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WORKING PAPER

Volatility Spillover Effects in European Equity Markets

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Volatility Spillover Effects in European Equity Markets

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Abstract

This paper quantifies the magnitude and time-varying nature of volatility spillovers from the aggregate European (EU) and US market to 13 local European equity markets. I develop a shock spillover model that decomposes local unexpected returns into a country specific shock, a regional European shock, and a global shock from the US. The innovation of the model is that regime switches in the shock spillover parameters are accounted for. I find that these regime switches are both statistically and economically important. While both the EU and US shock spillover intensity has increased over the 1980s and 1990s, the rise is more pronounced for EU spillovers. For most countries, the largest increases in shock spillover intensity are situated in the second half of 1980s and the first half of the 1990s. Increased trade integration, equity market development, and low inflation are shown to have contributed to the increase in EU shock spillover intensity. Finally, I find some evidence for contagion from the US market to a number of local European equity markets during periods of high world market volatility.

Keywords: Volatility Spillovers, Regime Switching, Contagion, EMU, Financial Integration

JEL Classification: C32, G12, G15

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1 Introduction

During the last two decades, Western Europe has gone through a period of extraordinary economic and monetary integration culminating in the introduction of the euro in January 1999. In addition, significant progress was made in strenghtening and deepening the various European capital markets[1]. In this paper, I investigate whether the efforts for more economic, monetary, and financial integration in Europe have fundamentally altered the intensity of shock spillovers from the US and aggregate European market to 13 European stock markets. A good understanding of the origins and transmission intensity of shocks is necessary for many financial decisions, including optimal asset allocation, the construction of global hedging strategies, as well as the development of various regulatory requirements, like capital requirements or capital controls.

There are several channels through which further integration may have affected the degree of interdependence in European equity markets. Further economic integration, boosted by the Single European Act in 1986, combined with the overall trend towards globalisation, should make the determinanants of cash flows more similar across countries. Recent evidence by Artis et al. (1999) and Peersman and Smets (2001) supports the hypothesis that the economic cycles of the various European countries have indeed become more and more synchronized. While further monetary and financial integration acted as important catalists for futher economic integration, they also contributed to a significant equalization of cross-country discount rates, defined as the sum of the riskfree rate and the equity risk premium. The significant convergence of inflation rates, exchange rate stability as well as further integration in the bond market resulted in a strong convergence of riskfree rates. The second component of the discount rate, the equity premium, is expected to equalize across countries because of two reasons. First, country-specific risk premia due to intra-European exchange rate risk decreased considerably in the second half of the 1990s, reflecting exchange rate stability and strong beliefs about what countries would participate to the euro. The remaining exhange rate premium, at least within the euro area, disappeared completely with the introduction of the euro in 1999. The determinants of the second component of the risk premium differ depending on whether equity markets are integrated or not. Under full integration, the equity risk premium is determined solely by risk factors common to all

[1] Various EU directives, including the second banking directive (1989), the capital adequacy directive (1993), and the investment services directive (1993) have played a crucial role in opening and deepening European financial markets.

countries, and no longer by a combination of local and global factors as is the case under partial integration. During the last two decades, various policy initiatives were taken in order to eliminate both direct and indirect barriers to international investment (see Licht (1998) for an overview). Remaining obstacles are currently being addressed by a battery of initiatives contained in the Financial Services Action Plan (FSAP). A number of recent empirical studies suggest that the degree of equity market integration is rising. Hardouvelis et al. (2002) show that the proportion of expected returns that is determined by common risk factors has increased dramatically in the run-up to the euro. Adam et al. (2002) and Adjaouté and Danthine (2002) argue that the strong increase in the share of assets invested in investment funds with an international investment strategy as well as in the proportion of non-domestic equity holdings suggests increasing European equity market integration. This may to some extent be attributed to the introduction of the single currency, which eliminated, at least within the euro area, the EU matching rule, which required insurance companies and pension funds, among others, to match liabilities in a foreing currency for a large percentage by assets in the same currency.

Apart from the focus on Europe, this paper distinguishes itself from other papers by extending the standard shock spillover model of Bekaert and Harvey (1997) and Ng (2000) to account for regime switches in the shock spillover intensity and variance-covariance parameters[2]. A number of recent papers have shown the importance of allowing for different regimes in both the conditional variance and covariance of equity returns. First, Diebold (1986) and Lamoureux and Lastrapes (1990) argued that the near integrated behavior of the volatility might be due to the presence of structural breaks, which are not accounted for by standard GARCH-models. Hamilton and Susmel (1994) and Cai (1994) were the first to allow for regime switches in the ARCH process; Gray (1996) extended their methodology to regime-switching GARCH-models. Using this methodology, several studies found the persistence in second moments to decrease significantly when different regimes are allowed for. The consequence of the spurious persistence in GARCH models is that volatility is underestimated in the high volatility state, typically during periods of low economic growth, and overestimated in the low volatility state. Second, there is considerable evidence that correla-

[2]Others who have studied information sharing between equity markets include Hamao et al. (1990), King and Wadhwani (1990), Koch and Koch (1991), King et al. (1994), Lin et al. (1994), Booth and Koutmos (1995), Karolyi (1995), Longin and Solnik (1995), Karolyi and Stulz (1996), Koutmos (1996), Booth et al. (1997), Fleming et al. (1998), Kanas (1998), Kroner and Ng (1998), and Ramchand and Susmel (1998).

tions are asymmetric: correlations are larger when markets move downwards than when they move upwards. This is especially true for extreme downside moves[3]. Recent work by Ang and Bekaert (2002b) shows that these asymmetric correlation asymmetries are well captured by a regime-swiching volatility model, but not by standard (asymmetric) GARCH models.

The main novelty of this paper is however that also the shock spillover intensities are made regime dependent. Previous studies typically used dummies to test whether important "events" (in case of EMU: acceptance of relevant directives, Single European Act, Treaty of Maastricht, official announcement of the participants to the third stage of EMU, etc.) had a significant impact on the intensity by which shocks are distributed through markets. An important problem of this approach is that these events may have been long anticipated, or may not be credible, or may just need time to become effective. Bekaert et al. (2002b) for instance look for a common, endogeneous break in a large number of financial and macroeconomic time series to determine the moment when an equity market becomes most likely integrated with world capital markets. They find that the "true" integration dates occur usually later than official liberalization dates. Clearly, this makes the use of dummy variables based on the official dates of certain important events flawed. Other studies have related shock spillover intensities to a small number of instruments. In practice however, there is considerable uncertainty both about the identity of the relevant instruments and the functional form that relates those instruments to the shock spillover intensities. Regime switching models do not have these disadvantages, as they allow the data to switch endogeneously from one state to another using a nonlinear filter.

The model presented below allows for shock spillovers from the aggregate European market, the regional market, and from the US market, a proxy for the world market. The time-variation in the sensitivities to EU and US shocks is driven by a latent regime variable. Three different regime dependent shock spillover models are estimated, each with a different interaction between the latent variables governing the EU and US shock spillover intensity. I find that regime switches in the spillover parameters are both statistically and economically important. For nearly all countries, both EU and US spillover intensities have increased significantly over the last two decades. The increase for EU shock spillover intensity is larger though, and is situated mainly in the second part of the 1980s and the first part of the

[3]For a discussion, see e.g. Ang and Bekaert (2002b), Longin and Solnik (2001), and Ang and Chen (2002).

1990s. Surprisingly, after the introduction of the euro in 1999, in many countries, sensitivities to EU and US shocks dropped considerably. On average, EU shocks explain about 15 percent of local variance, compared to 20 percent for US shocks. While the US is still the dominating market, the importance of EU shocks has increased proportionally more, hereby narrowing the gap with the US. To better understand the time-variation in the shock spillover intensity, I relate the latent regime variables to a large set of relevant economic/financial instruments. The results suggest that countries with an open economy, low inflation, and well developed financial markets share more information with the EU market. There is also some evidence that shock spillover intensity is related to the state of the business cycle. Finally, a test for market contagion is developed, similar to the one proposed by Bekaert et al (2002c). I find evidence for contagion effects from the US to the local European equity markets in times of high world market volatility.

The remainder of this paper is organized as follows. Section 2 describes the data and offers some descriptive statistics. Section 3 develops the regime dependent volatility spillover model, while section 4 reports the empirical results. The final section provides a summary and conclusions.

2 Data Analysis

I composed weekly total (dividend-adjusted) continuously compounded stock returns from 8 EMU countries[4] (Austria, Belgium, France, Germany, Ireland, Italy, the Netherlands, and Spain), three European Union (EU) countries that do not participate in EMU (Denmark, Sweden, and the UK), two countries from outside the EU (Norway, and Switzerland), and two regional markets (the aggregate European market, and the US). I take such a broad sample in order to compare shock spillover intensity between EMU, EU, and non-EU countries. The data are sampled weekly and cover the period January 1980 till August 2001, for a total of 1130 observations. For Spain and Sweden, the sample period is somewhat shorter due to data availability. I use the equity indices provided by Datastream[5], as they capture a larger share of the market and tend to be more homogeneous than other indices,

[4]The other four EMU countries, Finland, Greece, and Portugal have been left out due to limited data coverage. Luxemburg was left out because of its very tiny market capitalization.

[5]The regional European market index used here is the Datastream EU-15 index.

like those of MSCI. All returns are denominated in Deutschmark[6].

Table 1 presents some summary statistics on the weekly returns of the 13 markets under investigation, as well as for the US and EU aggregate market. There is considerable cross-sectional variation both in mean returns and standard deviations. The mean returns range from 0.235 percent for Austria to 0.34 percent for Ireland, while the returns in the Italian, Norwegian, and Swedish stock markets are the most volatile. The Jarque-Bera test rejects normality of the returns for all countries. This is caused mainly by the excess kurtosis, suggesting that any model for equity returns should accomodate this characteristic of equity returns. The ARCH test reveals that most returns exhibit conditional heteroskedasticity, while the Ljung–Box test (of fourth order) indicates significant autocorrelation in most markets.

Figure 1 plots the average 52-week moving correlation of all the individual countries with the aggregate European and US market. The estimated conditional correlations exhibit considerable variation through time, suggesting that correlations are not constant through time. The correlation with the aggregate European market is larger than with the US market. Notice also that while correlations increased in the run-up to the introduction of the euro, they decreased considerably after 1999.

3 A regime-switching volatility spillover model

The aim of this paper is to investigate the origins of time variation in correlations between 13 European equity markets and the US and EU. I allow for three sources of unexpected returns, being (1) a purely domestic shock, (2) a regional European shock, and (3) a global shock from the US. The model I propose is an extension of Bekaert and Harvey (1997), in a sense that I distinguish between two regional sources of shocks instead of one world shock, and of Ng (2000), Fratzscher(2001), and Bekaert et al. (2002c), as I allow for regime switches in the spillover parameters. The remainder of this section is organized as follows. In section 3.1, I describe a bivariate model for the US and European returns. The estimated innovations for the US and Europe are then used as inputs for the univariate volatility spillover model, which is described in section 3.2. In section 3.3, I discuss the estimation procedure as well as some specification tests.

[6] As from January 1999, for EMU countries, the fixed euro-deutschmark exchange rate is used to translate euro returns into deutschmark returns. Returns for non-EMU countries are first translated into euros, and then into deutschmarks.

3.1 A Bivariate model for the US and Europe

3.1.1 Standard model without Regime Switches

The joint process for European and US returns is governed by the following set of equations:

$$
\begin{aligned}
r_t &= \boldsymbol{\mu}_{t-1} + \boldsymbol{\varepsilon}_t \\
\boldsymbol{\mu}_{t-1} &= k_0 + \mathbf{K} r_{t-1} \\
\varepsilon_t | \Omega_{t-1} &\backsim N(0, \mathbf{H}_t)
\end{aligned}
\tag{1}
$$

where $r_t = [r_{eu,t}, r_{us,t}]'$ represents the weekly return of respectively the aggregate European and US market at time t, $\boldsymbol{\varepsilon}_t = [\varepsilon_{eu,t}, \varepsilon_{us,t}]'$ is a vector of innovations, $k_0 = [k_{eu}, k_{us}]'$, and $\mathbf{K} = [k_{eu}^{eu}, k_{eu}^{us}; k_{us}^{eu}, k_{us}^{us}]$ a two by two matrix of parameters linking lagged returns in the US and Europe to expected returns. Other studies have used more sophisticated information variables, like dividend yields, changes in the term structure, default spreads, and short term interest rates. I limit myself to lagged returns, as predictability of the other information variables is very low in weekly data, and because I want to focus on volatility spillovers rather than on spillovers in mean returns[7]. I provide four different (bivariate) specifications for the conditional variance-covariance matrix $\mathbf{H}_t$: a constant correlation model, a bivariate BEKK model, a regime-switching normal model, and a regime-switching GARCH model. Where appropriate, I will test whether there is evidence of asymmetry in the volatility equation.

Constant Correlation Model The constant correlation model was first proposed by Bollerslev (1990) and is the most restrictive of the models that is used here. It can be represented in the following way:

$$
\mathbf{H}_t = \digamma_t \Gamma \digamma_t \tag{2}
$$

$$
\begin{aligned}
\digamma_t &= \begin{bmatrix} h_{eu,t} & 0 \\ 0 & h_{us,t} \end{bmatrix} \\
\Gamma &= \begin{bmatrix} 1 & \rho \\ \rho & 1 \end{bmatrix}
\end{aligned}
$$

[7] As a robustness check, I will test whether the model's residuals are orthogonal to a set of information variables.

where ρ represents the correlation coefficient. I model the conditional variance $h_{i,t}$, where $i = \{eu, us\}$, as a simple GARCH(1,1)-model extended to allow for asymmetry (see Glosten et al.(1993)).

$$h_{i,t}^2 = \psi_{i,o} + \psi_{i,1}\varepsilon_{i,t-1}^2 + \psi_{i,2}h_{i,t-1}^2 + \psi_{i,3}\varepsilon_{i,t-1}^2 I\{\varepsilon_{i,t-1} < 0\} \tag{3}$$

where I is an indicator function for $\varepsilon_{i,t-1}$ and ψ_i a vector of parameters. Negative shocks increase volatility if $\psi_{i,3} > 0$.

Asymmetric BEKK Model I use the asymmetric version of the BEKK model of Baba et al. (1989), Engle and Kroner (1995), and Kroner and Ng (1998), which is given by

$$\mathbf{H}_t = \mathbf{C}'\mathbf{C} + \mathbf{A}'\boldsymbol{\varepsilon}_{t-1}\boldsymbol{\varepsilon}_{t-1}'\mathbf{A} + \mathbf{B}'\mathbf{H}_{t-1}\mathbf{B} + \mathbf{D}'\boldsymbol{\eta}_{t-1}\boldsymbol{\eta}_{t-1}'\mathbf{D} \tag{4}$$

where

$$\boldsymbol{\eta}_{t-1} = \boldsymbol{\varepsilon}_{t-1} \odot \mathbf{1}\{\boldsymbol{\varepsilon}_{t-1} < 0\} \tag{5}$$

The symbol $\odot$ is a Hadamard product representing an element by element multiplication, and $\mathbf{1}\{\boldsymbol{\varepsilon}_{t-1} < 0\}$ is a vector of individual indicator functions for the sign of the errors $\varepsilon_{eu,t}$ and $\varepsilon_{us,t}$. Matrix $\mathbf{C}$ is a 2 by 2 lower triangular matrix of coefficients, while $\mathbf{A}$, $\mathbf{B}$,and $\mathbf{D}$ are 2 by 2 matrices of coefficients. The advantage of this model is that $\mathbf{H}_t$ is guaranteed to be positive definite.

3.1.2 Model with Regime-Switches

I propose two models that make both the conditional expected return and the conditional variance regime-dependent: a regime-switching bivariate normal model, and a regime-switching GARCH model.

Regime Switching Bivariate Normal This model allows the returns r_t to be drawn from a mixture of two bivariate normal distributions. Which distribution is used at what time, depends on the regime the process is in. I distinguish between two different states, $S_t = 1$ and $S_t = 2$, and two bivariate normal distributions:

$$r_t|\Omega_{t-1} = \begin{cases} N(\boldsymbol{\mu}_{t-1}(S_t = 1), \mathbf{H}(S_t = 1)) \\ N(\boldsymbol{\mu}_{t-1}(S_t = 2), \mathbf{H}(S_t = 2)) \end{cases} \tag{6}$$

Both the conditional mean return $\boldsymbol{\mu}_{t-1}$ and the variance $\mathbf{H}$ are made regime dependent. To facilitate estimation, in the conditional mean specification, only the intercept k_0 is allowed to depend upon the latent regime variable S_t. The regimes follow a two-state Markov chain with transition matrix:

$$\Pi = \begin{pmatrix} P & 1-P \\ 1-Q & Q \end{pmatrix} \tag{7}$$

where the transition probabilities are given by $P = prob(S_t = 1|S_{t-1} = 1; \Omega_{t-1})$, and $Q = prob(S_t = 2|S_{t-1} = 2; \Omega_{t-1})$.

A regime-switching GARCH Model In the regime-switching bivariate normal model, volatility is restricted to be constant within a regime. The (generalized) regime-switching volatility models of Hamilton and Susmel (1994), Cai (1994), and Gray (1996) combine the advantages of a regime-switching model with the volatility persistence associated with GARCH effects. I allow the regime-dependent conditional volatility within a regime to follow an asymmetric GARCH(1,1) model:

$$r_t|\Omega_{t-1} = \begin{cases} N(\boldsymbol{\mu}_{t-1}(S_t = 1), \mathbf{H}_t(S_t = 1)) \\ N(\boldsymbol{\mu}_{t-1}(S_t = 2), \mathbf{H}_t(S_t = 2)) \end{cases} \tag{8}$$

and

$$\mathbf{H}(S_t = i) = \mathbf{C}(S_t = i)' \mathbf{C}(S_t = i) + \mathbf{A}\boldsymbol{\varepsilon}_{t-1}\boldsymbol{\varepsilon}'_{t-1}\mathbf{A} + \mathbf{B}'\mathbf{H}_{t-1}\mathbf{B} + \mathbf{D}'\boldsymbol{\eta}_{t-1}\boldsymbol{\eta}'_{t-1}\mathbf{D} \tag{9}$$

for $i = 1, 2$. The regime variable S_t follows the same two-state markov chain with transition probability Π as in equation (7). As even for two regimes this model requires a lot of parameters to be estimated, I restrict the matrices **A**, **B**, and **D** to be diagonal and regime independent, while only the intercept k_0 is allowed to switch in the mean equation. In addition, both the mean and volatility are forced to switch jointly. As in Gray (1996), $\boldsymbol{\varepsilon}_{t-1}$, $\mathbf{H}_{t-1}$, and $\boldsymbol{\eta}_{t-1}$were made regime independent by averaging over the ex-ante probabilities. For instance, $\mathbf{H}_{t-1}$ is calculated as follows:

$$\mathbf{H}_{t-1} = p_{1,t-1}\mathbf{H}_{t-1}(S_1) + (1 - p_{1,t-1})\mathbf{H}_{t-1}(S_2) \tag{10}$$

where $p_{1,t-1} = prob(S_{t-1} = 1|\Omega_{t-2})$.

