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Volatility-Spillover Effects in European Bond Markets

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Abstract

Volatility spillover from the US and aggregate European bond markets into individual European bond markets using a GARCH volatility-spillover model is analysed. Strong statistical evidence of volatility spillover from the US and aggregate European bond markets is found. For EMU countries, the US volatility-spillover effects are rather weak (in economic terms) whereas the European volatility-spillover effects are strong. The bond markets of EMU countries have become much more integrated after the introduction of the euro, and in recent years they have become close to being perfectly integrated. The main driver of the integration appears to be convergence in interest rates.

Keywords: *euro introduction*; *government bonds*; *integration of bond markets*; *international bond markets*; *volatility spillover*

JEL classification: *E*43; *F*36; *G*12; *G*15

1. Introduction

This paper is concerned with the effects of volatility spillover from the US and aggregate European bond markets to a number of European bond markets; Belgium, Denmark, France, Germany, Italy, the Netherlands, Spain, Sweden and the UK. A volatility-spillover model is applied to separate the shock to the individual country return into three effects; local (own country), regional (aggregate Europe) and global (US). In contrast to related volatility-spillover studies, we focus on international *bond* markets and thereby provide interesting new insights into the nature of the volatility of international bond markets. In particular, we document the proportion of the volatility of the individual country's bond market that the government can hope to exert political influence on. Furthermore, our analysis of bond market volatility spillover has strong implications

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for investors' optimal asset allocation decisions. We analyse *European* bond markets in an outstanding time period, namely around the introduction of the common currency, the euro. We are therefore also able to provide new information about the impact of the euro event. Finally, we add to the literature by analysing contagion effects into *European bond* markets.

Previous literature also analyses the interdependence of international bond markets but fails to separate the volatility into local, regional, and global effects as we do here. Ilmanen (1995) uses a linear regression model with local and global instruments to forecast the excess returns of long-term international bonds. The world factors turn out to be the most important factors. Clare and Lekkos (2000) investigate the interaction between the US, UK, and German bond markets in a VAR model for the short rates and the term-structure slopes. The variances of the term-structure slopes are found to be influenced by international factors as well as by local factors. The results of Ilmanen (1995) and Clare and Lekkos (2000) indicate that it is important to distinguish between local effects and other effects. Driessen *et al.* (2003) investigate the common factors in the US, German and Japanese bond markets using principal components analysis. They find that the positive correlation between bond markets is driven by the term structure levels (both world and local), not by the term structure slopes. As a consequence, we investigate interest rate levels not slopes. Laopodis (2004) applies a VAR model to describe the long-term interest rates of eight countries. He finds that markets have become more integrated through the 1990s, which has narrowed the scope for conducting monetary policy. We discuss this issue further.

The volatility-spillover analysis was initiated by Engle *et al.* (1990). Lin *et al.* (1994) investigate the volatility spillover between the US and Japanese stock markets. Daytime return and volatility in one market is correlated with overnight return and volatility in the other market. Bekaert and Harvey (1997), Ng (2000), Bekaert *et al.* (2005) and Baele (2005) investigate volatility-spillover effects on various equity markets using similar volatility-spillover models. They all find evidence of volatility-spillover effects. Bekaert and Harvey (1997) investigate the volatility of emerging stock markets. They distinguish between global and local shocks. Ng (2000) finds evidence of volatility-spillover effects to various Pacific Basin stock markets from Japan (regional effects) and the US (global effects). Baele (2005) investigates the volatility-spillover effects from the US (global effects) and aggregate European (regional effects) stock markets into various individual European stock markets.

We apply a volatility-spillover model which has been applied successfully to stock markets, cf. e.g. Bekaert *et al.* (2005). Modelling volatility-spillover effects for bond markets in this way has received no attention in the existing literature, which we hereby hope to make up for. First we model the returns of the US and European indices. The returns on the individual-country indices are assumed to follow AR-GARCH processes that are extended to include volatility-spillover effects. The conditional volatility of the unexpected return is divided into the proportion caused by US, European and own country effects. We use three model specifications that differ in how they describe the volatility-spillover functions. In the *constant spillover* model, the spillover parameters are constant. In the *euro spillover* model the spillover parameters are different before and after the introduction of the euro. In the *trend* spillover model, the spillover parameters change by a certain amount each year. Both the euro spillover model and the trend spillover model are new to the literature.

A novel finding is that all empirical results group the countries into the EMU-member countries plus Denmark (EMU countries) and the non-EMU member countries. In the constant spillover model, we find strong evidence of volatility-spillover effects from both the US and aggregate Europe to the individual bond markets. For the EMU countries the proportion of the variance of the unexpected return caused by US effects is rather small, whereas the proportion caused by European effects is quite large. The opposite applies to non-EMU countries. The own-country effects are larger for non-EMU countries than for EMU countries. This indicates that there is more room for success for economic policy that is directed at local bond markets for non-EMU countries than for EMU countries.

Subsequently, in the euro spillover model we account for the changes brought about by the introduction of the euro. For EMU countries the US volatility-spillover effects are stronger before the euro than after the euro whereas the European volatility-spillover effects are stronger after the euro than before. The local volatility effects are much stronger before the euro than after. Thus, EMU countries have become much more integrated after the euro. The results for non-EMU countries are not susceptible to the introduction of the euro. As expected, EMU countries are more integrated than non-EMU countries.

In the trend spillover model, we find even stronger evidence that EMU countries have become almost fully integrated during the sample period. Moreover, we find that the main reason for the integration of the bond markets is convergence in interest rates. Convergence in inflation rates and exchange rate stability also have some bearing on the integration. There is weak evidence that differences in liquidity is one of the reasons why EMU bond markets are not fully integrated, whereas differences in credit risk do not appear to be of any relevance.

Contagion effects in bond markets has not been analysed before. We find weak evidence of contagion effects from the US bond market into individual European bond markets, and only some evidence of contagion from the aggregate European bond market into individual European bond markets.

The remaining part of the paper is structured as follows: Section 2 provides a detailed description of the volatility-spillover model. In Section 3 the bond market data are described and the empirical findings are discussed in Section 4. Section 5 concludes.

2. The Volatility-Spillover Framework

Our empirical volatility-spillover model is based on the models specified by Bekaert and Harvey (1997), Ng (2000), Bekaert *et al.* (2005), and Baele (2005), who all consider volatility-spillover effects on international stock markets, whereas we consider international bond markets.

2.1. *The model*

The volatility-spillover model is estimated in three steps. We structure our presentation around these three steps. In the first step, a univariate AR-GARCH model is estimated for the US return. In the second step, an extended univariate AR-GARCH model is estimated for the aggregate European return. The US residual from the first-step regression is included as an explanatory variable. Finally, in the third step, we estimate a univariate extended AR-GARCH model for the return of the individual European country. Both the US and European residuals from the two first steps are included as explanatory variables. In this way, the model accommodates two sources of volatility-spillover effects. The multi-step estimation procedure provides consistent, but not necessarily efficient, estimates.

The three-step estimation procedure is equivalent to the method applied by Bekaert *et al.* (2005) which represents only practical differences to the two-step estimation procedure applied by Ng (2000) and Baele (2005).[1] The subsequent analysis (of what will be denoted variance ratios) relies on using US and European *idiosyncratic* shocks as explanatory variables. Thus, the recursive model structure applied in the previous literature is necessary.

We let the causality go from the US to Europe. Apart from the intuitive appeal of moving from world effects to regional effects, the fact that we include spillover effects from the US to Europe and not the other way round is supported by Granger causality tests with four lags, cf. Granger (1969). The European return fails to Granger cause the US return, whereas the US return Granger causes the European return.[2,3]

US bond returns. The conditional return on the US index is assumed to evolve according to an AR(1) process:

$$R_{US,t} = c_{0,US} + c_{1,US} R_{US,t-1} + e_{US,t}. \tag{1}$$

The idiosyncratic shock ($e_{US,t}$) is normally distributed with mean 0 and the conditional variance follows the GARCH(1,1) specification, cf. Engle (1982) and Bollerslev (1986):

$$\sigma^2_{US,t} = \omega_{US} + \alpha_{US} e^2_{US,t-1} + \beta_{US} \sigma^2_{US,t-1} \tag{2}$$

where $\omega_{US} > 0$ and $\alpha_{US}, \beta_{US} \geq 0$ to make sure the variance is positive, and $\alpha_{US} + \beta_{US} \leq 1$ to ensure stationarity.[4] We use a symmetrical variance process because we do not find any signs of asymmetry in this or in any other variance processes.

European bond returns. The return on the European index is assumed to be described by the following extended AR(1) specification:

$$R_{E,t} = c_{0,E} + c_{1,E} R_{E,t-1} + \gamma_{E,t-1} R_{US,t-1} + \phi_{E,t-1} e_{US,t} + e_{E,t}. \tag{3}$$

The conditional mean of the European return depends on its own lagged return as well as the lagged US return. The mean spillover effects are introduced by the lagged

[1] The two-step estimation procedure would be as follows: in the first step, a bivariate GARCH model for US and European returns is estimated. There are no volatility-spillover effects between the US and Europe. In an intermediate step, US and European residuals are orthogonalised: the European innovation is driven by a purely idiosyncratic shock and the US idiosyncratic shock whereas the US innovation is purely idiosyncratic. In the second step, the US and European idiosyncratic shocks are used as explanatory variables in the model for the return of the individual country. So, in fact, the two-step procedure also consists of three steps. The three-step estimation procedure distinguishes itself by allowing volatility-spillover effects from the US to Europe and avoiding orthogonalisation.

[2] Unless otherwise stated, a 5% level of significance is applied throughout.

[3] We also conduct the Granger causality test pairwise to the US return on the one side and the European indices each excluding one country on the other side. The Granger causality still goes from the US to Europe and not the other way; except for the index excluding the UK for which neither index Granger causes the other.

[4] As pointed out by an anonymous referee it is important to investigate potential structural volatility breaks. The US conditional volatility is not subject to a structural break at the introduction of the euro: ω_{US} is insignificantly different before and after the euro. Below we model other conditional variances. Neither of them is subject to structural breaks.

US return, $R_{US,t-1}$. The volatility spillover from the US to Europe takes place via the penultimate term, $e_{US,t}$. Thus, the European return depends on the US idiosyncratic shock. We return to why this represents a volatility-spillover effect in Section 2.2. In the practical estimation, the residual from equation (1) is used in place of $e_{US,t}$. The idiosyncratic shock ($e_{E,t}$) has mean 0 and the conditional variance evolves according to the GARCH(1,1) model described in equation (2).

