reintroduction of exchange rate risks.

In this paper we argue that the common factor found in euro area government bond spreads is beyond the control of national governments. To the extent that this common factor is a relevant component in government bond spreads the question of whether total government bond spreads have converged to zero and a "market for the same bond" has been created is not very relevant. What matters instead is whether national governments have succeeded in diminishing or even eliminating the country-specific or idiosyncratic components of the bond spreads. In standard international asset pricing models (see e.g. Harvey 1991) such components imply inefficiencies (e.g. illiquidity, taxation) since only systematic common risk should be priced. As a result, if the idiosyncratic components have converged towards zero during the transition to the euro, the government bond spreads have converged towards their common state and the efficiency of the government bond markets under investigation is said to have increased.

The contribution of this paper to the literature is both theoretical and methodological.

Theoretically, we use an international capital asset pricing (ICAPM) model as presented for instance by Harvey (1991) where a representative global investor invests in the bond markets of different countries. Acharya and Pedersen (2005) show that the pricing equations of a CAPM derived in a frictionless economy but expressed in net returns, i.e. returns minus some arbitrary idiosyncratic premium, are equivalent to those derived from a CAPM where frictions are explicitly incorporated. Thus we can allow for impediments and imperfections on the local bond markets simply by rewriting an ICAPM for net returns. The standard ICAPM is then obtained if the idiosyncratic premia equal zero. In the paper we assume that impediments and imperfections in the local bond markets disappear gradually over time. In particular, the idiosyncratic premia in the spreads of government bonds over a benchmark bond converge to zero asymptotically so that, in the limit, the standard asset pricing equations hold. Acharya and Pedersen's equivalence result thus allows us to nest a standard ICAPM into an ICAPM with market imperfections and to assume that the latter gradually converges to the former.

Methodologically, we use a linear state space approach to estimate the latent factor decomposition of the spreads that is implied by the theoretical model. In particular, we use weekly data over the period 1991-2006 to decompose the 10 year government bond spreads of Belgium, France, Italy and the Netherlands versus Germany into a common component and an idiosyncratic component.[1]

The country-specific time-varying impacts of the common factor on the bond spreads, the "betas" (i.e. the ratios of the conditional covariance of the common factor and the bond spread

[1] As such we avoid the use of proxies. The latter are often imperfectly capturing the common state. For instance, US risk proxies may be only a part of the common factor and, as Geyer et al. (2004) suggest, European factors also are part of the story.

over the conditional variance of the common factor), are also estimated. Del Negro and Otrok (2006) provide a Bayesian method to estimate a dynamic factor model with time-varying loadings where the factor loadings are assumed to follow driftless random walks. The approach that we follow, while less general in nature, is better suited for a CAPM model since it involves the direct parameterization of the conditional covariances and variances in the "betas". In other words, the time-varying loadings in our approach are not stochastic so that our model still fits into a standard linear state space framework.[2]

We then investigate whether the idiosyncratic components in government bond spreads have converged towards zero by multiplying the idiosyncratic premia in the bond spreads by convergence operators of the type suggested by Luginbuhl and Koopman (2004). We also investigate whether the idiosyncratic components are still significantly different from zero at the end of the sample period.

Our results suggest that our ICAPM with market imperfections converges to a standard ICAPM, i.e. the idiosyncratic components in the bond spreads converge towards zero for all four countries after the introduction of the euro implying that the efficiency of the euro area government bond markets under consideration has increased. The results imply a decrease in the relevance of market imperfections like illiquidity and taxation in the 10 year segment of the government bond markets of the euro area countries under investigation. As far as liquidity is concerned this supports Bernoth et al. (2004) who find that the introduction of the euro has decreased liquidity premia in euro area government bond spreads. Codogno et al. (2003) and Favero et al. (2007) find that, for the years 2002 and 2003, the liquidity component in bond spreads is not very important. While we find that in the years after the introduction of the euro the importance of the idiosyncratic component in bond spreads is strongly reduced we nevertheless conclude that full convergence to zero of this component has not yet occurred. Moreover, the reduction of the spreads is attributed to a decrease in local market impediments and imperfections rather than to a decrease in the country-specific exposure to international risk.

The outline of the paper is as follows. Section 2 presents the theoretical model. In section 3 we present the empirical specification and the estimation method. We also discuss data issues. Results from the estimations are reported in section 4, while the final section concludes.

[2] Our approach of dealing with conditional covariances in a state space framework is an extension of the state space models with time-varying conditional variances as studied by Harvey et al. (1992).

2 The model.

There is a representative international investor who maximizes expected utility by choosing a consumption path over an infinite lifetime. This investor invests in the government bond markets of N different countries ($i = 1, ..., N$), in a risk-free benchmark bond b, and in an international portfolio w. The period $t+1$ returns of the bonds i ($\forall i$), b, and the portfolio w are denoted by R_{it+1} ($\forall i$), R_{bt+1}, and R_{wt+1}. The variable α_{it+1} reflects the cost of impediments and imperfections (e.g. illiquidity, taxation) encountered on the bond market of country i ($\forall i$). We assume that $\lim_{t\to+\infty} \alpha_{it+1} = 0$, i.e. these costs converge to zero as time passes by. For the benchmark bond b and for the portfolio w we assume that these costs are zero, i.e. $\alpha_{wt+1} = \alpha_{bt+1} = 0$.[3] The period $t+1$ utility function for the international investor is denoted by $u(c_{t+1})$ where c_{t+1} is period $t+1$ consumption of the investor. The subjective rate of time preference of the investor is captured by the discount factor ρ (with $0 < \rho < 1$). The stochastic discount factor which drives the returns R_{it+1} ($\forall i$) and R_{wt+1} is defined as $m_{t,t+1} \equiv \rho \frac{u'(c_{t+1})}{u'(c_t)}$.

These assumptions lead to the following first-order conditions,

$$E_t \left[m_{t,t+1}(R_{it+1} - \alpha_{it+1}) \right] = 1 \tag{1}$$

$$E_t \left[m_{t,t+1} \right] R_{bt+1} = 1 \tag{2}$$

$$E_t \left[m_{t,t+1} R_{wt+1} \right] = 1 \tag{3}$$

where E_t is the expectations operator conditional on the period t information set and where eq.(1) holds $\forall i$. All first-order conditions reflect the fact that, in the optimum, the investor is indifferent between consuming an amount of 1 at time t or investing this amount into i ($\forall i$), b, or w and consuming $(R_{it+1} - \alpha_{it+1})$, R_{bt+1}, or R_{wt+1} at time $t+1$. The expected, discounted marginal utility of both decisions is equal. Since $\lim_{t\to+\infty} \alpha_{it+1} = 0$, eq.(1) converges asymptotically to the standard Euler equation. We thus nest a standard CAPM into a CAPM with market imperfections and we assume that the latter gradually converges to the former. This is possible because of Acharya and Pedersen's (2005) equivalence result which shows that the pricing equations of our CAPM derived for a frictionless economy with returns minus some arbitrary idiosyncratic premium α_{it+1}, are equivalent to those derived for a CAPM where frictions are explicitly incorporated.

[3] We thus assume that there is no common factor in the liquidity premium. See Acharya and Pedersen (2005) and Favero et al (2007) for evidence on this for respectively the US stock market and euro area bond markets.