3.2 Univariate spillover model

Similar in spirit to Bekaert and Harvey (1997), Ng (2000), and Fratzscher (2001), local unexpected returns are - apart from by a purely local component - allowed to be driven by innovations in US and European returns. I orthogonalize the innovations from the aggregate European market and the US, the residuals $\hat{\varepsilon}_{eu,t}$ and $\hat{\varepsilon}_{us,t}$ from the first step, assuming that the European return shock is driven by a purely idiosyncratic shock and by the US return shock. The orthogonalization process is outlined in Appendix 1. I denote the orthogonalized European and US innovations by $\hat{e}_{eu,t}$ and $\hat{e}_{us,t}$ and their variances by $\sigma^2_{eu,t}$ and $\sigma^2_{us,t}$. This section outlines several univariate volatility spillover models that take the orthogonalized EU and US innovations as given. In section 3.2.1., I describe the volatility spillover model that is standard in the literature. In section 3.2.2., a model is developed that allows for regime switches in the spillover parameters. Three different assumptions are made regarding the interaction between the switches in spillover intensity from the EU and US respectively.

3.2.1 A simple model without regime shifts

The univariate constant shock spillover model for country i is represented by the following set of equations:

$$\begin{aligned} r_{i,t} &= \mu_{i,t-1} + \varepsilon_{i,t} \\ \varepsilon_{i,t} &= e_{i,t} + \gamma_i^{eu}\hat{e}_{eu,t} + \gamma_i^{us}\hat{e}_{us,t} \\ e_{i,t}|\Omega_{t-1} &\backsim N(0, \sigma^2_{i,t}) \end{aligned} \quad (11)$$

where $e_{i,t}$ is a purely idiosyncratic shock which is assumed to follow a conditional normal distribution with mean zero and variance $\sigma^2_{i,t}$. For simplicity, the expected return $\mu_{i,t-1}$ is a function of lagged EU, US, and local returns only. The parameters γ_i^{eu} and γ_i^{us} govern the (constant) spillover effects from respectively European and US shocks on local return innovations. The conditional variance $\sigma^2_{i,t}$ is modeled as a simple asymmetric GARCH(1,1) process.

$$\sigma^2_{i,t} = \psi_{i,o} + \psi_{i,1}e^2_{i,t-1} + \psi_{i,2}\sigma^2_{i,t-1} + \psi_{i,3}\varepsilon^2_{i,t-1}I\{\varepsilon_{i,t-1} < 0\} \quad (12)$$

3.2.2 Model with regime shifts in the spillover parameters

As argued in the introduction, shock spillover intensities are however likely to change through time. In the model I propose here, I do not specify ex-ante the time-varying character of shock spillover intensity, but let it depend on a latent regime variable. I first rewrite equation (11) to include the possibility of switching between states:

$$\begin{aligned}
r_{i,t} &= \mu_{i,t-1} + \varepsilon_{i,t} \\
\varepsilon_{i,t} &= e_{i,t} + \gamma_i^{eu}(S_{i,t}^{eu})\hat{e}_{eu,t} + \gamma_i^{us}(S_{i,t}^{us})\hat{e}_{us,t} \\
e_{i,t}|\Omega_{t-1} &\backsim N(0, \sigma_{i,t}^2)
\end{aligned} \tag{13}$$

where $S_i^z = \{1, 2\}$, $z = \{eu, us\}$. To keep the analysis tractable, I limit the number of states to two. The spillover parameters are then given by

$$\gamma_{i,t}^{eu} = \begin{cases} \gamma_{i,t,1}^{eu} & \text{if} \quad S_{i,t}^{eu} = 1 \\ \gamma_{i,t,2}^{eu} & \text{if} \quad S_{i,t}^{eu} = 2 \end{cases}$$

and

$$\gamma_{i,t}^{us} = \begin{cases} \gamma_{i,t,1}^{us} & \text{if} \quad S_{i,t}^{us} = 1 \\ \gamma_{i,t,2}^{us} & \text{if} \quad S_{i,t}^{us} = 2 \end{cases}$$

Following Hamilton (1988, 1989, 1990), S_i^z evolves according to a first-order Markov chain. The conditional probabilities of remaining in/switching from state are then defined as:

$$\begin{aligned}
P(S_{i,t}^z &= 1|\mathrm{S}_{i,t-1}^z = 1) = P_i^z \\
P(S_{i,t}^z &= 2|S_{i,t-1}^z = 1) = 1 - P_i^z \\
P(S_{i,t}^z &= 2|S_{i,t-1}^z = 2) = Q_i^z \\
P(S_{i,t}^z &= 1|S_{i,t-1}^z = 2) = 1 - Q_i^z
\end{aligned}$$

Similar to Hamilton and Lin (1996), Susmel (1998), and Cappiello (2000), I distinguish between three possible variations of S_i^{eu} and S_i^{us}.

Common States In this case, the forces which govern shock spillover intensities from the US and regional European market are the same. Consequently, the latent variables S_i^{eu} and S_i^{us} are identical, or $S_{i,t}^{eu}$=$S_{i,t}^{us} = S_{i,t}$. This assumption yields the following simple transition matrix Π :

$$\Pi_i = \begin{bmatrix} P_i & 1 - P_i \\ 1 - Q_i & Q_i \end{bmatrix}$$

where $P_i = P(S_{i,t} = 1|S_{i,t-1} = 1)$, and $Q_i = P(S_{i,t} = 2|S_{i,t-1} = 2)$.

Independent States Shifts in shock spillover intensity from the US and regional European markets may be completely unrelated. For instance, shock spillovers from the regional European market may have shifted to a higher state with the evolution towards an Economic and Monetary Union (EMU), while shock spillovers from the US may be determined by the state of the US business cycle. The combination of $S_{i,t}^{eu}$ and $S_{i,t}^{us}$ yields a new latent variable $S_{i,t}$:

$$\begin{aligned}
S_{i,t} &= 1 \quad \text{if } S_{i,t}^{eu} = 1 \text{ and } S_{i,t}^{us} = 1 \\
S_{i,t} &= 2 \quad \text{if } S_{i,t}^{eu} = 2 \text{ and } S_{i,t}^{us} = 1 \\
S_{i,t} &= 3 \quad \text{if } S_{i,t}^{eu} = 1 \text{ and } S_{i,t}^{us} = 2 \\
S_{i,t} &= 4 \quad \text{if } S_{i,t}^{eu} = 2 \text{ and } S_{i,t}^{us} = 2
\end{aligned}$$

The assumption of independence between states significantly simplifies the transition matrix Π_i, which is now the product of the probabilities that drive $S_{i,t}^{eu}$ and $S_{i,t}^{us}$ (for a formal derivation, see appendix 2):

$$\Pi_i = \begin{bmatrix} P_i^{eu}P_i^{us} & (1-P_i^{eu})P_i^{us} & P_i^{eu}(1-P_i^{us}) & (1-P_i^{eu})(1-P_i^{us}) \\ (1-Q_i^{eu})P_i^{us} & Q_i^{eu}P_i^{us} & (1-Q_i^{eu})(1-P_i^{us}) & Q_i^{eu}(1-P_i^{us}) \\ P_i^{eu}(1-Q_i^{us}) & (1-P_i^{eu})(1-Q_i^{us}) & P_i^{eu}Q_i^{us} & (1-P_i^{eu})Q_i^{us} \\ (1-Q_i^{eu})(1-Q_i^{us}) & Q_i^{eu}(1-Q_i^{us}) & (1-Q_i^{eu})Q_i^{us} & Q_i^{eu}Q_i^{us} \end{bmatrix} \tag{14}$$

General case Instead of imposing a structure on the transition matrix, I can let the data speak for itself. Define the transition probabilities as $p_{jj'} = P(S_t = j'|S_{t-1} = j)$, for $j, j' = 1, ..., 4$ and the associated switching probability matrix Π_i as[8]:

$$\Pi_i = \begin{bmatrix} p_{11} & p_{12} & p_{13} & p_{14} \\ p_{21} & p_{22} & p_{23} & p_{24} \\ p_{31} & p_{32} & p_{33} & p_{34} \\ p_{41} & p_{42} & p_{43} & p_{44} \end{bmatrix} \tag{15}$$

The only constraints I have to impose is that the rows sum to one, or $\sum_{j'=1}^{4} p_{jj'} = 1$, for $j = 1, ..., 4$, and that all $p_{jj'} \geqslant 0$. This general specification nests the case of independent states, which allows me to test whether the added flexibility of the general model is statistically significant using a standard likelihood ratio test.

[8]For notational clarity, the country specific subscript i has been omitted from the transition probabilities $p_{jj'}$

3.2.3 Variance Ratios and Conditional Correlations

In this section, I decompose total local volatility $h_{i,t}$ in three components: (1) a component related to European conditional volatility, (2) a component related to US conditional volatility, and (3) purely local volatility. Recall the expression for shocks to local equity returns:

$$\varepsilon_{i,t} = e_{i,t} + \gamma_i^{eu}(S_{i,t}^{eu})\hat{e}_{eu,t} + \gamma_i^{us}(S_{i,t}^{us})\hat{e}_{us,t}$$

Assume now that the purely local shocks $e_{i,t}$ are uncorrelated across countries, $E\left[e_{i,t}e_{j,t}\right] = 0, \forall i \neq j$, and uncorrelated with the European and US benchmark index: $E\left[e_{i,t}\hat{e}_{eu,t}\right] = 0$, $E\left[e_{i,t}\hat{e}_{us,t}\right] = 0, \forall i$. Moreover, $\hat{e}_{eu,t}$ and $\hat{e}_{us,t}$ are orthogonalized in the first step. We obtain regime-independent shock spillover intensities by integrating over the states:

$$\tilde{\gamma}_i^{eu} = p_{1,t}\gamma_i^{eu}(S_{i,t}^{eu} = 1) + (1 - p_{1,t})\,\gamma_i^{eu}(S_{i,t}^{eu} = 2) \tag{16}$$

$$\tilde{\gamma}_i^{us} = p_{1,t}\gamma_i^{us}(S_{i,t}^{us} = 1) + (1 - p_{1,t})\,\gamma_i^{us}(S_{i,t}^{us} = 2) \tag{17}$$

This implies that:

$$E[\varepsilon_{i,t}^2|\Omega_{t-1}] = h_{i,t} = \sigma_{i,t}^2 + (\tilde{\gamma}_i^{eu})^2\,\sigma_{eu,t}^2 + (\tilde{\gamma}_i^{us})^2\,\sigma_{us,t}^2 \tag{18}$$

$$E[\varepsilon_{i,t}\hat{e}_{eu,t}|\Omega_{t-1}] = h_{i,eu,t} = (\tilde{\gamma}_i^{eu})\,\sigma_{eu,t}^2 \tag{19}$$

$$E[\varepsilon_{i,t}\hat{e}_{us,t}|\Omega_{t-1}] = h_{i,us,t} = (\tilde{\gamma}_i^{us})\,\sigma_{us,t}^2 \tag{20}$$

It follows that

$$\rho_{i,t}^{eu} = (\tilde{\gamma}_i^{eu})\,\frac{\sigma_{eu,t}}{\sqrt{h_{i,t}}} \tag{21}$$

$$\rho_{i,t}^{us} = (\tilde{\gamma}_i^{us})\,\frac{\sigma_{us,t}}{\sqrt{h_{i,t}}} \tag{22}$$

Equation (18) shows that the conditional volatility in market i is positively related to the conditional variance in the European and US market, as well as to the shock spillover intensity. Similarly, according to equations (21) and (22), the conditional correlations of local returns with US and European returns will depend upon the regime dependent shock spillover intensity, and the ratio of the volatility of the foreign market returns (US or EU) and the local returns. According to this model, correlations between local returns and EU and US returns will generally be high when the shock spillover parameters $\tilde{\gamma}_{i,t}^{eu}$ and $\tilde{\gamma}_{i,t}^{us}$ are high, or/and when European and US volatility is high relative to local volatility.

Finally, I also investigate the (relative) proportion of conditional variance that is explained by European and US market shocks. This ratio indicates what percentage of local shocks can be explained by EU and US shocks respectively. These ratios are computed as follows:

$$VR_{i,t}^{eu} = \frac{\left(\tilde{\gamma}_i^{eu}(S_{i,t}^{eu})\right)^2 \sigma_{eu,t}^2}{h_{i,t}} = \left(\rho_{i,t}^{eu}\right)^2 \quad (23)$$

$$VR_{i,t}^{us} = \frac{\left(\tilde{\gamma}_i^{us}(S_{i,t}^{us})\right)^2 \sigma_{us,t}^2}{h_{i,t}} = \left(\rho_{i,t}^{us}\right)^2 \quad (24)$$

3.3 Estimation and Specification Tests

3.3.1 Estimation

The natural procedure would be to estimate return models for the domestic country, the US, and Europe jointly. However, given the large number of parameters that would have to be estimated in this trivariate system, I follow a two-step procedure similar to the one followed by Bekaert and Harvey (1997) and Ng (2000). First, a bivariate model is estimated for the US and European returns. I estimate the different models outlined in section 3.1 and choose the best model based on the specification tests outlined below. I can however not use the European index as such, as shock spillovers from Europe to the individual countries may be spuriously high because the European index consists partly of the country under analysis. The bias may be especially high for the larger stock markets, like Germany, France, and especially the UK. Therefore, the EU and US innovations imposed on the different spillover models in the second step are different for each country, in a sense that the market-weighted EU index used consists of all country index returns except of the country under investigation.

In appendix 3, I show what conditions are needed to make this two-step procedure internally consistent in the general case of regime switches both in the first and second step. In both steps, I estimate the parameters by maximum likelihood, assuming a conditional normally distributed error term. I use the non-linear optimization algorithm of Broyden, Fletcher, Goldfarb, and Shanno (BFGS) to optimize the loglikelihood function. To avoid local maxima, all estimations are started at least from 10 different starting values. In order to avoid problems due to non-normality in excess returns, I provide Quasi-ML estimates (QML), as proposed by Bollerslev and Woolridge (1992).

3.3.2 Specification Tests

I use three tests to distinguish between the different models. First, if the model is correctly specified, the standardized residuals should be standard normally distributed. The latter hypothesis is tested using a GMM procedure. A second statistic investigates how well the regime-switching models can distinguish between regimes. Finally, we test for regimes using an empirical likelihood ratio test.

3.3.2.1. Test on Standardized Residuals

Bivariate Model To check whether the models are correctly specified, as well as to choose the best performing model, I follow a procedure similar to the one proposed by Richardson and Smith (1993), and used by Bekaert and Harvey (1997), Bekaert and Wu (2000), and Ng (2000) among others. I calculate standardized residuals, $\hat{z}_t = \hat{\mathbf{C}}_t^{\prime -1}\hat{\varepsilon}_t$, where $\mathbf{C}_t$ is obtained through a Choleski decomposition of $\mathbf{H}_t$. These standardized residuals should follow a multivariate standard normal distribution conditional on time $t-1$ information if the model is correctly specified. I then have the following orthogonality conditions to test:

(a) $E[\ \hat{z}_{i,t}\ \hat{z}_{i,t-j}] = 0,$ for $i = EU, US$
(b) $E[\ (\hat{z}_{i,t}^2 - 1)(\ \hat{z}_{i,t-j}^2 - 1)] = 0,$ for $i = EU, US$
(c) $E[\ (\hat{z}_{eu,t}\ \hat{z}_{us,t})(\hat{z}_{eu,t-j}\hat{z}_{us,t-j})] = 0$
(d) $E[\ \hat{z}_{i,t}^3] = 0$ for $i = EU, US$
(e) $E[\ \hat{z}_{eu,t}^2\hat{z}_{us,t}] = 0$
(f) $E[\ \hat{z}_{eu,t}\hat{z}_{us,t}^2] = 0$
(g) $E[\ \hat{z}_{i,t}^4 - 3] = 0$ for $i = EU, US$
(h) $E[\ (\hat{z}_{eu,t}^2 - 1)(\hat{z}_{us,t}^2 - 1)] = 0$

$j = 1, ..., \tau$. All moment restrictions are obtained using the generalized method of moments (Hansen (1982)). Conditions (a), (b), and (c) test whether there is any serial correlation left in $\{\hat{z}_{i,t}\}$, $\{\hat{z}_{i,t}^2 - 1\}$, and $\{\hat{z}_{eu,t}\hat{z}_{us,t}\}$. For four lags ($\tau = 4$), this yields three tests that are asymptotically distributed as a χ^2 distribution with four degrees of freedom. I also perform a joint test, which has 12 degrees of freedom. Conditions (d)-(h) test whether $\hat{z}_t$ follows a bivariate standard normal distribution. Equations (d) and (g) test whether the skewness and kurtosis are significantly different from those implied by a standard normal distrution. Both are asymptotically $\chi^2(1)$ distributed. Equations (e) and (f) test for cross-skewness; while equation (h) investigates whether there is any cross-kurtosis left in the residuals. These tests follow a χ^2 distribution with respectively two and one degrees of freedom. I also perform a joint test, which follows a χ^2 distribution with 7 degrees of freedom.

Univariate Model To check whether the models are correctly specified for country i, I investigate whether the standardized residuals $\hat{z}_{i,t} = \hat{e}_{i,t}/\hat{\sigma}_{i,t}$ violate the following orthogonality conditions, as implied by a standard normal distribution:

(a) $E[\ \hat{z}_{i,t}] = 0$
(b) $E[\ \hat{z}_{i,t},\ \hat{z}_{i,t-j}] = 0$
(c) $E[\ \hat{z}^2_{i,t} - 1] = 0$
(d) $E[(\ \hat{z}^2_{i,t} - 1)(\ \hat{z}^2_{i,t-j} - 1)] = 0$
(e) $E[\ \hat{z}^3_{i,t}] = 0$
(f) $E[\ \hat{z}^4_{i,t} - 3] = 0$

$j = 1, ..., \tau$. All moment restrictions are tested using the generalized method of moments. A test on the correct specification of the conditional mean is implicit in (b), which provides us for $\tau = 4$ with a χ^2-statistic with four degrees of freedom. A similar test is conducted on the conditional variance, using moment condition (d). The distributional assumptions of the model are tested by examining conditions (a), (c), (e), and (f). This results in a χ^2-statistic with four degrees of freedom. Finally, I jointly test all restrictions, which implies (again, for $\tau = 4$) a test with 12 degrees of freedom.

Notice that test statistics derived from this GMM procedure follow a χ^2 distribution asymptotically only. However, Bekaert and Harvey (1997) - in a similar setting - performed a Monte Carlo analysis to derive the small-sample distribution of this test statistic, and found that it is fairly close to a χ^2 distribution.

3.3.2.2. Regime Classification

Ang and Bekaert (2002a) developed a summary statistic which captures the quality of a model's regime qualification performance. They argue that a good regime-switching model should be able to clasify regimes sharply. This is the case when the smoothed (ex-post) regime probabilities $p_{j,t} = P(S_{i,t} = j|\Omega_T)$ is close to either one or zero. Inferior models however will exhibit p_j values closer to $1/k$, where k is the number of states. For $k = 2$, the regime classification measure (*RCM1*) is given by

$$RCM1 = 400 \times \frac{1}{T}\sum_{t=1}^{T} p_t(1 - p_t) \tag{25}$$

where the constant serves to normalize the statistic to be between 0 and 100. A perfect model will be associated with a RCM1 close to zero, while a model that cannot distinguish between regimes at all will produce a RCM1 close to 100. Ang and Bekaert (2002a)'s generalization of this formula to the multiple state case has many undesirable features[9]. I therefore propose the following adapted measure, denoted by RCM2:

$$RCM2 = 100 \times (1 - \frac{k}{k-1}\frac{1}{T}\sum_{t=1}^{T}\sum_{i=1}^{k}\left(p_{i,t} - \frac{1}{k}\right)^2) \tag{26}$$

RCM2 lies between 0 and 100, where the latter means that the model cannot distinguish between the regimes. However, contrary to the multi-state RCM proposed by Ang and Bekaert (2002a), this measure does only produce low values when the model consistently attaches a high probability to one state only. This *RCM2* also satisfies other ordering requirements: *RCM2* for instance prefers a model that is able to eliminate 2 states relative to one state only. In addition, in the two state case, it is easy to show that *RCM2* is identical to *RCM1*. Table 2 reports the *RCM2* for different number of states and for different probability structures.

