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1. Introduction

On January 1, 1999, an epochal event took place in the arena of international finance: 11 European countries replaced their national currencies and introduced a single European currency, the Euro.† The creation of the euro was preceded by a long convergence process starting in 1979 with the establishment of the European Monetary System. The introduction of the euro as a single currency apparently reduced investors' transaction and information costs among the stock markets of Eurozone countries.‡ It also eliminated intra-European currency risk (Hardouvelis *et al.* 2006). Detken and Hartmann (2002) propose that the Euro became one of the three major currencies in the world after its introduction, taking its place alongside the U.S. dollar and the Japanese yen. When the euro took over the functions of national currencies, all aspects of country-specific risk connected with exchange rates were removed. Therefore, the euro's introduction should facilitate capital movement. Broadened investment opportunities across the Eurozone countries were therefore expected. Coupled with a

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†The original euro 11 includes Austria, Belgium, Finland, France, Germany, Ireland, Italy, Luxembourg, the Netherlands, Portugal, and Spain. Four member countries of the European Union—Denmark, Greece, Sweden, and the United Kingdom—did not join the first wave. Greece, however, joined the euro club in 2001 when it could satisfy the convergence criteria. For more details, consult the website of the European Central Bank and the textbook by Eun and Resnick (2007).

‡On May 1, 2004, the EU expanded by admitting 10 new member countries: Cyprus, Czech Republic, Estonia, Hungary, Latvia, Lithuania, Malta, Poland, Slovak Republic, and Slovenia. As of January 1, 2007, Bulgaria and Romania were also admitted into the EU.

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regulatory convergence of macroeconomic fundamentals of the European economies, the launch of the euro was anticipated to stabilize the volatilities of the associated stock markets.

Many studies have explored the consequences of the interesting experiment of the euro's advent on European capital markets (for example, Rouwenhorst 1999, Billio and Pelizzon 2003, Kim *et al.* 2005 and Hardouvelis *et al.* 2006). In particular, Morana and Beltratti (2002) formally evaluate whether or not the euro decreased the stock volatility of four EMU member countries, including France, Germany, Italy, and Spain. Using the data of daily stock returns over the period 1 January 1988 to 26 May 2000, they conduct a simple *F*-test for the equality of the unconditional variance and a GARCH(1, 1) model before (January 1988–December 1998) and after (January 1999–May 2000) the introduction of the euro. Their empirical implications for *F*-tests and GARCH approaches are different. The GARCH model with a dummy variable suggests no effect of the euro on the volatility of European stock markets, while the *F*-tests confirm a reduction in unconditional variance of the stock markets, except for Italy. As mentioned by Morana and Beltratti (2002), the inconsistency of the two approaches may result from the erroneous calendar timing of the historical event on sub-sampling the data since the possible adjustment in volatilities is not actually known. They finally propose a more flexible approach for depicting structural breaks in stock markets, the switching three-regime Markov model, to characterize the varying levels of volatilities and correlations to support their hypothesis of a reduced conditional variance of European stock markets due to the introduction of the euro.

Using more data, our empirical results presented in this paper, however, suggest that the reduction in volatilities of these EMU member countries after the launch of the euro is not the end of the story. We propose an alternative and more flexible smooth transition GARCH model that considers endogenously determined smooth structural changes. The model was first developed by Granger and Teräsvirta (1993) and Lin and Teräsvirta (1994) and thereafter extended to GARCH specifications by Lundbergh and Teräsvirta (2002). The LM tests of parameter constancy for all four countries suggest an inverted U-shaped structural-change pattern in the volatilities of the stock markets. Estimation of the model reveals that the structural shifts in volatilities of the four EMU member countries' stock markets began earlier than the calendar date of the euro's launch. The estimated results also show that the stock markets of France and Germany experienced two sharp switches. They moved from a regime of low volatility to a regime of high volatility and then went back to a regime of low volatility. In contrast, the processes of market regime switch of Italy and Spain were much smoother. They gradually changed from a regime of low volatility to a regime of high volatility and then went back to the regime of low volatility during the sample period.

The rest of this paper is organized as follows. In section 2 we discuss related GARCH models for stock market volatility. In section 3, the ST-GARCH model and corresponding tests are presented. For the purpose of demonstrating the problem of ill-dating on sub-samples, we also estimate a standard GARCH(1, 1) model with two sub-samples and a GARCH(1, 1) model with a dummy variable. The associated tests and estimation results of the ST-GARCH as well as their insights are then illustrated in section 4. Finally, concluding remarks are given in section 5.

2. GARCH model for stock market volatility

The GARCH model proposed by Engle (1982) and generalized by Bollerslev (1986) has been used to discuss the stochastic behavior of various financial time series and, in particular, to explore the changing behavior of volatility over time.†

The general approach adopted for volatility analysis concerning the occurrence of a specific event is to compare the difference in price volatility before and after an artificial date. This is typically achieved through the inclusion of a dummy variable in the volatility structure. Alternatively, the dynamics of the volatility process for the major European stock markets are discussed with structural changes for the second moment in GARCH-type models. Specifically, an elegant study by Lundbergh and Teräsvirta (2002) shows the statistical properties for structural changes embedded in the lagged squared residual term in the volatility equation of the GARCH model.

We will extend the concept of nonlinear structural changes to the lagged conditional variance terms of the GARCH framework. By modeling the volatility patterns of the main stock markets of the euro member states, we examine whether the introduction of the euro induced volatility dynamic changes in the markets.

First, consider a simple GARCH(1, 1) specification,

$$\begin{aligned} R_t &= \varepsilon_t, \\ h_t &= \alpha_0 + \alpha_1 \varepsilon_{t-1}^2 + \beta_1 h_{t-1}, \\ &\varepsilon_t \mid \Omega_{t-1} \sim N(0, h_t), \end{aligned} \tag{1}$$

where h_t is the conditional volatility at time t, α_1 is a coefficient that relates the past value of the squared residuals (ε_{t-1}^2) to the current volatility, β_1 is a coefficient that relates the current volatility to the past period of volatility, and Ω_{t-1} denotes the information set at time $t-1$. Essentially, the GARCH(1, 1) model uses three parameters to characterize the evolution of the volatility process. The parameters α_0, α_1, and β_1 jointly determine the long-run uncertainty level. The parameter α_1 is the coefficient of the lagged residual squared, meaning the

†See Bollerslev *et al.* (1992) for an earlier extensive literature review and Teräsvirta (2009) for a more recent survey.

short-run impact of volatility shocks. The parameter β_1 is the coefficient of the lagged conditional variance, meaning the long-run effect of volatility shocks.

