Who Leads Financial Markets?

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Abstract

The present paper embarks on an analysis of interactions between the US and Euroland in the capital, foreign exchange, money and stock markets from 1994 until 2006. Estimating multivariate EGARCH processes for the structural financial innovations determines causality-in-variance effects and provides a solution to the simultaneity problem of identifying the contemporaneous impacts between the daily variables. Structural mean equations can therefore give answers to the question of financial markets leadership: Generally speaking, the US effects on Europe still dominate, but the special econometric methodology is able to uncover otherwise neglected spillovers in the reverse direction.

Keywords: Structural EGARCH, Financial Markets, United States, Euro Zone
JEL classification: C32, G15

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1 Introduction

Every day, the whole professional financial world looks at New York and Washington: The latest stock developments at the Wall Street are taken as indicators for the future behaviour of other stock exchanges; decisions by the Federal Reserve Bank on the US monetary policy rate determine the perception of the current state of the business cycle and the development of economic growth; the US dollar is the only real world currency, its exchange rate is of outstanding importance for the international economy; bonds of the Federal Government are the epitome of capital assets, their yields serve as secure benchmark returns.

In the Old World, the European Union (EU) and the Economic and Monetary Union (EMU) have formed a regional bloc, which could be a match for the USA in terms of economic figures. Nonetheless, neither its currency, nor its central bank or stock exchanges have yet attained the same significance as their US counterparts. This seeming contradiction brings us to a simple research question: *Who leads financial markets?*

In detail, the underlying paper focuses on spillover dynamics between four of these markets: Answering the question for the short-term money market implies to pinpoint the relative strength and international autonomy of the Federal Reserve and the European Central Bank. The long-term interest rates, which are crucial for investment and business cycles, are determined in the capital market open to both domestic and foreign influences. In the foreign exchange market, one has to ask for the world's leading currency. At last, for the equity market the task is to figure out, where the effective trend-setting impulses are generated and how they transmit "across the Atlantic".

The US-European interactions have been comprehensively analysed by Ehrmann et al. (2005), who state an American predominance in the international financial markets. Ehrmann and Fratzscher (2002, 2004), Chinn and Frankel (2003) as well as Weber (2007a) have focused specifically on interest rate relations. As a main result, predominantly the European markets are found depending on US influences, even if a shift in the run-up to the third stage of the EMU in 1999 might be perceptible. The US-European foreign exchange market has for example been considered by Galati and Ho (2001), who investigate exchange rate effects of macroeconomic news with overall mixed results. An examination of the leading equity indices of New York and Frankfurt is given in Baur and Jung (2006), paying special attention to the sequence of trading segments.

The current analysis features a special econometric approach in the spirit of identification through heteroscedasticity, which is able to determine the contemporaneous effects between the daily financial variables. Thereby, on the one hand, the results confirm that the US still holds on to the leading position. On the other hand, applying the particular estimation technique, important influences originating on the part of Euroland can be uncovered. In contrast, neglecting the simultaneous causalities for example in the capital and equity markets would erroneously indicate far-reaching independence of the US.

Methodological wise, I specify bivariate time series models for the conditional means and variances of the daily European and US short-term interest rates, long-term bond yields, exchange rates and stock indices. The corresponding financial markets are processing new economic signals in very short time windows, even within hours of the same day. Therefore, it proves crucial to identify the contemporaneous effects between the variables. To meet this necessity, I adopt the approach proposed in Weber (2007b), proceeding in three steps: For establishing systematic responses in the conditional mean, I first estimate reduced-form VAR models, thereby taking regard of possible cointegrating relations. The heteroscedasticity in the residuals is then picked up in multivariate exponential generalised autoregressive conditional heteroscedasticity (EGARCH) models. The exponential form assures the covariance matrix to be positive definite and additionally allows for asymmetries.