3.3.2.3 Testing for Regimes

While the specification tests and the regime classification measure may indicate whether the data generating process exhibits regimes or not, they do not constitute a formal test. Unfortunately, there is no straightforward test for regimes as the usual χ^2 asymptotic tests do not apply because of the presence of nuisance parameters under the null. This means that starting e.g. in regime 1, all parameters under regime two are not identified under the null hypothesis of no regimes. Hansen (1996) developed an asymptotic test that overcomes this problem. As this procedure is difficult to implement, in this paper, I will use an empirical likelihood ratio test similar to Ang and Bekaert (2002b). In a first step, the likelihood ratio statistic of the regime-switching model against the null of one regime is calculated. Second, N series (of lenght T, the sample length) are generated based upon a model with no regime switches. Examples of such models in this paper include the constant correlation model in section 3.1.1 and the constant volatility spillover model of section 3.2.1. For each of the N series, both the model with and without regime switches is estimated. The likelihood values are

[9] More specifically, their measure produces small *RCM*'s as soon as one state has a very low probability, even if the model cannot distinguish between the other states.

stored in respectively L_{RS} and L_{NRS}. For each simulated series, as well as for the data in sample, the Likelihood Ratio (LR) test is calculated as

$$LR_{NRS \leftrightarrow RS} = -2 \log \left(L_{NRS} - L_{RS} \right)$$

Finally, the significance of the LR test statistic is obtained by calculating how many of the LR test values on the similated series are larger than the LR statistic for the data in sample.

4 Empirical Results

This section summarizes the main empirical results of the paper. Section 4.1. discusses the estimation results for the different bivariate models for the US and European returns. Based on various specification tests, I choose the best performing model. The EU and US innovations are then estimated using the best model and a European index that excludes the country under investigation. Section 4.2. reports the results from four volatility spillover models: a constant volatility spillover model, and three models exhibiting regime switches in the spillover parameters. In section 4.3., I relate the shock spillover intensities to a large set of economic variables. Finally, in section 4.4., I test for contagion effects.

4.1 Bivariate Model for Europe and US

In order to have a good specification for the EU and US shocks, I estimate and compare four different bivariate models: (1) a constant correlation model, (2) a BEKK model, (3) a regime-switching normal model, and (4) a regime-switching GARCH model. Table 3 presents the specification tests as outlined in section 3.3.2.

The univariate specification tests of Panel A show no evidence against any of the variance specifications, and neither against the specification for the US mean equation. There is however evidence against zero autocorrelations in $\{z_{eu,t}\}$ and $\{z_{eu,t} z_{us,t}\}$ in most cases. The test statistics for the joint test are all far above their critical values. Notice however that the test statistics for both regime-switching models are slightly lower (about 52 versus about 66). The last column of Table A reports a Wald test for asymmetry in the variance specifications of models (1), (2), and (4). The results suggest that there are strong asymmetric effects in the variance-covariance matrix.

In Panel B of Table 3, I tests whether the standardized residuals of the four different models exhibit excess (cross-) skewness and kurtosis relative to the bivariate normal distribution. The results indicate that there is skewness, kurtosis, cross-skewness, and cross-kurtosis left in the standardized residuals. Here, the test statistics for the joint test are much lower for the regime-switching models than for the constant correlation and BEKK model. In particular, the regime-switching volatility models perform much better in the tests for kurtosis and cross-kurtosis, which suggests that regime-switching models do better in proxying for the fat tails in the return's distribution. Moreover, the regime classification measure (RCM) outlined in section 3.3.2.2. equals 28.87, implying that on average, the most likely regime has a probability of more than 90 percent (see Table 2). This means that the regimes are well distinguished.

An empirical likelihood ratio test strongly supports a model with regime switches. More specifically, we test the regime-switching normal against the constant correlation model following the procedure outlined in Section 3.3.2.3. The LR statistic amounts to 55.8. Only 0.4 percent of the 500 simulated LR statistics is larger than 55.8, hereby rejecting the null hypothesis of no regimes at a 1 percent level. Finally, the residuals from the four models were regressed upon a set of information variables[10]. The hypothesis that all instruments had a zero influence could not be rejected for any of the four models.

While all models seem to give relatively similar results, I take the residuals from the regime-switching normal as input for the second-step estimation, as this model produced the lowest test statistic for both the univariate and bivariate joint test for normality, as the null of one regime is rejected, and as the regime classification performance is satisfactory. The estimation results for the bivariate regime-switching normal model are given in Table 4. The results suggest that the European and US equity markets are both at the same time in high and low volatility states. The volatility in Europe and the US is respectively about 2.1 and 1.7 times higher in the high volatility regime. Notice also that on average the volatility in the US is higher than in Europe, while the correlation between both series is significantly higher in the high volatility regime (0.80 versus 0.56 in the low volatility regime). A Wald test shows this difference to be statistically significant at the 5%

[10]The following information variables were used (all lagged once): the world dividend yield in excess of the one month US T-Bill rate, the change in the US term structure, and the change in the three-month US interest rate, and the change in the US default spread. For Europe, the same variables were constructed based on German series.

level[11]. In addition, the mean returns are negative or insignificant in times of high volatility, but significantly positive in the low volatility state. This confirms that correlations between equity markets are high exactly when the diversification effects from low correlations are most needed. Finally, Figure 2 plots the filtered probability of being in the high volatility regime. Most of the time, both the EU and US market are in the low volatility regime, and switch for short periods of time to the high regime. Peaks coincide with the debt crisis in 1982, the October 1987 stock market crash, and the economic crisis at the beginning of the 1990s. Similarly, the financial crises in Asia and Russia, the LTCM debacle, and the start of a market downturn since the end of 2000 did have a strong impact on market volatility at the end of the sample.

As argued before, shock spillovers from the aggregate European market to the local countries may be overestimated, as the EU index consists partly of the local returns. Therefore, for each country, I construct a matching European index that is a market-weighted average of the returns of all countries except those of the country under investigation. The country-specific EU and US innovations are then obtained by estimating the regime-switching normal model on the US and adapted EU returns.

4.2 Univariate Volatility Spillover Model

In this paragraph, I report and discuss the estimation results of the univariate volatility spillover models with and without regime shifts in the shock spillover parameters. The (adapted) EU and US innovations obtained in the first step estimation are orthogonalized, assuming that the EU innovations are driven by a purely idiosyncratic shock and by a US shock (see appendix 1). The US return innovation as well as the orthogonalized EU innovations serve as input for the univariate volatility model presented here. Notice that this orthogonalization procedure has consequences for the expected absolute value of the spillover parameters. This can be seen as follows. Suppose I know that the US and EU volatility explain about the same proportion of the total volatility of a particular country and that shocks are on average of the same magnitude. Then the volatility spillover parameters will generally not be equal unless EU and US shocks are unrelated. If they are related, part of EU shocks will be explained by US shocks, and the pure EU shocks - this is, the full EU shocks orthogonal to the US shocks - will be small relative to the full EU shocks. Consequently, to have the same impact on the local

[11] The test statistic is 4.0497, which has a probability value of 4.42%.

volatility, the EU spillover parameter at time t will have to be higher than the one of the US. How much higher depends upon the conditional variance-covariance matrix at time t. Therefore, one should only compare the size of EU and US spillover parameters between countries, but not with each other. To compare the relative importance of EU and US news, one may look at the proportion of the local variance each of them explains separately.

4.2.1 A Volatility Spillover Model without Regime Shifts

Table 5 shows the estimation results of the univariate volatility spillover model without regime switches. All spillover parameters are significant at a 1 percent level. The EU shock spillover parameter is on average 0.64, but shows considerable cross-sectional variation (ranging from 0.35 for Austria to 0.93 for Spain). In addition, the shock spillover intensity is on average higher for EMU countries than for non-EMU countries (averages of respectively 0.67 and 0.60). Shock transmission from the US amounts on average to 0.41. This means that a 1 percent decrease in the US stock markets leads *ceteris paribus* to an average decrease of 0.41 percent in the different European equity markets. Contrary to the estimates of γ^{eu}, the US spillover intensity parameters are fairly similar across countries. Interestingly however, the average γ^{us} is higher for non-EMU countries than for EMU countries (0.46 versus 0.39), even though this result is to a large extent driven by the high spillover intensity for Sweden. Finally, it is worth noting that in none of the countries there is evidence of asymmetry in the conditional volatility specification for the idiosyncratic shocks once EU and US shocks are included.

4.2.2 A Volatility Spillover Model with Regime Shifts in the Spillover Parameters

This section discusses the estimation results for the three univariate volatility spillover models with regime shifts in the spillover parameters, and compares those with the standard constant spillover model (CSM). For each country, the best performing model is chosen on the basis of the three criteria discussed in Section 3.3.2.: (1) by comparing the values of a standard normality test on the standardized error terms, (2) by comparing their regime classification performance, and (3) by an empirical likelihood ratio test.

The performance statistics are reported in Table 6. Panel A of Table 6 reports the results from a normality test on the standardized residuals of the different models. I only report the joint test for normality, this is the hypothesis of mean zero, unit variance, no autocorrelation (up to order 4) in both the standardized and squared standardized residuals, no skewness, and no excess kurtosis[12]. One can directly see that the models with regime-switching in the spillover parameters perform much better than the single regime model. On average, the test statistic is 11.2 times lower for the models with regime-switching[13]. While the single regime model is rejected for all countries, the regime-switching models are only rejected for three countries[14]. The regime-switching models do overall slightly better on modelling the mean and variance of the local returns. The large difference in test statistic with the constant spillover case is largely due to a much lower test statistic for excess kurtosis (and to some extent also for skewness). This suggests that the regime-switching models perform much better in modelling the tails of the distribution. The distinction between the different regime-switching models is less clear-cut. While the model with joint switches in the spillover parameters (JRS) produces on average the lowest test statistics, it only performs best in three of the thirteen cases, compared to five times for the model with independent regime switches (IRS) and the fully flexible model (FULL).

In panel B of Table 6, I calculate (empirical) likelihood ratio tests to see whether the different models are significantly different from each other. Column 1 and 2 compare the constant spillover model with the models with joint and independent regime switches using an empirical likelihood ratio test. Similarly, columns 3 and 4 compare the fully flexible model with those with joint and independent regime switches. While the model with independent regime switches is nested in the full model, the specification which assumes joint switches is not. Given the highly nonlinear character of the full model however, the reported probability values are in both cases taken from a standard χ^2 distribution. As a consequence, these probabilities should be seen as an indication of significance only. In all countries, the single regime model is rejected in favor of the JRS or IRS model, confirming previous results that regime switches in shock spillover intensity are important. There is no easy test statistic available to compare the JRS and IRS model. However, one can get a feeling for the statistical difference between the two models by comparing their LR test statistic against the single regime model.

[12]The reported test statistics follow a χ^2 distribution with 12 degrees of freedom.

[13]11.2 is calculated as the ratio of the average test statistic for the constant spillover model, and the average of those of the three regime switching models

[14]As a rough indication of the relevance of regime switches, I reject regime switching in the spillover parameters if none of the three regime switching models has a probability value of more than 5 percent.

In eight of the 13 cases, the LR test statistic is substantially higher for the IRS than for the JRS model. The JRS model seems to perform best only in case of Austria, France, Denmark, and Sweden. These results suggest that for these countries, the EU and US shock spillover intensity are governed by the same underlying factors, while for the other countries, the factors may be very different. For most countries, the fully flexible model (FULL) does not perform statistically better than the best of the JRS or IRS model. The (informal) χ^2 test statistic is only statistically significant for Germany, the Netherlands, and Switzerland.

In panel C of Table 6, I analyse the regime qualification performance of the different regime-switching models. Column one till three report the regime classification measure (RCM) derived in section 3.3.2.2. and the associated probability of the most likely regime, assuming that the other states share the remaining probability mass between them. As was clear from Table 2, the RCM's are only comparable if the number of states is equal. Therefore, to compare the JRS model with the IRS and FULL model (both have 4 states), in column four and five, I calculate what the RCM would be in the two state case. To do so, I allocate the probability of the most likely regime to state 1, and the probability of the three remaining states to state 2. Finally, the last column reports the model that performs best for this metric. In nine of the thirteen cases, the JRS model distinguishes best between the different regimes: on average, it allocates 85.8 percent to the most likely regime, compared to 77.4 and 75.2 percent for the IRS and FULL model respectively. In addition, in eight cases, the most likely regime in the JRS model has a probability of more than 85 percent, compared to only three and zero times for the IRS and FULL model. The relatively worse regime classification performance for the IRS and FULL models does not come as a surprise, as these models allow for more flexibility. Overall, it is fair to say that all models distinguish relatively well between the different states, as nearly always, the most likely regime has a probability above 75 percent.

In conclusion, all tests indicate strongly in favor of regime-switching shock spillover intensities. While in most cases the different performance statistics for the regime-switching models point in the same direction, I choose the best model based upon the (empirical) likelihood ratio test statistics. The last column of panel B of Table 6 shows for each country the model with the highest LR test statistic (versus the NRS model). However, given its large number of parameters, the $FULL$ model is only choosen if it performs statistically better than the JRS and IRS model. In what follows, the regime-switching shock spillover intensities are those estimated using the best performing model.

Table 7 investigates whether the shock spillover parameters are statistically different across regimes. The Wald tests are distributed as a χ^2 distribution with one degree of freedom. Nearly all EU and US shock spillover parameters are significantly different across regimes. Only for Denmark and Switzerland, the results seem to suggest that there is no significant difference between the high and low spillover regime.

To get an understanding of the magnitude and evolution of shock spillover intensity through time and across countries, Table 8 reports average shock spillover intensities over different subperiods, while Figure 3 plots the shock spillover intensities through time for the different countries. The latter are calculated using equations 16 and 17.

Let us first inspect the spillover intensity from the EU market (left hand side of figure 3, Panel A of Table 8). In all countries, the sensitivity to EU shocks is considerably larger during the 1990s. The largest increases are found in Austria (+153%) and Denmark (+152%), the lowest in the Netherlands (-8%), and Norway (-1%). Looking at shorter subperiods, I find that the largest increases were observed in the second half of the 1980s and the first half of the 1990's. Sensitivities stay more or less the same during the 1996-1999 period, to decrease again after 1999. This result is surprising, given that during 1996-1999, Europe was going through a period of monetary integration and exchange rate stability, culminating in the introduction of a single currency in the EMU member countries. These results suggest that the economic integration (boosted by the Single European Act (1986)) as well as efforts to further liberalize European capital markets were more important in bringing markets closer together than the process towards monetary integration and the introduction of the single currency. While the sensitivity to EU shocks has increased substantially (on average +31%), the rise in US shock spillover intensity was not so pronouned (on average, +14%). Exceptions are Austria (+233%), Germany (+46%), Netherlands (-8%), Denmark (-12%), and the UK (-7%). These results show that the importance of EU shocks is rising relative to those of US shocks.

Table 9 reports the proportion of total return variance that can be attributed to EU (panel A) and US shock spillovers (panel B). Over the full sample, EU shocks explain about 15 percent of local variance, while US shocks account for about 20 percent. While the US - as a proxy for the world market - is still the dominant force, the proportion of variance attributed to EU shocks has increased substantially more: from about 10% during the 1980s to about 20% during the nineties (increase of 97%) for

Europe; for the US from about 17% to 25% (increase of 51% only)[15]. The EU variance proportion is on average higher for EMU than for non-EMU countries (17% versus 13%). However, while for EMU countries the EU variance proportion increased with 81%, it increased with 146% for non-EMU countries. Over the 1990s, the largest EU variance ratios were observed in France and Spain (27%); the lowest in Austria (12%), Norway (15%), the UK (15%), and Switzerland (16%). While the US variance proportions are on average similar across countries (20%), the evolution has been different: during the 1980s, the fraction of local shocks explained by US shocks was higher for non-EMU countries (18% versus 15%). However, the picture has flipped during the 1990s, in which 25% of EMU equity market shocks are explained by US shocks compared to 23% only for non-EMU countries. The Dutch index has a very high US variance ratio of 44%, as it is dominated by companies who have high proportions of their cash flows outside Europe. Also the UK (41%), France (31%), Germany (27%), Sweden (28%), and Switzerland (28%) have high US variance ratios, while Austria (10%) and Denmark (14%) are relatively isolated from the US market.

4.3 Economic Determinants of Shock Spillover Intensity

The advantage of regime-switching models is at the same time also their weakness: they let the data decide in what state the economy is at a specific time. As argued in the introduction, little is known about what factors determine shock spillover intensity and in what fashion. In this section, I relate the latent state variable $S_{i,t}^{eu}$ to a large set of economic and financial variables that may influence shock spillover intensity. The dependent variable $S_{i,t}^{eu}$ takes on the value of one when the ex-post probability of being in the high spillover state is larger than 50 percent, and zero otherwise[16]. I focus on the the EU shock spillover intensity as to investigate the effect of the intense efforts aimed at opening European capital markets, and at strengthening the economic and monetary integration in the EU. Many of the explanatory variables I use are standard in the literature, and have been used - even though not all at the same time - by e.g. Bekaert and Har-

[15]The proportionally larger increase in EU (US) variance ratios compared to EU (US) shock spillover intensity is due to an increase in the average ratio of EU (US) market volatility to total local volatility. Notice that this in itself is an indication of larger correlations between countries.

[16]Alternatively, I estimated the relationship between the different information variables and the probability weighted EU shock spillover intensity using GMM (with correction for autocorrelation in de dependent variable). Results are qualitatively very similar to the logit analysis presented here. Results are available upon request.

vey (1995, 1997), Chen and Zhang (1997), Beck et al. (2000), Ng (2000), Bekaert et al (2002a), and Fratzscher (2001). The instruments used can be grouped under the following headers.

Market Development and Integration[17] More developed financial markets are likely to share information more intensively, as they are, on average, more liquid, more diversified, and better integrated with world financial markets than smaller markets. In addition, Demirgüç-Kunt and Levine (1996) found that countries with larger stock markets are usually better institutionally developed, with strong information disclosure laws, international accounting standards, and unrestricted capital flows, hence with lower asymmetric information costs. In addition, a gradual shift from segmentation to financial integration implies a shift from a local to a common global discount factor, and hence a more homogeneous valuation of equity. An often used proxy for stock market development is the ratio of equity market capitalisation to GDP ($MCAP/GDP$). Bekaert and Harvey (1995) and Ng (2000) among others found that countries with a higher $MCAP/GDP$ are on average better integrated with world capital markets. Therefore, we expect a positive influence from market development on shock spillover intensities.

Economic Integration through Trade To the extent that economic and financial integration go hand in hand, variables that proxy for trade integration may be useful in explaining the time-varying nature of shock spillover intensity. The more economies are linked, the more they will be exposed to common shocks, and the more companies' cash flows will be correlated. This argument is particularly valid for European Union countries, as these countries went through a period of significant trade integration. Much of this progress was made in the aftermath of the Single European Act (1986). Chen and Zhang (1997) found that countries with heavier bilateral trade with a region also tend to have higher return correlations with that region. Similarly, Bekaert and Harvey (1995) found that countries with open economies are generally better integrated with world capital markets. An often used proxy for trade integration is the size of the trade sector. More specifically, I test whether the ratio of import plus export of country i

[17] Lack of high quality data prevent me from looking at more detailed characteristics of market development. Turnover data (relative to market capitalization or GDP) for instance may provide information about the efficiency or liquidity - and hence attraction - of the equity market. For most markets however, turnover data is only available from the beginning of the 1990s.

with the EU to GDP is significantly positivily correlated with the EU shock spillover intensity.