In analysing the impact of the euro, an intuitive approach is to introduce a dummy variable in addition to (1). Such a model is represented by the following modification for the volatility structure:

$$h_t = \alpha_0 + \delta_0 D_t + \alpha_1 \varepsilon_{t-1}^2 + \delta_1 D_t \varepsilon_{t-1}^2 + \beta_1 h_{t-1} + \delta_2 D_t h_{t-1}, \quad (2)$$

where D_t is a dummy variable taking on the value 0 pre-euro and 1 post-euro. In addition to D_t, we also add two cross-product terms for all regressors to the variance equation with dummy variables. The parameter δ_0 measures the level of time-varying conditional volatility and long-run effect after the euro's introduction. Parameters δ_1 and δ_2 are used to measure the additional impact after the euro's launch of the short-run and long-run effect of volatility distributions.

3. The smooth transition GARCH model

When a dummy variable is used to model structural changes due to an external event, it is required to know a certain point in time *a priori* for a sharp switch. However, even if the timing of the target event can be confirmed, the question remains of whether or not these changes exactly coincide with the calendar timing of the event. This was a formal policy of the euro member states and had been declared long before. Thus, it is reasonable to assume that the participating countries prepared in many respects for the new currency. Their investment decision would be made in anticipation of the introduction of the euro before January 1, 1999. Furthermore, there are large diversities in the economic situations of the member states of the euro and investors are also likely to make heterogeneous timing decisions when they participate in the markets.

Hence, it is difficult to find an *event date* to define the dummy variable that will work for all countries. Incorrect timing on specifying the dummy variable would lead to insignificant coefficients of the associated dummy variables and false statistical inference. In other words, simple GARCH models with dummies may be insufficient to capture such a regime change due to the forward-looking nature of market participants.

The smooth transition regression models proposed by Granger and Teräsvirta (1993) are suitable for depicting the gradual changes among different regimes associated with exogenous events. The Markov regime switching model (for example, Morana and Beltratti 2002) is another ingenious way to demonstrate the structural changes if these changes occur discretely. Many studies have extended the idea posed by Granger and Teräsvirta (1993) from the mean equation to the variance equation in order to capture the gradual changes in the parameters versus the assumption of parameter constancy in the GARCH specification (Hagerud 1996, Lee and Degennaro 2000, Lundbergh and Teräsvirta 2002). We generally term these models as ST-GARCH-type models. The ST-GARCH model provides a more flexible representation of the aggregate behavior in volatility, while a single structural break model does not.

Specifically, Lundbergh and Teräsvirta (2002) developed a generalized framework for testing the adequacy of an estimated ST-GARCH-type model. In general, they consider a model as follows:

$$\begin{aligned} y_t &= f(\mathbf{w}_t; \boldsymbol{\varphi}) + \varepsilon_t, \\ \varepsilon_t &= z_t(h_t + g_t)^{1/2}, \end{aligned} \quad (3)$$

where $h_t = \boldsymbol{\eta}' \boldsymbol{s}_t$, $g_t = \boldsymbol{\lambda}' \boldsymbol{s}_t F(\tau_t; \gamma, \mathbf{c})$, $\mathbf{w}_t$ is a regressor vector in the mean equation, $\boldsymbol{\varphi}$ is the coefficient vector, $\{z_t\}$ is a sequence of i.i.d. random variables with zero mean and unit variance, $\boldsymbol{s}_t = (1, \varepsilon_{t-1}^2, \ldots, \varepsilon_{t-q}^2, h_{t-1}, \ldots, h_{t-p})'$, $\boldsymbol{\eta} = (\alpha_0, \alpha_1, \ldots, \alpha_q, \beta_1, \ldots, \beta_p)'$, and $\boldsymbol{\lambda} = (\bar{\alpha}_0, \bar{\alpha}_1, \ldots, \bar{\alpha}_q, \bar{\beta}_1, \ldots, \bar{\beta}_p)'$. In particular,

$$F(\tau_t; \gamma, \mathbf{c}) = \left(1 + \exp\left(-\gamma \prod_{i=1}^{k} (\tau_t - c_i)\right)\right)^{-1}, \quad (4)$$

where τ_t is the transition variable at time t, γ is the slope parameter ($\gamma > 0$), $\mathbf{c} = (c_1, c_2, \ldots, c_k)$ is a location vector where $c_1 \le c_2 \le \cdots \le c_k$, and k is the number of transitions.

The advantage of this model is threefold. First, the timing of the shifts in parameters is determined endogenously in the estimation rather than artificially chosen *a priori*. Second, the dummy-variables-type sharp switches in parameters become a special case when γ goes to infinity. Finally, the transition function in (4) provides flexible specifications in modeling to determine the patterns of structural changes. For example, equation (4) becomes a step function, the value of which equals 1 for $\tau_t \ge c_1$ and zero otherwise if the slope parameter $\gamma \to \infty$ with $k = 1$. In addition, when $\gamma \to \infty$ and $k = 2$, equation (4) turns out to be a double-step function. For details, see Lundbergh and Teräsvirta (2002).

Nevertheless, the approach that allows time-varying unconditional variance in the ST-GARCH model is one way. The approach of van Bellegem and von Sachs (2004) and Engle and Gonzalo Rangel (2008) is another alternative.† Unlike the ST-GARCH model, which parametrically species the dating of structural shifts in the GARCH framework, these authors non-parametrically decompose the volatility process to describe the low-frequency changes in volatility.

Following the suggestions of Lundbergh and Teräsvirta (2002), we first test the hypothesis of parameter constancy in the GARCH model before estimating the ST-GARCH model. Formally, assume the null model is $g_t = 0$ and let $\bar{\mathbf{x}}_t' = \hat{h}_t^{-1} \partial \hat{h}_t / \partial \boldsymbol{\eta}'$ under the null. In this paper, we consider the transition variable to be time, i.e. $\tau_t = t$, in order to evaluate the impact of the launch of the euro on the

†We acknowledge one of the anonymous referees for pointing out this alternative.