Country i bond returns: The last step consists of providing a model for the individual country returns. The mean specification for the European return in equation (3) is extended even further. For country i ($i = 1, \ldots, N$):

$$R_{i,t} = c_{0,i} + c_{1,i} R_{i,t-1} + \gamma_{i,t-1} R_{US,t-1} + \delta_{i,t-1} R_{E,t-1} + \phi_{i,t-1} e_{US,t} + \psi_{i,t-1} e_{E,t} + e_{i,t} \tag{4}$$

The conditional country i mean return depends on the lagged US, European, and own return. This specification allows mean spillover effects from both the US and European returns to the individual countries by the lagged returns $R_{US,t-1}$ and $R_{E,t-1}$. Volatility-spillover effects from the US and Europe to the individual countries are introduced by the idiosyncratic US and European shocks, $e_{US,t}$ and $e_{E,t}$. Shortly, it will become clear exactly why this corresponds to volatility-spillover effects, cf. equation (8) below. In the estimation, the residuals from equations (1) and (3) are applied as explanatory variables. The idiosyncratic country shocks are subject to the same distributional assumptions as the other idiosyncratic shocks; they have mean 0 and the conditional volatilities follow the GARCH(1,1) specification, cf. equation (2).

2.2. *Variance ratios*

The idiosyncratic shocks $e_{US,t}$, $e_{E,t}$, and $e_{i,t}$ (for $i = 1, \ldots, N$) are assumed to be independent. This, however, does not apply to the unexpected returns:

$$\epsilon_{US,t} = e_{US,t} \tag{5}$$

$$\epsilon_{E,t} = \phi_{E,t-1} e_{US,t} + e_{E,t} \tag{6}$$

$$\epsilon_{i,t} = \phi_{i,t-1} e_{US,t} + \psi_{i,t-1} e_{E,t} + e_{i,t} \tag{7}$$

It follows from equation (7) that the conditional variance of the unexpected return of country i based on the information available at time $t-1$ (I_{t-1}) is given as follows:

$$h_{it} = E\left(\epsilon_{i,t}^2 \mid I_{t-1}\right) = \phi_{i,t-1}^2 \sigma_{US,t}^2 + \psi_{i,t-1}^2 \sigma_{E,t}^2 + \sigma_{i,t}^2. \tag{8}$$

The conditional variance of the unexpected return for country i depends on the variance of the contemporary US, European, and own idiosyncratic shocks. When for example the US idiosyncratic volatility is large, the volatility of the unexpected returns for country i also tends to be large if $\phi_{i,t-1}$ is significant. This is what we denote volatility-spillover effects. So, the significance of the parameters $\phi_{i,t-1}$ and $\psi_{i,t-1}$ determine whether volatility-spillover effects from the US and Europe, respectively, are present in country i. The conditional variance of the European unexpected return depends only on the US and its own idiosyncratic volatility, cf. equation (6). So, there is volatility-spillover from the US to the aggregate European bond market. The conditional variance of the US unexpected return is equal to the variance of the US idiosyncratic shock, cf. equation (5).

We measure the proportion of the variance of the unexpected return of country i, cf. equation (8), that is caused by the US and European volatility-spillover effects, respectively. To this end, we define the following variance ratios:

$$VR_{i,t}^{US} = \frac{\phi_{i,t-1}^2 \sigma_{US,t}^2}{h_{i,t}} \tag{9}$$

$$VR_{i,t}^{E} = \frac{\psi_{i,t-1}^2 \sigma_{E,t}^2}{h_{i,t}}. \tag{10}$$

The variance ratios take on values between 0 and 1. The remaining part of the variance of the unexpected return for country i is caused by pure local effects;

$$VR_{i,t}^{i} = 1 - VR_{i,t}^{US} - VR_{i,t}^{E} = \frac{\sigma_{i,t}^2}{h_{i,t}}. \tag{11}$$

The variance ratios provide a measure of the impact of global, regional, and local effects on the local variance.

The specification of the functions for the spillover parameters, γ_t, δ_t, ϕ_t, and ψ differentiates the volatility-spillover models. We use three different specifications. First, the spillover parameters are assumed to be constant throughout the entire sample period; the *constant spillover* model:

$$\begin{aligned} \gamma_{i,t} &= \gamma_i \ \forall t \\ \delta_{i,t} &= \delta_i \ \forall t \\ \phi_{i,t} &= \phi_i \ \forall t \\ \psi_{i,t} &= \psi_i \ \forall t. \end{aligned} \tag{12}$$

Second, the spillover parameters are assumed to be constant before and after the introduction of the euro; the *euro spillover* model:

$$\begin{aligned} \gamma_{i,t} &= \gamma_{0,i} + \gamma_{1,i} D_t \\ \delta_{i,t} &= \delta_{0,i} + \delta_{1,i} D_t \\ \phi_{i,t} &= \phi_{0,i} + \phi_{1,i} D_t \\ \psi_{i,t} &= \psi_{0,i} + \psi_{1,i} D_t. \end{aligned} \tag{13}$$

The dummy variable D_t equals 0 before 1 January 1999 (i.e. before the introduction of the euro) and 1 after 1 January 1999. Using a dummy variable is the simplest way to capture the changes produced by the introduction of the euro. We contribute to the general volatility spillover modelling literature by introducing the euro spillover model.

Third, the spillover parameters undergo a gradual transition by taking on a different value each year in the sample, the *trend spillover* model:

$$\begin{aligned} \gamma_{i,t} &= \gamma_{0,i} + \gamma_{1,i} DT_t \\ \delta_{i,t} &= \delta_{0,i} + \delta_{1,i} DT_t \\ \phi_{i,t} &= \phi_{0,i} + \phi_{1,i} DT_t \\ \psi_{i,t} &= \psi_{0,i} + \psi_{1,i} DT_t. \end{aligned} \tag{14}$$

The function DT_t equals 1 for 1988 observations, 2 for 1989 observations, etc.[5] The trend spillover model is novel to the literature.

3. The Data

We apply the JPMorgan total return government bond indices extracted from DataStream for the US, Europe, six EU EMU-member countries (Belgium, France, Germany, Italy, the Netherlands and Spain), and three EU non-EMU member countries (Denmark, Sweden and the UK). Total return implies that the received coupons are re-invested into the bonds of the index. Log-returns are calculated as the logarithmic growth rate of the indices.[6] All returns are measured in a common currency, namely DEM.[7] The weekly data (sampled on Wednesdays) cover the period from 6 January 1988 to 27 November 2002, thus providing us with a total of 777 observations. The average duration of the 11 indices over the sample period is 4.86 years. The mean durations of all indices are within plus/minus one standard error of the overall average duration.

Using weekly data (compared to for example daily data) partially overcomes the potential problem of non-synchronous data, cf. Burns and Engle (1998). In stable periods, close-to-close (i.e. non-synchronous) daily returns on international stock markets tend to underestimate correlations, cf. Martens and Poon (2001). If this carries over to weekly returns on international bond markets, it might suggest that the hypothesis of no spillover effects is accepted too often. This would mean that the proposition of spillover effects is not wrongly accepted on this account.

The European index is a (market value) weighted average of the indices of the nine European countries. Thus, there is spurious correlation between the individual country return and the European return caused by the same bonds being included in both indices. Therefore, for each country, we use the return of an artificial European index excluding the country itself, cf. Bekaert *et al.* (2005). We use the weights that JPMorgan applied in January 2002, cf. Diamond *et al.* (2002).[8] The artificial European indices are strongly correlated with the JPMorgan European index (the correlation coefficients are in excess of 95%). The empirical results are robust to using the artificial European indices and the JPMorgan European index. For the European equation we only show the JPMorgan European index results. Yet, the individual country equations rely on the artificial European indices.

Table 1 contains summary statistics for the returns of the nine country indices as well as the US and European indices. The average weekly returns fall within a fairly tight range, from 0.13% to 0.19%. The variability of the returns is much more dispersed across the indices; the standard deviation of the weekly return falls in the interval between 0.48% and 1.53%. Generally, the return of the index tends to be more variable the higher its average return. Apart from the US, the return distributions are skewed to the left, and all the distributions show excess kurtosis. Therefore the Jarque and

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[5] We use the yearly step function to facilitate intuitive interpretations of the results compared to the ordinary trend which equals 1 for observation 1, 2 for observation 2, etc.

[6] We follow the previous literature in applying log-returns of total return government bond indices, cf. e.g. Bodart and Reding (1999).

[7] Volatility-spillover analysis generally applies common currency returns, cf. e.g. Bekaert *et al.* (2005).

[8] We only have access to this set of weights.

Table 1

Summary statistics

The table reports the summary statistics for the weekly returns (in %) of the JPMorgan government bond indices for the US, Europe (Eu), Belgium (Be), Denmark (De), France (Fr), Germany (Ge), Italy (It), the Netherlands (Ne), Spain (Sp), Sweden (Sw) and the UK. The following statistics are reported: mean, standard deviation, skewness, kurtosis, autocorrelation (order 1 and 4), and autocorrelation of the squared time series (order 1 and 4). * (§) [♯] indicates that the Ljung and Box (1978) test statistic is significant at a 1% (5%) [10%] level of significance.

	Mean	Stdev.	Skew.	Kurt.	AC(1)	AC(4)	$AC^2(1)$	$AC^2(4)$
US	0.18	1.53	0.13	3.91	−0.05	0.04	0.03	0.05*
Eu	0.15	0.59	−0.57	5.38	−0.02	−0.00	0.08§	0.09*
Be	0.16	0.58	−0.37	5.21	−0.05	−0.09*	0.20*	0.19*
De	0.17	0.69	−0.39	6.18	0.05	0.01	0.18*	0.14*
Fr	0.16	0.66	−0.11	4.65	−0.02	−0.04	0.11*	0.13*
Ge	0.13	0.48	−0.65	4.65	0.03	−0.01§	0.14*	0.11*
It	0.18	1.23	−1.18	15.0	0.06	−0.02*	0.20*	0.11*
Ne	0.13	0.53	−0.63	4.63	0.00	0.06§	0.11*	0.14*
Sp	0.17	0.97	−0.73	9.78	−0.04	−0.04	0.14*	0.12*
Sw	0.15	1.39	−0,42	5.66	−0.06	0.10§	0.19*	0.07*
UK	0.19	1.34	−0.12	3.94	−0.02	0.01	0.05	−0.02

Bera (1980) test rejects normality for all the series (test statistics not shown). Some of the return series are significantly autocorrelated (Ljung and Box (1978) tests for first order and up to fourth order autocorrelation). Thus, there are indications that the AR specification is useful. The squared return series are significantly autocorrelated, again, with the exception of the UK. The presence of heteroskedasticity motivates the GARCH specifications.