In appendix A we show that we can write,

$$E_t\left[R_{it+1}\right] - R_{bt+1} = E_t\left[\alpha_{it+1}\right] + \beta_{it+1}(E_t\left[R_{wt+1}\right] - R_{bt+1}) \tag{4}$$

where $\beta_{it+1} = \frac{cov_t[R_{wt+1}, R_{it+1} - \alpha_{it+1}]}{V_t[R_{wt+1}]}$. This equation states that the expected excess return of bond i over the benchmark bond return depends on the expected costs of market imperfections on market i, on the expected excess return of the international portfolio ($R_{wt+1} - R_{bt+1}$), and on β_{it+1} which reflects the conditional covariance of the country-specific net bond return and the return on the global portfolio.

Assuming eq.(4) holds for ex post returns we can write it as,

$$\widetilde{R}_{it+1} = \alpha_{it+1} + \beta_{it+1}\widetilde{R}_{wt+1} \tag{5}$$

where the " $\sim$ " over a variable denotes excess returns, i.e. $\widetilde{R}_{it+1} = R_{it+1} - R_{bt+1}$, $\widetilde{R}_{wt+1} = R_{wt+1} - R_{bt+1}$, and where $\beta_{it+1} = \frac{cov_t\left[\widetilde{R}_{wt+1}, \widetilde{R}_{it+1} - \alpha_{it+1}\right]}{V_t\left[\widetilde{R}_{wt+1}\right]}$.[4]

The remainder of the paper deals with the identification and the estimation of α_{it+1}, $\widetilde{R}_{wt+1}$, and β_{it+1}, and with the question of whether the variable α_{it+1} has converged towards zero. If it has, the efficiency of the government bond market under investigation is said to have increased.

3 Empirical specification, data, and estimation method.

3.1 Empirical specification.

We estimate the following system (where $i = 1, ..., N$),

$$\widetilde{R}_{it+1} = \alpha_{it+1} + \beta_{it+1}\widetilde{R}_{wt+1} \tag{6}$$

$$\alpha_{it+1} = \kappa_{it+1}\mu_i + \pi_i\alpha_{it} + \kappa_{it+1}\varepsilon_{it+1} \tag{7}$$

$$\beta_{it+1} = \phi_{it+1}\beta_{it} + \chi_{it+1} \tag{8}$$

$$\widetilde{R}_{wt+1} = \mu_w + \pi_w\widetilde{R}_{wt} + \varepsilon_{wt+1} \tag{9}$$

[4] We have also estimated eq.(5) with an additional error term $\epsilon_{it+1} = (R_{it+1} - E_t\left[R_{it+1}\right]) - (\alpha_{it+1} - E_t\left[\alpha_{it+1}\right]) - \beta_{it+1}(R_{wt+1} - E_t\left[R_{wt+1}\right])$ where $E_t\left[\epsilon_{it+1}\right] = 0$, $cov_t(R_{wt+1}, \epsilon_{it+1}) = 0$, and $cov_t(\alpha_{it+1}, \epsilon_{it+1}) = 0$. We find that (the variance of) this error term is very small. The estimations are available from the authors upon request.

In section 2 we assume that $\lim_{t\to+\infty} \alpha_{it+1} = 0$. We therefore model the idiosyncratic component α_{it+1} as an $AR(1)$ process, i.e. $\alpha^*_{it+1} = \mu_i + \pi_i \alpha^*_{it} + \varepsilon_{it+1}$ with $0 < \pi_i < 1$, multiplied by a deterministic convergence operator κ_{it+1} where $\lim_{t\to+\infty} \kappa_{it+1} = 0$. Thus, $\alpha_{it+1} = \alpha^*_{it+1}\kappa_{it+1}$. We can write $\alpha^*_{it+1} = \mu_i/(1-\pi_i L)+\varepsilon_{it+1}/(1-\pi_i L)$ where L is the lag operator. Multiplication by κ_{it+1} then gives $\alpha_{it+1} = \kappa_{it+1}\mu_i/(1-\pi_i L) + \kappa_{it+1}\varepsilon_{it+1}/(1-\pi_i L)$. On multiplication of both sides of the latter expression by $(1-\pi_i L)$ we obtain eq.(7). For κ_{it} we use the following specification (see Luginbuhl and Koopman 2004),

$$\kappa_{it} = \exp\left[\xi_i(t-\tau_i)\right]/(1+\exp\left[\xi_i(t-\tau_i)\right]) \tag{10}$$

where $\xi_i < 0$ is the rate of convergence. Since $\xi_i < 0$ we have $\kappa_{it} = 0$ for $t \to +\infty$ and $\kappa_{it} = 1$ for $t \to -\infty$. In a sample of size T the fact that $\xi_i < 0$ implies that $\kappa_{it} \approx 0$ for $t >>> \tau_i$ and that $\kappa_{it} \approx 1$ for $t <<< \tau_i$. The parameter τ_i with $1 < \tau_i < T$ determines the mid-point of the change.

The error term ε_{it+1} is white noise and follows a $GARCH(1,1)$ process,

$$\varepsilon_{it+1} = h^{1/2}_{it+1}\nu_{it+1} \tag{11}$$

where $\nu_{it+1} \sim i.i.d(0,1)$ and where

$$h_{it+1} = V_t\left[\varepsilon_{it+1}\right] = \delta^a_i + \delta^b_i\varepsilon^2_{it} + \delta^c_i h_{it} \tag{12}$$

with $\delta^a_i > 0$, $0 < \delta^b_i < 1$, $0 < \delta^c_i < 1$, and $0 < \delta^b_i + \delta^c_i < 1$.[5] Note that the unconditional variance of ε_{it+1} is given by $\delta^a_i/(1-\delta^b_i-\delta^c_i)$.

As can be seen in eq.(9) we assume that the common component $\widetilde{R}_{wt+1}$ follows an $AR(1)$ process with $0 < \pi_w < 1$ and where ε_{wt+1} is white noise and follows a $GARCH(1,1)$ process,

$$\varepsilon_{wt+1} = h^{1/2}_{wt+1}\nu_{wt+1} \tag{13}$$

where $\nu_{wt+1} \sim i.i.d(0,1)$ and where

$$h_{wt+1} = V_t\left[\varepsilon_{wt+1}\right] = \delta^a_w + \delta^b_w\varepsilon^2_{wt} + \delta^c_w h_{wt} \tag{14}$$

with $\delta^a_w > 0$, $0 < \delta^b_w < 1$, $0 < \delta^c_w < 1$, and $0 < \delta^b_w + \delta^c_w < 1$. Note that the unconditional variance of ε_{wt+1} is given by $\delta^a_w/(1-\delta^b_w-\delta^c_w)$.

[5] Note that Dungey et al. (2000) model the shocks in weekly government bond spreads as $GARCH(1,1)$ processes.