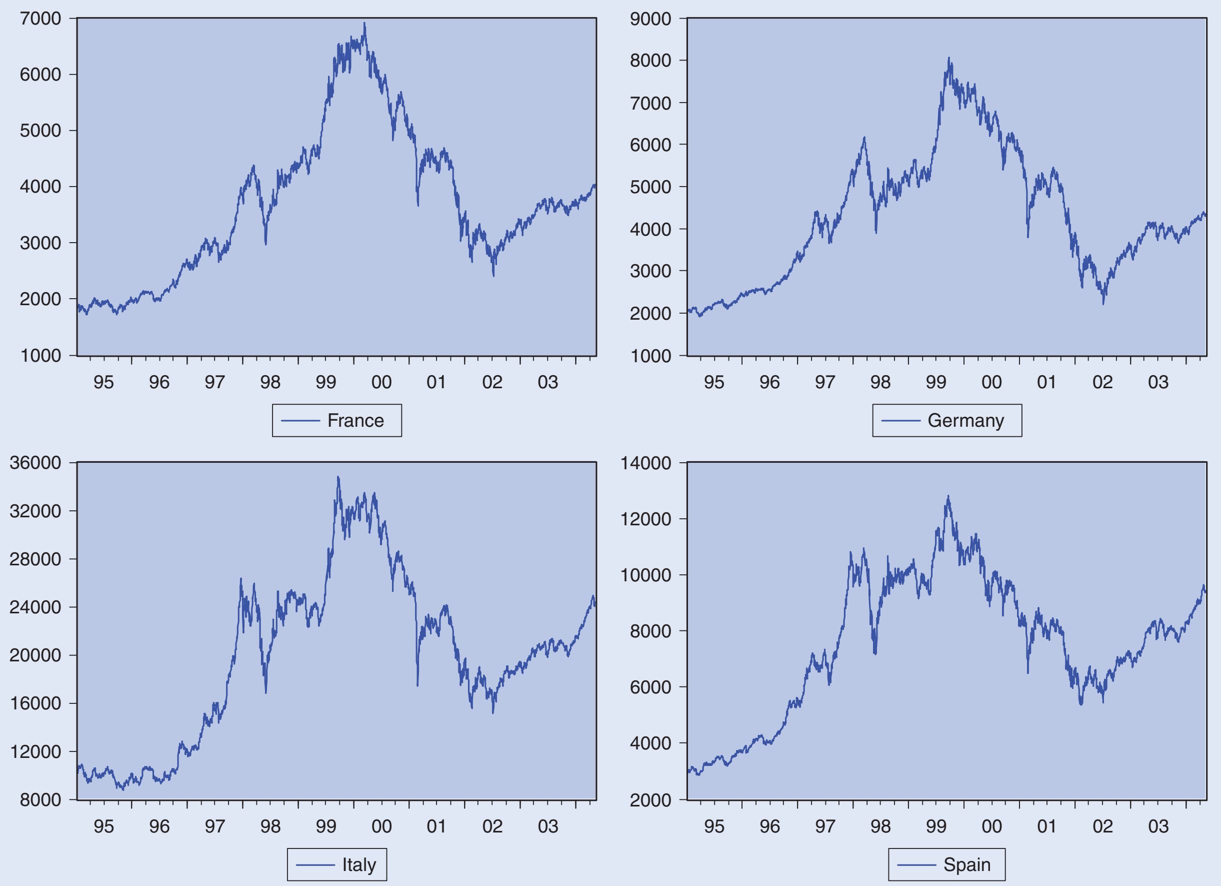

Figure 1. Daily closing prices for the stock markets of France, Germany, Italy and Spain over the period 1 January 1995 to 10 May 2004.

volatilities of the stock indexes among major EU countries. Let $\mathbf{v}_{it} = t^i \mathbf{s}_t$, $\hat{\mathbf{v}}_{it} = t^i \hat{\mathbf{s}}_t$, and $\hat{\mathbf{v}}_t = (\hat{\mathbf{v}}_{1t}, \hat{\mathbf{v}}_{2t}, \hat{\mathbf{v}}_{3t})'$ for $i = 1, 2, 3$.

The statistical test procedure can be carried out rigorously using an artificial regression as follows. First, estimate the parameters of the conditional model under the null. Let $SSR_0 = \sum_{t=1}^{T} (\hat{\varepsilon}_t^2 / \hat{h}_t - 1)^2$. Then regress $(\hat{\varepsilon}_t^2 / \hat{h}_t - 1)$ on $\bar{x}_t', \hat{v}_t'$ and collect the sum of squared residuals, SSR_1. The LM-version test statistic can be computed by $\text{LM} = T(SSR_0 - SSR_1)/SSR_0$. Or, the alternative F-version test statistic can be computed by $F = ((SSR_0 - SSR_1)/k / SSR_1/(T - p - q - 1 - k))$. The statistics can help us to determine an appropriate k to specify the ST-GARCH model. We choose k with the smallest p-values, as suggested by Lundbergh and Teräsvirta (2002).

4. Data and empirical results

All data are taken from the data set of Yahoo! Finance (http://finance.yahoo.com/). Daily stock price indices of the four European countries for the period 1 January 1995 to 10 May 2004 are adopted.† Figure 1 shows the plots for these series for France, Germany, Italy and Spain. The continuously compounded returns of stock indices are constructed by taking the first difference of the logarithmic prices. These returns processes are shown in figure 2.

Table 1 contains the summary statistics of the sample. The first sub-sample period (1 January 1995 to 31 December 1998) indicates the pre-euro phase and the second sub-sample (1 January 1999 to 10 May 2004) is the post-euro phase. We note that the sign of the average returns from all countries appears positive during the pre-euro period, but becomes negative for the post-euro period.

As shown in table 1, the standard deviation of the returns seems to be larger in the pre-euro stage. This is true for all countries except Italy. The coefficient of kurtosis is larger than 3 in all cases, indicating a fat-tailed unconditional distribution. The time series plot of the returns for all countries in figure 2 also shows that volatility clustering is evident in the pre- and post-euro

†An anonymous referee suggests extending the time period to include the current episodes of stock market volatility. We thus extend the sample period to the end of 2007 and re-estimate the whole model. The results are available upon request from the corresponding author. In general, the location parameters and shape of the estimated unconditional variances are similar between the original sample and the extended sample.