Business Cycle There is considerable evidence that correlations between equity markets are higher during recessions than during growth periods (see e.g. Erb et al. (1994)) . In the model proposed here, this may be due to an increase in the volatility ratio ($\sigma_{eu,t}/\sqrt{h_{i,t}}$) or to an increase in the EU shock spillover intensity. As I focus on the latter, I focus on the relationship between γ_i^{eu} and the evolution of the business cycle. Previous authors have investigated correlations conditional on the ex-post state of the business cycle (e.g. the NBER official "recession dates"). However, given the forward looking character of equity prices, it seems more natural to use a forward looking business cycle indicator. Here, I will relate the OECD leading indicator for the aggregate EU market- more specifically, the deviation from its (quadratic) trend[18] - to the EU shock spillover intensity. By separating conditional correlations into a spillover and volatility component, I am able to test whether the increase in correlations during recessions is due to an increase in shock spillover intensity or not.

Shock spillover intensity may not only depend upon the state of the EU business cycle, but also on how the local economy is expected to perform relative to the aggregate business cycle. Erb et al. (1994)) for instance also found that correlations are generally lower when business cycles are out of phase. This may be especially relevant for countries whose business cycle moves asymmetrically relative to the EU. Similarly, a reduction of cross-country business cycle asymmetries will increase return correlations through a convergence of cash flow expectations. To test whether countries and/or periods with large business cycle asymmetries are characterized by lower return correlations, I include a dummy that records whether or not the economy of country i is out of phase with the European economy[19].

Monetary Integration and Exchange Rate Stability Monetary integration, started with the Maastricht Treaty (1992) did not only make infla-

[18]Results are robust to the use of a linear trend, as well as of Hodrick-Prescott filtered series.

[19]This dummy is calculated as follows. First, a (quadratic) trend is fitted for the OECD leading indicator of each country, as well as for the EU. Second, deviations from this trend are generated. Positive deviations indicate a boom; negative deviations a recession. Third, for each country, an "out-of-phase" dummy is created. This dummy has a value of one when the deviation of the OECD leading indicator from its trend has a different sign for the EU and the country under investigation, and zero otherwise.

tion and nominal interest rates converge across countries, it also created an environment of exchange rate stability[20]. The convergence in real interest rates as well as the reduction (elimination) of currency risk premia resulted in a convergence of cross-country discount rates, and hence a more homogeneous valuation of equity. Notice moreover that the introduction of the euro eliminated an important impediment to cross-border investment, more specifically the EU matching rule, which prevented insurance companies, pension funds, and other financial institutions with liabilities denominated in Euro from fully exploiting diversification benefits within the euro area. The lower currency hedging costs and the elimination of this barrier should induce investors to increase their holdings of pan-European assets, leading to an increase in information sharing across European capital markets. As a measure of monetary policy convergence, I use the difference between local inflation and the EU15 inflation average[21]. The effect of exchange rate stability is determined by fitting a GARCH(1,1) model on the exchange rate returns of country i vis-a-vis the ECU, and using the estimated conditional variance as an explanatory variable for γ_i^{eu}.

A potential problem is that some of the explanatory variables are highly correlated. This is especially relevant for the trade variable and the market development variable. Therefore, I use the trade variable, and the part of market capitalization over GDP that is orthogonal to the trade variable. A univariate logit regression is used to relate the binary dependent variable $S_{i,t}^{eu}$ to the explanatory variables. Robust standard errors are computed using quasi-maximum likelihood. Results are reported in Table 10. Many of the explanatory variables enter significantly. The trade integration variable is positive and significant in all countries except for Austria, Ireland, and Norway. This suggests that trade has been an important catalyst for increased information sharing between equity markets. Inflation enters negatively and significantly for all countries, except for Austria, Germany, and Switzerland, indicating that equity markets share more information in a low-inflation environment. The deviating result for Germany may be explained by the surge in inflation after the German reunification, a period that coincided with a

[20] Exchange rate stability was the aim of the Exchange Rate Mechanism (ERM) started in 1979. Several devaluations in the summer of 1992 and 1993 resulted in a widening of the fluctuation marging from 2.25 percent to 15 percent around parity. However, exchange rate stability was a prerequisite for countries willing to participate to the euro. Necessarily, all countries now part of the euro zone had stable exchange rates vis-a-vis each other's currency in the second half of the 1990s. However, exchange rate volatility was also low for the other EU countries.

[21] Long-term nominal or real interest rates could not be included because these series were not available over the full sample.

rapid increase in spillover intensity. The similar result for Austria is likely to be explained by the high degree of correlation between the German and Austrian equity market. Finally, Switzerland had fairly low inflation levels all over the 1990s. While Austria and Belgium appear to be negatively affected by sudden increases in currency volatility, for most other countries, the spillover intensities are positively or insignificantly related to currency volatility. This somehow confirms the empirical regularity that correlations between markets increase in times of turmoil, more specifically during a currency crisis[22]. This was especially apparent during the exchange rate turmoil in the summer of 1992 and 1993. In addition, these results show that this effect persists even if one corrects for the changes in the ratio of EU market volatility to local volatility. The market development indicator - market capitalization over GDP - is positive and signficant in 7 cases and insignificant (at a 5 percent level) for the other countries. The "high world volatility dummy" is nearly always positive, but only significant for The Netherlands, Spain, and Sweden. In most countries, shock spillover intensity is significantly related to the state of the European business cycle. In Germany, Ireland, Denmark, and Switzerland, the shock spillover intensity increases in times of recessions. This result is consistent with the results of Erb et al. (1994). However, in Austria, Belgium, Italy, The Netherlands, Spain, and Sweden, the opposite seems true. The same mixed results prevail when looking at the dummy measuring whether the local business cycle is out of phase with the European cycle. While for some countries it is the case that their spillover intensity decreases when they are out of phase with the European business cycle, the opposite seems true for other countries.

4.3.1 A simple test for Contagion

The model used here allows for a simple test of contagion. This test is very similar to the test proposed by Bekaert et al. (2002c). They define contagion as "correlation over and above what one would expect from economic fundamentals". The authors estimate a two-factor model that is similar to the one used here, in a sense that the two factors are the regional market - in my case, the European market - and the US market. Correlations will change when the volatility of the factors changes; by how much is determined by the factor sensitivities. While both models allow for time-variation in the factor sensitivities, the way to do so is quite different. In their model, the evolu-

[22]Theoretically, one would expect that the gradual decrease of currency risk in Europe and the consequent decrease in currency risk premium would increase correlations between markets. This is however not what I find.

tion of the sensitivities is governed by a bilateral trade variable, compared to a latent regime variable in my model. I believe that the model used here has some advantages. First, as shown in the previous section, the variation of the sensitivities through time is influenced by more factors than trade alone. Second, as argued by Ang and Bekaert (2002b), regime-switching models may do better in capturing asymmetric correlations.

The contagion test of Bekaert et al. (2002c) is based on the argument that in the case of no contagion, there should not be any correlation left between the error terms. They investigate this hypothesis during various crisis periods by estimating the following regression:

$$\hat{e}_{i,t} = b_1 + (b_2 + b_3 D_{i,t})\hat{e}_{m,t} + u_{i,t}$$

over different crisis periods, where $\hat{e}_{i,t}$ is the local market's residual, $\hat{e}_{m,t}$ the residual from a benchmark index (here, EU and US index), and $D_{i,t}$ a dummy variable that takes the value of one if it coincides with the crisis being looked at, and zero otherwise. Clearly, much of the power of this test will depend upon the ability of the factor specification to model conditional correlations. To the extend that the specification used here is a better model for conditional correlations, also the test for contagion should be more powerful. I estimate the following specification by GMM[23]

$$\hat{e}_{i,t} = b_1 + (b_2 + b_3 D_t)\hat{e}_{eu,t} + (b_4 + b_5 D_t)\hat{e}_{us,t} + u_{i,t}$$

where $\hat{e}_{eu,t}$ and $\hat{e}_{us,t}$ are the orthogonalized residuals from the bivariate model for EU and US returns. Contrary to Bekaert et al. (2000), I let the data decide when world equity markets are going through a crisis period. Therefore, D_t takes on a one when the EU and US are jointly in a high volatility state, and zero otherwise. So this is a test of whether there has been contagion over the full sample and during the crisis moments occuring during the sample (I do not look a crises individually). Overall contagion from the aggregate European market (excluding the country being looked at) would be there if b_2 and b_3 are jointly different from zero, while b_3 measures the extra contagion during crisis periods. Similarly, we cannot reject contagion from the US markets when b_4 and b_5 are jointly different from zero; here b_5 measures the extra contagion during crisis periods.

Results are contained in Table 11. There is some evidence of contagion from the EU market to the German equity market (at a 5 percent level). However, for all other countries, the hypothesis of no contagion cannot be

[23] with heteroskedasticity and autocorrelation consistent standard errors (Barlett kernel, Newey-West bandwith selection (6))

rejected. The evidence is stronger for contagion from the US market. For France, Germany, Ireland, Italy, Spain, Sweden, and Switzerland, the parameters b_4 and b_5 are jointly significant. Looking more into detail, one can see that this is mainly due to the high significance of b_5, which measures whether correlation between local and US residuals is higher during crisis periods.

5 Conclusion

This paper investigates whether the efforts for more economic, monetary, and financial integration in Europe have fundamentally altered the intensity of shock spillovers from the US and aggregate European equity markets to 13 European stock markets. The innovation of the paper is that the EU and US shock spillover intensity is allowed to switch between a high and low state according to a latent regime variable. Three regime-switching shock spillover models are derived that differ in the way switches in the EU and US spillover intensity interact. I find that regime switches in the spillover intensities are both statistically and economically important. For nearly all countries, the probability of a high EU and US shock spillover intensity has increased significantly over the 1980s and 1990s, even though the increase is more pronounced for the sensitivity to EU shocks. It may be surprising to some that the increase in EU shock spillover intensity is mainly situated in the second part of the 1980s en the first part of the 1990s, an not during the period directly before and after the introduction of the single currency. In fact, in many countries, the sensitivity to EU shocks dropped considerably after 1999. Over the full sample, EU shocks explain about 15 percent of local variance, compared to 20 percent for US shocks. While the US - as a proxy for the world market - continues to be the dominating influence in European equity markets, the importance of the regional European market is rising considerably.

Next, I look for the factors that have contributed to this increased information sharing. I consider instruments related to equity market development, economic integration, monetary integration, exchange rate stability, and to the state of the business cycle. Results suggest that countries with an open economy, low inflation, and well developed financial markets share more information with the regional European market. There is some evidence that shock spillover intensity is related to the business cycle.

Finally, a test for contagion is derived similar to the one proposed by Bekaert et al. (2002c). They define contagion as "correlation over and above what one would expect from economic fundamentals". If the model used here is correct (what the specification tests tend to suggest), changes in conditional correlaton will be entirely driven by changes in the conditional EU (US) and local market volatility, and switches in the EU (US) shock spillover intensity. The hypothesis of no contagion can be tested by investigating the correlation between the model's residuals. I discover a statistically significant correlation between local residuals and those of the US market during crisis periods. There is however no evidence of contagion from the regional European market to the local equity markets.

The methodology developed in this paper may prove useful in analyzing many other interesting issues. First, it may be useful to investigate the capacity of this type of models in capturing asymmetric correlations between markets. A model that features both regime switches in spillover intensity and the level of volatility seems promising in this context. Second, this methodology is especially appropriate in analyzing the interaction between different asset markets, in casu the foreign exchange, money, bond, and equity markets. Third, instead of applying this methodology to shock spillover models, it may be interesting to investigate whether also prices of risk are subject to regime switches, and if so, what economic factors force these to switch.

References

[1] Ang, Andrew, and Geert Bekaert (2002a), Regime Switches in Interest Rates, Journal of Business and Economic Statistics, 20, 2, 163-182.

[2] Ang, Andrew, and Geert Bekaert (2002b), International Asset Allocation with Regime Shifts, Review of Financial Studies 15, no. 4, pp. 1137-1187.

[3] Ang, Andrew, and Joe Chen (2001), Asymmetric Correlations in Equity Portfolios, Journal of Financial Economics 63, no. 3, pp. 443-494.

[4] Artis, M, H.-M Krolzig, and J. Toro (1998), The European Business Cycle, CEPR discussion paper 2242.

[5] Baba, Y., Robert Engle, D. Kraft, and K.F. Kroner (1989), Multivariate Simulatneous Generalized ARCH, Discussion Paper 89-57, University of California, San Diego.

[6] Beck, Thorsten, Asli Demirgüç-Kunt, and Ross Levine (2000), A New Database on Financial Development and Structure, World Bank Economic Review

[7] Bekaert, Geert, Campbell R. Harvey, and Christian Lundblad (2002a), Does Financial Liberalization Spur Growth?, Working Paper.

[8] Bekaert, Geert, Campbell R. Harvey, and Robin L. Lumsdaine (2002b), Dating the Integration of World Equity Markets, Journal of Finance, forthcoming. Also published as NBER Working Paper No.6724, September 1998

[9] Bekaert, Geert, Campbell R. Harvey, and Angela Ng (2002c), Market Integration and Contagion, Journal of Business, Forthcoming.

[10] Bekaert, Geert, and Guojun Wu (2000), Asymmetric Volatility and Risk in Equity Markets,Review of Financial Studies 13, pp. 1-42.

[11] Bekaert, Geert, and Campbell R. Harvey (1997), Emerging Equity Market Volatility, Journal of Financial Economics 43, pp. 29-77.

[12] Bekaert, Geert, and Campbell R. Harvey (1995), Time-Varying World Market Integration, Journal of Finance 50, pp. 403-444.

[13] Bollerslev, Tim (1990), Modelling the coherence in short-run nominal exchange rates: a multivariate generalized ARCH approach, Reveiw of Economics and Statistics, Vol. 772, pp. 498-505.

[14] Bollerslev, Tim, and Jeffrey M.Woolridge (1992), Quasi-maximum likelihood estimation and inference in dynamic models with time-varying covariances Econometric Reviews 11, pp. 143-172.

[15] Booth, G. Geoffrey, Teppo Martikainen, and Yiuman Tse (1997), Price and Volatility Spillovers in Scandinavian Stock Markets, Journal of Banking and Finance 21: pp. 811-823.

[16] Booth, G. G. and G. Koutmos.(1995), Asymmetric volatility transmission in international stock markets, Journal of International Money and Finance, 14, pp. 747-62.

[17] Cai, J. (1994), A Markov Model of Switching-Regime ARCH, Journal of Business and Economic Statistics, 12, 309-316.

[18] Cappiello, Lorenzo (2000), A Conditional multi-asset intertemporal CAPM with switching prices of risk, The Graduate Institute of International Studies, University of Geneva.

[19] Chen, Nai-fu, and Feng Zhang (1997), Correlations, Trades and Stock Returns of the Pacific Rim Markets, Pacific Basis Finance Journal 5, pp.559-577.

[20] Das, S.R. and R. Uppal (1999), The effect of systemic risk on international portfolio choice, Working Paper, Harvard University.

[21] Demirgüç-Kunt, Asli, and Ross Levine (1996), Stock Markets, Corporate Finance, and Economic Growth: An Overview, The World Bank Economic Review 10, no. 2, pp.223-239.

[22] Diebold, F.X. (1986), Comment on Modelling the persistence of Conditional Variance, Econometric Reviews, Vol. 5, pp.51-56.

[23] Engle, Robert F. and K.F. Kroner (1995), Multivariate Simulaneous Generalized ARCH, Econometric Theory, 11, pp. 122-150.

[24] Erb, Claude B., Campbell Harvey, and Tadas Viskanta (1994), Forecasting International Equity Correlations, Financial Analysts Journal, November/december, pp. 32-45.

[25] Fleming, Jeff, Chris Kirby, and Barbara Ostdiek (1998), Information and volatility linkages in the stock, bond, and money markets, Journal of Financial Economics, Vol. 49, pp. 111-137.

[26] Forbes, Kristin, and Roberto Rigobon (2001), No Contagion, Only Interdependence: Measuring Stock Market Co-movements, Journal of Finance, forthcoming.

[27] Fratzscher, Marcel (2001), Financial Market Integration in Europe: on the effects of EMU on Stock Markets, International Journal of Finance and Economics, forthcoming. Also published as ECB Working Paper 48.

[28] Glosten, Lawrence R., Ravi Jagannathan, and David Runkle (1993), On the Relation between the Expected Value and the Volatility of the Nominal Excess Returns on Stocks, Journal of Finance 48, no. 5, pp. 1770-1801.

[29] Gray, S. (1996), Modeling the Conditional Distribution of Interest Rates as a Regime Switching Process, Journal of Financial Economics, 42, 27-62

[30] Hamao, Y., Masulis, R.W., and V. Ng (1990), Correlations in Price Changes and Volatility across International Stock Exchanges, The Review of Financial Studies, Volume 3. number 2, pp. 281-307.

[31] Hamilton, J. D. and R. Susmel (1994), ARCH and Changes in Regime, Journal of Econometrics, 64, 307-333.

[32] Hamilton, James D. (1988), Rational Expectations Econometric Analysis of Changes in Regime: an Investigation of the Term Structure of Interest Rates, Journal of Economic Dynamics and Control, Vol. 12, pp.385-423.

[33] Hamilton, James D. (1989), A New Approach to the Economic Analysis of Nonstationary Time Series and the Business Cycle, Econometrica 57, pp. 357-384.

[34] Hamilton, James D. (1990), Analysis of Time Series Subject ot Changes in Regime, Journal of Econometrics, Vol. 4, pp.357-384.

[35] Hamilton, James D. and Gang Lin (1996), Stock Market Volatility and the Business Cycle, Journal of Applied Econometrics 11, no. 5, pp. 573-593.

[36] Hansen, Lars-Peter (1982), Large Sample Properties of Generalized Method of Moments Estimators, Econometrica 50, no. 4, pp.1029-1054.

[37] Hansen, B.E. (1992), The likelihood ratio test under nonstandard conditions: testing the Markov Switching model of GDP, Journal of Applied Econometrics, Vol. 7, pp. 61-82.

[38] Hardouvelis, Gikas A., Dimitrios Malliaropulos, and Richard Priestley (2002), EMU and European Stock Market Integration, Working Paper. Also published as CEPR Discussion Paper No. 2124 (1999).

[39] Kanas, Angelos (1998), Volatility spillovers across equity markets : European Evidence, Applied Financial Economics, Vol 8, 1998, pp. 245-256.

[40] Karolyi, G.A., and Rene M. Stulz (1996), Why do markets move together? An investigation of US-Japan stock return comovements using ADRs, Journal of Finance, Vol. 51, pp. 951-986.

[41] Karolyi, G. (1995), A multivariate GARCH model of international transmission of stock returns and volatility: the case of the United States and Canada, Journal of Business and Economic Statistics, Volume 13. pp. 11-25.

[42] King, M.A., E. Sentana, and S. Wadhwani (1994), Volatility and links between national stock markets, Econometrica, Vol. 62, pp. 901-933.

[43] King, M.A. and S. Wadhwani (1990), Transmission of volatility between stock markets, Review of Financial Studies, Volume 3. No. 1, pp.5-33.

[44] Koch, P.D and T.W. Koch (1991), Evolution in dynamic linkages across daily national stock indices, Journal of International Money and Finance, Volume 10, pp.231-251.