Eq.(8) represents the law of motion for $\beta_{it+1.}$ In Del Negro and Otrok's (2006) factor model β_{it+1} is assumed to follow a driftless random walk, i.e. $\phi_{it+1} = 1$ and χ_{it+1} is white noise. The approach that we follow in this paper is based on a direct parameterization of the conditional covariance and variance in $\beta_{it+1.}$ Remember from section 2 that $\beta_{it+1} = \frac{cov_t\left[\widetilde{R}_{wt+1},(\widetilde{R}_{it+1}-\alpha_{it+1})\right]}{V_t\left[\widetilde{R}_{wt+1}\right]} = \frac{cov_t\left[\varepsilon_{wt+1},\varepsilon^*_{it+1}\right]}{V_t\left[\varepsilon_{wt+1}\right]} = \frac{g_{it+1}}{h_{wt+1}}$ where $\varepsilon^*_{it+1} = \varepsilon^R_{it+1} - \kappa_{it+1}\varepsilon_{it+1}$ with $\varepsilon^R_{it+1} = \widetilde{R}_{it+1} - E_t\left[\widetilde{R}_{it+1}\right]$ and where h_{wt+1} is given by eq.(14). We parameterize the conditional covariance g_{it+1} as a function of the lagged cross-product of the errors ε_{wt+1} and ε^*_{it+1} and of its own lag (i.e. the diagonal bivariate VEC(1,1) of Bollerslev et al. 1988),

$$g_{it+1} = \gamma_i^a + \gamma_i^b \varepsilon_{wt}\varepsilon^*_{it} + \gamma_i^c g_{it} \tag{15}$$

with $\gamma_i^a > 0$, $0 < \gamma_i^b < 1$, $0 < \gamma_i^c < 1$, and $0 < \gamma_i^b + \gamma_i^c < 1$ for $i = 1, ..., N$. Note that from eq.(6) we can write $\varepsilon^*_{it+1} = \varepsilon^R_{it+1} - \kappa_{it+1}\varepsilon_{it+1} = \beta_{it+1}\varepsilon_{wt+1}$ so that eq.(15) can be rewritten as,

$$g_{it+1} = \gamma_i^a + \gamma_i^b \beta_{it}\varepsilon^2_{wt} + \gamma_i^c g_{it} \tag{16}$$

Substituting eqs.(14) and (16) into $\beta_{it+1} = \frac{g_{it+1}}{h_{wt+1}}$ we obtain eq.(8) with $\phi_{it+1} = \frac{\gamma_i^b \varepsilon^2_{wt} + \gamma_i^c h_{wt}}{h_{wt+1}}$ and $\chi_{it+1} = \frac{\gamma_i^a}{h_{wt+1}}$(where we have used $g_{it} = \beta_{it}h_{wt}$). Contrary to the specification for β_{it+1} considered by Del Negro and Otrok our specification for β_{it+1} can still be estimated in a linear state space framework with maximum likelihood. Our approach of dealing with conditional covariances in a state space framework is an extension of the state space models with time-varying conditional variances as studied by Harvey et al. (1992).

3.2 A look at the data.

To investigate whether the idiosyncratic components α_{it+1} of the government bond spreads have converged towards zero we need long enough time series. For 4 countries, Belgium, France, Italy, and the Netherlands (i.e. $N = 4$), and for the benchmark country Germany data are available from the early nineties onward. All data are taken from Datastream/Thomson Financial. We average the available daily data to weekly data because with weekly data we avoid day-of-the-week effects and we average out other noise present in daily spreads (see Dungey et al. 2000). The available dataset covers the period 28-06-1991 to 04-08-2006 providing 789 weekly observations. For $\widetilde{R}_{it+1}$ we use corrected government bond spreads given by $r_{it} - r_{ft} - (s_{it} - s_{ft})$ where r_{it} and r_{if} are the yields to maturity of 10 year government bonds issued by country i (i.e. Belgium, France, Italy, and the Netherlands) and by the benchmark country f (i.e. Germany) respectively and where s_{it}

and s_{ft} are 10 year fixed interest rates on swaps denominated in the currency of country i (euro after 01-01-1999) and in DEM (euro after 01-01-1999) respectively.[6] The government bond spreads are corrected with the spread $s_{it} - s_{ft}$ to remove expected exchange rate changes and exchange rate risk from the government bond spreads before the introduction of the euro (i.e. before 01-01-1999). We refer to Favero et al. (1997, 2007) and Codogno et al. (2003) for more on this. Note that after the introduction of the euro $s_{it} - s_{ft}$ equals zero for all i and t.

In Figure 1 we report the spreads against Germany for the government bonds of Belgium, France, Italy, and the Netherlands. From this figure we note, first, that the spreads move together. This suggests that they are driven by a common component as our theoretical model predicts and as reported for instance by Geyer et al. (2004). Moreover, from table 1, we note that the correlation between the spreads is much higher after the introduction of the euro which indicates that after 1998 the idiosyncratic components of the spreads have become less relevant. Second, while the spreads undoubtedly converge after 1998 they remain positive reflecting significant liquidity and/or risk premia. After the introduction of the euro the spreads are on average 8 basispoints for France and the Netherlands, 16 basispoints for Belgium and 24 basispoints for Italy. The magnitude of the spreads has non-trivial consequences for the public finances of these countries, especially for the highly indebted ones. Third, the decrease in the volatility of the spreads of all countries in the nineties is more pronounced than the decrease in the means. For Italy the sharp peaks in the spread series in the beginning of the nineties may reflect currency issues. Even though our calculated spread is corrected for exchange rate factors and should reflect only liquidity and/or risk premia, to the extent that exchange rate factors are correlated with default or credit risk, these factors may have an impact on the corrected spreads nonetheless.

3.3 Estimation.

3.3.1 Method.

We obtain estimates for the unobserved states α_{it+1} and $\widetilde{R}_{wt+1}$, for the conditional variance series h_{it+1} and h_{wt+1}, for the conditional covariance series g_{it+1} and thus for β_{it+1}, for the convergence operator series κ_{it+1}, and for the parameters in the model by putting the model described in section

[6] While German bond yields may not be totally risk-free they are believed to approximate this condition fairly well in the 10 year segment of the government bond market. Using German bonds as a benchmark is common practice in this line of research, see e.g. Bernoth et al. (2004).

3.1 in state space form. In particular, we estimate a Gaussian linear state space system including time-varying conditional variances (see Harvey et al. 1992 and Kim and Nelson 1999, chapter 6) and covariances. In Appendix B we report the state space representation of the model. Estimates of the state vector are obtained with the Kalman filter and smoother. Given the assumption of stationarity the initialization of the system is non-diffuse.

The time-varying conditional variances and covariances complicate the otherwise standard state space framework. To deal with this we follow the approach by Harvey et al. (1992) and augment the state vector with the shocks ε_{it+1} and ε_{wt+1}. The Kalman filter then provides estimates of the conditional variance of the shocks, i.e. estimates for h_{it+1} and h_{wt+1}. This also allows us to calculate g_{it+1} and β_{it+1}. We refer to appendix B for more details on the approach followed.

To deal with potential computational difficulties that are caused by the relatively large dimension of the observation vector we follow the univariate approach to multivariate filtering and smoothing as presented by Koopman and Durbin (2000) and Durbin and Koopman (2001, chapter 6). A major advantage of this approach is that it avoids taking the inverse of the variance matrix of the one-step-ahead prediction errors. We refer to Koopman and Durbin (2000) for the filtering and smoothing recursions and for the calculation of the likelihood.

3.3.2 Identification.

As is standard in factor models, first, to identify the factor loadings β_{it+1} we impose an unconditional variance of unity on ε_{wt+1}, i.e $\sigma_w^2 = 1$. This amounts to setting $\delta_w^a = 1 - \delta_w^b - \delta_w^c$ (see Dungey et al. 2000). Second, implicitly factor loadings are also estimated on the states α_{it+1} since eq.(6) can be written as $\widetilde{R}_{it+1} = \kappa_{it+1}\alpha^*_{it+1} + \beta_{it+1}\widetilde{R}_{wt+1}$. We impose $\kappa_{i1} = 1$ $(\forall i)$. Third, the sign of the factor loadings β_{it+1} cannot be identified because of the sign invariance of the factor variance decompositions of the spreads. Therefore, we impose $\beta_{it+1} > 0$. Since $\beta_{it+1} = \frac{g_{it+1}}{h_{wt+1}}$ we need $g_{it+1} > 0$ and $h_{wt+1} > 0$. It is straightforward to show that under the processes and parameter restrictions discussed in section 3.1 these conditions hold. Fourth, our model contains constants (i.e. the data we use are not in deviations from the mean). Since we can only estimate N constants but the model contains $N + 1$ constants (i.e. μ_w and μ_i for $i = 1, ..., N$ where $N = 4$) we impose the identifying restriction $\sum_{i=1}^N \mu_i = 0$.