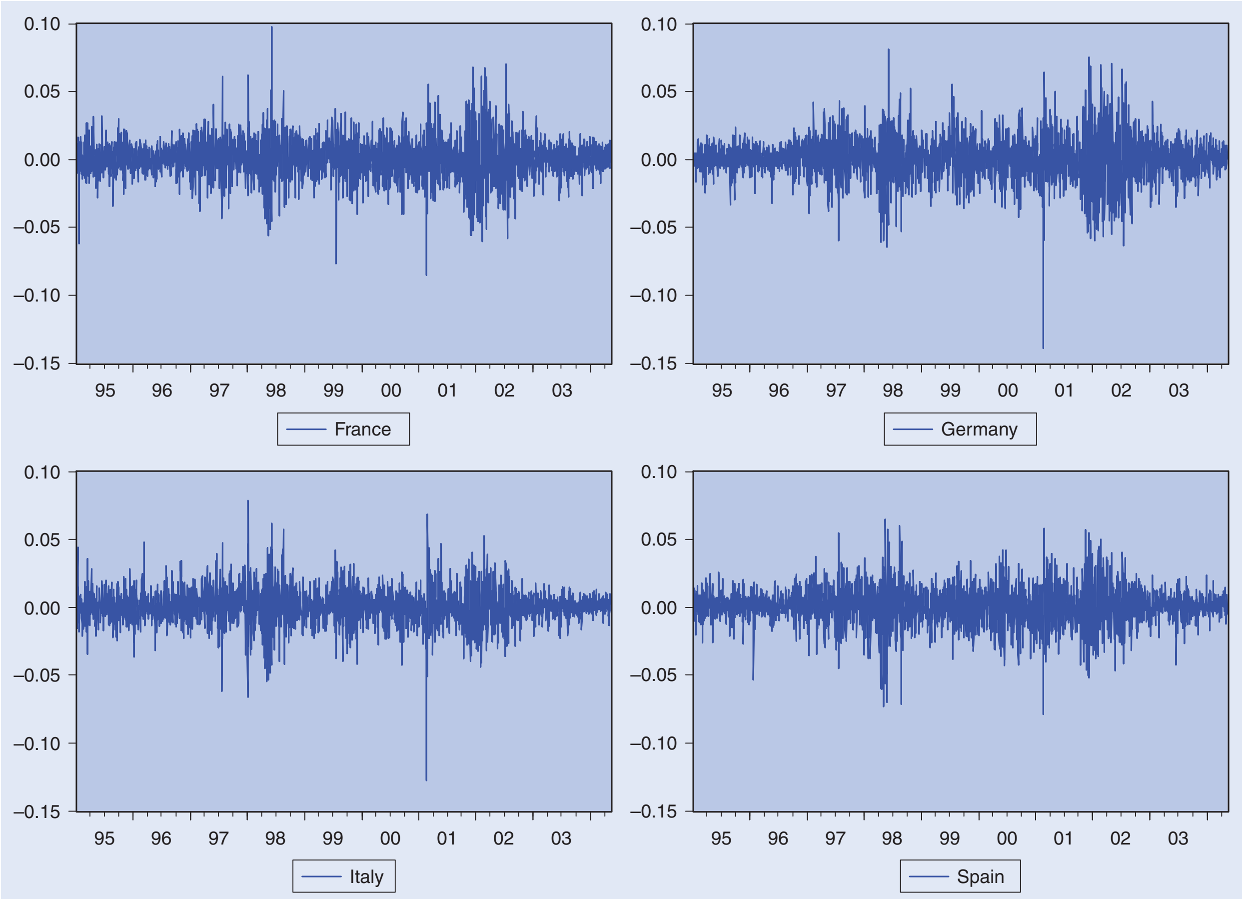

Figure 2. Daily returns for the stock markets of France, Germany, Italy and Spain over the period 1 January 1995 to 10 May 2004.

Table 1. Summary statistics (daily stock returns).

	Mean	S.D.	Skewness	Kurtosis	Maximum	Minimum
Pre-euro (1 January 1995 to 31 December 1998)						
France CAC	0.081	1.350	0.133	7.758	9.757	−6.185
Germany DAX	0.093	1.355	−0.394	6.902	8.116	−6.450
Italy MIBTEL	0.091	1.454	−0.027	5.713	7.861	−6.634
Spain IBEX 35	0.122	1.374	−0.555	7.048	6.468	−7.327
Post-euro (1 January 1999 to 10 May 2004)						
France CAC	−0.023	1.653	−0.071	5.195	7.002	−8.543
Germany DAX	−0.016	1.887	−0.195	6.173	7.553	−13.919
Italy MIBTEL	−0.082	1.358	−0.527	9.841	6.832	−12.737
Spain IBEX 35	−0.015	1.555	−0.049	4.456	6.028	−7.909

Notes: This table reports the basic descriptive statistics for the logarithmic stock returns before and after the introduction of the euro. The return is defined as $100 \times [\log(p_t) - \log(p_{t-1})]$.

period for all countries. The GARCH model is well-suited for these empirical regularities.

Sub-sample estimation results are given in table 2. There are some noticeable changes in the coefficients after the euro's introduction. Both the coefficients $\hat{\alpha}_0$ and $\hat{\alpha}_1$ clearly decreased in the second sub-sample period for Italy and Spain. The coefficient $\hat{\beta}_1$, on the contrary, increases after the euro's launch for these two countries. Changes in the coefficients for France and Germany are not apparent. It is interesting to note that the volatility persistence parameter $\hat{\alpha}_1 + \hat{\beta}_1$ is higher, approaching unity for all four countries in the second sub-sample. The unconditional variance computed using the formula $\hat{\alpha}_0/(1 - \hat{\alpha}_1 - \hat{\beta}_1)$ hence unanimously increases for all four countries after the euro's introduction. This result seems to be at odds with that of Morana and Beltratti (2002).

It is important to note that the observation in table 2 is merely descriptive. It provides no formal statistical inference for the purpose of distinguishing differences between the pre-euro and post-euro periods. Dummy variables will allow us to construct empirical models to identify changes in parameters statistically. The empirical results of the modified GARCH(1, 1) model with dummy

Table 2. GARCH(1, 1) model for the pre-euro and post-euro period.

$$R_t = \varepsilon_t$$
$$h_t = \alpha_0 + \alpha_1 \varepsilon_{t-1}^2 + \beta_1 h_{t-1}$$
$$\varepsilon_t \mid \Omega_{t-1} \sim N(0, h_t).$$