4. Empirical Findings

We estimate the model using the Quasi Maximum Likelihood (QML) method with (univariate) Gaussian likelihood functions. The estimation is conducted using a combination of the Berndt *et al.* (1974) and the Newton-Raphson numerical optimization algorithm.

4.1. *Constant spillover model*

Table 2 reports the results from estimating the constant spillover model.[9] The US results are shown in the first row of Table 2. The AR(1) parameter ($\hat{c}_{1,US}$) is small, negative, and insignificant, which implies no (or weak negative) first-order autocorrelation, consistent with the summary statistics reported in Table 1. The volatility process is quite persistent in that $\hat{\alpha}_{US} + \hat{\beta}_{US}$ roughly equals 0.9.

The second row of Table 2 reports the results for the European index.[10] Own lagged and US lagged returns are of minor importance to the conditional mean; $\hat{c}_{1,E}$ and $\hat{\gamma}_E$

[9] The estimation is conducted in GAUSS using the Constrained Maximum Likelihood module.

[10] Notice that only the results for the European index including all nine countries are reported. The results for the nine artificial European indices excluding one country each are similar.

Table 2

Constant spillover model

The table reports the results from estimating the constant spillover model. US return: $R_{US,t} = c_{0,US} + c_{1,US}R_{US,t-1} + e_{US,t}$ where $e_{US,t}$ has mean 0 and conditional variance: $\sigma^2_{US,t} = \omega_{US} + \alpha_{US}e^2_{US,t-1} + \beta_{US}\,\sigma^2_{US,t-1}$. European return: $R_{E,t} = c_{0,E} + c_{1,E}R_{E,t-1} + \gamma_E R_{US,t-1} + \phi_E e_{US,t} + e_{E,t}$ where $e_{E,t}$ has mean 0 and conditional variance: $\sigma^2_{E,t} = \omega_E + \alpha_E e^2_{E,t-1} + \beta_E\sigma^2_{E,t-1}$. Country i return (the European index excluding country i is applied): $R_{i,t} = c_{0,i} + c_{1,i}R_{i,t-1} + \gamma_i R_{US,t-1} + \delta_i R_{E,t-1} + \phi_i e_{US,t} + \psi_i e_{E,t} + e_{i,t}$ where $e_{i,t}$ has mean 0 and conditional variance: $\sigma^2_{i,t} = \omega_i + \alpha_i e^2_{i,t-1} + \beta_i\,\sigma^2_{i,t-1}$. Bollerslev and Wooldridge (1992) robust standard errors in parentheses. * (§) [♯], indicates that the value is significant at a 1% (5%) [10%] level of significance.

	$c_{0,i}$	$c_{1,i}$	γ_i	δ_i	ϕ_i	ψ_i	$100\omega_i$	α_i	β_i
US	0.165*	−0.040					23.003	0.047♯	0.854*
	(0.055)	(0.036)					(18.238)	(0.026)	(0.091)
Eu	0.141*	0.015	0.007		0.086*		0.888♯	0.069*	0.903*
	(0.019)	(0.037)	(0.013)		(0.018)		(0.493)	(0.022)	(0.031)
Be	0.136*	−0.124*	0.008	0.127*	0.052*	0.863*	0.097	0.105§	0.895*
	(0.014)	(0.045)	(0.008)	(0.047)	(0.011)	(0.041)	(0.114)	(0.049)	(0.049)
De	0.144*	−0.062	0.002	0.131*	0.057*	0.805*	0.213	0.157§	0.843*
	(0.011)	(0.052)	(0.008)	(0.048)	(0.008)	(0.028)	(0.217)	(0.075)	(0.075)
Fr	0.137*	−0.104§	0.002	0.128*	0.042*	0.893*	0.386	0.241	0.759*
	(0.013)	(0.043)	(0.009)	(0.046)	(0.009)	(0.045)	(0.503)	(0.148)	(0.148)
Ge	0.111*	0.037	0.019♯	−0.001	0.028§	0.568*	0.290	0.125♯	0.868*
	(0.014)	(0.040)	(0.010)	(0.035)	(0.012)	(0.074)	(0.284)	(0.069)	(0.063)
It	0.161*	−0.036	0.003	0.029	0.066*	0.933*	0.103	0.151*	0.849*
	(0.011)	(0.051	(0.009)	(0.050)	(0.009)	(0.026)	(0.739)	(0.042)	(0.042)
Ne	0.119*	−0.108♯	0.013♯	0.125♯	0.046*	0.902*	0.079	0.120*	0.880*
	(0.018)	(0.061)	(0.008)	(0.066)	(0.014)	(0.038)	(0.072)	(0.031)	(0.031)
Sp	0.151*	−0.122§	−0.003	0.147*	0.047*	0.992*	0.080	0.130*	0.870*
	(0.010)	(0.052)	(0.009)	(0.057)	(0.006)	(0.022)	(0.064)	(0.039)	(0.039)
Sw	0.183*	−0.063	0.016	0.060	0.266*	1.071*	2.405♯	0.103*	0.882*
	(0.033)	(0.042)	(0.027)	(0.092)	(0.023)	(0.087)	1.305)	(0.025)	(0.028)
UK	0.182*	−0.021	−0.022	0.031	0.413*	0.721*	2.098	0.036	0.946*
	(0.039)	(0.040)	(0.034)	(0.072)	(0.052)	(0.092)	(3.283)	(0.029)	(0.052)

are small, positive and insignificant.[11] In contrast, the contemporaneous US residual is significant in explaining the current mean value ($\hat{\phi}_E$ is significant). Thus, there is evidence of volatility spillover from the US to the European bond market but no evidence of mean spillover. The robust joint Wald test for no US-spillover effects at all, $H_0: \gamma_E = \phi_E = 0$, is strongly rejected. The volatility process is highly persistent ($\hat{\alpha}_E + \hat{\beta}_E = 0.97$).

Lastly, the models for the individual countries are estimated. The models include mean and volatility-spillover effects from both the US and Europe. The bottom rows of

[11] As a robustness check, we redo the analysis excluding any insignificant parameters that enter into the first two steps of the estimation. In the first step, we impose the restriction that $c_{1,US} \equiv 0$, which does not alter the subsequent results. In the second step, we let $c_{1,E} \equiv \gamma_E \equiv 0$. This does not alter the empirical results for the individual countries (i.e. the third step).

Table 3

Variance ratios – constant spillover model

The table reports the mean and standard deviation of the US, European, and own variance ratios for the constant spillover model: $VR^{US}_{i,t} = \frac{\phi_i^2\sigma^2_{US,t}}{h_{i,t}}$, $VR^{E}_{i,t} = \frac{\psi_i^2\sigma^2_{E,t}}{h_{i,t}}$, and $VR^{i}{}_{i,t} = 1 - VR^{US}{}_{i,t} - VR^{E}{}_{i,t}$. $\sigma_{US,t}(\sigma_{E,t})$ is the conditional volatility of the US (European) idiosyncratic shock and $h_{i,t}$ is the conditional variance of the unexpected return of country i.

	$VR^{US}_{i,t}$		$VR^{E}_{i,t}$		$VR^{i}_{i,t}$	
	Mean	Stdev.	Mean	Stdev.	Mean	Stdev.
Be	0.019	0.011	0.605	0.172	0.375	0.178
De	0.024	0.014	0.551	0.192	0.425	0.200
Fr	0.012	0.006	0.615	0.178	0.373	0.181
Ge	0.008	0.005	0.459	0.118	0.532	0.119
It	0.022	0.016	0.441	0.281	0.536	0.295
Ne	0.015	0.008	0.653	0.170	0.332	0.175
Sp	0.011	0.008	0.544	0.259	0.445	0.265
Sw	0.119	0.060	0.221	0.077	0.659	0.113
UK	0.231	0.047	0.100	0.066	0.668	0.082

Table 2 provide the results. For all countries, the returns show negative or no first-order autocorrelation. The conditional volatility processes are highly persistent. In fact, in some cases the restriction that $\alpha_i + \beta_i \leq 1$ is binding (the exceptions are Germany, Sweden and the UK). This implies that some of the conditional variances evolve according to Integrated GARCH processes. Still, the conditional volatilities are not explosive.

The US mean-spillover parameter, $\hat{\gamma}_i$, is insignificant everywhere. For about half of the countries, there is evidence of European mean-spillover effects, i.e. $\hat{\delta}_i$ is positive and significant. The robust Wald test for the joint hypothesis of no mean-spillover effects ($H_0 : \gamma_i = \delta_i = 0$) leaves the conclusions unaltered; for about half of the countries (the same as before) there are significant mean-spillover effects. For all the countries there are significant volatility-spillover effects from both the US and Europe, i.e. $\hat{\phi}_i$ and $\hat{\psi}_i$ are individually significant. The robust Wald tests for the joint hypothesis of no volatility-spillover effects, $H_0 : \phi_i = \psi_i = 0$, is rejected for all countries. The results also lead us to reject the null hypotheses of no US-spillover effects ($H_0 : \gamma_i = \phi_i = 0$) as well as no European-spillover effects ($H_0 : \delta_i = \psi_i = 0$) for all countries. Finally, the null hypothesis of no spillover effects at all is rejected in all cases: $H_0 : \gamma_i = \delta_i = \phi_i = \psi_i = 0$.

To summarise, there are very strong indications of volatility-spillover effects from both the US and European bond markets into all the individual bond markets, less strong indications of mean-spillover effects from the European market, and no signs of mean-spillover effects from the US market.