4 Results.

In table 2 we present the results of estimating the system given by eqs.(6) to (15). In figures 2-5 we present the factor decompositions of the spreads for each country. The smoothed estimates for α_{it+1} are contrasted with the smoothed estimates for $\widetilde{R}_{wt+1}$. The latter series is multiplied by the estimated country-specific time-varying factor loadings β_{it+1} which are presented seperately in figure 7. The estimates of the idiosyncratic states α_{it+1} are presented in figure 6. In figure 8 the estimates for the convergence operators series κ_{it+1} are presented for every country while in figure 9 the estimated country-specific GARCH series h_{it+1} are presented.

From table 2 we note, first, that standard confidence intervals for the $AR(1)$ coefficients π of the common state and of the idiosyncratic states for Belgium, France, and the Netherlands do not contain the unit root case. For Italy however we find a near unit root. Imposing $\pi = 1$ for Italy and estimating the resulting state space model (with semi-diffuse initialization) does not affect the results however.

Second, the common constant μ_w is positive and significant. Since we impose the identifying restriction $\mu_4 = -\mu_1 - \mu_2 - \mu_3$ some idiosyncratic constants are positive (Italy and Belgium) and others are negative (France and the Netherlands).

Third, we note that there are significant convergence effects since ξ is significantly lower than zero for all countries. The estimates for ξ and τ imply an estimated series κ_{it+1} for each country. These series are presented in figure 8. It is obvious from this figure that the convergence of the government bond spreads to their common state has occurred after the introduction of the euro for all countries under investigation. The importance of the idiosyncratic components of the spreads is now much lower than in the nineties. However, full convergence to zero has not yet occurred. While the magnitude of α_{it+1} is much lower after the introduction of the euro, tests of the hypothesis $\alpha_{it+1} = 0$ (at the 5% level) for each i and t during the years 2005 and 2006 reveal that $\alpha_{it+1} = 0$ is still rejected in a large percentage of cases. Note that we also report the Akaike Information Criterion (AIC) as a goodness of fit statistic (see Durbin and Koopman 2001, p.152). According to a comparison of this statistic for our model with convergence effects and the statistic calculated for the same model but without convergence effects (i.e. when setting $\xi_i = 0 \; \forall i$) we find that the model with convergence effects is clearly preferred.

Fourth, the persistence in the volatility is very high as reflected by the large estimates found for δ_i^c and δ_w^c. This high persistence in the volatility of bond spreads is also reported by Dungey et al. (2000) for Australia, Japan, Germany, Canada, and the UK (versus the US). Moreover, we find integrated $GARCH$ estimates for the common state, i.e. the sum $\delta_w^b + \delta_w^c$ is very close to 1, i.e. h_{wt} contains a unit root. The same is true for the estimated idiosyncratic conditional covariance series g_{it} since we find that the sum $\gamma_i^b + \gamma_i^c$ is very close to 1 for all countries. Hence, the estimated series for β_{it} which are approximated by dividing the estimated series for g_{it} by the estimated series for h_{wt} are also non-stationary as can be seen in figure 7. While non-stationary the estimated β_{it} shows a slight tendency to decline after the introduction of the euro in Belgium and France but not in the Netherlands nor in Italy.

Summarizing, our results suggest that the idiosyncratic components have converged towards zero for all four countries after the introduction of the euro implying that the efficiency of the euro area government bond markets under consideration has increased. The results imply a decrease in the relevance of market imperfections like illiquidity and taxation in the 10 year segment of the government bond markets of the euro area countries under investigation. As far as liquidity is concerned this supports Bernoth et al. (2004) who find that the introduction of the euro has decreased liquidity premia in euro area government bond spreads. Codogno et al. (2003) and Favero et al. (2007) find that, for the years 2002 and 2003, the liquidity component in bond spreads is not very important. While we find that in the years after the introduction of the euro the importance of the idiosyncratic component in bond spreads is strongly reduced we nevertheless conclude that full convergence to zero of this component has not yet occurred. Moreover, our results suggest that it is α_{it}, not β_{it}, that is responsible for the decrease in government bond spreads. Thus the reduction of the spreads is attributed to a decrease in local market impediments and imperfections rather than to a decrease in the country-specific exposure to international risk.

5 Conclusions.

In this paper we derive a model in which a standard international capital asset pricing model is nested within an ICAPM model with market imperfections. In the latter model an idiosyncratic stochastic factor affects the return of risky assets (over a risk-free rate) on top of the systematic component that is common to all countries (and that is interacted with a time-varying idiosyncratic "beta"). We introduce asymptotic convergence from the full ICAPM model with imperfections to the standard model by multiplying the idiosyncratic factor by convergence operators.

Methodologically, we use a linear state space approach to estimate the latent factor decomposition of the spreads that is implied by the theoretical model. In particular, we use weekly data over the period 1991-2006 to decompose the 10 year government bond spreads of Belgium, France, Italy, and the Netherlands versus Germany into a common component and an idiosyncratic component. The country-specific time-varying impacts of the common factor on the bond spreads, the "betas" (i.e. the ratios of the conditional covariance of the common factor and the bond spread over the conditional variance of the common factor), are also estimated. We investigate whether the idiosyncratic components in government bond spreads have converged towards zero and we also investigate whether the idiosyncratic components are still significantly different from zero at the end of the sample period.

Our results suggest that the idiosyncratic components have converged towards zero for all four countries after the introduction of the euro implying that the efficiency of the euro area government bond markets under consideration has increased. While we find that in the years after the introduction of the euro the importance of the idiosyncratic components in bond spreads is strongly reduced we nevertheless conclude that full convergence to zero of these components has not yet occurred. Moreover, the reduction of the spreads is attributed to a decrease in local market impediments and imperfections rather than to a decrease in the country-specific exposure to international risk.

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Appendix A. Derivation of eq.(4).

Write eq.(1) as,

$$E_t\left[m_{t,t+1}\right]E_t\left[R_{it+1}-\alpha_{it+1}\right]+cov_t(m_{t,t+1},R_{it+1}-\alpha_{it+1})=1 \tag{A1}$$

From eq.(2) we have $E_t[m_{t,t+1}] = 1/R_{bt+1}$. By using this into eq.(A1) and re-arranging we obtain,

$$E_t[R_{it+1}] - R_{bt+1} = E_t[\alpha_{it+1}] - cov_t(m_{t,t+1}, R_{it+1} - \alpha_{it+1})R_{bt+1} \tag{A2}$$

Since $\alpha_{wt+1} = 0$ eq.(A2) written for R_{wt+1} is,

$$E_t[R_{wt+1}] - R_{bt+1} = -cov_t(m_{t,t+1}, R_{wt+1})R_{bt+1} \tag{A3}$$

(alternatively, use eq.(3) with $E_t[m_{t,t+1}] = 1/R_{bt+1}$ to obtain this).