In the conventional approach, such models are specified for the *reduced-form* residuals, which can be seen as linear combinations of the underlying structural shocks. In contrast, the present methodology addresses directly the conditional variances of these *structural* innovations, thereby giving a solution to the problem of identifying the mutual simultaneous impacts between the variables. This enables me to estimate structural-form mean equations in the last step without imposing constraints on the parameters, all of which are necessarily questionable in a financial markets context. Rigobon and Sack (2003) have proposed a related variant of "structural GARCH", but these authors still estimated a model with reduced-form residuals, even if characterised by structural restrictions. An according application to Asian and US stock markets is given in Lee (2006).

The underlying analysis is set up as follows: In the subsequent section, economic theories on international financial spillovers are treated. Thereafter, section 3 introduces the methodology of structural EGARCH estimation. The empirical results of the application on the transatlantic interrelations will be presented in section 4, which is followed by a summary with concluding considerations.

2 Economic Foundation

This paper sets out analysing the mutual dependences between important US and European financial markets. Common macroeconomics provides useful pieces of theory for international interlinkages of different asset types as well as certain common-sense explanations, shortly sketched in the following paragraphs. While therein, the focus naturally lies on interactions in the conditional mean, spillovers of volatility can be ascribed the role of a proxy for information flows between markets (Ross 1989).

Capital and Money Markets

In international economics, both short- and long-term interest rates are typically modelled by the uncovered interest rate parity (UIP). The economic rationale of the UIP is the *ex ante* arbitrage condition between domestic and foreign capital markets: Interest differentials between assets with equal maturity measured in local currencies with otherwise similar characteristics must be offset by corresponding expectations on the exchange rate development. This leads to the logarithmic UIP version

$$i_{t,m} - i^*_{t,m} = \frac{\text{days per year}}{m}(E_t(s_{t+m}) - s_t) + \rho_{t,m} \ , \qquad (1)$$

where $i_{t,m}$ and $i^*_{t,m}$ are the annualised domestic and foreign interest rates with m day maturity, s_t is the logarithm of the spot exchange rate (in terms of domestic currency units per foreign currency unit) and E_t the conditional expectations operator. $\rho_{t,m}$ denotes the logarithm of a risk premium, reflecting risk aversion, differences in credit worthiness and such.

In line with the relevant literature, assume the exchange rate integrated of order one (I(1)), or more special a random walk. Then, under rational expectations, the change on the right hand side of (1) should at least be stationary. In this case, a valid linkage following the UIP relation depends on interest differentials to be stationary, too. Hence, provided that interest rates represent I(1) processes, domestic and foreign bond yields should be cointegrated with the vector $(1, -1)$.

While such a result would direct at strong market integration, the focus in this research will be on causal effects in long-run adjustment and in short-run dynamics, including contemporaneous impacts. At the short end of the yield curve, such influences can be connected to central bank behaviour, because transmission from monetary policy rates is relatively direct. Accordingly, international dependences address the degree of autonomous power of central banks, or more likely the market expectations on this subject. Opposingly, the long-term bond market tends to be governed far more by private arbitrage between different capital assets. Here, a leading role therefore indicates that the burden of adjustment to international equilibrium deviations can be shifted onto foreign economies.

Stock Market

The stock market interlinkages might be characterised by a concept quite similar to the UIP: Equilibrium requires equalised overall expected returns, which again are determined by the assets themselves as well as the underlying currencies. The uncovered equity return parity (URP, see Cappiello and De Santis 2005) in its logarithmic form results as

$$E_t(r_{t+1} - r^*_{t+1}) = E_t(s_{t+1}) - s_t + \rho_t \ , \tag{2}$$

with r_t and r^*_t being daily returns on domestic and foreign stocks, and ρ_t again denoting a risk premium. While the arbitrage mechanism is similar to the UIP (1), unlike the interest rates, the stock returns in the URP are not known *ex ante*. As a second difference, only stationary variables are included in the URP, which so does not allow deducing any cointegrating relations. Consequently, equation (2) is not of direct importance for the empirical assessment of impulse transmission in section 4.