So far, we have only discussed the significance of the spillover parameters. The relative size of the parameters is not particularly relevant for evaluating the quantitative influence of the US and European bond markets on the individual bond markets. To assess the importance of the US and European volatility-spillover effects on the variance of the unexpected return of country i, the time series of the variance ratios $VR^{US}_{i,t}$, $VR^{E}_{i,t}$ and $VR^{i}_{i,t}$ from equations (9)–(11) are calculated. Table 3 reports the mean of the variance

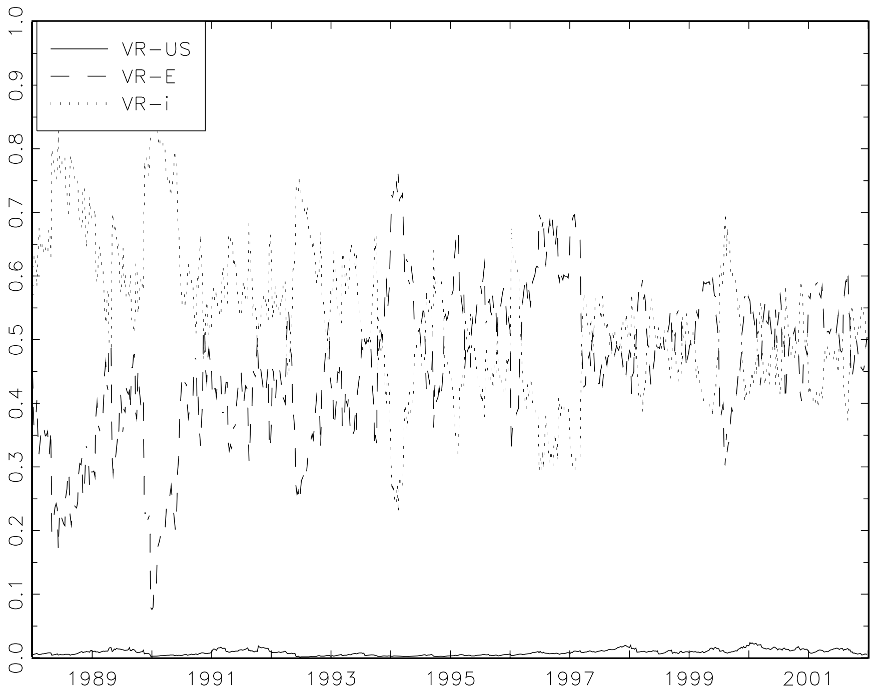

Fig. 1. Variance ratios – Germany – constant spillover model

ratios for each country. Over the period, on average, the US volatility-spillover effects make up between 0.8% and 23.1% of the conditional variance of the unexpected return of country *i*. For most countries the mean of the US variance ratio is around 1–2%, only in two cases is it considerably higher; Sweden (12%) and the UK (23%). It is remarkable that the US volatility-spillover effects are particularly strong for two out of three non-EMU member countries (Denmark is the exception). The average European variance ratios range between between 44% and 65% for the EMU-member countries plus Denmark, whereas the European variance ratios for Sweden and the UK are much smaller (means of 22% and 10%, respectively). Finally, the purely local volatility effects are larger for Sweden and the UK (means of 66% and 67%) than for the other countries (the means range between 33% and 54%). To all intents and purposes Denmark behaves like an EMU-member country because the Danish central bank pegs the Danish Krone to the euro.

Figures 1 and 2 show the time series evolution of the variance ratios for Germany and the UK (they are representative of EMU and non-EMU countries). The European variance ratio generally increases over the sample period for all countries except (again) Sweden and the UK, for which it appears to be stable.

The variance ratios show that it makes sense for governments to try to influence their local bond market variance by their policies. Moreover, governments might also be able to influence the bond market variance (indirectly) through the policies conducted by the European Union on behalf of EMU countries. Only a small fraction of the variability is outside the reach of government policy, namely the minor part that is influenced by US effects. Non-EMU countries are able to influence their bond markets variance even further than EMU countries.

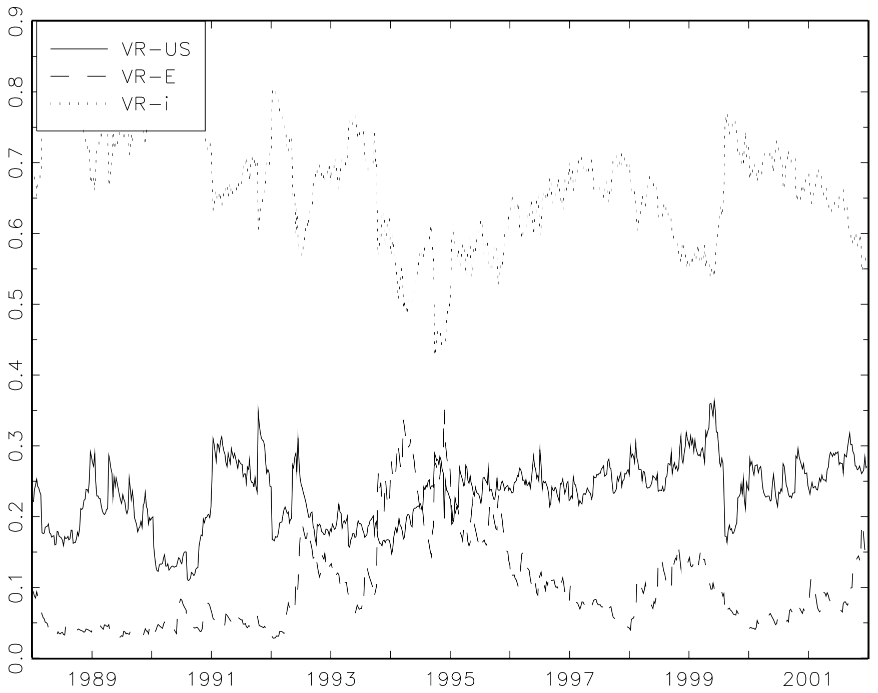

Fig. 2. Variance ratios – UK – constant spillover model

The results regarding the relative size of volatility-spillover effects from the US and aggregate European bond markets into EMU and non-EMU bond markets compare well with the results for the corresponding stock markets, cf. Baele (2005). Our findings appear to be in contrast to those of Ilmanen (1995) who finds that world factors are more important than local factors. Our results appear to be in accordance with those of Clare and Lekkos (2000) who find that both international and local factors influence bond returns.

4.2. *Euro spillover model*

The euro spillover model accounts for the changes brought about by the introduction of the euro by letting the spillover parameters take on different values before and after the introduction of the euro, cf. equation (13). Table 4 contains the results from estimating the euro spillover model. The estimates of the GARCH parameters (ω_i, α_i, and β_i) are not reported because they are similar to the results of the constant spillover model.

The univariate model for the US return is identical to that of the constant spillover model; for convenience the results are repeated in the first row of Table 4.

The second step of the estimation concerns the return of the European bond index, cf. the second row of Table 4. The joint hypothesis of no spillover changes after the euro is borderline to not being rejected; the robust Wald test for the hypothesis $H_0 : \gamma_{1,E} = \phi_{1,E} = 0$ results in a p-value equal to 4.7%. The subsequent results are robust to including or excluding euro changes in the spillover parameters. Thus, we continue with the dummy variable in the specification for the European return.

Table 4

Euro spillover model

The table reports the results from estimating the euro spillover model. US return: $R_{US,t} = c_{0,US} + c_{1,US}R_{US,t-1} + e_{US,t}$ where $e_{US,t}$ has mean 0 and conditional variance: $\sigma^2_{US,t} = \omega_{US} + \alpha_{US}e^2_{US,t-1} + \beta_{US}\ \sigma^2_{US,t-1}$. European return: $R_{E,t} = c_{0,E} + c_{1,E}R_{E,t-1} + (\gamma_{0,E} + \gamma_{1,E}\ D_{t-1})R_{US,t-1} + (\phi_{0,E} + \phi_{1,E}\ D_{t-1})e_{US,t} + e_{E,t}$ where $e_{E,t}$ has mean 0 and conditional variance: $\sigma^2_{E,t} = \omega_E + \alpha_E e^2_{E,t-1} + \beta_E\ \sigma^2_{E,t-1}$. Country i return (the European index excluding country i is applied): $R_{i,t} = c_{0,i} + c_{1,i}R_{i,t-1} + (\gamma_{0,i} + \gamma_{1,i}\ D_{t-1})R_{US,t-1} + (\delta_{0,i} + \delta_{1,i}\ D_{t-1})R_{E,t-1} + (\phi_{0,i} + \phi_{1,i}\ D_{t-1})e_{US,t} + (\psi_{0,i} + \psi_{1,i}\ D_{t-1})e_{E,t} + e_{i,t}$ where $e_{i,t}$ has mean 0 and conditional variance: $\sigma^2_{i,t} = \omega_i + \alpha_i e^2_{i,t-1} + \beta_i\ \sigma^2_{i,t-1}$. D_t equals 0 before January 1, 1999 and 1 hereafter. $\hat{\omega}_i$, $\hat{\alpha}_i$, and $\hat{\beta}_i$ not reported. Bollerslev and Wooldridge (1992) robust standard errors in parentheses. * (§) [♯], indicates that the value is significant at a 1% (5%) [10%] level of significance.

	$c_{0,i}$	$c_{1,i}$	$\gamma_{0,i}$	$\gamma_{1,i}$	$\delta_{0,i}$	$\delta_{1,i}$	$\phi_{0,i}$	$\phi_{1,i}$	$\psi_{0,i}$	$\psi_{1,i}$
US	0.165*	−0.040								
	(0.055)	(0.036)								
Eu	0.145*	0.009	0.013	−0.020			0.105*	−0.077§		
	(0.019)	(0.038)	(0.016)	(0.025)			(0.023)	(0.031)		
Be	0.150*	−0.147*	0.018	−0.033§	0.092♯	0.083§	0.075*	−0.042♯	0.556*	0.466*
	(0.011)	(0.047)	(0.016)	(0.016)	(0.052)	(0.042)	(0.023)	(0.024)	(0.055)	(0.058)
De	0.145*	−0.059	−0.003	0.002	0.138§	−0.025	0.080*	−0.048*	0.782*	0.038
	(0.011)	(0.053)	(0.013)	(0.015)	(0.058)	(0.041)	(0.015)	(0.017)	(0.059)	(0.064)
Fr	0.144*	−0.088§	0.013	−0.023	0.101§	0.022	0.076*	−0.039§	0.693*	0.352*
	(0.011)	(0.041)	(0.018)	(0.020)	(0.051)	(0.043)	(0.017)	(0.018)	(0.063)	(0.068)
Ge	0.135*	−0.005	0.032♯	−0.031	−0.014	0.040	0.037♯	0.000	0.365*	0.627*
	(0.009)	(0.046)	(0.019)	(0.021)	(0.037)	(0.043)	(0.022)	(0.023)	(0.054)	(0.057)
It	0.167*	−0.033	0.008	−0.015	−0.018	0.054	0.136*	−0.103*	0.732*	0.242§
	(0.011)	(0.053)	(0.028)	(0.028)	(0.086)	(0.077)	(0.030)	(0.031)	(0.116)	(0.119)
Ne	0.128*	−0.121§	0.028	−0.033	0.097	0.049	0.078	−0.048	0.578*	0.446*
	(0.009)	(0.060)	(0.026)	(0.028)	(0.061)	(0.055)	(0.051)	(0.051)	(0.089)	(0.090)
Sp	0.151*	−0.107§	−0.016	0.004	0.165§	−0.039	0.092*	−0.060*	0.923*	0.089
	(0.010)	(0.052)	(0.025)	(0.025)	(0.075)	(0.061)	(0.020)	(0.021)	(0.057)	(0.061)
Sw	0.172*	−0.051	0.021	−0.003	0.014	0.171	0.289*	−0.069	1.238*	−0.683*
	(0.036)	(0.041)	(0.030)	(0.061)	(0.112)	(0.167)	(0.026)	(0.058)	(0.107*)	(0.170)
UK	0.177*	−0.007	0.000	−0.102♯	−0.012	0.155	0.318*	0.274*	0.843*	−0.323§
	(0.039)	(0.040)	(0.040)	(0.056)	(0.091)	(0.134)	(0.049)	(0.063)	(0.103)	(0.162)