Note that eq.(A2) can be rewritten as,

$$\begin{aligned} E_t[R_{it+1}] - R_{bt+1} &= E_t[\alpha_{it+1}] \\ &\quad - cov_t(m_{t,t+1}, R_{it+1} - \alpha_{it+1})R_{bt+1}\frac{cov_t(m_{t,t+1}, R_{wt+1})}{cov_t(m_{t,t+1}, R_{wt+1})} \end{aligned} \tag{A4}$$

By using eq.(A3) into this we obtain,

$$\begin{aligned} E_t[R_{it+1}] - R_{bt+1} &= E_t[\alpha_{it+1}] \\ &\quad + \frac{cov_t(m_{t,t+1}, R_{it+1} - \alpha_{it+1})}{cov_t(m_{t,t+1}, R_{wt+1})}\left(E_t[R_{wt+1}] - R_{bt+1}\right) \end{aligned} \tag{A5}$$

The returns $(R_{it+1} - \alpha_{it+1})$ and R_{wt+1} are driven by the stochastic discount factor $m_{t,t+1}$ as in the following equations,

$$R_{it+1} - \alpha_{it+1} = \tau^R_{it+1} m_{t,t+1} + u^R_{it+1} \tag{A6}$$

$$R_{wt+1} = \tau^R_{wt+1} m_{t,t+1} + u^R_{wt+1} \tag{A7}$$

where u^R_{it+1} and $u^R_{wt+1}$are $i.i.d.$shocks and where $\tau^R_{it+1} = \frac{cov_t(m_{t,t+1}, R_{it+1} - \alpha_{it+1})}{V_t[m_{t,t+1}]}$ and $\tau^R_{wt+1} = \frac{cov_t(m_{t,t+1}, R_{wt+1})}{V_t[m_{t,t+1}]}$. Note also that,

$$V_t[R_{wt+1}] = \tau^R_{wt+1} cov_t(m_{t,t+1}, R_{wt+1}) \tag{A8}$$

Then we can write,

$$\begin{aligned} cov_t(R_{it+1} - \alpha_{it+1}, R_{wt+1}) &= \tau^R_{it+1}\tau^R_{wt+1} V_t[m_{t,t+1}] \\ &= \frac{cov_t(m_{t,t+1}, R_{it+1} - \alpha_{it+1})}{V_t[m_{t,t+1}]}\tau^R_{wt+1} V_t[m_{t,t+1}] \\ &= cov_t(m_{t,t+1}, R_{it+1} - \alpha_{it+1})\tau^R_{wt+1} \\ &= \frac{cov_t(m_{t,t+1}, R_{it+1} - \alpha_{it+1})}{cov_t(m_{t,t+1}, R_{wt+1})} V_t[R_{wt+1}] \end{aligned}$$

where the last line uses eq.(A8). This gives,

$$\frac{cov_t(m_{t,t+1}, R_{it+1} - \alpha_{it+1})}{cov_t(m_{t,t+1}, R_{wt+1})} = \frac{cov_t(R_{it+1} - \alpha_{it+1}, R_{wt+1})}{V_t\left[R_{wt+1}\right]} = \beta_{it+1} \tag{A9}$$

By substituting eq.(A9) into eq.(A5) we obtain eq.(4) in the text.

Note that this derivation is obtained without imposing a specific process for α_{it+1}. Any process can be assumed, i.e. α_{it+1} can be driven by $m_{t,t+1}$ (see Acharya and Pedersen 2005) but this need not be the case. In the empirical section we implicitly assume that α_{it+1} is independent of $m_{t,t+1}$ and thus of R_{wt+1} as we need to impose this as an identifying restriction to estimate the latent factor decomposition implied by the model.

Note further that the solution method presented in this section is by no means unique. We refer to Cochrane (2005, chapter 9) for alternative solution methods.

Appendix B. State space representation of the model.

The state space system with state vector S_{t+1} is,

$$y_{t+1} = Z_{t+1} S_{t+1} \tag{B1}$$

$$S_{t+1} = T_{t+1} S_t + K_{t+1} \eta_{t+1} \tag{B2}$$

with

$$\eta_{t+1|t} \sim N(0, Q_{t+1}) \tag{B3}$$

$$S_1 \sim N(A_1, P_1) \tag{B4}$$

Since $N = 4$ we have $y_{t+1} = \begin{bmatrix} \widetilde{R}_{1t+1} & \widetilde{R}_{2t+1} & \widetilde{R}_{3t+1} & \widetilde{R}_{4t+1} \end{bmatrix}'$,

$S_{t+1} = \begin{bmatrix} 1 & \alpha_{1t+1} & \alpha_{2t+1} & \alpha_{3t+1} & \alpha_{4t+1} & \widetilde{R}_{wt+1} & \varepsilon_{1t+1} & \varepsilon_{2t+1} & \varepsilon_{3t+1} & \varepsilon_{4t+1} & \varepsilon_{wt+1} \end{bmatrix}'$,

$\eta_{t+1} = \begin{bmatrix} \varepsilon_{wt+1} & \varepsilon_{1t+1} & \varepsilon_{2t+1} & \varepsilon_{3t+1} & \varepsilon_{4t+1} \end{bmatrix}'$,

$A_1 = \begin{bmatrix} 1 & \frac{\mu_1}{1-\pi_1} & \frac{\mu_2}{1-\pi_2} & \frac{\mu_3}{1-\pi_3} & \frac{\mu_4}{1-\pi_4} & \frac{\mu_w}{1-\pi_w} & 0 & 0 & 0 & 0 & 0 \end{bmatrix}'$

where $\mu_4 = -\mu_1 - \mu_2 - \mu_3$,

$diag(P_1) = \begin{bmatrix} 0 & \frac{\sigma^2_{\alpha 1}}{1-\pi_1^2} & \frac{\sigma^2_{\alpha 2}}{1-\pi_2^2} & \frac{\sigma^2_{\alpha 3}}{1-\pi_3^2} & \frac{\sigma^2_{\alpha 4}}{1-\pi_4^2} & \frac{\sigma^2_{w}}{1-\pi_w^2} & \sigma^2_{\alpha 1} & \sigma^2_{\alpha 2} & \sigma^2_{\alpha 3} & \sigma^2_{\alpha 4} & \sigma^2_{w} \end{bmatrix}'$

where $\sigma^2_{\alpha i} = \delta^a_i/(1 - \delta^b_i - \delta^c_i)$ (for $i = 1, ..., 4$) and $\sigma^2_w = \delta^a_w/(1 - \delta^b_w - \delta^c_w) = 1$,

$diag(Q_{t+1}) = \begin{bmatrix} h_{wt+1} & h_{1t+1} & h_{2t+1} & h_{3t+1} & h_{4t+1} \end{bmatrix}'$

where $h_{it+1} = \delta_i^a + \delta_i^b \varepsilon_{it}^2 + \delta_i^c h_{it}$ (for $i = 1, ..., 4$) and $h_{wt+1} = \delta_w^a + \delta_w^b \varepsilon_{wt}^2 + \delta_w^c h_{wt}$ with $\delta_w^a = 1 - \delta_w^b - \delta_w^c$

(note that the non-diagonal elements of the matrices P_1 and Q_{t+1} are zero),

$$Z_{t+1} = \begin{bmatrix} 0 & 1 & 0 & 0 & 0 & \beta_{1t+1} & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & \beta_{2t+1} & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 & \beta_{3t+1} & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 & \beta_{4t+1} & 0 & 0 & 0 & 0 & 0 \end{bmatrix}$$

where $\beta_{it+1} = \frac{g_{it+1}}{h_{wt+1}}$ with $g_{it+1} = \gamma_i^a + \gamma_i^b \beta_{it} \varepsilon_{wt}^2 + \gamma_i^c g_{it}$ (for $i = 1, ..., 4$) ,

$$K_{t+1} = \begin{bmatrix} 0 & 0 & 0 & 0 & 0 & 1 & 0 & 0 & 0 & 0 & 1 \\ 0 & \kappa_{1t+1} & 0 & 0 & 0 & 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & \kappa_{2t+1} & 0 & 0 & 0 & 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & \kappa_{3t+1} & 0 & 0 & 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 & \kappa_{4t+1} & 0 & 0 & 0 & 0 & 1 & 0 \end{bmatrix}' ,$$

$$T_{t+1} = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ \mu_1 \kappa_{1t+1} & \pi_1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ \mu_2 \kappa_{2t+1} & 0 & \pi_2 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ \mu_3 \kappa_{3t+1} & 0 & 0 & \pi_3 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ \mu_4 \kappa_{4t+1} & 0 & 0 & 0 & \pi_4 & 0 & 0 & 0 & 0 & 0 & 0 \\ \mu_w & 0 & 0 & 0 & 0 & \pi_w & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \end{bmatrix}$$

where $\mu_4 = -\mu_1 - \mu_2 - \mu_3$ and where $\kappa_{it+1} = \exp\left[\xi_i(t+1-\tau_i)\right] / (1 + \exp\left[\xi_i(t+1-\tau_i)\right])$ (for $i = 1, ..., 4$). Note that $\kappa_{i1} = 1$ (for $i = 1, ...4$).