[45] Koutmos, G. (1996), Modeling the dynamic interdependence of major European stock markets, Journal of Business, Finance, and Accounting, Volume 23. pp. 975-988.

[46] Kroner, K.F. and V.K. Ng (1998), Modeling asymmetric comovement of asset returns, Review of Financial Studies, Volume 11, pp. 817-844.

[47] Lamoreux, C.G. and W.D. Lastrapes (1990), Persistence in Variance, Structural Change and the GARCH model, Journal of Business and Economic Statistics, Vol. 5, pp.121-129.

[48] Licht, A.N. (1998), Stock Market Integration in Europe, CAER II discussion Paper, no. 15, Harvard Institute for International Development, Cambridge, M.A.

[49] Lin, W.L., Robert F. Engle, and T. Ito (1994), Do bulls and bears move across borders? International transmission of stock returns and volatility, The Review of Financial Studies, Vol. 7, pp. 507-538.

[50] Longin, F. and B, Solnik (2001), Extreme Correlations of International Equity Markets, Journal of Finance, Vol. 56, no.2, pp.649-674.

[51] Longin, F. and B, Solnik (1995), Is the correlation in international equity returns constant: 1960-1990?, Journal of International Money and Finance, Vol. 14, pp. 3-26.

[52] Ng, Angela (2000), Volatility Spillover effects from Japan and the US to the Pacific-Basin, Journal of International Money and Finance, Vol. 19, pp.207-233.

[53] ∞

[54] Ramchmand, L. and R. Susmel (1998), Volatility and cross correlation across major stock markets, Journal of Empirical Finance, Vol.3. pp.119-131.

[55] Richardson, M., and T. Smith (1993), A Test for Multivariate Normaility in Stock Returns, Journal of Business, Vol. 66, pp. 267-321.

[56] Susmel, Rauli (2000), Switching Volatility in International Markets, International Journal of Finance and Economics, Vol. 5.

Appendix 1: Derivation of orthogonalized innovations

The univariate shock spillover model allows us to quantify the relative importance of EU and US shocks on the different European equity markets. However, as is likely that a common news component drives part of the EU and US returns, I make the local European returns conditional on the innovations from the US and the orthogonalized innovations from the European aggregate market. The innovations from Europe and the US are orthogonalized by assuming that the EU return is driven by a purely idiosyncratic shock and by the US return shock. Similar to Ng (2000), the orthogonalized European and US innovations are denoted respectively by $e_{eu,t}$ and $e_{us,t}$,and are given by:

$$\varepsilon_t = \begin{bmatrix} \varepsilon_{eu,t} \\ \varepsilon_{us,t} \end{bmatrix} = \begin{bmatrix} 1 & k_{t-1} \\ 0 & 1 \end{bmatrix} \begin{bmatrix} e_{eu,t} \\ e_{us,t} \end{bmatrix} = \mathbf{K}_{t-1} e_t$$

$$\varepsilon_t | \Omega_{t-1} \sim N(0, \mathbf{H}_t)$$

$$e_t | \Omega_{t-1} \sim N(0, \Sigma_t)$$

$$\Sigma_t = \begin{bmatrix} \sigma^2_{eu,t} & 0 \\ 0 & \sigma^2_{us,t} \end{bmatrix}$$

where k_{t-1}, σ^2_{eu}, and σ^2_{us} are calculated such that $\mathbf{H}_t = \mathbf{K}_{t-1} \mathbf{\Sigma}_t \mathbf{K}'_{t-1}$. Under these assumptions, it is straightforward to show that k_{t-1} is determined by the ratio of the covariance between EU and US innovations and the variance of the latter:

$$k_{t-1} = \frac{Cov_{t-1}\left(\varepsilon_{eu,t}, \varepsilon_{us,t}\right)}{Var_{t-1}\left(\varepsilon_{us,t}\right)} = \frac{H_{eu,us,t}}{H_{us,t}}$$

The orthogonalized innovations can then be calculated as:

$$e_t = \begin{bmatrix} e_{eu,t} \\ e_{us,t} \end{bmatrix} = \begin{bmatrix} 1 & -k_{t-1} \\ 0 & 1 \end{bmatrix} \begin{bmatrix} \varepsilon_{eu,t} \\ \varepsilon_{us,t} \end{bmatrix} = \mathbf{K}^{-1}_{t-1} \varepsilon_t$$

while the variance-covariance matrix is given by:

$$\Sigma_t = \mathbf{K}^{-1}_{t-1} \mathbf{H}_t \mathbf{K}'^{-1}_{t-1}$$

Appendix 2: Derivation of transition matrix in case of independent states.

By way of example, we derive the third column of the transition matrix given in equation (14), supposing that the states variables $S_{i,t}^{eu}$ and $S_{i,t}^{us}$ are independent.

$$\begin{aligned}
P(S_{i,t}=3|S_{i,t-1}=1) &= P(S_{i,t}^{eu}=1 \text{ and } S_{i,t}^{us}=2|S_{i,t-1}^{eu}=1 \text{ and } S_{i,t-1}^{us}=1) \\
&= P(S_{i,t}^{eu}=1|S_{i,t-1}^{eu}=1)P(S_{i,t}^{us}=2|S_{i,t-1}^{us}=1) \\
&= P^{eu}(1-P^{us})
\end{aligned}$$

$$\begin{aligned}
P(S_{i,t}=3|S_{i,t-1}=2) &= P(S_{i,t}^{eu}=1 \text{ and } S_{i,t}^{us}=2|S_{i,t-1}^{eu}=2 \text{ and } S_{i,t-1}^{us}=1) \\
&= P(S_{i,t}^{eu}=1|S_{i,t-1}^{eu}=2)P(S_{i,t}^{us}=2|S_{i,t-1}^{us}=1) \\
&= (1-Q^{eu})(1-P^{us})
\end{aligned}$$

$$\begin{aligned}
P(S_{i,t}=3|S_{i,t-1}=3) &= P(S_{i,t}^{eu}=1 \text{ and } S_{i,t}^{us}=2|S_{i,t-1}^{eu}=1 \text{ and } S_{i,t-1}^{us}=2) \\
&= P(S_{i,t}^{eu}=1|S_{i,t-1}^{eu}=1)P(S_{i,t}^{eu}=2|S_{i,t-1}^{eu}=2) \\
&= P^{eu}Q^{us}
\end{aligned}$$

$$\begin{aligned}
P(S_{i,t}=3|S_{i,t-1}=4) &= P(S_{i,t}^{eu}=1 \text{ and } S_{i,t}^{us}=2|S_{i,t-1}^{eu}=2 \text{ and } S_{i,t-1}^{us}=2) \\
&= P(S_{i,t}^{eu}=1|S_{i,t-1}^{eu}=2)P(S_{i,t}^{eu}=2|S_{i,t-1}^{eu}=2) \\
&= (1-Q^{eu})Q^{us}
\end{aligned}$$

Appendix 3: The likelihood function for the spillover model in case of regime-switching

The aim of this appendix is to show what assumptions are needed in order to make the two-step estimation procedure used in this paper fully internally consistent.We consider the most general case, i.e. when there are regime-switches both in the first and second step. Define $r_t = [r_{eu,t}, r_{us,t}, r_{1,t}, ..., r_{L,t}]'$ and $r_{l,t} = [r_{1,t}, r_{2,t}, ..., r_{L,t}]'$. We summarize all information variables used in the estimation of the spillover models in a matrix $\mathbf{Z}_t = [\mathbf{X}'_{eu,t}, \mathbf{X}'_{us,t}, \mathbf{X}'_{1,t}, ..., \mathbf{X}'_{L,t}]$. The information set Ω_{t-1} consists of all information available up to time $t-1$, which is represented by $\{\mathbf{q}_{t-1}, \mathbf{q}_{t-2}, ..., \mathbf{q}_1, \mathbf{q}_0\}$, where $\mathbf{q}_t = [r_t, \mathbf{Z}_t]$. The whole dataset that we will use in the estimation procedure can then be summarized by $\tilde{\mathbf{q}}_T = [\mathbf{q}'_T, \mathbf{q}'_{T-1}, ..., \mathbf{q}'_1, \mathbf{q}'_0]'$. We denote the parameters to be estimated by ϑ. The aim is to maximize a density function $f(\tilde{\mathbf{q}}_t, \vartheta)$, which can be split up in the following way:

$$\begin{aligned} f(\tilde{\mathbf{q}}_T, \vartheta) &= \prod_{t=1}^{T} f(\mathbf{q}_t | \Omega_{t-1}; \vartheta) \\ &= \prod_{t=1}^{T} f(\mathbf{Z}_t | \mathbf{r}_t, \Omega_{t-1}; \vartheta) f(r_t | \Omega_{t-1}; \vartheta) \end{aligned}$$

Ideally, we would estimate $f(\tilde{\mathbf{q}}_t, \vartheta)$ using full-maximum likelihood, which is however infeasible given the large number of parameters. To reduce the complexity, we split up the parameter vector ϑ into $[\theta'_z, \theta']'$ in such a way that

$$f(\tilde{\mathbf{q}}_T, \vartheta) = \prod_{t=1}^{T} f(\mathbf{Z}_t | \mathbf{r}_t, \Omega_{t-1}; \theta_z) f(r_t | \Omega_{t-1}; \theta)$$

We will ignore the information in $f(\mathbf{Z}_t | \mathbf{r}_t, \Omega_{t-1}; \theta_z)$, which means that our estimation yields consistent but possibly inefficient estimates, relative to full information maximum likelihood. In the paper, we limited the models to two states only. We then further decompose $f(r_t, S_t|\Omega_{t-1};\theta)$:

$$\begin{aligned} f(\tilde{\mathbf{q}}_T, \vartheta) &= \prod_{t=1}^{T} f(r_{eu,}r_{us,}r_{l,t}|\Omega_{t-1};\theta) \\ &= \prod_{t=1}^{T} \left\{ \sum_{i^{eu}}^{2} \sum_{i^{us}}^{2} \sum_{i^{1}}^{2} ... \sum_{i^{L}}^{2} f(r_{eu,}r_{us}, r_1, ..., r_L, S_t^{eu} = i^{eu}, S_t^{us} = i^{us}, S_t^1 = i^1, ..., S_t^L = i^L|\Omega_{t-1};\theta) \right\} \\ &= \prod_{t=1}^{T} \left\{ \sum_{i^{eu}}^{2} \sum_{i^{us}}^{2} \sum_{i^{1}}^{2} ... \sum_{i^{L}}^{2} f(r_{eu,}r_{us}, r_1, ..., r_L|S_t^{eu} = i^{eu}, S_t^{us} = i^{us}, S_t^1 = i^1, ..., S_t^L = i^L, \Omega_{t-1};\theta) \right. \\ &\qquad \left. f(S_t^{eu} = i^{eu}, S_t^{us} = i^{us}, S_t^1 = i^1, ..., S_t^L = i^L|\Omega_{t-1};\theta) \right\} \\ &= \prod_{t=1}^{T} \left\{ \sum_{i^{eu}}^{2} \sum_{i^{us}}^{2} \sum_{i^{1}}^{2} ... \sum_{i^{L}}^{2} f(r_1, ..., r_L|r_{eu,}r_{us}, S_t^{eu} = i^{eu}, S_t^{us} = i^{us}, S_t^1 = i^1, ..., S_t^L = i^L, \Omega_{t-1};\theta) \right. \\ &\qquad f(r_{eu,}r_{us}|S_t^{eu} = i^{eu}, S_t^{us} = i^{us}, S_t^1 = i^1, ..., S_t^L = i^L, \Omega_{t-1};\theta\} \\ &\qquad \left. f(S_t^{eu} = i^{eu}, S_t^{us} = i^{us}, S_t^1 = i^1, ..., S_t^L = i^L|\Omega_{t-1};\theta) \right\} \end{aligned}$$

The structure of our model allows to further divide the parameter vector in two seperate parts: $\theta = [\theta_{eu}^{us}, \theta_l]$, where θ_{eu}^{us} contains the parameters determining $f(r_{eu,}r_{us}|S_t, \Omega_{t-1};\theta)$, and θ_l the remaining ones. If we further assume that the states of the individual countries - summarized by $S_{lt}-$ are independent of the states of the European and US market - represented by $S^{us}_{eu,t}$, we get the following:

$$\begin{aligned} f(\tilde{\mathbf{q}}_t, \vartheta) &= \prod_{t=1}^{T} \left\{ \sum_{i^1}^{2} ... \sum_{i^L}^{2} f(r_1, ..., r_L | r_{eu,} r_{us}, S_t^1 = i^1, ..., S_t^L = i^L, \Omega_{t-1}, \theta) f(S_t^1 = i^1, ..., S_t^L = i^L, \Omega_{t-1}; \theta) \right. \\ &\quad \left. \sum_{i^{eu}}^{2} \sum_{i^{us}}^{2} f(r_{eu,} r_{us} | S_t^{eu} = i^{eu}, S_t^{us} = i^{us}, \Omega_{t-1}, \theta)) f(S_t^{eu} = i^{eu}, S_t^{us} = i^{us}, \Omega_{t-1}; \theta) \right\} \end{aligned}$$

We have now split the full likelihood in two sub-likelihoods, to make a two-step estimation possible. By doing so, we sacrifice some efficiency. It is however the only way to make the estimation procedure feasible.

Now define $\theta_l = [\theta_1, \theta_2, ..., \theta_L]$, $\varepsilon_{l,t} = [\varepsilon_{1,t}, \varepsilon_{2,t}, ..., \varepsilon_{L,t}]$, and $e_{l,t} = [e_{1,t}, e_{2,t}, ..., e_{L,t}]$. We can then do the following simplifications to the model:

$$\begin{aligned} f(r_1, ..., r_L | r_{eu,} r_{us}, S_t^1 = i^1, ..., S_t^L = i^L, \Omega_{t-1}, \theta) = &\quad f(\varepsilon_{1,t}, ..., \varepsilon_{L,t} | \hat{\varepsilon}_{eu,} \hat{\varepsilon}_{us}, S_t^1 = i^1, ..., S_t^L = i^L, \Omega_{t-1}, \theta) \\ = &\quad f(e_{1,t}, ..., e_{L,t} | \hat{\varepsilon}_{eu,} \hat{\varepsilon}_{us}, S_t^1 = i^1, ..., S_t^L = i^L, \Omega_{t-1}, \theta) \end{aligned}$$

The first step stems from the definition of the information set, while the second step follows (1) from the definition of the individual elements of $\varepsilon_{l,t}$ ($\varepsilon_{i,t} = e_{i,t} + \gamma^{eu}_{i,t,S^{eu}_{i,t}} e_{eu,t} + \gamma^{us}_{i,t,S^{us}_{i,t}} e_{us,t}$) and (2) from the fact that we condition on $\hat{\varepsilon}_{eu,}$ and $\hat{\varepsilon}_{us}$. We now make the following additional assumptions:

$$\begin{aligned} E[e_{i,t} e_{j,t}] &= 0, & \forall i \neq j, \quad i, j = 1...L \\ E[e_{i,t} \hat{\varepsilon}_{eu,t,}] &= 0 & \forall i, \quad i, j = 1...L \\ E[e_{i,t} \hat{\varepsilon}_{us,t}] &= 0 & \forall i, \quad i, j = 1...L \end{aligned}$$

In addition, we suppose that the states $S_{i,t}$ are independent across the individual countries. It then follows that

$$\begin{aligned}
&\sum_{i^1}^{2} ... \sum_{i^L}^{2} f(r_1, ..., r_L | r_{eu,} r_{us}, S_t^1 = i^1, ..., S_t^L = i^L, \Omega_{t-1}; \theta) f(S_t^1 = i^1, ..., S_t^L = i^L, \Omega_{t-1}; \theta) \\
= &\sum_{i^1}^{2} f(r_1 | r_{eu,} r_{us}, S_t^1 = i^1) f(S_t^1 = i^1 | \Omega_{t-1}; \theta) .... \sum_{i^L}^{2} f(r_L | r_{eu,} r_{us}, S_t^L = i^L) f(S_t^L = i^L | \Omega_{t-1}; \theta) \\
= &\sum_{i^1}^{2} f(r_1 | r_{eu,} r_{us}, S_t^1 = i^1) f(S_t^1 = i^1 | \Omega_{t-1}; \theta_1) .... \sum_{i^L}^{2} f(r_L | r_{eu,} r_{us}, S_t^L = i^L) f(S_t^L = i^L | \Omega_{t-1}; \theta_L)
\end{aligned}$$

The first step comes directly from the independency assumptions, while the second step is possible as the individual country specifications do not share any parameters. Having done this, the parameters θ_i, for $i = 1...L$, can be determined by maximizing N different univariate likelihoods, each of them conditional on $\hat{\varepsilon}_{eu}$ and $\hat{\varepsilon}_{us}$:

$$\sum_{t=1}^{T} \log \left[\sum_{i^i=1}^{2} f(e_{i,t} | \hat{\varepsilon}_{eu,} \hat{\varepsilon}_{us}, S_t^i = i^i, \Omega_{t-1}; \theta_i) f(S_t^i = i^i | \Omega_{t-1}; \theta_i) \right]$$

Appendix 4: Derivation of the second-step loglikelihood in case of regime-switching

In the previous paragraph, the trivariate likelihood for US, EU, and local returns was split into a bivariate system for EU and US returns, and a univariate likelihood for the local returns, where the latter model is conditional on the return innovations from the bivariate model. In this appendix, I show how the univariate likelihood is constructed in case of regime-switching in the two spillover parameters. The case where both the spillover parameters and the local volatility can switch, is analogous. For the sake of generality, we focus on the case where no structure is added to the transition matrix. We start from the following model:

$$\begin{aligned} r_{i,t} &= \mu_{i,t} + \varepsilon_{i,t} \\ \varepsilon_{i,t} &= e_{i,t} + \gamma^{eu}_{i,S^{eu}_{i,t}} \hat{e}_{eu,t} + \gamma^{us}_{i,S^{us}_{i,t}} \hat{e}_{us,t} \\ e_{i,t}|\Omega_{t-1} &\sim N(0, \sigma^2_{i,t}) \end{aligned}$$

We suppose that the spillover parameters γ^{eu}_i and γ^{us}_i can be in two states only:

$$\gamma^{eu}_{i,t} = \left\{ \begin{array}{lll} \gamma^{eu}_{i,1} & \text{if} & S^{eu}_{i,t} = 1 \\ \gamma^{eu}_{i,2} & \text{if} & S^{eu}_{i,t} = 2 \end{array} \right\}$$

and

$$\gamma^{us}_{i,t} = \left\{ \begin{array}{lll} \gamma^{us}_{i,1} & \text{if} & S^{us}_{i,t} = 1 \\ \gamma^{us}_{i,2} & \text{if} & S^{us}_{i,t} = 2 \end{array} \right\}$$

The combination of $S^{eu}_{i,t}$ and $S^{us}_{i,t}$ creates a new state variable $S_{i,t}$, which is defined as:

$$\begin{aligned} S_{i,t} &= 1 \quad \text{if } S^{eu}_{i,t} = 1 \text{ and } S^{us}_{i,t} = 1 \\ S_{i,t} &= 2 \quad \text{if } S^{eu}_{i,t} = 2 \text{ and } S^{us}_{i,t} = 1 \\ S_{i,t} &= 3 \quad \text{if } S^{eu}_{i,t} = 1 \text{ and } S^{us}_{i,t} = 2 \\ S_{i,t} &= 4 \quad \text{if } S^{eu}_{i,t} = 2 \text{ and } S^{us}_{i,t} = 2 \end{aligned}$$

We make no assumptions about the interaction between the different states. In this general case, the transition matrix Π is given by

$$\Pi = \begin{bmatrix} p_{11} & p_{12} & p_{13} & p_{14} \\ p_{21} & p_{22} & p_{23} & p_{24} \\ p_{31} & p_{32} & p_{33} & p_{34} \\ p_{41} & p_{42} & p_{43} & p_{44} \end{bmatrix}$$

where $p_{ij} = P(S_t = j|S_t = i)$. The only structure imposed is that $p_{i1} + ... + p_{i4} = 1$, for $i = 1, ..., 4$, and that all $p_{ij} \geqq 0$. As shown in appendix 4, we want to maximize $f(r_{i,t}|\hat{e}_t^{eu}, \hat{e}_t^{us}, \Omega_{t-1}; \theta_i)$. For notational ease, we will omit $\hat{e}_t^{eu}$ and $\hat{e}_t^{us}$ from the formulas. Recall that $r_{i,t}$ is subject to switches between four regimes:

$$\begin{aligned}
f(r_{i,t}|\Omega_{t-1}; \theta_i) &= \sum_{j=1}^{4} f(r_{i,t}, S_{i,t} = j|\Omega_{t-1}; \theta_i) \\
&= \sum_{j=1}^{4} f(r_{i,t}|S_{i,t} = j, \Omega_{t-1}; \theta_i) P(S_{i,t} = j|\Omega_{t-1}; \theta_i) \\
&= \sum_{j=1}^{4} f(r_{i,t}|S_{i,t} = j, \Omega_{t-1}; \theta_i) p_{jt}
\end{aligned}$$

where $p_{jt} = P(S_{i,t} = j|\Omega_{t-1}; \theta_i)$ is the ex-ante probability of being in state j.