In the third step of the estimation, we investigate the effect of the euro on the mean-spillover effects and the volatility-spillover effects from the US and European bond markets to the individual European bond markets, cf. the bottom rows of Table 4. As for the constant spillover model, we only find scattered evidence of mean spillover effects. For the period before the euro there are only mean-spillover effects for a few countries, $\hat{\gamma}_{0,i}$ and $\hat{\delta}_{0,i}$ are only significantly positive for a few countries. Moreover, there is no evidence that the mean-spillover effects are different after the euro; for all countries we reject the hypothesis $H_0 : \gamma_{1,i} = \delta_{1,i} = 0$. After the euro, there are still almost no signs of mean-spillover effects; we cannot reject $H_0 : \gamma_{0,i} = \gamma_{1,i} = \delta_{0,i} = \delta_{1,i} = 0$ for all i save Belgium for which the p-value equals 0.4%.

There are strong indications of both US and European volatility-spillover effects. For the period before the euro the US volatility-spillover effects as well as the European volatility-spillover effects are significant, i.e. $\phi_{0,i}$ and $\psi_{0,i}$ are significant. In addition,

Table 5

Variance ratios – euro spillover model

The table reports the mean and standard deviation of the US, European, and own variance ratios for the euro spillover model: $VR_{i,t}^{US} = \frac{(\phi_{0,i}+\phi_{1,i}D_{t-1})^2\sigma_{US,t}^2}{h_{i,t}}$, $VR_{i,t}^{E} = \frac{(\psi_{0,i}+\psi_{1,i}D_{t-1})^2\sigma_{E,t}^2}{h_{i,t}}$, and $VR_{i,t}^{i} = 1 - VR_{i,t}^{US} - VR_{i,t}^{E}$. $\sigma_{US,t}(\sigma_{E,t})$ is the conditional volatility of the US (European) idiosyncratic shock and $h_{i,t}$ is the conditional variance of the unexpected return of country *i*. D_t equals 0 before 1 January 1999 and 1 hereafter.

	$VR_{i,t}^{US}$		$VR_{i,t}^{E}$		$VR_{i,t}^{i}$	
	Mean	Stdev.	Mean	Stdev.	Mean	Stdev.
Be	0.041	0.034	0.496	0.249	0.462	0.239
De	0.033	0.026	0.542	0.200	0.424	0.202
Fr	0.032	0.023	0.550	0.227	0.418	0.220
Ge	0.017	0.010	0.436	0.274	0.547	0.271
It	0.061	0.061	0.392	0.306	0.548	0.306
Ne	0.045	0.035	0.530	0.241	0.425	0.228
Sp	0.027	0.023	0.523	0.267	0.450	0.268
Sw	0.122	0.061	0.227	0.121	0.651	0.153
UK	0.223	0.140	0.120	0.096	0.670	0.124

the volatility-spillover effects are significantly different after the euro; the robust Wald test for the hypothesis $H_0 : \phi_{1,i} = \psi_{1,i} = 0$ is rejected for all countries. There are also significant volatility-spillover effects after the euro; $H_0 : \phi_{0,i} = \phi_{1,i} = \psi_{0,i} = \psi_{1,i} = 0$ is rejected for all *i*.

We provide novel insights into how the introduction of the euro has significantly changed the volatility spillover effects into European bond markets. For the EMU countries plus Denmark the following commonalities are observed. The US volatility-spillover effects before the euro are found to be stronger than the effects estimated by the constant spillover model, i.e. $\hat{\phi}_{0,i} > \hat{\phi}_i$. The US volatility-spillover effects are dampened by the euro ($\hat{\phi}_{1,i} < 0$) and are weaker than estimated by the constant spillover model, $\hat{\phi}_{0,i} + \hat{\phi}_{1,i} < \hat{\phi}_i$. In contrast, the European volatility-spillover effects before the euro are found to be weaker than estimated by the constant spillover model; $\hat{\psi}_{0,i} < \hat{\psi}_i$. After the euro, the effects of European volatility spillover are strengthened ($\hat{\psi}_{1,i} > 0$) and are stronger than estimated by the constant spillover model, $\hat{\psi}_{0,i} + \hat{\psi}_{1,i} > \hat{\psi}_i$. Not taking the introduction of the euro into account when considering volatility-spillover effects appears to be identical to averaging across the two time periods. These observations, however, do not apply to Sweden and the UK. The US volatility-spillover effects are unaltered (Sweden) or strengthened (UK) by the euro, whereas the European volatility-spillover effects are dampened by the euro.

Table 5 presents the mean of the variance ratios. Sweden and the UK are again exceptions that we will return to. In the first sub period the average US variance ratios are smaller than in the second sub period. Compared to the constant spillover model, the mean of the US variance ratio has increased from 1–2% to 2–6%, because the sample period before the euro is longer than the sample period after the euro. The average European variance ratios are smaller in the first sub period than in the last sub

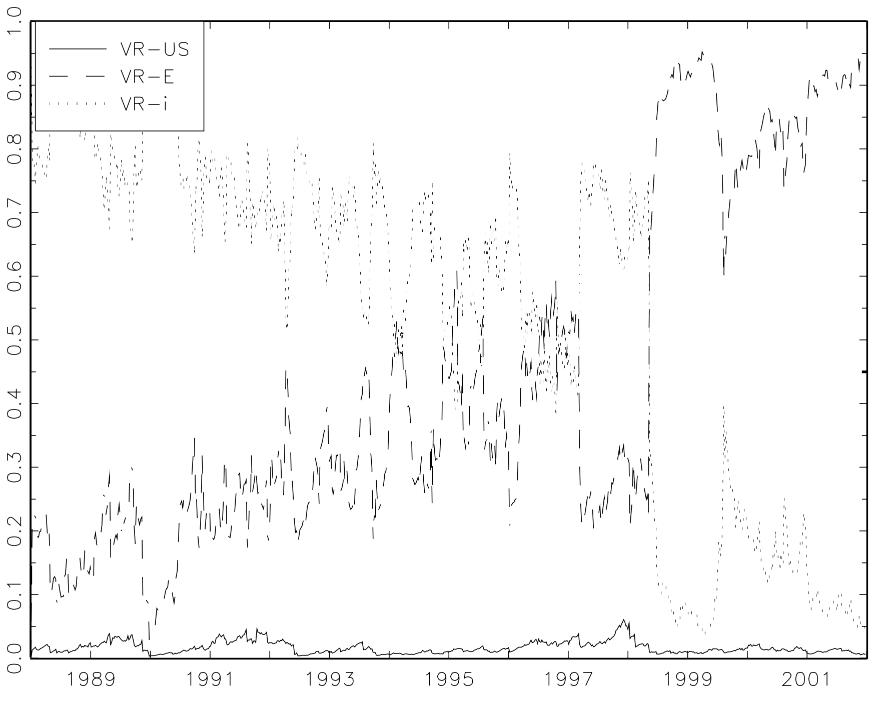

Fig. 3. Variance ratios – Germany – euro spillover model

period. Therefore, the means of the European variance ratios tend to be smaller in the euro spillover model than in the constant spillover model. The purely local volatility effects are smaller after the introduction of the euro. The average local variance ratio has decreased from 58% to 15%. The tendencies are also seen in the plots of the variance ratios for the example country Germany in Figure 3. Thus, after the introduction of the euro there is much less room for conducting policy directed at own bond markets and much more scope for common European policy making.

Our results are in line with Laopodis (2004) who finds that conducting monetary policy has become more difficult in recent years due to the integration of international bond markets. To support this, Dewachter *et al.* (2004) show that the German yield curve has changed significantly after the euro due to differences in monetary policies of the German Bundesbank and the European Central Bank. Moreover, Allen and Song (2005) find that the euro has also helped to integrate the financial services industry in Europe.[12]

For Sweden and the UK the means of the US and European variance ratios are almost unaltered. This is not surprising given that the euro appears to be an insignificant event for the spillover processes for these non-EMU countries. The tendencies are also seen in the plots of the variance ratios for the UK in Figure 4.

[12] Using data from before the euro (1985–97), Berger and Udell (2001) find that the European financial services industry does not consolidate across borders but within countries.

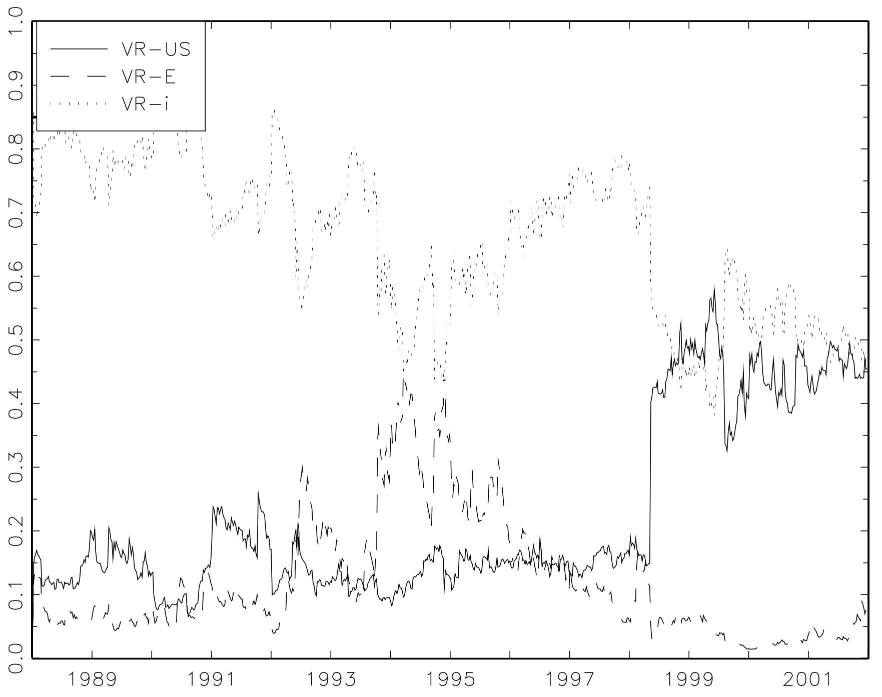

Fig. 4. Variance ratios – UK – euro spillover model

4.3. *Trend spillover model*

The new trend spillover model allows the spillover parameters to increase or decrease with a constant value each year, cf. equation (14). Thus, the spillover parameters may change gradually during the sample period. Table 6 shows the results arising from estimating the trend spillover model. Table 6 is structured like Table 4 and the results for the artificial European indices are not shown. For all the artificial European indices, the mean spillover parameter is not significantly time varying, so in the results presented $\gamma_{1,E} \equiv 0$. For the artificial European indices the volatility spillover from the US increases significantly over time ($\widehat{\phi_{1E}} > 0$ but small, around 0.006), except for the index excluding the UK.[13]

As in both the constant and the euro spillover model the US as well as the European mean spillover effects are fairly weak. Only very few of the mean spillover parameters are significant.