Some technical notes:

1. To apply the method proposed by Harvey et al. (1992) the conditional distribution of the error η_{t+1} is assumed to be Gaussian. The unconditional distribution is of course not normal (see Hamilton 1994, p662).

2. Consistent estimation necessitates system matrices Z_{t+1}, T_{t+1}, K_{t+1}, and Q_{t+1} that are either constant, exogenous or predetermined (see Hamilton 1994, chapter 13) which is the case in our model.

3. The system is initialized with the matrices A_1 and P_1 which, given the assumption of stationary states, contain the unconditional means and variances of the states.

4. The time-varying conditional variances h_{it+1} and h_{wt+1} complicate the linear Gaussian state space framework. To deal with this we follow the approach by Harvey et al. (1992) and we include the shocks ε_{it+1} and ε_{wt+1} in the state vector. We note then that h_{it+1} (for $i = 1, ..., N$) and h_{wt+1} and therefore Q_{t+1} are functions of the unobserved states ε_{it} and ε_{wt}. Harvey et al. (1992) replace h_{it+1} and h_{wt+1} in the system by $h^*_{it+1} = \delta^a_i + \delta^b_i \varepsilon^{*2}_{it} + \delta^c_i h^*_{it}$ and $h^*_{wt+1} = \delta^a_w + \delta^b_w \varepsilon^{*2}_{wt} + \delta^c_w h^*_{wt}$ where the unobserved ε^2_{it} and ε^2_{wt} are replaced by their conditional expectations $\varepsilon^{*2}_{it} = E_t \varepsilon^2_{it}$ and $\varepsilon^{*2}_{wt} = E_t \varepsilon^2_{wt}$. Note that $E_t \varepsilon^2_{it} = [E_t \varepsilon_{it}]^2 + \left[E_t(\varepsilon_{it} - E_t \varepsilon_{it})^2\right]$ and $E_t \varepsilon^2_{wt} = [E_t \varepsilon_{wt}]^2 + \left[E_t(\varepsilon_{wt} - E_t \varepsilon_{wt})^2\right]$ where the quantities between square brackets are period t Kalman filter output (conditional means and variances of the states ε_{it} and ε_{wt}). Thus, given h^*_{it} and h^*_{wt} (which are initialized by the unconditional variances of ε_{it} and ε_{wt}, i.e. $\sigma^2_{\alpha i}$ and σ^2_w) and given the Kalman filter output from period t, namely $E_t(S_t)$ and $V_t(S_t)$, we can calculate h^*_{it+1} and h^*_{wt+1} and the system matrix Q_{t+1} which makes it possible to calculate $E_t(S_{t+1})$, $V_t(S_{t+1})$ and $E_{t+1}(S_{t+1})$, $V_{t+1}(S_{t+1})$, and so on... .

5. The time-varying conditional covariances g_{it+1} and time-varying β_{it+1} further complicate the linear Gaussian state space framework. Note that g_{it+1} and β_{it+1} can be replaced by $g^*_{it+1} = \gamma^a_i + \gamma^b_i \beta^*_{it} \varepsilon^{*2}_{wt} + \gamma^c_i g^*_{it}$ and $\beta^*_{it+1} = \frac{g^*_{it+1}}{h^*_{wt+1}}$ where ε^{*2}_{wt} and $h^*_{wt+1}$are calculated as reported in 4. The variable h^*_{wt} is initialized by the unconditional variance of ε_{wt}, i.e. σ^2_w, while g^*_{it} is initialized by $\frac{\gamma^a_i}{1-\gamma^b_i-\gamma^c_i}$.

Tables and Figures

Table 1: Correlation matrix of corrected 10 yr government bond spreads versus Germany $\widetilde{R}_t$ (weekly data).

	Full sample				Sample after introduction euro			
	29-06-1991 to 04-06-2006				01-01-1999 to 04-08-2006			
	BE	FR	NL	IT	BE	FR	NL	IT
Belgium	1	-	-	-	1	-	-	-
France	0.2884	1	-	-	0.8780	1	-	-
the Netherlands	0.6106	0.1987	1	-	0.8891	0.8193	1	-
Italy	0.3095	-.4663	-.0735	1	0.8608	0.7978	0.7264	1

Table 2: Maximum likelihood estimation of the common factor model with GARCH errors and convergence effects (eqs. 6-15).

	Country-specific parameters				Common parameters
	Belgium	France	Netherlands	Italy	
π	0.9683	0.9830	0.7110	0.9938	0.9729
	(0.0084)	(0.0060)	(0.0565)	(0.0043)	(0.0065)
μ	0.0030	-0.0007	-0.0035	0.0012 [a]	0.0787
	(0.0010)	(0.0008)	(0.0017)	(0.0014)	(0.0348)
ξ	-0.0063	-0.0181	-0.0090	-0.0081	-
	(7.4E-8)	(0.0042)	(0.0035)	(0.0017)	-
τ	788.99	715.70	687.56	615.19	-
	(4.5E-5)	(22.815)	(121.32)	(0.168)	-
δ^a	1.9E-5	0.0001	3.3E-5	9.9E-6	0.0008 [b]
	(7.0E-6)	(2.9E-5)	(1.2E-5)	(1.1E-5)	(0.0006)
δ^b	0.1882	0.3601	0.3681	0.1129	0.0923
	(0.0518)	(0.0643)	(0.0919)	(0.0133)	(0.0144)
δ^c	0.7701	0.5492	0.5950	0.8848	0.9068
	(0.0489)	(0.0629)	(0.0819)	(0.0142)	(0.0143)
γ^a	9.4E-6	4.0E-9	1.9E-5	3.1E-5	-
	6.2E-6	8.3E-9	9.4E-6	1.6E-5	-
γ^b	0.0776	0.0777	0.0884	0.0967	-
	0.0123	0.0121	0.0114	0.0125	-
γ^c	0.9220	0.9222	0.9107	0.9016	-
	0.0122	0.0121	0.0113	0.0123	-
			Goodness of fit		
AIC^c			-18.2940		
AIC^n			-18.2408		

Note: Hessian based standard errors between brackets. [a] For Italy the point estimate and standard error of μ are obtained from the restriction $\mu_4 = -\mu_1 - \mu_2 - \mu_3$. [b] For the common state the point estimate and standard error of δ^a are obtained from the restriction $\delta^a = 1 - \delta^b - \delta^c$. AIC^c denotes the Akaike information criterion for the model with convergence effects while AIC^n denotes the value of this statistic if it assumed that there is no convergence, i.e. $\xi = 0 \ \forall i$. A model with a smaller AIC is preferred. For the calculation we refer to Durbin and Koopman (2001, p.152).