Nevertheless, sources and functioning of typical stock market transmission shall be briefly explained: The most direct way is the appearance of shocks, which are believed not to be regionally limited and thus to have further reaching relevance. A more distinctly defined channel works through trade relations: In case positive equity developments indicate accelerating growth in one country, others can participate according to their export shares. Similar effects arise from the presence of transnational companies or production networks: Innovations in one part of the world can easily affect the balance of the parent firm listed in the domestic stock exchange. Moreover, expectations rather than manifested economic processes themselves inherently govern financial pricing. By the same reason, *a priori* one should not rule out purely psychological reasons of spillovers as an outcome of a simple signalling function of stock indices.

Foreign Exchange Market

Economic theories concerning the foreign exchange market normally consider the bilateral exchange rate between the countries of interest directly. The present approach however aims at finding evidence for interactions *between* the US and European financial market variables. Therefore, I adopt the British pound as numeraire for both the euro and US dollar exchange rate. To me, this choice seems appropriate, because the UK maintains intensive relations with both Euroland and the US, and the pound is a liquid and much traded currency. However, one should keep in mind that the EUR/GBP and USD/GBP rates naturally share the common UK component. Since this obviously triggers a trivial link, interpreting interaction effects is not as straightforward as with other assets.

Causal effects between different exchange rates have for example been addressed in the contagion literature (for a review, see for example Pericoli and Sbracia 2003). Therein, spillovers of sudden depreciation and volatility in periods of crisis are subject of interest. In this, normally the US dollar serves as base currency. The present analysis follows the same principle, differing however in focusing on mutual dependences in non-crisis times.

3 Methodological Proceeding

3.1 Models for the Mean Process

The basic generating process of the data in the econometric procedure is assumed to be approximated by the structural-form VAR (SVAR) with lag length $q+1$

$$Ay_t = \mu_0^* + \mu_1 d_t + \sum_{j=1}^{q+1} B_j^* y_{t-j} + \varepsilon_t \ , \tag{3}$$

where y_t contains n (here 2) endogenous variables, the B_j^* denote $n \times n$ coefficient matrices, and the deterministic terms are a constant and centred daily seasonal dummies d_t, which control for possible day-of-the-week effects. ε_t represents an n-dimensional vector of uncorrelated heteroscedastic residuals. Uncorrelatedness comes as a standard assumption in the structural VAR literature. In the present empirical application one might think of all news shocks either belonging to the US or to the European side; the whole contemporaneous correlation is logically captured by mutual spillover effects in the matrix A (with normalised diagonal), what might be an acceptable postulation in case of the two largest economic world powers.

Given the presence of unit roots in the data, according to Johansen (1995), the commonness of $n-r$ stochastic trends is reflected by a reduced rank of $B^*(1)$, with $B^*(L) = A - \sum_{j=1}^{q+1} B_j^* L^j$. Consequently, one can write $B^*(1) = -\alpha\beta'$, where β spans the space of the r cointegrating vectors, and α contains the corresponding adjustment coefficients. Granger's representation theorem then leads to the structural vector error correction model (SVECM)

$$A\Delta y_t = \alpha(\beta' y_{t-1} + \mu_0) + \mu_1 d_t + \sum_{j=1}^{q} B_j \Delta y_{t-j} + \varepsilon_t \ , \tag{4}$$

with $B_j = -\sum_{k=j+1}^{q+1} B_k^*$, $j = 1, \ldots, q$. This representation assumes the constant absorbed in the cointegrating relation.

As (4) for itself is not identified, the reduced-form VECM is derived:

$$\Delta y_t = \alpha^r(\beta' y_{t-1} + \mu_0) + \mu_1^r d_t + \sum_{j=1}^{q} B_j^r \Delta y_{t-j} + u_t \ . \tag{5}$$

All coefficients are obtained by premultiplying A^{-1} in (4), therefore being marked by the superscript r for "reduced". Accordingly, the new residuals are given by $u_t = A^{-1}\varepsilon_t$.