There are strong US and European volatility-spillover effects at play. In order to have significant volatility spillover from the US bond market either $\phi_{0,i}$ or $\phi_{1,i}$ should be significant. We find that this is the case for all countries except the Netherlands. Moreover, for most countries we cannot reject that the US volatility-spillover parameter is constant (the hypothesis that $\phi_{1,i} = 0$ is not rejected). There are two exceptions. First, for the UK the volatility-spillover effect increases strongly during the sample

[13] For the JPMorgan European index we cannot reject that the volatility spillover parameter is constant, $\gamma_{1,E}$ is insignificant.

Table 6

Trend spillover model

The table reports the results from estimating the trend spillover model. US return: $R_{US,t} = c_{0,US} + c_{1,US} R_{US,t-1} + e_{US,t}$ where $e_{US,t}$ has mean 0 and conditional variance: $\sigma^2_{US,t} = \omega_{US} + \alpha_{US}\, e^2_{US,t-1} + \beta_{US}\, \sigma^2_{US,t-1}$. European return: $R_{E,t} = c_{0,E} + c_{1,E} R_{E,t-1} + \gamma_{0,E} R_{US,t-1} + (\phi_{0,E} + \phi_{1,E}\, DT_{t-1}) e_{US,t} + e_{E,t}$ where $e_{E,t}$ has mean 0 and conditional variance: $\sigma^2_{E,t} = \omega_E + \alpha_E\, e^2_{E,t-1} + \beta_E\, \sigma^2_{E,t-1}$. Country i return (the European index excluding country i is applied): $R_{i,t} = c_{0,i} + c_{1,i} R_{i,t-1} + (\gamma_{0,i} + \gamma_{1,i}\, DT_{t-1}) R_{US,t-1} + (\delta_{0,i} + \delta_{1,i}\, DT_{t-1}) R_{E,t-1} + (\phi_{0,i} + \phi_{1,i}\, DT_{t-1}) e_{US,t} + (\psi_{0,i} + \psi_{1,i}\, DT_{t-1}) e_{E,t} + e_{i,t}$ where $e_{i,t}$ has mean 0 and conditional variance: $\sigma^2_{i,t} = \omega_i + \alpha_i\, e^2_{i,t-1} + \beta_i\, \sigma^2_{i,t-1}$. DT_t equals 1 in 1988, 2 in 1989, etc. $\hat{\omega}_i$, $\hat{\alpha}_i$, and $\hat{\beta}_i$ not reported. Bollerslev and Wooldridge (1992) robust standard errors in parentheses. * (§) [♯], indicates that the value is significant at a 1% (5%) [10%] level of significance.

	$c_{0,i}$	$c_{1,i}$	$\gamma_{0,i}$	$\gamma_{1,i}$	$\delta_{0,i}$	$\delta_{1,i}$	$\phi_{0,i}$	$\phi_{1,i}$	$\psi_{0,i}$	$\psi_{1,i}$
US	0.165*	−0.040								
	(0.055)	(0.036)								
Eu	0.141*	0.015	0.007				0.089*	−0.000		
	(0.019)	(0.037)	(0.013)				(0.029)	(0.003)		
Be	0.142*	−0.104§	0.008	−0.000	0.102	0.001	0.046♯	0.003	0.261§	0.053*
	(0.013)	(0.050)	(0.020)	(0.002)	(0.080)	(0.006)	(0.028)	0.002	(0.108)	(0.008)
De	0.144*	−0.058	−0.031	0.003§	0.167§	−0.004	0.054§	0.002	0.764*	0.004
	(0.011)	(0.053)	(0.020)	(0.002)	(0.081)	(0.005)	(0.026)	(0.002)	(0.135)	(0.010)
Fr	0.142*	−0.091§	0.006	−0.000	0.059	0.006	0.063§	0.001	0.509*	0.036*
	(0.012)	(0.043)	(0.028)	(0.002)	(0.078)	(0.006)	(0.026)	(0.002)	(0.125)	(0.009)
Ge	0.132*	0.033	0.015	0.001	−0.001	−0.001	−0.007	0.006§	−0.210	0.083*
	(0.010)	(0.042)	(0.031)	(0.002)	(0.056)	(0.005)	(0.038)	(0.003)	(0.152)	(0.011)
It	0.162*	−0.030	0.017	−0.001	−0.004	0.003	0.111*	−0.001	0.486*	0.036*
	(0.012)	(0.054)	(0.034)	(0.002)	(0.130)	(0.009)	(0.039)	(0.003)	(0.150)	(0.011)
Ne	0.124*	−0.093♯	0.011	0.000	0.127	−0.001	0.073	0.000	0.320♯	0.005*
	(0.010)	(0.054)	(0.051)	(0.004)	(0.101)	(0.007)	(0.093)	(0.007)	(0.190)	(0.014)
Sp	0.149*	−0.111§	−0.048	0.004	0.375*	−0.018§	0.107§	−0.002	0.901*	0.007
	(0.010)	(0.054)	(0.051)	(0.004)	(0.127)	(0.009)	(0.044)	(0.003)	(0.138)	(0.010)
Sw	0.178*	−0.063	0.016	0.000	0.063	−0.000	0.278*	−0.003	1.253*	−0.023
	(0.035)	(0.041)	(0.041)	(0.006)	(0.145)	(0.015)	(0.036)	(0.005)	(0.160)	(0.017)
UK	0.185*	−0.012	0.026	−0.007	−0.139	0.020	0.011	0.048*	1.020*	−0.035
	(0.038)	(0.040)	(0.064)	(0.006)	(0.187)	(0.019)	(0.078)	(0.007)	(0.212)	(0.021)

period, $\widehat{\phi_{1,i}} = 0.05$. So for each year that elapses the volatility-spillover parameter increases by 0.05. Second, for Germany the $\phi_{1,i}$ parameter is also significant but almost negligible, $\widehat{\phi_{1,i}} = 0.006$. For EMU countries the results are different from those in the euro spillover model. In the trend spillover model, we find only weak signs of time variation in US volatility-spillover effects, whereas in the euro spillover model we find evidence of changes. For the non-EMU countries, the results from the two models are comparable.

The European volatility-spillover effects are significant for all countries without exception; either $\psi_{0,i}, \psi_{1,i}$, or both are significant for all countries. For the EMU countries the European volatility-spillover effects generally increase during the sample period, $\psi_{1,i}$ is significantly positive. For Spain and Denmark the point estimates are also positive but insignificant. For Sweden and the UK, the European volatility-spillover effects seem to be constant during the sample period, $\psi_{1,i}$ is insignificant. The point estimates are negative which might indicate a slight tendency to be decreasing over

Table 7

Variance ratios – trend spillover model

The table reports the mean and standard deviation of the US, European, and own variance ratios for the trend spillover model: $VR_{i,t}^{US} = \frac{(\phi_{0,i}+\phi_{1,i}DT_{t-1})^2\sigma_{US,t}^2}{h_{i,t}}$, $VR_{i,t}^{E} = \frac{(\psi_{0,i}+\psi_{1,i}DT_{t-1})^2\sigma_{E,t}^2}{h_{i,t}}$, and $VR_{i,t}^{i} = 1 - VR_{i,t}^{US} - VR_{i,t}^{E}$. $\sigma_{US,t}(\sigma_{E,t})$ is the conditional volatility of the US (European) idiosyncratic shock and $h_{i,t}$ is the conditional variance of the unexpected return of country *i*. DT_t equals 1 in 1988, 2 in 1989, etc.

	$VR_{i,t}^{US}$		$VR_{i,t}^{E}$		$VR_{i,t}^{i}$	
	Mean	Stdev.	Mean	Stdev.	Mean	Stdev.
Be	0.040	0.023	0.488	0.258	0.472	0.268
De	0.038	0.025	0.537	0.190	0.425	0.207
Fr	0.038	0.020	0.546	0.230	0.416	0.237
Ge	0.022	0.022	0.414	0.319	0.563	0.337
It	0.059	0.041	0.370	0.289	0.561	0.311
Ne	0.049	0.026	0.530	0.253	0.421	0.253
Sp	0.042	0.023	0.520	0.254	0.438	0.269
Sw	0.113	0.059	0.224	0.087	0.668	0.124
UK	0.243	0.186	0.112	0.081	0.645	0.165

time. The European volatility-spillover results complement those from the euro spillover model. For the EMU countries, the European volatility-spillover parameters are time varying and the European volatility spillover increases during the sample period. For non-EMU countries the European volatility-spillover effects do not increase over time.

Table 7 shows the average variance ratios. Compared to the constant spillover model, the average US effects have become stronger, the average European effects weaker, and the average local effects have become much larger.

The most interesting information from the trend spillover model is shown in the graphs of the time series of the variance ratios. Figures 5 and 6 concern the example countries Germany and the UK. For fully integrated bond markets the European variance ratio should tend to one, and the local and global variance ratios should tend to zero. We expect EMU countries to be, if not fully integrated, then on their way to becoming integrated. In other words, we expect the European variance ratios to be increasing during the sample period and to be very large at the end of the sample period. For non-EMU countries we do not expect full integration with the European bond markets.