Country	σ^2	$\hat{\alpha}_0$	$\hat{\alpha}_1$	$\hat{\beta}_1$	$\hat{\alpha}_1 + \hat{\beta}_1$	$Q(25)$	$Q^2(25)$
Sample period: 1995/1/1–1998/12/31							
Pre-euro							
France	1.907	0.014 [0.063]	0.052 [0.000]	0.941 [0.000]	0.993	23.209 [0.565]	18.179 [0.835]
Germany	2.121	0.023 [0.000]	0.095 [0.000]	0.894 [0.000]	0.989	21.552 [0.661]	18.528 [0.819]
Italy	1.988	0.135 [0.000]	0.155 [0.000]	0.777 [0.000]	0.932	28.801 [0.272]	28.507 [0.285]
Spain	1.973	0.050 [0.000]	0.121 [0.000]	0.854 [0.000]	0.975	30.751 [0.197]	13.181 [0.974]
Sample period: 1999/1/1–2004/5/10							
Post-euro							
France	3.404	0.019 [0.019]	0.084 [0.000]	0.910 [0.000]	0.994	24.855 [0.470]	8.735 [0.999]
Germany	4.918	0.027 [0.038]	0.099 [0.000]	0.896 [0.000]	0.995	22.661 [0.597]	19.241 [0.785]
Italy	8.353	0.017 [0.001]	0.119 [0.000]	0.879 [0.000]	0.998	26.370 [0.388]	23.098 [0.572]
Spain	2.856	0.014 [0.012]	0.077 [0.000]	0.918 [0.000]	0.995	21.101 [0.687]	16.196 [0.909]

Notes: The numbers in square brackets are p-values. $Q(25)$ is the Ljung–Box (1978) test for serial correlation up to the 25th order in the standardized residuals. $Q^2(25)$ is the Ljung–Box test for serial correlation up to the 25th order in the squared standardized residuals. σ^2 represents the unconditional variance that can be calculated using the formula $\hat{\alpha}_0/(1-\hat{\alpha}_1-\hat{\beta}_1)$ from the GARCH model.

Table 3. Modified GARCH(1, 1) model with dummy variable.

$$R_t = \varepsilon_t$$
$$h_t = \alpha_0 + \delta_0 D_t + \alpha_1 \varepsilon_{t-1}^2 + \delta_1 D_t \varepsilon_{t-1}^2 + \beta_1 h_{t-1} + \delta_2 D_t h_{t-1}$$
$$\varepsilon_t \mid \Omega_{t-1} \sim N(0, h_t).$$

Country	$\hat{\alpha}_0$	$\hat{\alpha}_1$	$\hat{\beta}_1$	$\hat{\delta}_0$	$\hat{\delta}_1$	$\hat{\delta}_2$	$Q(25)$	$Q^2(25)$	$\log L$
Sample period: 1995/1/1–2004/5/10									
France	0.014 [0.057]	0.053 [0.000]	0.940 [0.000]	0.005 [0.645]	0.030 [0.017]	−0.027 [0.056]	28.153 [0.301]	13.072 [0.976]	−4155.591
Germany	0.025 [0.000]	0.099 [0.000]	0.890 [0.000]	0.001 [0.925]	−0.003 [0.874]	0.009 [0.647]	23.514 [0.548]	17.464 [0.864]	−4247.478
Italy	0.136 [0.000]	0.155 [0.000]	0.775 [0.000]	−0.119 [0.001]	−0.037 [0.139]	0.104 [0.004]	35.450 [0.080]	20.498 [0.720]	−3880.061
Spain	0.050 [0.000]	0.122 [0.000]	0.853 [0.000]	−0.037 [0.014]	−0.048 [0.055]	0.068 [0.018]	29.382 [0.248]	16.790 [0.889]	−4033.635

Notes: Figures in square brackets are p-values. $Q(25)$ is the Ljung–Box (1978) test for serial correlation up to the 25th order in the standardized residuals. $Q^2(25)$ is the Ljung–Box test for serial correlation up to the 25th order in the squared standardized residuals. $\log L$ denotes the log likelihood value. Before 31 December 1998, the dummy variable (D_t) is 0. After 1 January 1999, the dummy variable (D_t) is 1.

variables embedded in the variance equation are illustrated in table 3.

It seems that Germany's volatility structure of stock returns is not affected by the euro's introduction regardless of the dummy variable used to model the possible shifts in the intercept term, lagged squared residual term and lagged conditional variance term for the volatility equation. For Italy and Spain, most of the coefficients of the dummy variables embedded in the modified GARCH model are significantly different from zero. It would appear that the dynamic volatility structures have crucially changed only for these two stock markets after the euro launch. Since the only significant coefficient of the dummy variables is that of the lagged squared residual term according to the conventional significant level of 5%, one would suggest that the impact of the euro on France's stock market seems minor. However, in the following we will show that the above inferences may be false due to incorrect dating on the dummy variable or smooth changes rather than sharp shifts in the parameters of the GARCH specification.

Table 4. LM tests of parameter constancy for $k = 1$, 2, and 3.

	k		
	1	2	3
France	1.910 [0.591]	9.399 [0.152]	13.689 [0.401]
Germany	5.186 [0.159]	19.668 [0.003]	20.811 [0.020]
Italy	3.841 [0.279]	18.185 [0.006]	20.953 [0.033]
Spain	5.255 [0.154]	19.525 [0.003]	23.175 [0.021]
United Kingdom	3.842 [0.28]	15.739 [0.02]	17.675 [0.07]
Japan	8.992 [0.03]	11.555 [0.07]	16.961 [0.24]

Note: The figures in square brackets are p-values.

To avoid the problem caused by incorrect dating using dummy variables, we now discuss the analysis using the ST-GARCH model. For the purpose of comparison, two non-euro countries, e.g. the United Kingdom and Japan, are added to the analysis. We first use the LM-version statistics (for $k = 1$, 2, and 3) proposed by Lundbergh and Teräsvirta (2002) to test for parameter constancy. The null model we adopt in the tests is GARCH(1, 1).