The unit root behaviour of the series is checked by ADF tests (see e.g. Dickey and Fuller 1979), including a constant and centred seasonal dummies as deterministic terms. Here, as well as in all subsequent models, the lag length is set following the usual information criteria (maximum lag 10) and Lagrange multiplier (LM) autocorrelation tests. Simulated critical values for the null hypothesis of non-stationarity are taken from MacKinnon (1996).

For finding out the number of common stochastic trends, Johansen (1994, 1995) provides the likelihood ratio (LR) trace test. The test statistic for the null hypothesis of at most r cointegrating relations is given by:

$$\Lambda(r) = -T \sum_{j=r+1}^{n} log(1 - \hat{\lambda}_j) \ , \tag{6}$$

where n is the number of endogenous variables and T the number of observations. $\hat{\lambda}_j$ denotes the j-th largest squared sample canonical correlation between Δy_t and the respective cointegrating relation, both corrected for the influence of the remaining regressors. Critical values are obtained by computing the response surface in Doornik (1998).

In case even $r = 0$ is not rejected, equations (4) and (5) are specified without any cointegration terms, simply leaving VARs in first differences. Consequently, the constants are then considered outside the cointegrating relations.

3.2 Identification through Heteroscedasticity

In the structural VAR equation (3), the matrix A of contemporaneous effects cannot be identified without further constraints. However, the research target of determining *mutual* financial impacts between two "large countries" does not suggest any sensible zero restrictions. By the same token, it proves impossible to recover the structural parameters from the reduced form given by (4): In the matrix A with normalised diagonal, $n(n-1)$ simultaneous impacts have to be estimated, but due to its symmetry the covariance-matrix of the reduced-form residuals delivers only $n(n-1)/2$ equations for simultaneous covariances.

Instead of reducing the number of parameters by restrictions, the present approach aims on principle at augmenting the number of determining equations (for instance, see Rigobon (2003) for a lengthy discussion): If one can identify for example several regimes with differing volatility, the necessary shifts in the covariance-matrix yield additional equations for uncovering the structural parameters, which for their part have to be assumed constant over time.

The next section describes the set-up and estimation of a so-called "structural EGARCH" model. In this, I basically follow the intuition of identification through volatility regimes. A multivariate GARCH however practically defines a distinct variance state for every single observation. This can be thought of as modelling a continuum of regimes, which is reflected in the estimated variance processes.

3.3 Structural EGARCH

Since the variance model shall be specified for the structural residuals, according to (5), these are recovered by

$$\varepsilon_t = Au_t \; . \tag{7}$$

Furthermore, define the conditional variances of the elements in ε_t by

$$\mathrm{Var}(\varepsilon_{jt}|\Omega_{t-1}) = \mathrm{Var}_{t-1}(\varepsilon_{jt}) = h_{jt} \qquad j = 1, \ldots, n \; , \tag{8}$$

where Ω_{t-1} denotes the whole set of available information at time $t-1$.

Then, stack the h_{jt} in the vector $H_t = \begin{pmatrix} h_{1t} & h_{2t} & \ldots & h_{nt} \end{pmatrix}'$.

At last, denote the standardised white noise residuals by

$$\tilde{\varepsilon}_{jt} = \varepsilon_{jt}/\sqrt{h_{jt}} \qquad j = 1, \ldots, n \; . \tag{9}$$

The multivariate EGARCH(1,1)-process is then given by

$$\log H_t = C + G \log H_{t-1} + D|\tilde{\varepsilon}_{t-1}| + F\tilde{\varepsilon}_{t-1} \;, \tag{10}$$

where C is an $n \times 1$ vector of constants, and G, D and F are $n \times n$ coefficient matrices. The absolute value operation is to be applied element by element.