For Germany we see that the European variance ratio generally increases over time. Starting at a very low level in the beginning of the sample period it increases so much that it is around 0.9 at the end of the sample period. The US effects are comparatively low throughout. The local effects decrease over time and in 2002 the local variance ratio is around 0.05. The graphs of the variance ratios document that EMU countries are becoming more and more integrated during the sample period; aggregate European effects increase in importance and local effects decrease in importance. For the last six months of the sample, the average European variance ratio for the EMU countries (plus Denmark) equals 0.87. The corresponding average US variance ratio equals 0.05 and the average local variance ratio equals 0.08. So, although EMU countries are not yet fully integrated they are well on their way to becoming so.

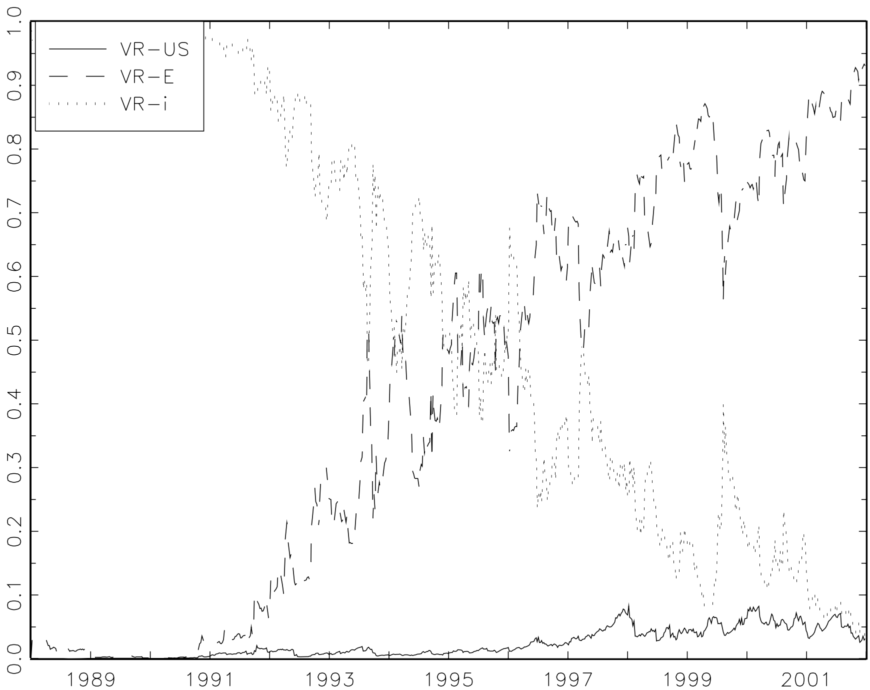

Fig. 5. Variance ratios – Germany – trend spillover model

For the UK bond market the US effects increase over time to a level of about 0.5. The local effects decrease but to a much higher level than for Germany. In 2002 the local variance ratio equals around 0.40. The aggregate European effects have decreased in importance. As expected, the non-EMU countries – the UK and Sweden – have not become more integrated with other European bond markets.

Overall, the graphs of the time series of the variance ratios corroborate and strengthen the conclusions from the euro spillover model.

4.3.1. *Sources of integration.* Above we document that the bond markets of EMU countries have become much more integrated by observing that the European variance ratios have increased during the sample period. Now, what are the economic sources behind the integration? To answer this question, we analyse how the size of the European variance ratio for country *i* is explained by different variables using multiple OLS regressions. Here we provide a new way of illuminating what the economic sources behind the integration of the European bond markets are, and therefore our answers are also new. The explanatory variables are the following:[14]

[14] The data are obtained from DataStream. The EU15 inflation rate is only available from February 1990. For Germany, the inflation rate is only available from January 1996. For Sweden the inflation rate is only available from September 1996. We shrink the sample periods accordingly.

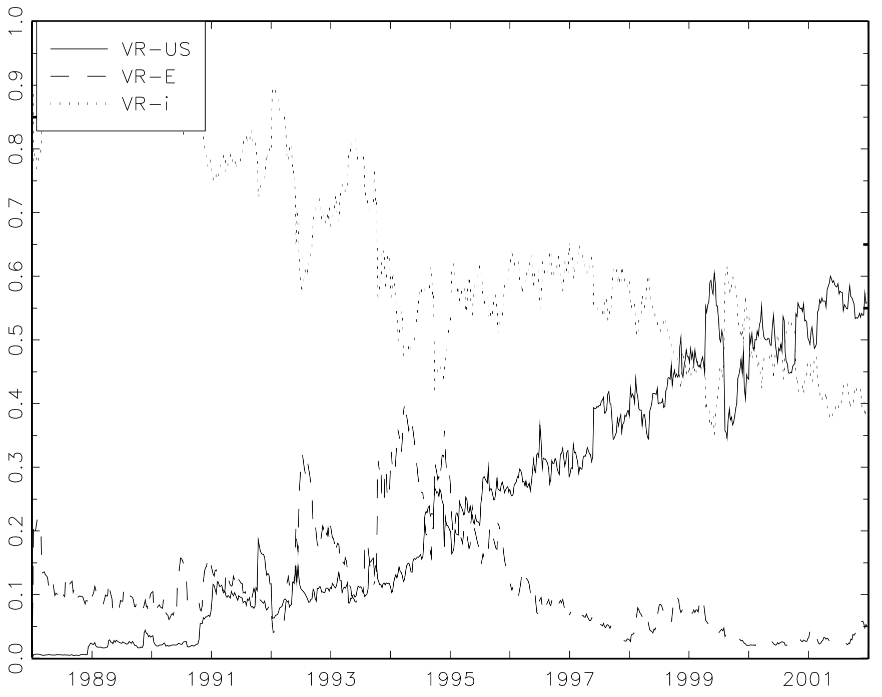

Figure 6. Variance ratios – UK – trend spillover model

1. The absolute difference between the EU15 inflation rate and country *i*'s inflation rate.
2. The absolute difference between the (unweighted) average 1-month IBOR rate across the 9 countries and country *i*'s 1-month IBOR rate.
3. The GARCH(1,1) variance from the log-return on the USD exchange rate to country *i*.[15]

The first explanatory variable measures whether integration has occurred due to convergence in inflation rates. We conjecture that the smaller the inflation difference, the stronger the integration. Thus, we expect a negative coefficient. The second explanatory variable measures whether integration has occurred due to convergence in interest rates. We presume that the smaller the interest rate difference, the stronger the integration. Thus, we expect a negative coefficient. The third explanatory variable measures the exchange rate stability. We expect that the more stable the exchange rate (low variance), the stronger the integration. This implies a negative coefficient.

The results from the regressions are contained in Table 8. For the EMU countries (plus Denmark) the interest rate differential plays a leading role. In accordance with our expectations, the estimated coefficients are all negative and significant. Only for some of the countries is the inflation differential significant, namely for Denmark, France and Spain. In these cases the relation is negative as hypothesised. The exchange rate stability is only significant for France, Germany and the Netherlands. For these countries the

[15] The exchange rate for the UK is given as USD per GBP.

Table 8

Sources of integration – trend spillover model

The table reports the results from regressing the European variance ratio for country *i* on a constant, (1) the absolute difference between the EU15 inflation rate and country *i*'s inflation rate, (2) the difference between the average 1-month IBOR rate and country *i*'s 1-month IBOR rate, and (3) the GARCH(1,1) variance (scaled by 1000) for country *i*'s US exchange rate log-return. Based on Newey and West (1987) standard errors * (§) [♯], indicates that the value is significant at a 1% (5%) [10%] level of significance.

	Inflation	Interest rate	Exchange volatility	R^2
Be	−0.063	−0.139*	0.240	0.32
De	−0.105*	−0.054§	0.336	0.33
Fr	−0.203*	−0.170*	−0.780*	0.50
Ge	−0.033	−0.173*	−0.376§	0.65
It	−0.044	−0.175*	0.108♯	0.73
Ne	0.003	−0.168*	−0.293§	0.73
Sp	−0.064§	−0.145*	−0.370	0.59
Sw	0.023	0.005	−0.656*	0.46
UK	−0.051♯	−0.034*	0.483*	0.12

relation is negative as expected. The explanatory power of the regressions are large; the R^2s range from 0.32 to 0.73.

The picture is different for non-EMU countries: for Sweden, only exchange rate stability is of importance and its coefficient has the expected sign. For the UK, the interest rate differential also has explanatory power, although the coefficient is much smaller than for EMU countries. For the UK, the European variance ratio depends positively on the exchange rate variance, the opposite of what is the case for EMU countries.

Due to data limitations, credit risk and liquidity risk are not included as explanatory variables in the above regressions. Differences in credit risk and differences in liquidity risk are potentially why the EMU bond markets are not fully integrated.[16] With a couple of exceptions, all countries have credit rating AAA for sovereign issuers of long term local currency debt. Belgium and Italy are one notch below at AA.[17] The average European variance ratios across the last six months of the sample do not differ across rating categories for EMU countries plus Denmark, (the average equals 0.868 for the AAA bond markets and 0.870 for the AA bond markets). So, differences in credit ratings do not appear to cause bond markets not to be fully integrated. Favero *et al.* (2005) show that for some EMU countries the spread in government bond yield above Germany depends upon liquidity differences. We measure the liquidity by the size of the government bond market in the last year of the sample.[18] We find a small negative

[16] The author is most indebted to an anonymous referee for pointing towards the importance of credit risk and liquidity risk in this respect.

[17] The credit ratings are obtained from the Fitch Ratings home page.

[18] The data stem from Table 16 A of Bank for International Settlements (2005).

correlation equal to -0.097 between the size of the government bond market and the average European variance ratio for the last six months of the sample for the EMU countries (plus Denmark). There is a very weak tendency that the more liquid the bond market is, the less integrated it is. This again provides a very weak indication that larger bond markets work autonomously from the aggregate European bond market. Thus, differences in liquidity appear to play a minor role in explaining why the bond markets are not fully integrated. Policy makers could try to remove these differences in liquidity risk to obtain complete bond markets.

In summary, the convergence of EMU bond markets appears to be caused mainly by convergence in interest rates. To a smaller degree it is also caused by convergence in inflation rates and exchange rate stability. There is very weak evidence that differences in liquidity is one of the reasons why the EMU bond markets are not fully integrated.

4.4. *Specification tests*

In this section we compare the appropriateness of the three model specifications. To this end we study the properties of the standardised residuals; the US standardised residuals, the nine artificial European standardised residuals and the nine country i standardised residuals.