Table 5. Estimation results of the GARCH and ST-GARCH models.

$$R_t = \varepsilon_t$$
$$h_t = \omega + \alpha_1 \varepsilon_{t-1}^2 + \beta_1 h_{t-1} + \left[\bar{\omega} + \bar{\alpha}_1 \varepsilon_{t-1}^2 + \bar{\beta}_1 h_{t-1}\right] F(t)$$
$$\text{ST-GARCH}(1,1), \text{ if } F(t) = \left(1 + \exp(-\gamma \Pi_{i=1}^{k} (\tau_t - c_i))\right)^{-1}$$
$$\text{GARCH}(1,1), \text{ if } F(t) = 0$$

	France	Germany	Italy	Spain	UK	Japan
GARCH						
ω	0.018 [0.00]	0.022 [0.00]	0.030 [0.00]	0.026 [0.00]	0.011 [0.00]	0.074 [0.00]
α_1	0.072 [0.00]	0.097 [0.00]	0.126 [0.00]	0.092 [0.00]	0.079 [0.00]	0.070 [0.00]
β_1	0.921 [0.00]	0.897 [0.00]	0.864 [0.00]	0.897 [0.00]	0.913 [0.00]	0.898 [0.00]
$\log L$	−4156.782	−4248.418	−3889.182	−4036.856	−3349.243	−4125.728
$Q(25)$	28.169 [0.30]	23.504 [0.55]	35.905 [0.07]	29.042 [0.26]	28.638 [0.28]	23.298 [0.56]
$Q^*(25)$	15.647 [0.92]	18.850 [0.80]	20.387 [0.73]	17.265 [0.87]	22.353 [0.62]	15.824 [0.92]
ST-GARCH						
k	2	2	2	2	2	1
γ	124.353 [0.06]	1746.735 [0.88]	24.230 [0.01]	46.666 [0.86]	18.569 [0.00]	35.522 [0.17]
c_1	0.193 [0.00]	0.317 [0.00]	0.086 [0.09]	0.908 [0.00]	0.593 [0.00]	0.147 [0.01]
c_2	0.886 [0.00]	0.741 [0.00]	0.754 [0.00]	0.077 [0.01]	−a	
ω	0.058 [0.00]	0.120 [0.01]	0.118 [0.00]	0.038 [0.00]	0.108 [0.01]	1.199 [0.00]
α_1	0.077 [0.00]	0.124 [0.00]	0.190 [0.00]	0.102 [0.00]	0.307 [0.00]	0.255 [0.00]
β_1	0.901 [0.00]	0.843 [0.00]	0.772 [0.00]	0.886 [0.00]	0.697 [0.00]	—b
$\bar{\omega}$	–0.058 [0.00]	–0.104 [0.02]	–0.111 [0.00]	–0.020 [0.00]	–0.096 [0.03]	–1.101 [0.00]
$\bar{\alpha}_1$	–0.096 [0.00]	–0.056 [0.00]	–0.188 [0.00]	–0.195 [0.00]	–0.306 [0.00]	–0.180 [0.00]
$\bar{\beta}_1$	0.113 [0.00]	0.077 [0.01]	0.196 [0.00]	0.151 [0.00]	0.264 [0.00]	0.887 [0.00]
$\log L$	–4126.930	–4237.707	−3862.864	−4014.367	−3332.341	−4122.644
$Q(25)$	28.841 [0.27]	24.765 [0.48]	31.949 [0.16]	27.696 [0.32]	31.552 [0.17]	23.261 [0.56]
$Q^*(25)$	17.332 [0.87]	20.772 [0.71]	20.469 [0.72]	17.986 [0.84]	27.332 [0.34]	21.950 [0.64]
Persistence						
GARCH(1, 1)	0.993	0.994	0.990	0.989	0.992	0.968
ST-GARCH						
In upper regime	0.995	0.988	0.969	0.944	0.962	0.962
In lower regime	0.978	0.968	0.962	0.988	0.983	0.255

Notes: The figures in square brackets are p-values. $Q(25)$ is the Ljung–Box statistic. $Q^*(25)$ is the statistic for testing 'no remaining ARCH' up to the 25th orders proposed by Lundbergh and Teräsvirta (2002).
[a]Estimation of the ST-GARCH model for the UK with $c_1 \neq c_2$ leads to non-convergence; $c_1 = c_2$ is then estimated instead.
[b]β_1 is statistically insignificant and thus excluded.

Not surprisingly, most of the test statistics in table 4 strongly reject parameter constancy for all countries at conventional significance levels, except for France. This indicates that there exist structural changes in the corresponding GARCH models. It can also be seen that $k=2$ yields the smallest p-value for all European countries, although the statistic is not significant at the 5% level for France. In contrast, the LM statistic with $k=1$ has the smallest p-value for Japan.

Therefore, we choose to estimate the following ST-GARCH(1, 1) model with $k=2$ for all European countries and $k=1$ for Japan:

$$\begin{aligned} R_t &= \varepsilon_t, \\ \varepsilon_t &= z_t (h_t + g_t)^{1/2}, \\ h_t &= \omega + \alpha_1 \varepsilon_{t-1}^2 + \beta_1 h_{t-1}, \\ g_t &= \left(\bar{\omega} + \bar{\alpha}_1 \varepsilon_{t-1}^2 + \bar{\beta}_1 h_{t-1}\right) F(t; \gamma, \mathbf{c}), \end{aligned} \tag{5}$$

where $F(t; \gamma, \mathbf{c}) = (1 + \exp(-\gamma \Pi_{i=1}^{k}))^{-1}$. Note that the unconditional variance in this specification of the ST-GARCH model is time-varying.†

The estimated results for the null model and the smooth-transition GARCH model as well as the corresponding diagnostic statistics for each country are shown in table 5. The estimated results of the GARCH models under the null (with whole samples) are also shown in table 5 for the purpose of comparison.

It can be seen that the statistics to test for no serial correlation up to 25th order in the standardized residuals and residuals squared, i.e. $Q(25)$ and $Q^*(25)$ for the GARCH and ST-GARCH models, are all insignificant at conventional levels for all four countries.‡

The estimated transition function $F(t)$ for each country is plotted in figure 3. It can be seen that the graphs for $F(t)$ are U-shaped for $k=2$ and S-shaped for $k=1$. It is thereby termed the upper regime when $F(t)=1$ and the

†We acknowledge the two anonymous referees for pointing this out.
‡An anonymous referee suggests that the $Q^2(25)$ statistic does not have an asymptotic χ^2 distribution under the current situation. We instead report the $Q^*(25)$ statistic to test for 'no remaining ARCH' up to 25th order proposed by Lundbergh and Teräsvirta (2002), who showed that $Q^*(25)$ is asymptotically equal to the portmanteau statistic introduced by Li and Mak (1994).