If conditional normality is assumed, we get that

$$f(r_{i,t}|S_{i,t} = j, \Omega_{t-1}; \theta_i) = \frac{1}{\sqrt{2\pi h_{it}}} \exp \left\{ \frac{-(r_{it} - \mu_{it} - \gamma^{eu}_{i,S^{eu}_{i,t}} \hat{e}_{eu,t} - \gamma^{us}_{i,S^{us}_{i,t}} \hat{e}_{us,t})^2}{2h_{it}} \right\}$$

We can decompose the ex-ante probability of being in state j as follows:

$$P(S_{i,t} = j|\Omega_{t-1}; \theta_i) = \sum_{k=1}^{4} P(S_{i,t} = j|S_{i,t-1} = k) P(S_{i,t-1} = k|\Omega_{t-1}; \theta_i), \quad j = 1, ..., 4.$$

where $P(S_{i,t-1} = k|\Omega_{t-1}; \theta_i)$ represents the filtered probability, which indicates the state of the economy at time t, using all the information available at time t. The latter can be further decomposed:

$$\begin{aligned}
P(S_{i,t-1} = k|\Omega_{t-1}; \theta_i) &= P(S_{i,t-1} = k|r_{i,t-1}, \Omega_{t-2}; \theta_i) \\
&= \frac{P(S_{i,t-1} = k, r_{i,t-1}|\Omega_{t-2}; \theta_i)}{P(r_{i,t-1}|\Omega_{t-2})} \\
&= \frac{f(r_{i,t-1}|S_{i,t-1} = k, \Omega_{t-2}; \theta_i) P(S_{i,t-1} = 1|\Omega_{t-2}; \theta_i)}{\sum_{j=1}^{4} f(r_{i,t-1}|S_{i,t-1} = j, \Omega_{t-2}; \theta_i) P(S_{i,t-1} = j|\Omega_{t-2}; \theta_i)}
\end{aligned}$$

Given initial values for the ex-ante probabilities, one can construct the likelihood iteratively. The parameters of the model are estimated by maximizing the loglikelihood function with respect to θ_i :

$$\pounds(r_{i,t}|\Omega_{t-1};\theta_i) = \sum_{t=1}^{T} \ln \phi(r_{i,t}|\Omega_{t-1};\theta_i)$$

The model and the construction of the information set imply that the loglikelihood function can be written as

$$\pounds(r_{i,t}|\Omega_{t-1};\theta_i) = \sum_{t=1}^{T} \ln \phi(e_{i,t}|\Omega_{t-1};\theta_i)$$

where the density function $\phi(.)$ is a weighted average of the four state dependent densities, where the weights are determined by the ex-ante probabilities:

$$\phi(e_{i,t}|\Omega_{t-1};\theta_i) = \sum_{j=1}^{4} f(e_{i,t}|S_{i,t-1} = j, \Omega_{t-1};\theta_i)P(S_{i,t-1}|\Omega_{t-2};\theta_i)$$

Table 1: Summary Statistics

All data are weekly Deutschmark denominated total returns over the period January 1980-August 2001, for a total of 1130 observations. SK stands for skewness, KURT for kurtosis, JB is the Jarque-Bera test for normality, ARCH(4) is a standard LM test for autoregressive conditional heteroskedasticity of order 4, Q(4) tests for fourth-order autocorrelation, and nobs is the number of observations. ***, **, and * means signficant at a 1, 5, and 10 percent level respectively.

	mean (%)	min (%)	max (%)	stdev (%)	SK	KURT	JB	ARCH(4)	Q(4)	Nobs
AUSTRIA	0.2350	-13.91	17.12	2.573	0.598	10.99	3059***	138.9***	45.58***	1130
BELGIUM	0.2703	-15.31	14.85	2.178	-0.068	7.47	934***	49.0***	29.14***	1130
FRANCE	0.3037	-18.11	11.82	2.667	-0.629	6.55	665***	120.7***	17.15***	1130
GERMANY	0.2403	-14.12	8.23	2.263	-0.729	5.86	481***	119.4***	13.76***	1130
IRELAND	0.3461	-21.89	10.45	2.754	-0.581	9.02	759***	138.9***	24.46***	1130
ITALY	0.3216	-16.23	12.76	3.508	-0.060	4.28	77***	82.8***	26.27***	1130
NETHERLANDS	0.3365	-12.39	10.16	2.132	-0.425	6.41	576***	108.7***	10.2**	1130
SPAIN	0.2437	-17.63	7.77	2.912	-0.677	5.50	246***	21.2***	11.33**	756
DENMARK	0.3212	-9.32	14.49	2.341	0.284	5.35	273***	3.7	12.95**	1130
NORWAY	0.2651	-18.46	18.66	3.399	0.030	5.96	409***	53.5***	20.55***	1130
SWEDEN	0.3269	-14.62	13.92	3.497	-0.342	4.73	167***	121.1***	7.46	1025
SWITZERLAND	0.2791	-17.04	8.08	1.989	-1.007	10.25	2656***	155.4***	26.75***	1130
UK	0.3195	-15.89	13.38	2.424	-0.415	6.44	587***	91.1***	12.72**	1130
US	0.3395	-14.61	8.71	2.711	-0.446	4.98	219***	52.3***	5.55	1130
EU	0.2884	-14.42	6.52	1.972	-0.929	7.55	1131***	166.4***	27.48***	1130

Table 2: Regime Classification Measure

This Table reports the Regime Classification Measure (RCM) under different number of regimes (two till four) and for different probabilities for the first state, assuming that the other states evenly divide the probability mass left between them. The RCM is given by

$$RCM2 = 1 - \frac{k}{k-1}\frac{1}{T}\sum_{t=1}^{T}\sum_{i=1}^{k}\left(p_{i,t} - \frac{1}{k}\right)^2$$

where k is the number of states, and $p_{i,t} = P\left(S_t = i|\Omega_T\right).$

	RCM		
	4 States	3 States	2 States
0.250	100.000	-	-
0.346	98.356	99.963	-
0.385	96.778	99.408	-
0.423	94.675	98.188	-
0.462	92.045	96.302	-
0.500	88.889	93.750	100.000
0.539	85.207	90.533	99.408
0.577	80.999	86.649	97.633
0.615	76.266	82.101	94.675
0.654	71.006	76.886	90.533
0.673	68.179	74.029	88.018
0.712	62.130	67.816	82.101
0.750	55.556	60.938	75.000
0.789	48.455	53.393	66.716
0.827	40.828	45.183	57.249
0.865	32.676	36.307	46.598
0.904	23.997	26.766	34.763
0.923	19.461	21.746	28.402
0.942	14.793	16.559	21.746
0.962	9.993	11.206	14.793
0.981	5.063	5.686	7.544
1.000	0.000	0.000	0.000

Table 3: Estimation Results for the Bivariate Models for EU and US returns

This table reports estimation results from a bivariate constant correlation model, a bivariate BEKK model, a regime-switching normal model, and a regime-switching GARCH model for the EU and US returns over the period January 1980 - August 2001. All returns are weekly Deutschmark denominated total returns. In Panel A, we report a battery of specification tests for the different models. First, it is tested whether the standardized residuals violate the orthogonality conditions implied by a standard normal distribution. "Mean" and "Variance" tests whether there is fourth-order autocorrelation left in the standardized and squared standardized residuals. "Covariance" tests whether the product of the standardized EU and US residuals is autocorrelated up to order 4. These test statistics are chi-square distributed with four degrees of freedom. "Joint" tests the mean, variance, and covariance jointly, and is $\chi^2(12)$ distributed. "Asym" tests whether the (co-)variance reacts asymmetrically to return innovations (Wald test on parameters in matrix D in the variance specification). In panel B, it is investigated whether the standardized residuals violate the conditions of the bivariate standard normal distribution. Specifically, it is tested whether the skewness, excess kurtosis, cross-skewness, and cross-kurtosis are significantly diferent from zero. These tests are all $\chi^2(1)$ distributed. The "joint" statistic tests the conditions jointly, and is $\chi^2(6)$ distributed. Probability levels are reported in squared brackets.

Panel A: Univariate Specification Tests for Bivariate Models for EU and US returns

UNIVARIATE TESTS	Mean		Variance		Covariance	Joint	Asym
	EU	US	EU	US			
Constant Correlation Model	8.898	3.202	5,615	2.415	58.189	65,647	13.289
	[0.064]	[0.525]	[0.229]	[0.659]	[0.000]	[0.000]	[0.001]
BEKK model	11.045	3.217	6.321	0.859	61.317	66,197	15.891
	[0.026]	[0.526]	[0.176]	[0.930]	[0.000]	[0.000]	[0.000]
Markov-Switching Normal	15,092	3.379	6,971	1,765	45,348	51.762	-
	[0.005]	[0.497]	[0.137]	[0.779]	[0.000]	[0.000]	-
Markov-Switching GARCH	10.198	3.4606	7.8483	4.082	45.0608	51.8548	6.223
	[0.037]	[0.484]	[0.097]	[0.544]	[0.000]	[0.000]	[0.045]

Panel B: Bivariate Specification Tests for Bivariate Models for EU and US returns

BIVARIATE TESTS	Skewness		Kurtosis		Cross-Skewness	Cross-Kurtosis	Joint
	EU	US	EU	US			
Constant Correlation Model	2,976	0,366	22,851	4,006	6,518	91,154	1029
	[0.085]	[0.366]	[0.000]	[0.045]	[0.038]	[0.000]	[0.000]
BEKK model	5,948	1,016	8,718	28,373	7,494	61,304	742,711
	[0.016]	[0.314]	[0.003]	[0.000]	[0.024]	[0.000]	[0.000]
Markov-Switching Normal	2,907	3,714	3,065	1,895	7,262	4,341	80,732
	[0.088]	[0.054]	[0.080]	[0.169]	[0.027]	[0.037]	[0.000]
Markov-Switching GARCH	4,191	5,044	4,554	4,282	9,372	7,997	105,049
	[0.041]	[0.025]	[0.033]	[0.039]	[0.009]	[0.005]	[0.000]

Table 4: Estimation Results for the Bivariate regime-switching Normal Model for EU and US returns

This table reports estimation results for the bivariate regime-switching normal model for EU and US returns. The model allows the returns $r_t = [r_{eu,t}, r_{us,t}]$ to be drawn from two different bivariate normal distributions:

$$r_t|\Omega_{t-1} = \begin{cases} N(\boldsymbol{\mu}_{t-1}(S_1), \mathbf{H}(S_1)) \\ N(\boldsymbol{\mu}_{t-1}(S_2), \mathbf{H}(S_2)) \end{cases} \tag{27}$$

The regimes follow a two-state Markov chain with transition matrix:

$$\Pi = \begin{pmatrix} P & 1-P \\ 1-Q & Q \end{pmatrix} \tag{28}$$

where the transition probabilities are given by $P = prob(S_t = 1|S_{t-1} = 1; \Omega_{t-1})$, and $Q = prob(S_t = 2|S_{t-1} = 2; \Omega_{t-1})$. In the mean equation, only the intercepts α_0 are made regime dependent:

$$\boldsymbol{\mu}_t = \boldsymbol{\mu}_{t-1} = \alpha_0 + \mathbf{A}r_{t-1}$$

where $\alpha_0 = [\alpha_{eu}, \alpha_{us}]'$, and $A = [\alpha_{eu}^{eu}, \alpha_{eu}^{us}; \alpha_{us}^{eu}, \alpha_{us}^{us}]$. Probability levels are

reported in squared brackets.

	EUROPEAN RETURS		US RETURNS	
	state 1	state 2	state 1	state 2
Volatility	0.0327	0.0156	0.0404	0.0236
	[0.0161]	[0.0000]	[0.0090]	[0.0000]
Correlation	0.8062	0.5605		
	[0.0498]	[0.0523]		
Constant	-0.0052	0.0039	-0.0041	0.0048
	[0.0339]	[0.0001]	[0.1516]	[0.0000]
P	0.9297 [0.0086]			
Q	0.9871 [0.0031]			

Table 5: Univariate model with constant shock spillover intentsity

This table reports estimation results for the univariate model with constant shock spillover intensity. This model is given by

$$\begin{aligned}
r_{i,t} &= \mu_{i,t-1} + \varepsilon_{i,t} \\
\varepsilon_{i,t} &= e_{i,t} + \gamma_i^{eu}\hat{e}_{eu,t} + \gamma_i^{us}\hat{e}_{us,t} \\
\mu_{i,t} &= \beta_{i0} + \beta_{i1}r_{i,t-1} + \beta_{i2}r_{eu,t-1} + \beta_{i3}r_{us,t-1} \\
e_{i,t}|\Omega_{t-1} &\backsim N(0, \sigma_{i,t}^2) \\
\sigma_{i,t}^2 &= \psi_{i0} + \psi_{i1}e_{i,t-1}^2 + \psi_{i2}\sigma_{i,t-1}^2 + \psi_{i3}\max\left(-e_{i,t-1}, 0\right)^2
\end{aligned}$$

where $\hat{e}_{eu}$ and $\hat{e}_{us}$ are respectively the European and US market shocks obtained from a first step estimation. Notice that $\hat{e}_{eu}$ is different for every country, as the European market portfolios are formed using the returns from all countries but the country that is being looked at. ***, **, and * means significant at a 1, 5, and 10 percent level respectively.

	ψ_{i0}	ψ_{i1}	ψ_{i2}	ψ_{i3}	β_{i0}	β_{i1}	β_{i2}	β_{i3}	γ_i^{eu}	γ_i^{us}
Austria	0.0032***	0.8310***	0.1352***	0.0690	0.0009*	0.0972**	0.1470***	-0.0733***	0.3418***	0.1672***
Belgium	0.0017***	0.9272***	0.0418***	0.0521	0.0024***	-0.0087	0.1439***	0.0109	0.5556***	0.3063***
France	0.0025***	0.8779***	0.1071***	-0.0009	0.0034***	-0.0607	0.0702	0.0046	1.0104***	0.4608***
Germany	0.0028***	0.8736***	0.0832***	0.0233	0.0027***	0.0344	0.0326	0.0127	0.8015***	0.3996***
Ireland	0.0021***	0.9247***	0.0502***	0.0412	0.0032***	0.0080	0.2469***	-0.0395	0.7186***	0.4118***
Italy	0.0031***	0.9245***	0.0712***	-0.0172	0.0026***	0.0225	0.0988	-0.0139	1.0686***	0.4347***
Netherlands	0.0017***	0.9188***	0.0602***	0.0052	0.0035***	-0.1067	0.1327***	0.0404	0.6473***	0.4568***
Spain	0.0051***	0.8444***	0.0613**	0.0445	0.0022***	0.0145	0.0285	0.0514	1.0071***	0.4603***
Denmark	0.0025***	0.9398***	0.0466**	-0.0011	0.0029***	0.0199	0.0776	0.0305	0.5111***	0.3032***
Norway	0.0028***	0.9225***	0.0535***	0.0354	0.0025***	0.0364	0.1369**	-0.0126	0.7139***	0.4698***
Sweden	0.0059***	0.8407***	0.0837***	0.0634	0.0033***	0.0232	0.1227	-0.0921	0.8991***	0.6281***
Switzerland	0.0063***	0.7138***	0.0574***	0.0953	0.0027***	0.0541	0.0221	0.0063	0.4496***	0.3551***
UK	0.0040***	0.7202***	0.1161***	0.0826	0.0035***	-0.0286	0.0623	0.0099	1.1608***	0.4412***

Table 6: Comparison of Different Univariate Spillover Models with Switches in the Spillover Parameters

Panel A of Table 6 reports the critical value of the joint test for normality of the residuals of four univariate volatility spillover models with different assumptions about the regime-switching properties of the spillover parameters: (1) No Regimes (NRS), (2) Common Regime Switches (JRS), (3) Independent Regime-Switches (IRS), and (4) a fully flexible transition probability matrix for the spillover parameter states (FULL). More detailed information about this statistic is provided in paragraph 3.3.2.1. Panel B reports the Regime Classification Measure for the different regime-switching shock spillover models, as described in paragraph 3.3.2.2. and Table 2. Panel C tests whether the models are significantly different from one another. Column one tests whether the model with joint regime switches perform statistically better than the constant spillover model, column two whether the IRS and JRS are statistically different, and column three JRS against the FULL model As in these tests the alternative hypothesis is not specified, the likelihood ratio test statistics are compared with their empirical distribution, obtained by a Monte Carlo analysis. In column two, likelihood ratio test investigates whether the FULL model can be simplified to the IRS model. Given that the latter model is nested in the former, the significance of the test statistic can be obtained from a χ^2 distribution with 1 degree of freedom. The values between brackets represent probability values. Finally, panel D tests whether the spillover parameters are significantly different across regimes. The Wald test is distributed as a χ^2 distribution with 1 degree of freedom. Probability levels are reported in squared brackets.