Figure 1: Corrected 10 yr government bond spreads versus Germany

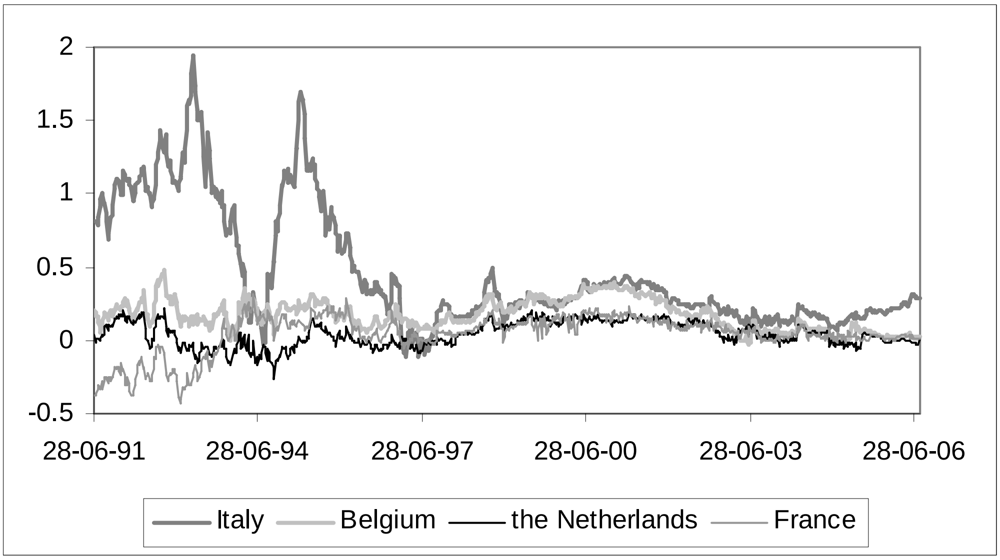

Figure 2: Spread for Belgium: idiosyncratic state α_{it} and common state $\widetilde{R}_{wt}$ (multiplied by β_{it})

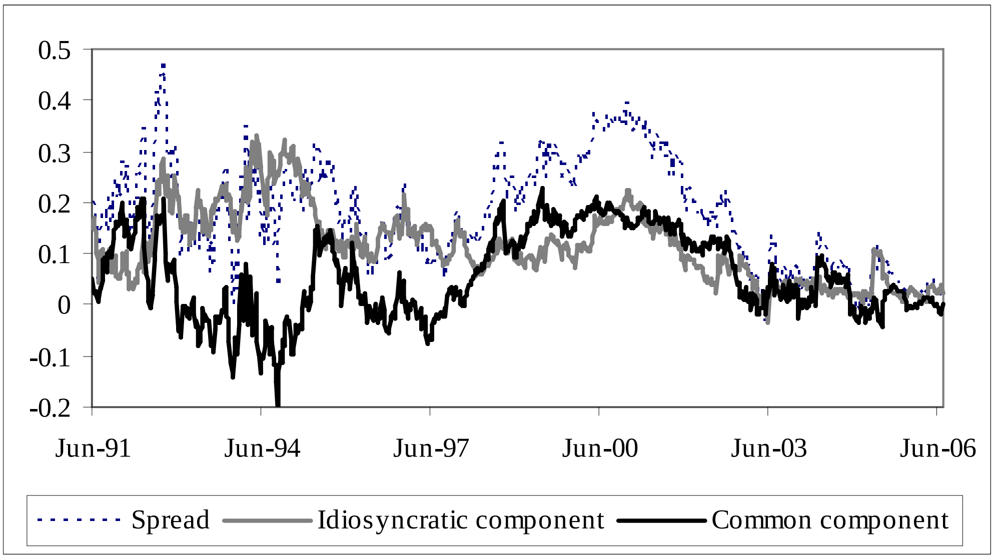

Figure 3: Spread for France: idiosyncratic state α_{it} and common state $\widetilde{R}_{wt}$ (multiplied by β_{it})

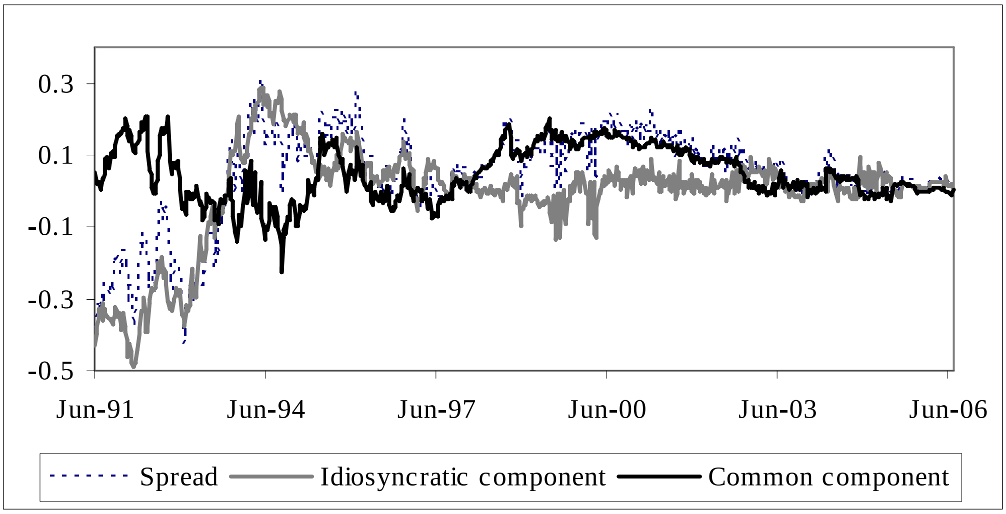

Figure 4: Spread for the Netherlands: idiosyncratic state α_{it} and common state $\widetilde{R}_{wt}$ (multiplied by β_{it})

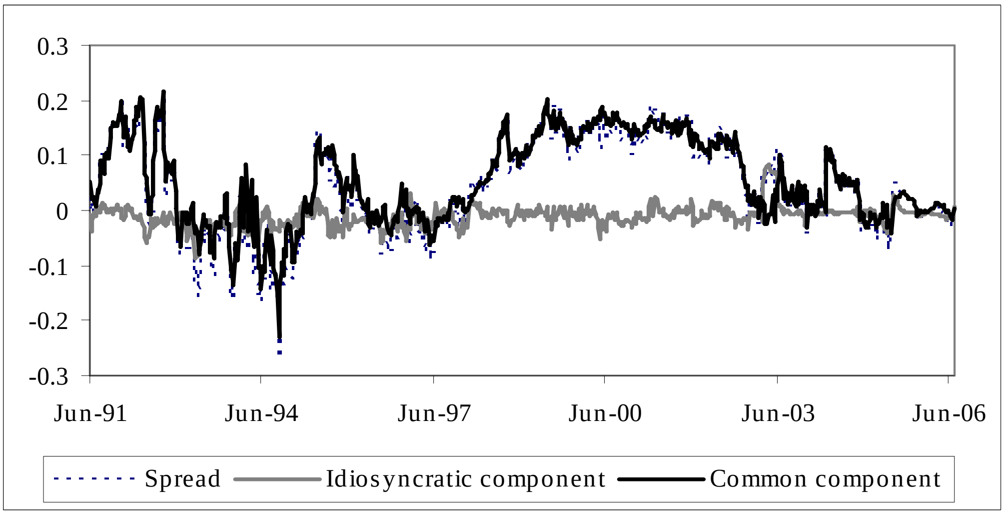

Figure 5: Spread for Italy: idiosyncratic state α_{it} and common state $\widetilde{R}_{wt}$ (multiplied by β_{it})