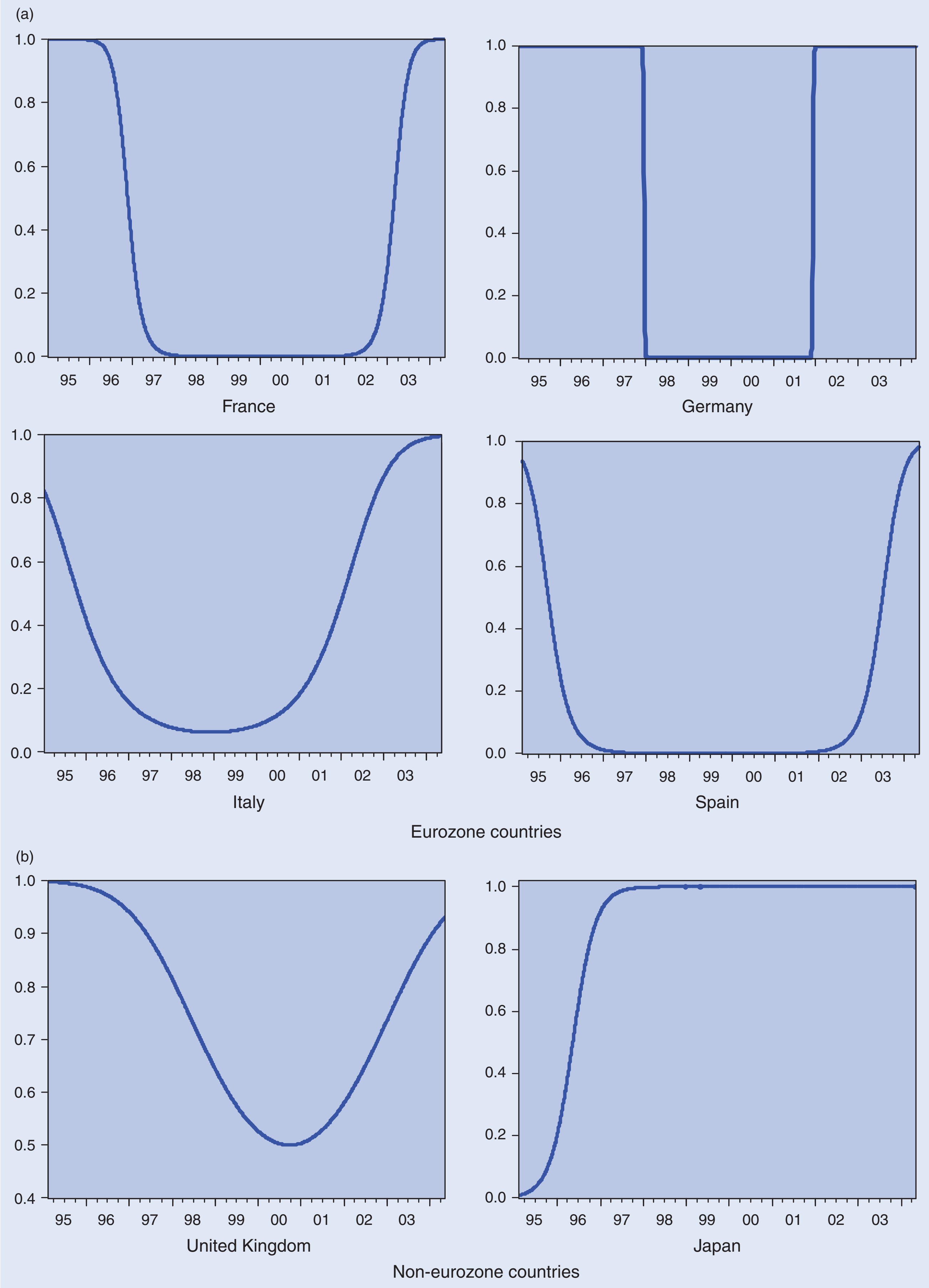

Figure 3. Estimated smooth transition functions for eurozone and non-eurozone countries.