Panel A: Specification Tests

	NRS	JRS	IRS	FULL
Austria	1106.37	32.48	46.06	39.91
	[0.000]	[0.001]	[0.000]	[0.000]
Belgium	74.92	15.70	10.75	11.85
	[0.000]	[0.205]	[0.550]	[0.458]
France	224.12	18.47	16.11	14.07
	[0.000]	[0.102]	[0.186]	[0.296]
Germany	203.72	30.95	18.84	19.47
	[0.000]	[0.002]	[0.092]	[0.078]
Ireland	97.02	16.80	17.99	19.44
	[0.000]	[0.157]	[0.116]	[0.078]
Italy	75.47	32.84	29.25	24.75
	[0.000]	[0.001]	[0.004]	[0.016]
Netherlands	36.86	13.21	15.48	17.43
	[0.000]	[0.354]	[0.216]	[0.134]
Spain	101.99	24.97	20.29	17.49
	[0.000]	[0.015]	[0.062]	[0.132]
Denmark	32.79	11.97	11.16	12.18
	[0.001]	[0.448]	[0.515]	[0.431]
Norway	131.34	13.40	8.98	7.90
	[0.000]	[0.341]	[0.705]	[0.793]
Sweden	201.00	19.83	19.48	20.08
	[0.000]	[0.070]	[0.078]	[0.064]
Switzerland	306.00	39.50	27.32	26.61
	[0.000]	[0.000]	[0.007]	[0.009]
UK	233.26	18.76	15.07	16.42
	[0.000]	[0.094]	[0.238]	[0.173]

Panel B: Likelihood Ratio tests

	JRS>NRS[1]	IRS>NRS[1]	FULL>JRS[2]	FULL>IRS[3]	CHOICE[4]
Austria	81.38	72.91	6.90	15.37	JRS
	[0.000]	[0.000]	[0.735]	[0.052]	
Belgium	9.92	20.52	17.99	7.39	IRS
	[0.250]	[0.096]	[0.055]	[0.495]	
France	42.24	27.90	2.42	16.76	JRS
	[0.000]	[0.048]	[0.992]	[0.033]	
Germany	106.65	126.15	45.96	26.46	FULL
	[0.000]	[0.000]	[0.000]	[0.001]	
Ireland	27.23	40.14	14.96	2.06	IRS
	[0.000]	[0.000]	[0.133]	[0.979]	
Italy	33.49	42.12	13.16	16.96	IRS
	[0.000]	[0.000]	[0.215]	[0.031]	
Netherlands	51.02	82.30	48.08	16.80	FULL
	[0.000]	[0.000]	[0.000]	[0.032]	
Spain	10.40	18.60	15.84	7.64	(FULL)
	[0.424]	[0.133]	[0.104]	[0.469]	
Denmark	44.14	20.60	2.68	26.22	JRS
	[0.000]	[0.119]	[0.988]	[0.001]	
Norway	3.36	36.20	35.64	2.80	IRS
	[0.798]	[0.031]	[0.000]	[0.946]	
Sweden	59.86	55.48	14.88	19.26	JRS
	[0.000]	[0.000]	[0.136]	[0.014]	
Switzerland	70.92	69.40	20.80	22.32	JRS
	[0.000]	[0.000]	[0.023]	[0.004]	
UK	26.26	31.57	7.39	15.7	IRS
	[0.049]	[0.008]	[0.688]	[0.123]	

[1]probability values obtained through a Monte-Carlo analysis

[2]assumed to be distributed as a χ^2 distribution with 12 degrees of freedom

[3]distributed as a χ^2 distribution with 10 degrees of freedom

[4]model with highest LR test statistic. FULL is only choosen if it does statistically better than IRS/JRS.

Panel C: Regime Classification Measure

				Implied JRS		
	JRS	IRS	FULL	IRS	FULL	Best
Austria	32.90	52.71	29.74	70.41	42.08	JRS
	89.2%	72.3%	84.8%			
Belgium	63.75	33.82	54.83	48.10	77.25	IRS
	76.9%	84.3%	71.2%			
France	28.18	64.74	37.19	80.00	48.30	JRS
	89.9%	62.8%	80.6%			
Germany	31.09	44.28	49.25	60.71	62.18	JRS
	89.7%	77.4%	72.7%			
Ireland	25.92	43.03	41.05	60.03	58.37	JRS
	91.1%	79.3%	79.9%			
Italy	34.95	30.93	44.50	45.10	60.81	JRS
	87.7%	85.0%	77.5%			
Netherlands	38.13	33.67	41.07	47.93	61.03	JRS
	87.5%	83.5%	77.90%			
Spain	74.32	2.74	53.43	3.39	72.23	IRS
	72.4%	98.4%	71.4%			
Denmark	35.83	14.39	39.34	21.38	56.44	IRS
	88.7%	92.9%	83.0%			
Norway	21.25	52.71	61.27	70.41	88.77	JRS
	94.1%	72.3%	64.1%			
Sweden	57.55	64.08	35.27	81.09	51.38	FULL
	81.6%	64.4%	81.5%			
Switzerland	51.32	43.65	55.91	58.76	71.83	JRS
	82.2%	58.8%	69.3%			
UK	49.60	48.52	62.22	66.06	78.40	JRS
	84.8%	75.4%	63.8%			

Table 7: Wald Test for difference in shock spillover parameters

This table investigates whether the shock spillover intensities for the EU and US are statistically different across regimes. More specifically,we test the null hypotheses that $\gamma_i^{eu}(S_t = 1) = \gamma_i^{eu}(S_t = 1)$, and that $\gamma_i^{us}(S_t = 1) = \gamma_i^{us}(S_t = 1)$ against the alternative hypothesis that they are statistically different. Both Wald test statistics are χ^2 distributed with 1 degree of freedom.

	EU	US
Austria	3.985	5.325
	[0.046]	[0.021]
Belgium	22.121	8.624
	[0.000]	[0.003]
France	5.572	140.629
	[0.018]	[0.000]
Germany	38.369	49.325
	[0.000]	[0.000]
Ireland	26.701	11.308
	[0.000]	[0.001]
Italy	14.108	19.569
	[0.000]	[0.000]
Netherlands	3.714	4.235
	[0.054]	[0.040]
Spain	4.587	6.493
	[0.032]	[0.010]
Denmark	0.576	9.466
	[0.448]	[0.002]
Norway	19.243	1.256
	[0.000]	0.262
Sweden	5.698	36.150
	[0.017]	[0.000]
Switzerland	0.017	1.319
	[0.895]	[0.251]
UK	3.889	2.032
	[0.049]	[0.154]

Table 8: Shock Spillover Intensity through Time

This table reports shock spillover intensities from the EU (table A) and the US (table B) equity markets to the different local European equity markets, based upon the best performing regime-switching model (see Table 6, Panel B). The time-varying shock spillover intensities are obtained by weighting the shock spillover intensities by their filtered probability.

Panel A: Shock Spillover Intensity from EU

	Austria	Belgium	France	Germany	Ireland	Italy	Netherl.	Spain	Denmark	Norway	Sweden	Switz.	UK	Mean
FULL	0.33	0.51	0.91	0.75	0.69	1.03	0.65	0.97	0.45	0.74	0.84	0.44	0.54	0.68
80's	0.18	0.48	0.77	0.58	0.64	0.89	0.68	0.88	0.25	0.75	0.55	0.37	0.53	0.58
90's	0.46	0.52	1.02	0.90	0.74	1.15	0.63	1.00	0.62	0.74	1.04	0.50	0.56	0.76
%	152.7%	7.8%	31.4%	55.1%	15.0%	29.4%	-7.7%	13.6%	151.9%	-1.2%	90.2%	36.0%	5.0%	30.8%
80-85	0.11	0.46	0.70	0.39	0.60	0.93	0.70	-	0.24	0.74	0.47	0.32	0.56	0.52
86-90	0.31	0.52	0.88	0.88	0.70	0.82	0.64	0.89	0.33	0.75	0.66	0.46	0.49	0.64
91-95	0.50	0.52	1.02	0.84	0.74	1.13	0.58	1.02	0.63	0.75	0.95	0.47	0.59	0.75
96-98	0.53	0.55	1.04	0.83	0.75	1.21	0.67	1.00	0.64	0.73	1.31	0.53	0.57	0.80
99-01	0.31	0.49	1.04	1.04	0.73	1.21	0.69	0.97	0.60	0.73	0.98	0.52	0.50	0.75
86-90	174.7%	13.2%	27.0%	128.5%	15.1%	-11.8%	-7.4%	-	36.8%	0.5%	38.7%	43.6%	-11.0%	23.8%
91-95	58.7%	-1.6%	15.6%	-4.8%	6.4%	37.3%	-10.3%	14.7%	92.7%	0.1%	44.3%	1.4%	18.7%	16.8%
96-98	6.6%	7.4%	1.3%	-0.6%	0.8%	7.4%	15.4%	-1.8%	1.8%	-2.0%	38.1%	14.3%	-3.1%	6.4%
99-01	-41.1%	-11.6%	0.4%	24.7%	-1.5%	-0.3%	2.7%	-2.9%	-6.4%	-0.8%	-25.1%	-2.3%	-11.6%	-5.3%

Panel B: Shock Spillover Intensity from US

	Austria	Belgium	France	Germany	Ireland	Italy	Netherl.	Spain	Denmark	Norway	Sweden	Switzerl.	UK	Mean
FULL	0.15	0.29	0.40	0.34	0.36	0.42	0.47	0.45	0.31	0.46	0.61	0.35	0.49	0.39
80's	0.06	0.27	0.33	0.27	0.29	0.37	0.49	0.42	0.33	0.46	0.57	0.32	0.51	0.36
90's	0.22	0.30	0.46	0.39	0.42	0.46	0.45	0.46	0.29	0.46	0.63	0.37	0.48	0.41
%	232.5%	10.1%	39.0%	46.3%	44.8%	24.9%	-8.0%	7.4%	-12.1%	0.3%	11.8%	15.9%	-6.6%	13.9%
80-85	0.03	0.26	0.29	0.20	0.24	0.39	0.50	-	0.33	0.46	0.54	0.30	0.53	0.34
86-90	0.14	0.30	0.39	0.38	0.36	0.35	0.46	0.43	0.32	0.46	0.61	0.36	0.48	0.39
91-95	0.24	0.30	0.46	0.34	0.42	0.46	0.41	0.46	0.29	0.46	0.63	0.36	0.47	0.41
96-98	0.25	0.32	0.47	0.36	0.43	0.49	0.48	0.46	0.29	0.46	0.53	0.38	0.46	0.41
99-01	0.14	0.28	0.47	0.52	0.41	0.48	0.49	0.45	0.29	0.46	0.74	0.38	0.52	0.43
86-90	384.7%	17.1%	34.6%	87.7%	51.4%	-10.1%	-7.6%	-	-2.8%	-0.1%	13.2%	18.7%	-11.0%	14.1%
91-95	73.3%	-2.0%	18.9%	-11.3%	16.8%	31.2%	-10.6%	8.0%	-10.0%	0.0%	3.8%	1.0%	-1.3%	5.3%
96-98	7.5%	9.3%	1.5%	7.5%	2.0%	6.4%	16.1%	-1.0%	-0.4%	0.5%	-16.6%	5.0%	-1.9%	1.5%
99-01	-46.6%	-14.4%	0.4%	44.9%	-3.6%	-0.3%	2.8%	-1.7%	1.5%	0.2%	39.7%	0.3%	13.0%	4.9%

Table 9: Variance Proportions through Time

This table reports what proportion of the variance of the different local European countries is explained by EU (table A) and US (table B) shocks respectively. These are calculated using estimates from the best performing regime-switching shock spillover model (see table 7, panel B).

Panel A: Proportion of Variance explained by EU market

	Austria	Belgium	France	Germany	Ireland	Italy	Netherlands	Spain	Denmark	Norway	Sweden	Switzerland	UK	Average
FULL	0.09	0.17	0.21	0.15	0.17	0.16	0.20	0.17	0.12	0.13	0.12	0.14	0.12	0.15
80's	0.04	0.12	0.14	0.05	0.13	0.10	0.16	0.22	0.03	0.10	0.04	0.11	0.08	0.10
90's	0.12	0.21	0.27	0.24	0.21	0.20	0.24	0.27	0.20	0.15	0.19	0.16	0.15	0.20
%	179.2%	71.2%	90.7%	398.7%	67.5%	93.5%	46.8%	20.2%	662.1%	54.2%	317.4%	43.9%	90.8	97.9%
80-85	0.03	0.12	0.14	0.02	0.12	0.08	0.15	-	0.02	0.10	0.02	0.09	0.08	0.08
86-90	0.05	0.12	0.15	0.10	0.15	0.14	0.19	0.23	0.05	0.10	0.09	0.13	0.09	0.12
91-95	0.13	0.23	0.28	0.24	0.22	0.15	0.26	0.25	0.19	0.14	0.15	0.11	0.15	0.19
96-98	0.17	0.25	0.24	0.26	0.27	0.19	0.25	0.28	0.24	0.17	0.29	0.20	0.17	0.23
99-01	0.09	0.15	0.33	0.23	0.13	0.33	0.19	0.29	0.18	0.16	0.14	0.21	0.15	0.20
86-90	50.3%	-0.8%	5.0%	414.4%	24.6%	83.9%	27.2%	-	133.2%	5.5%	315.5%	46.2%	8.9%	51.1%
91-95	157.0%	86.1%	84.1%	143.1%	49.9%	9.4%	40.0%	12.1%	255.0%	38.6%	60.8%	-17.4%	73.5%	57.6%
96-98	30.3%	9.7%	-12.4%	5.5%	22.9%	26.1%	-6.3%	11.3%	21.9%	22.1%	89.1%	78.5%	15.2%	19.3%
99-01	-46.1%	-41.5%	36.5%	-8.9%	-51.8%	71.7%	-22.7%	3.8%	-25.3%	-7.0%	-50.3%	8.4%	-12.7%	-13.5%

Panel B: Proportion of Variance explained by US market

	Austria	Belgium	France	Germany	Ireland	Italy	Netherlands	Spain	Denmark	Norway	Sweden	Switzerland	UK	Average
FULL	0.07	0.18	0.23	0.18	0.18	0.14	0.42	0.24	0.13	0.19	0.23	0.26	0.34	0.21
80's	0.03	0.13	0.13	0.07	0.15	0.11	0.40	0.21	0.12	0.15	0.19	0.25	0.25	0.17
90's	0.10	0.22	0.31	0.27	0.21	0.18	0.44	0.24	0.14	0.22	0.28	0.28	0.41	0.25
%	273.4%	72.4%	138.1%	273.5%	40.0%	68.7%	9.7%	18.0%	15.5%	52.2%	47.7%	11.0%	63.7%	50.7%
80-85	0.01	0.13	0.12	0.05	0.16	0.11	0.37	-	0.10	0.14	0.15	0.25	0.23	0.15
86-90	0.04	0.13	0.15	0.12	0.15	0.13	0.45	0.22	0.14	0.16	0.24	0.26	0.29	0.19
91-95	0.11	0.24	0.30	0.27	0.20	0.12	0.41	0.23	0.13	0.19	0.25	0.29	0.38	0.24
96-98	0.14	0.25	0.33	0.29	0.24	0.22	0.42	0.24	0.16	0.25	0.26	0.24	0.47	0.27
99-01	0.08	0.16	0.35	0.28	0.20	0.23	0.48	0.27	0.14	0.25	0.35	0.28	0.45	0.27
86-90	181.3%	3.5%	33.8%	158.7%	-6.6%	16.9%	22.8%	-	42.2%	14.1%	52.6%	6.2%	25.0%	25.6%
91-95	164.9%	80.7%	97.5%	121.3%	33.9%	-6.7%	-9.8%	2.2%	-10.4%	20.2%	7.7%	12.3%	30.7%	25.8%
96-98	33.5%	4.5%	8.9%	7.7%	21.7%	83.3%	3.9%	3.3%	22.5%	31.3%	0.7%	-18.5%	22.9%	12.4%
99-01	-46.4%	-35.6%	5.4%	-3.5%	-16.6%	5.9%	14.7%	13.0%	-13.1%	-1.9%	35.5%	17.3%	-3.2%	0.4%

Table 10: Economic Determinants of Shock Spillover Intensity

In this table, I investate the link between the EU shock spillover intensity and a number of economic and financial indicators. First, a dummy variable is created that takes on a value of one when the filtered probability that the EU shock spillover intensity is in the high spillover state is large than 50 percent, and zero otherwise. Second, I investigate what variables explain why the shock spillover intensities switch from a low to a high state, or otherwise, using a simple logit model. The explanatory variables are (a) the degree of trade integration, measured by the sum of local exports to and imports from the EU over local GDP, (b) excess inflation, calculated as the local inflation in excess of the EU-15 inflation average, (c) exchange rate volatility, obtained by fitting a GARCH(1,1) to local exchange rate returns with respect to the ECU, (d) an equity market development indicator, proxied by the ratio of local equity market capitalization over local GDP, (e) a recession dummy, which is a dummy variable equaling one when the OECD leading indicator is below its quadratic trend, and zero otherwise, and finally (f) an "out-of-phase" dummy that records whether the local economy is out-of-phase with the aggregate EU economy. This dummy equals one if the recession dummy (see (e)) for the local economy and the EU are not equal.

	Austria	Belgium	France	Germany	Ireland	Italy	Netherlands	Spain	Denmark	Sweden	Switzerland	UK
C	27.26***	-3.30***	-2.06***	-4.15**	12.94***	-13.94***	-24.29**	-3.78***	-55.38***	-15.91***	-3.01***	0.83
	(2.48)	(0.61)	(0.50)	(1.95)	(2.41)	(2.01)	(10.64)	(0.57)	(10.20)	(1.45)	(0.44)	(1.47)
Trade Integration	53.41***	43.76***	17.90***	29.48***	-10.05***	81.22***	27.60*	14.36***	182.33***	42.43***	-	10.40
	(5.08)	(0.64)	(2.80)	(8.29)	(2.51)	(10.06)	(14.81)	(1.57)	(31.30)	(4.06)	-	(10.50)
Excess inflation	2609.3***	-2352.46***	-317.95*	3092.05***	-1999.62***	-15.61	-1793.27**	-450.00*	-10740.88***	-636.39***	622.62***	285.98
	(306.14)	(281.19)	(191.46)	(312.05)	(267.16)	(85.95)	(1026.68)	(242.76)	(2994.11)	(83.09)	(75.47)	(300.77)
Currency Volatility	-1219.49***	-166.80*	110.47**	813.49***	265.53***	76.77	1225.15**	89.72**	-90.88	12.04	186.54***	71.10
	(259.11)	(95.38)	(46.931)	(134.89)	(95.97)	(49.205)	(485.50)	(42.41)	(296.11)	(11.22)	(60.44)	(69.56)
MCAP/GDP	42.35***	23.14***	0.87	7.96**	-4.30*	0.07	10.04***	3.62*	23.88***	3.70***	0.68***	-3.65***
	(4.12)	(1.89)	(0.88)	(3.37)	(2.22)	(1.62)	(2.12)	(2.01)	(8.35)	(0.34)	(0.10)	(1.26)
Business Cycle	-0.46***	-0.19***	-0.09*	0.43***	0.79***	-0.55***	-1.55***	-0.23***	0.72**	-0.48***	0.19***	-0.42***
	(0.11)	(0.05)	(0.05)	(0.05)	(0.16)	(0.08)	(0.40)	(0.04)	(0.30)	(0.05)	(0.04)	(0.12)
Out of Phase	1.05	1.41***	-0.60**	-0.58***	30.96***	-1.30***	-47.31***	0.22	11.89***	-0.25	-	0.09
	(0.73)	(0.35)	(0.24)	(0.47)	(0.71)	(0.30)	(2.72)	(0.34)	(2.83)	(0.26)	-	(0.19)
R^2	69.5%	56.9%	15.9%	68.4%	69.8%	18.9%	56.9%	18.2%	92.0%	26.1%	19.8%	20.9%