In a well specified model, the standardised residuals (*i*) have mean zero, (*ii*) have unit variance, (*iii*) are serially uncorrelated at order 1 (*iv*) are serially uncorrelated up to order 4 (*v*) no ARCH(1) effects remain, (*vi*) the series of the squared standardised residuals are serially uncorrelated at order 1, and (*vii*) the series of the squared standardised residuals are serially uncorrelated up to order 4. We apply the Ljung and Box (1978) Q-test for no autocorrelation and the Lagrange Multiplier test for no ARCH effects, cf. Engle (1982). As there are 19 relevant series and 7 tests for each model the results are not tabulated. We apply a level of significance of 1%.

The US standardised residuals are identical for all three models. There are no signs of misspecification in the model for the US index. Also, for the European indices (that each exclude one country), we find no evidence of misspecification in either of the models. In addition, the cross products of the US and artificial European standardised residuals are not significantly autocorrelated (neither up to lag 1 nor up to lag 4) and there are no ARCH(1) effects remaining in the cross products.

We analyse the diagnostics for the nine individual countries in each of the three models. In all three models we find only a few instances where the misspecification tests are rejected (details provided shortly). Thus, the country models appear to be well specified. There are fewest rejections for the trend spillover model (3) followed by the constant model (4) and most for the euro spillover model (6). For each model, the rejections concern three different countries.

The following hypotheses introduced above are rejected. Constant spillover model: France (*iv* and *vii*), Italy (*ii*) and the Netherlands (*iv*). Euro spillover model: Germany (*iv* and *vii*), Italy (*ii*) and the Netherlands (*ii, iii*, and *iv*). Trend spillover model: Belgium (*v*), Italy (*ii*) and the Netherlands: (*iv*).

Overall, we conclude that based on the properties of the standardised residuals, the spillover models are well specified and that the trend spillover model is slightly preferred to the other specifications. Thus, we find it safe to trust the results.

Table 9

Contagion regressions – constant spillover model

The table reports the results from running the following regression $\widehat{e_{i,t}} = a_{1i} + (a_{2i} + a_{3i}DC_t)\widehat{e_{US,t}} + (a_{4i} + a_{5i}DC_t)\widehat{e_{E,t}} + resid_{it}$ where DC_t is a crisis indicator which equals one from 1 January 1998 to 31 December 1999. $\widehat{e_{i,t}}$, $\widehat{e_{US,t}}$, and $\widehat{e_{E,t}}$ are the estimated idiosyncratic shock for country i, the US, and aggregate European from the constant spillover model. The last two columns show the p-values from the Wald test of $a_{2i} = a_{3i} = 0$ and $a_{4i} = a_{5i} = 0$. Based on Newey and West (1987) standard errors * (§) [♯], indicates that the value is significant at a 1% (5%) [10%] level of significance.

	a_1	a_2	a_3	a_4	a_5	$a_2 = a_3 = 0$	$a_4 = a_5 = 0$
Be	0.019	0.007	−0.017	−0.317*	0.540*	0.0%	78.0%
De	0.015	0.027	0.029	−0.001	0.056	4.5%	45.5%
Fr	0.024	0.049§	−0.042	−0.216§	0.404*	9.6%	0.0%
Ge	0.010	0.004	0.009	−0.204*	0.590*	53.7%	0.0%
It	0.013	0.131§	−0.122§	0.264§	−0.180	4.3%	1.4%
Ne	0.009	0.001	−0.018	−0.381*	0.581*	47.1%	0.0%
Sp	0.011	0.081*	−0.068*	0.09	−0.065	0.0%	46.8%
Sw	−0.041	−0.019	−0.124	0.172♯	−0.271	21.2%	18.7%
UK	−0.009	−0.055	0.175*	0.088	−0.480*	1.4%	1.0%

4.5. *Contagion effects*

The volatility-spillover framework allows us to test for contagion effects, that is excess correlation between the idiosyncratic shocks of the model. The contagion methodology is introduced in Bekaert *et al.* (2005) and expanded in Baele (2005). Here, the contagion methodology is used for the first time on international bond markets. If there is no contagion from the US and aggregate Europe into country i, then their idiosyncratic shocks are not correlated. Moreover, it is hypothesised that there is (more) contagion during times of crisis. Thus we run the following regression for each country:

$$\widehat{e_{i,t}} = a_{1i} + (a_{2i} + a_{3i}DC_t)\widehat{e_{US,t}} + (a_{4i} + a_{5i}DC_t)\widehat{e_{E,t}} + resid_{it} \tag{15}$$

DC_t is a crisis indicator function which equals one in the period surrounding the introduction of the euro, more precisely from one year before to one year after (1 January 1998 to 31 December 1999), and zero otherwise. The crisis indicator that we apply differs from those used previously so that it accommodates the specific time period under investigation. If $a_{2i} = a_{3i} = 0$ there is no contagion from the US and if $a_{4i} = a_{5i} = 0$ there is no contagion from the aggregate European bond market into country i's bond market. a_{3i} and a_{5i} measure any additional contagion during the crisis period. Tables 9–11 show the results of estimating the contagion regression shown in (15) for each of the three volatility-spillover models.

At the 1% level of significance, there are US contagion effects in the constant spillover model, but not in the euro spillover model nor in the trend spillover model ($a_{2i} = a_{3i} = 0$ is tested). In all models, European contagion effects are present; for around half of the countries we cannot reject that $a_{4i} = a_{5i} = 0$. In the constant spillover model it occurs during crisis as well as in non-crisis periods. In contrast, for the euro spillover model and the trend spillover model, the contagion effects from aggregate Europe mainly occur during the crisis period (a_{5i} is significant and a_{4i} is not). Only for Italy is a_4 significant.

Table 10

Contagion regressions – euro spillover model

The table reports the results from running the following regression $\widehat{e_{i,t}} = a_{1i} + (a_{2i} + a_{3i}DC_t)\widehat{e_{US,t}} + (a_{4i} + a_{5i}DC_t)\widehat{e_{E,t}} + resid_{it}$ where DC_t is a crisis indicator which equals one from 1 January 1998 to 31 December 1999. $\widehat{e_{i,t}}$, $\widehat{e_{US,t}}$, and $\widehat{e_{E,t}}$ are the estimated idiosyncratic shock for country *i*, the US, and aggregate European from the euro spillover model. The last two columns show the p-values from the Wald test of $a_{2i} = a_{3i} = 0$ and $a_{4i} = a_{5i} = 0$. Based on Newey and West (1987) standard errors * (§) [♯], indicates that the value is significant at a 1% (5%) [10%] level of significance.

	a_1	a_2	a_3	a_4	a_5	$a_2 = a_3 = 0$	$a_4 = a_5 = 0$
Be	0.019	−0.002	−0.018	−0.070	0.290*	69.0%	0.3%
De	0.015	0.017	0.040	0.015	0.040	10.8%	39.9%
Fr	0.022	0.024	−0.020	−0.058	0.221§	54.1%	0.0%
Ge	0.002	−0.001	−0.007	−0.064	0.236*	86.3%	1.5%
It	0.019	0.087	−0.103♯	0.415*	−0.268§	16.4%	0.0%
Ne	0.011	−0.017	−0.016	−0.115	0.340*	15.8%	0.4%
Sp	0.011	0.050§	−0.050§	0.149♯	−0.099	3.8%	9.6%
Sw	−0.037	−0.029	−0.096	0.096	0.105	30.7%	28.8%
UK	−0.008	−0.022	0.095	0.069	−0.427§	41.8%	2.7%

In closing, we hardly find any evidence of contagion from the US bond market into the individual European bond markets, and only some evidence of contagion from the aggregate European bond market into the individual bond markets, and that mainly in the period surrounding the introduction of the euro.

Table 11

Contagion regressions – trend spillover model

The table reports the results from running the following regression $\widehat{e_{i,t}} = a_{1i} + (a_{2i} + a_{3i}DC_t)\widehat{e_{US,t}} + (a_{4i} + a_{5i}DC_t)\widehat{e_{E,t}} + resid_{it}$ where DC_t is a crisis indicator which equals one from 1 January 1998 to 31 December 1999. $\widehat{e_{i,t}}$, $\widehat{e_{US,t}}$, and $\widehat{e_{E,t}}$ are the estimated idiosyncratic shock for country *i*, the US, and aggregate European from the trend spillover model. The last two columns show the p-values from the Wald test of $a_{2i} = a_{3i} = 0$ and $a_{4i} = a_{5i} = 0$. Based on Newey and West (1987) standard errors * (§) [♯], indicates that the value is significant at a 1% (5%) [10%] level of significance.

	a_1	a_2	a_3	a_4	a_5	$a_2 = a_3 = 0$	$a_4 = a_5 = 0$
Be	0.015	−0.007	0.005	−0.098♯	0.314*	89.8%	0.0%
De	0.015	0.017	0.046	0.017	0.032	5.2%	51.5%
Fr	0.021	0.018	0.000	−0.079	0.238*	28.1%	0.2%
Ge	−0.006	−0.007	0.011	−0.031	0.224*	84.8%	0.0%
It	0.013	0.092♯	−0.084	0.451*	−0.336*	20.5%	0.0%
Ne	0.006	−0.028	0.013	−0.160§	0.372*	16.9%	0.0%
Sp	0.001	0.030	−0.022	0.147♯	−0.114	27.8%	16.1%
Sw	−0.038	−0.012	−0.084	0.165♯	−0.185	52.5%	23.2%
UK	−0.012	0.011	−0.034	0.045	−0.327§	83.3%	10.1%

5. Conclusion

In this paper we provide novel findings about how volatility in a number of European bond markets is affected by volatility in the US and European bond markets. We distinguish between global, regional, and local volatility effects. We apply a GARCH-type model that allows for both mean and volatility spillover from the US and aggregate European bond markets into the individual countries. Mean-spillover effects are almost negligible, whereas volatility-spillover effects are essential. For EMU countries (plus Denmark) regional effects are most important, followed by local effects. Global effects are almost inconsequential. For non-EMU countries own country effects are stronger, European effects smaller and US effects larger. EMU countries have become much more integrated after the introduction of the euro. Moreover, EMU countries' bond markets have become close to being perfectly integrated during the sample period. The main reason for the integration is convergence in interest rates.

We find evidence of substantial differences between the nature of the volatility of the bond markets of EMU-member countries and of non-EMU member countries. Still, we only consider old European bond markets, not the emerging bond markets. Thus, one can only speculate as to how the new European bond markets differ from the established bond markets in this respect. At the moment, data constraints prevent us from conducting this interesting exercise.

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