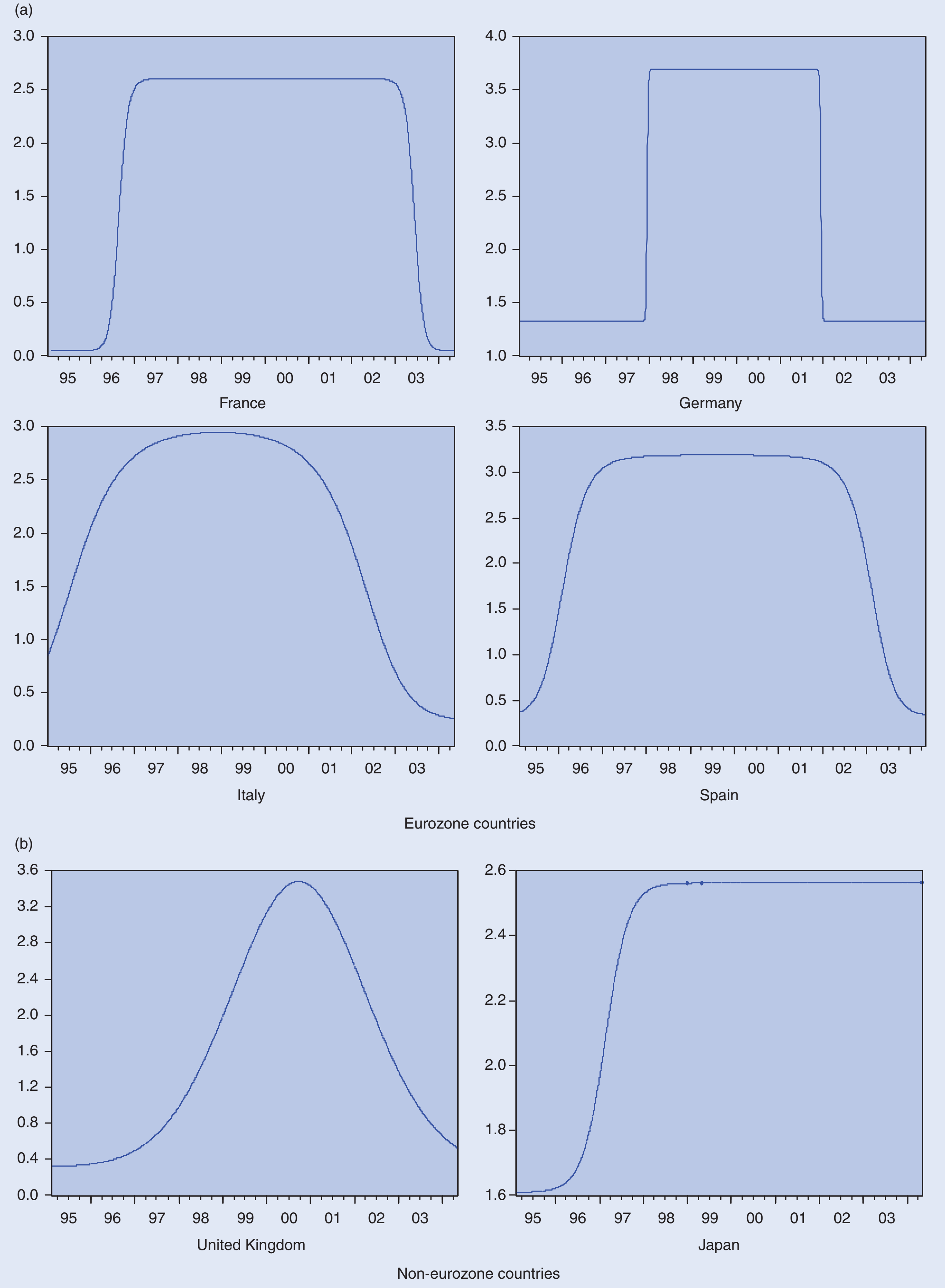

Figure 4. Estimated unconditional variance under ST-GARCH models for eurozone and non-eurozone countries.

Table 6. Estimated location parameters and corresponding calendar dates.

Country	c_1	Date	c_2	Date
France	0.193	October 22, 1996	0.886	April 16, 2003
Germany	0.317	December 18, 1997	0.741	December 16, 2001
Italy	0.086	October 20, 1995	0.754	January 21, 2002
Spain	0.077	September 21, 1995	0.930	July 2, 2003
UK	0.593	September 25, 2000		
Japan	0.147	May 16, 1996		

lower regime when $F(t)$ reaches its minimum value. In particular, the minimum values of the estimated $F(t)$ are zero or close to zero for France, Germany, Italy, and Spain. As might be expected, the United Kingdom also has a U-shaped $F(t)$ since it is one of the EU members, although is not in the eurozone. Not surprisingly, Japan's $F(t)$ is different from those of European countries.

To gain a more intuitive interpretation of the results from ST-GARCH models, we also plot the estimated time-varying unconditional variance in figure 4 and report the calendar dates corresponding to the estimated location parameters in table 6.

From figure 3 and 4 it can be seen that, for these four eurozone countries as well as the United Kingdom, the estimated volatilities switch from a lower level in the pre-euro period (before the transition period) when $F(t) = 1$ to a higher level when $F(t)$ shifts from 1 to 0 during the transition period. Finally, the volatilities move back to a lower level as $F(t)$ shifts from 0 to 1 at the end of the transition period.

It appears that France and Germany experience sharper changes while the other countries show smoother changes, as indicated in figures 3 and 4. This is because the estimated slope parameter γ (table 5) is 124.35, 1746.73, 24.23, and 44.67 for France, Germany, Italy, and Spain, respectively. The large estimated values of γ for France and Germany indicate $F(t)$ is close to a double-step function.

Note that the values of $F(t)$ for France and Germany reach their minimum value in late 1997. France's $F(t)$ then switches back to 1 from mid 2002. Interestingly, the $F(t)$ of Germany switches back to 1 earlier than that of France. One of the most interesting facts shown by the $F(t)$ of the eurozone countries is that the change in volatility occurred before the actual introduction date of the euro. The early response of the volatility, as indicated by $F(t)$, in Italy is an example. By contrast, the estimated $F(t)$ of the United Kingdom suggests a late response. Furthermore, the estimated transition function for Japan may not have anything to do with the introduction of the euro because the timing of the change may suggest no connection to the euro event.†

These earlier responses of the volatility changes for the four eurozone countries might partially explain the insignificance of the coefficients of the dummy variable in the GARCH models estimated in the previous section due to the possibly incorrect specification in timing on the dummy variable. The dating in table 6 also excludes the possibility of misidentification of the event causing volatility changes in our sample period. For example, the 'dotcom bubble' in 2001–2002 may have caused the stock volatilities to increase. Furthermore, it may have prolonged the duration of the high volatility episode. Hence, a more delicate specification separating the impacts of the euro event from the bubble may be useful for future studies.‡

In summary, the estimated results of the ST-GARCH model presented here show that the volatilities all switch from a lower level before the launch of the euro to a higher level after the euro was introduced for four eurozone countries. Specifically, for France and Germany, the volatility was low at the beginning of the sample and shifted sharply to a higher level earlier than the euro's launch. It eventually moved back to a lower level. In a similar but smoother way, the volatilities of the other two eurozone countries also shifted from a low level to a high level and then switched back to a low level. With the estimation of a more recent data set, this finding is different from those of earlier studies on the impacts of the euro. Our finding is consistent with that of Cappiello *et al.* (2006), although their focus is on co-movements among European stock markets.

The results have some useful implications. First, the successful modeling of volatility as a smooth-transition GARCH model is reasonable. In contrast, the conventional simple GARCH model with a pre-specified dummy variable or sub-samples is inappropriate. Furthermore, the sensitivity of stock market volatility to the same event suggests that the country-specific effect is important when analysing stock volatility.

5. Conclusions

The purpose of this study was to examine the changes in the dynamics of volatilities on the introduction of the euro. For this purpose we compare the dynamics of the conditional volatility process estimated by several

†We thank an anonymous referee for expressing this possible viewpoint.
‡We thank the referees for pointing this out.

models: a GARCH model, a GARCH with dummy variable model, and a smooth-transition GARCH model characterized by structural changes. We find interesting implications from the results of these models. Conventional modeling of structural changes with dummy proxies according to calendar historical events might not accurately capture the actual timing of some unobserved changes in markets. Evidence from the ST-GARCH model in this paper appears to be much more satisfactory. It can detect underlying patterns of volatility among four European markets. We obtain the clear picture that the volatility processes for the stock markets of the euro states contain structural changes before January 1, 1999, when the euro was formally introduced. Volatilities gradually increase prior to the euro's launch date and shrink after the advent of the euro. The timing for the transition point is triggered about two to three years earlier than the introduction of the euro.

The approach that allows the time-varying unconditional variance in the ST-GARCH model is one way. The non-parametric approach of van Bellegem and von Sachs (2004) and Engle and Gonzalo Rangel (2008) is an alternative. The ST-GARCH parametrically estimates the dating of structural changes in the GARCH framework, whereas in the non-parametric approach the volatility process is decomposed into two components: a deterministic, low-frequency component and a GARCH component. In addition, changes in correlations among the stocks of eurozone countries is an interesting topic, as studied by Cappiello *et al.* (2006) and Bartram *et al.* (2007). The extension of the ST-GARCH model to correlations of eurozone stock markets is left for future research.

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