Table 11: Test for Contagion

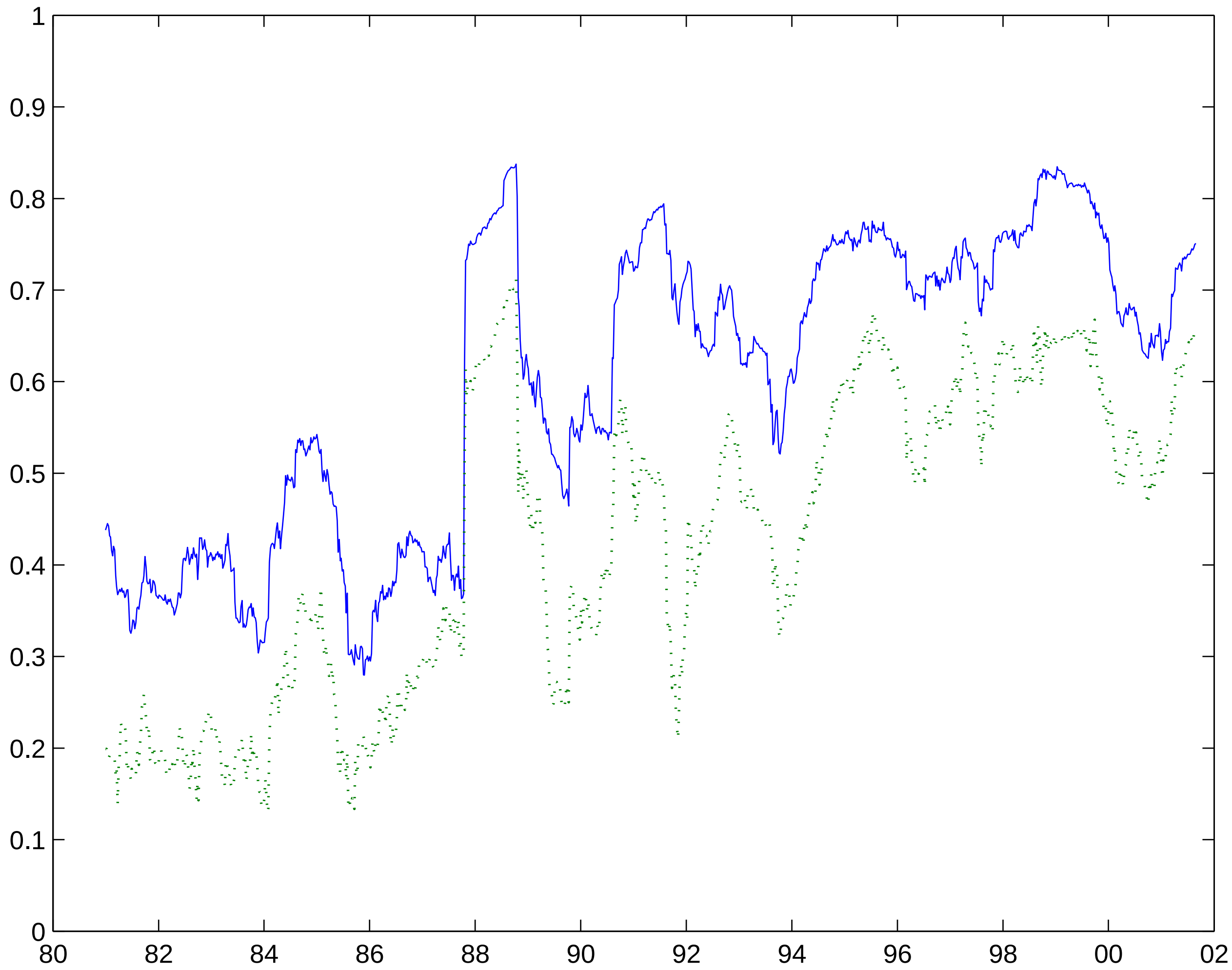
In this table, I investigate whether there is evidence for contation effects from the EU and US markets to the local European equity markets. I estimate the following model using GMM:

$$\hat{e}_{i,t} = b_1 + (b_2 + b_3 D_t)\hat{e}_{eu,t} + (b_4 + b_5 D_t)\hat{e}_{us,t} + u_{i,t}$$

where $\hat{e}_{eu,t}$ and $\hat{e}_{us,t}$ are the orthogonalized residuals from the bivariate model for EU and US returns, and D_t is a "crisis" dummy that takes on a one when the EU and US are jointly in a high volatility state, and zero otherwise. Overall contagion from the aggregate European market (excluding the country being looked at) would be there if b_2 and b_3 are jointly different from zero, while b_3 measures the extra contagion during crisis periods. Similarly, we cannot reject contagion from the US markets when b_4 and b_5 are jointly different from zero; here b_5 measures the extra contagion during crisis periods. Probability values are based upon heteroskedasticity and autocorrelaton consistent standard errors. Both the EU and US wald tests for no contagion are χ^2 distributed with two degrees of freedom.

	Austria	Belgium	France	Germany	Ireland	Italy	Netherlands	Spain	Denmark	Norway	Sweden	Switzerland	UK	joint
b1	0.0013	-0.0002	-0.0005	-0.0005	-0.0004	0.0004	-0.0001	-0.0001	0.0001	-0.0003	-0.0004	-0.0003	0.0003	0.0003
	(0.0008)	(0.0006)	(0.0007)	(0.0006)	(0.0007)	(0.0010)	(0.0004)	(0.00078)	(0.0006)	(0.0009)	(0.0008)	(0.0005)	(0.0014)	(0.0002)
b2	0.092	-0.068	-0.077	0.040	0.020	0.059	0.010	0.013	-3.58E-02	0.014	0.047	0.043	0.042	-0.004
	(0.062)	(0.044)	(0.053)	(0.043)	(0.053)	(0.080)	(0.031)	(0.064)	(0.048)	(0.072)	(0.068)	(0.036)	(0.046)	(0.018)
b3	0.041	0.252	0.170	0.189**	0.025	-0.261	0.147	0.061	-0.003	0.018	0.195	-0.027	-0.082	0.029
	(0.167)	(0.161)	(0.157)	(0.086)	(0.201)	(0.217)	(0.109)	(0.135)	(0.124)	(0.177)	(0.203)	(0.111)	(0.084)	(0.038)
b4	0.002	-0.012	-0.043	0.023	-0.024	-0.056	-0.001	-0.028	-0.030	0.018	-0.001	-0.001	-0.022	-0.027***
	(0.030)	(0.022)	(0.027)	(0.025)	(0.030)	(0.035)	(0.017)	(0.032)	(0.024)	(0.038)	(0.035)	(0.019)	(0.039)	(0.009)
b5	0.134	0.148**	0.290**	0.207***	0.306***	0.261***	0.063*	0.212***	0.096	0.161**	0.286***	0.175**	0.083*	0.092***
	(0.107)	(0.074)	(0.116)	(0.052)	(0.079)	(0.086)	(0.033)	(0.074)	(0.079)	(0.079)	(0.058)	(0.070)	(0.059)	(0.019)
Wald EU	3.03	4.09	2.49	8.82	0.16	1.48	2.43	0.42	6.74E-01	0.07	1.96	1.42	0.05	0.60
	[0.22]	[0.13]	[0.29]	[0.01]	[0.92]	[0.48]	[0.30]	[0.81]	[0.71]	[0.97]	[0.38]	[0.49]	[0.98]	[0.74]
Wald US	1.75	4.03	7.15	24.48	15.10	9.31	2.43	8.76	2.41	5.38	32.67	6.32	36.98	25.14
	[0.42]	[0.13]	[0.03]	[0.00]	[0.00]	[0.01]	[0.10]	[0.01]	[0.30]	[0.07]	[0.00]	[0.04]	[0.00]	[0.00]

Figure 1: 52-week average moving correlation of local returns with EU and US*.



* dotted line represents US

Figure 2: Filtered Probability of being in High Volatility Regime
(for bivariate regime-switching normal model for EU and US returns)

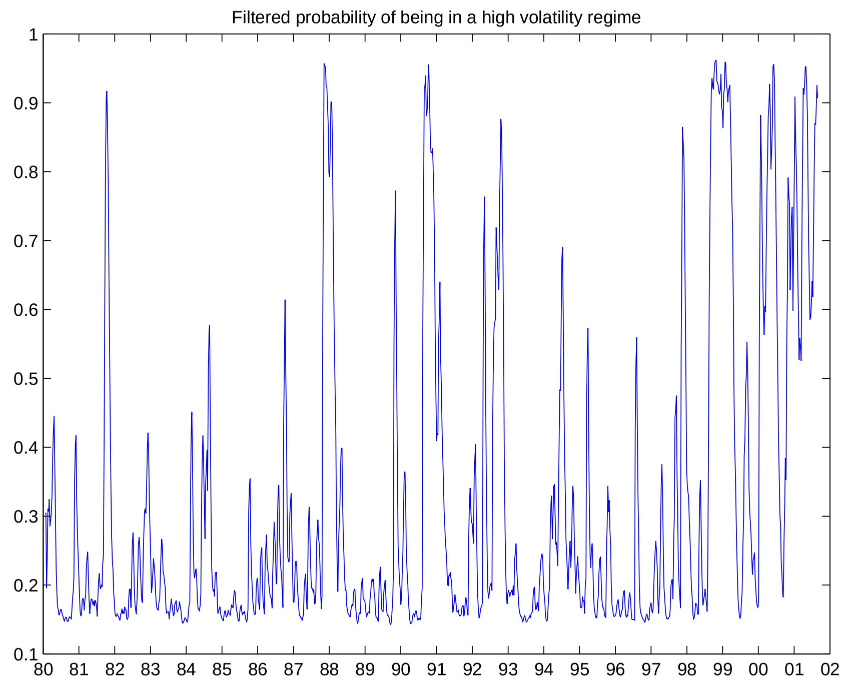
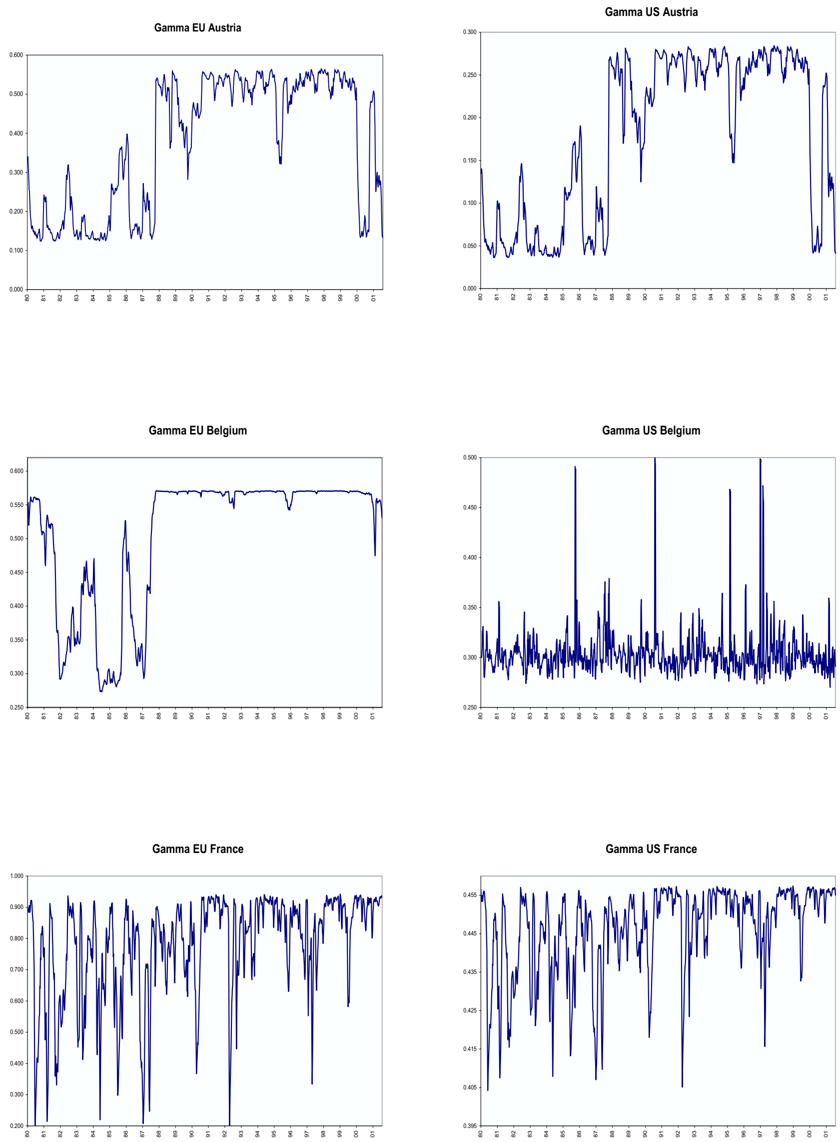


Figure 3: Time-Varying Volatility Spillovers from EU and US Markets to the Local European Equity Markets



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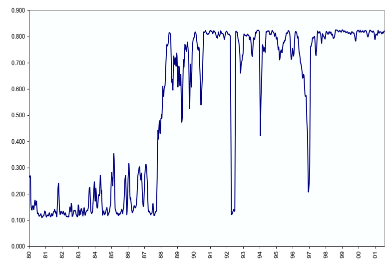
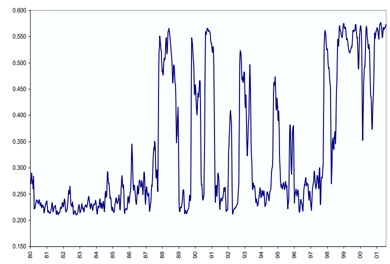


Gamma US Germany



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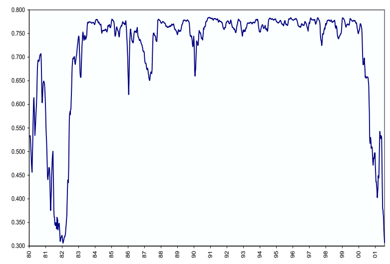

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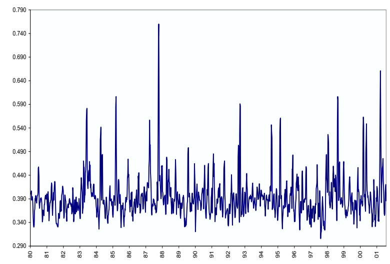

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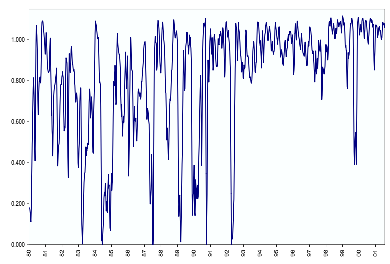

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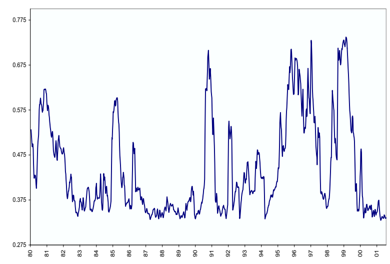

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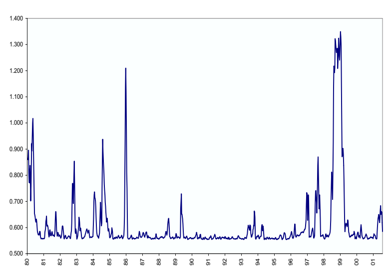

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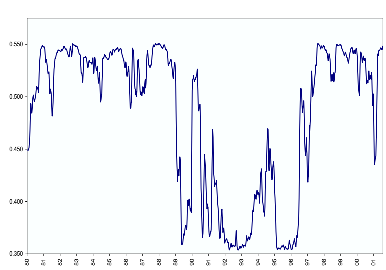

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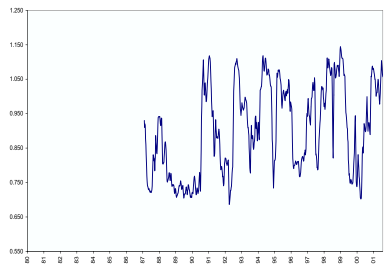

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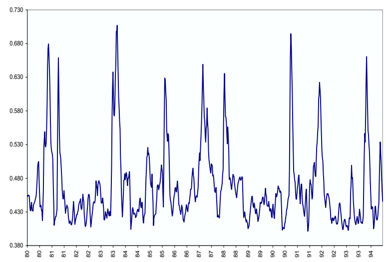

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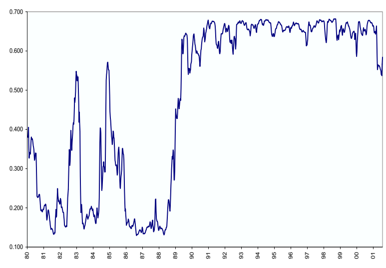

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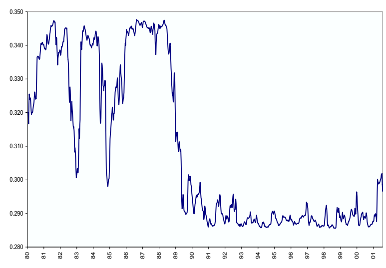

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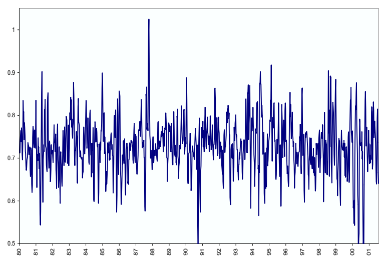

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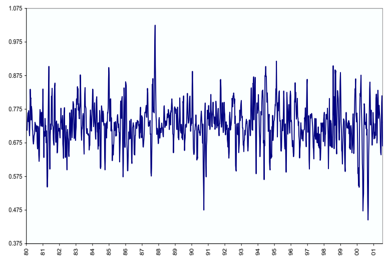

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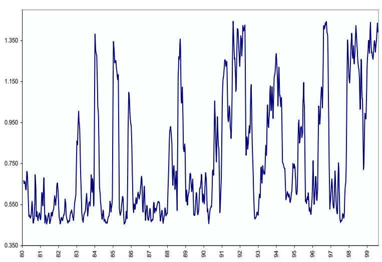

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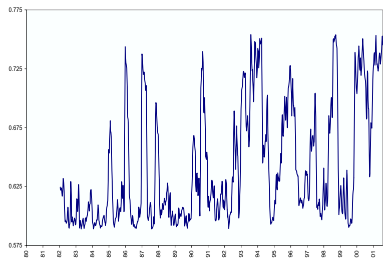
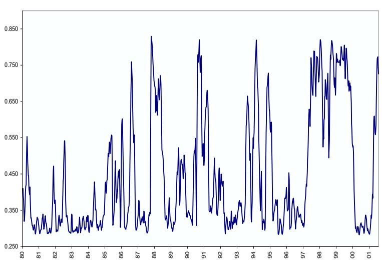


Gamma EU Switzerland



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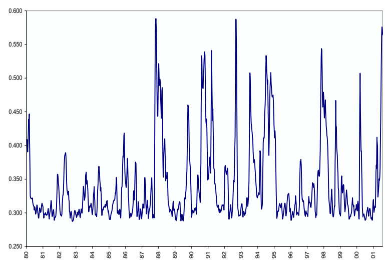

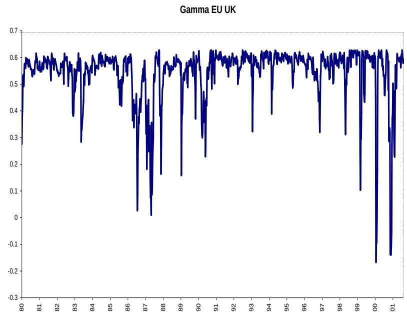
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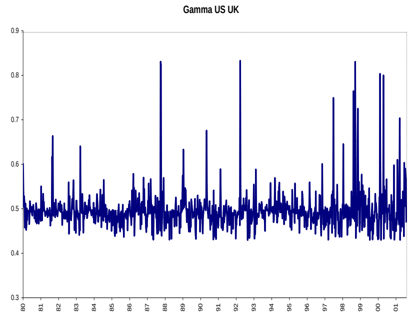
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