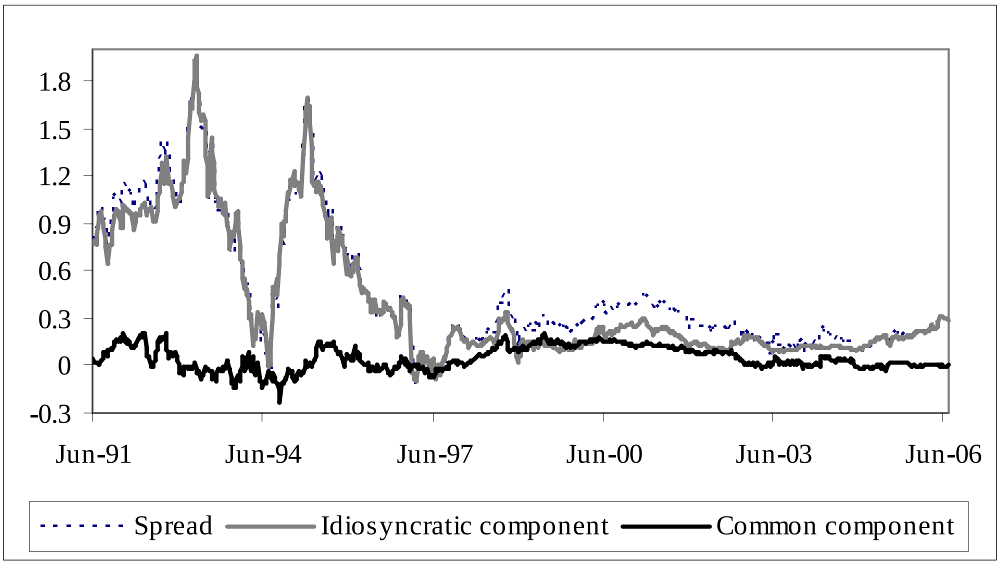

Figure 6: Idiosyncratic state α_{it} for all countries

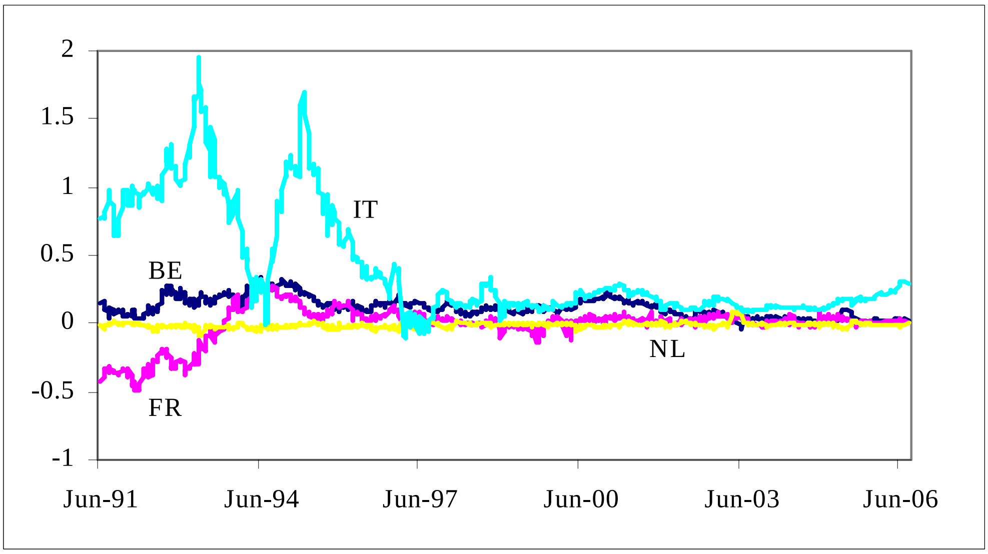

Figure 7: Time-varying β_{it} for all countries

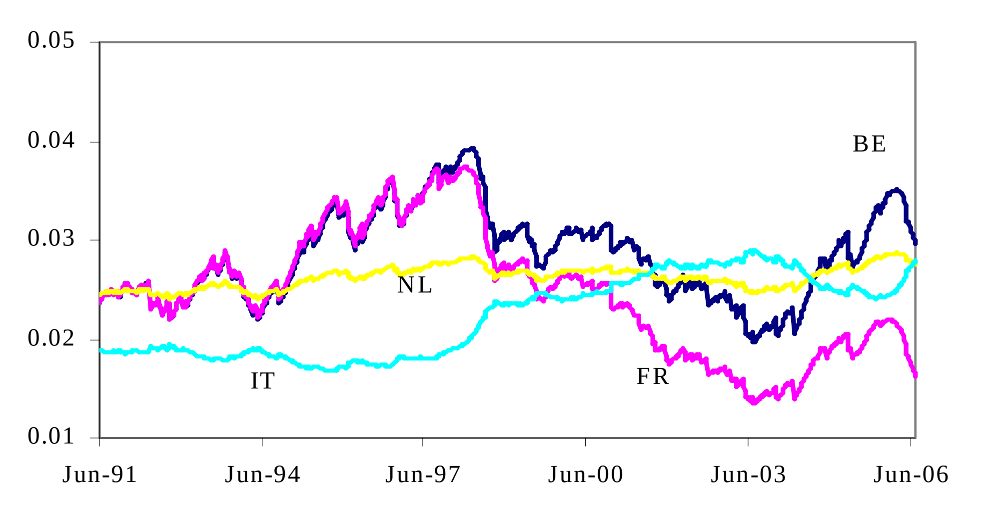

Figure 8: Convergence dynamics (convergence operators κ_{it}) of the idiosyncratic components α_{it}

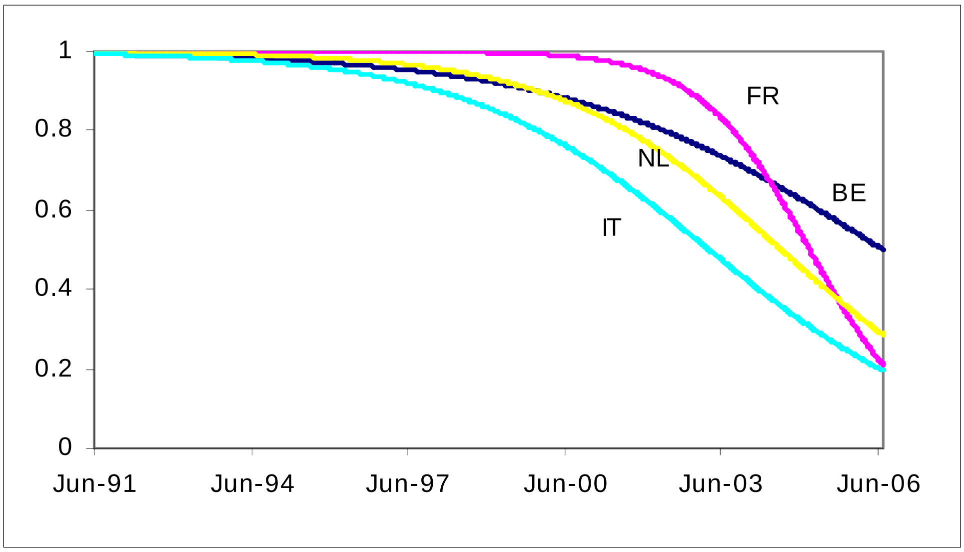

Figure 9: GARCH series h_{it} for idiosyncratic states α_{it}