The univariate EGARCH has been proposed by Nelson (1991). Due to the conditional uncorrelatedness of the structural residuals, the multivariate extension (10) only comprises the conditional variances. The exponential form, linearised by taking logarithms, guarantees these variances to be positive. Together with the zero correlations assumption, this is sufficient for positive definite covariance matrices. Furthermore, asymmetric effects are incorporated by including $\tilde{\varepsilon}_t$ without taking absolute values: Any parameters in F differing from zero indicate that besides the magnitude of a shock its sign contains valuable information for forecasting the conditional variances. EGARCH(1,1) seems to be appropriate for most series, what will be shown by ARCH-LM tests, has additionally been checked in univariate models and is quite usual in financial econometrics. Apart from that, higher-order lags would considerably complicate the likelihood optimisation.

Let Σ_t denote the conditional covariance-matrix of ε_t including the h_{jt} on the main diagonal and zeros off-diagonal. Then the log-likelihood under the assumption of conditional normality results as

$$L(A, C, G, D, F) = L(\theta) = \sum_{t=1}^{T} \log l_t(\theta) = -\frac{1}{2} \sum_{t=1}^{T} (n \log 2\pi + \log|\Sigma_t| + \tilde{\varepsilon}_t' \tilde{\varepsilon}_t) \;. \tag{11}$$

Since assuming conditional normality is often problematic using financial markets data, the estimation relies on Quasi-Maximum-Likelihood (QML), see Bollerslev and Wooldridge (1992). While excess kurtosis may be taken as an argument for adopting in example a Student-t-distribution, QML has the advantage of being robust against violations of the distributional assumption. Although consistency and asymptotic normality are not proven for my particular model, results from the MGARCH literature suggest that

$$\sqrt{T}(\hat{\theta} - \theta) \xrightarrow{d} N(0, M_1^{-1} M_0 M_1^{-1}) \;, \tag{12}$$

where $M_1 = -E(\frac{\partial^2 \log l_t(\theta)}{\partial\theta\partial\theta'})$ and $M_0 = E(\frac{\partial \log l_t(\theta)}{\partial\theta} \frac{\partial \log l_t(\theta)}{\partial\theta'})$ (see Comte and Lieberman 2003). Note that the usual ML covariance matrix estimator $\hat{M}_1^{-1}$ would not be consistent under non-normality.

The likelihood optimisation is done using the BHHH algorithm (Berndt et al. 1974) in the Gauss Maxlik procedure, the code is written by the author. The parameter starting values are obtained from univariate EGARCH estimates, the structural and cross-effect coefficients are set to zero, and the variance process is started at the sample moments. The choice of the starting values did not prove crucial, supported by the common result in the multivariate GARCH literature that in any case most cross-coefficients are insignificant.

3.4 Combining Reduced-Form and Structural Estimates

For concluding, remember that the first two steps comprise estimating the reduced form (5) and the structural EGARCH (10). At last, the obtained A-matrix is substituted in equation (4), which is then re-estimated to gain values and standard errors of the remaining parameters. This procedure allows the determination of dynamic effects running over a certain time span in addition to the contemporaneous impacts from the identified structural coefficient matrix.

4 Empirical Evidence

4.1 Data

I consider the various markets under interest using the following daily data for the US and the euro zone: For the capital markets, I chose the yields of 10-year constant maturity government bonds, the most common benchmark assets. The European yield is calculated as GDP-weighted average from the EU-11 countries. The money markets are represented by the 3-month LIBOR fixing rates, which should be close enough to monetary policy decisions on the one hand, but should exhibit at the same time sufficient continuous market-driven variation on the other. For the equity markets, I decided to include the Dow Jones Industrial Average and the Euro Stoxx 50 Index. The latter is taken as a certain European summary measure, which indeed differs only negligibly from the main single stock exchanges in Frankfurt or Paris. The exchange rates are measured as the amount of euro (ECU before 1999) respectively US dollar per British pound sterling, see section 2.

Naturally, dealing with daily data brings up the problematic of non-synchronous trading. The time difference between Europe and the US is about six hours, leading to an overlap of trading periods of roughly one half (for example, refer to the discussion in Baur and Jung 2006). Close-to-close data is used nonetheless, since this study emphasises general cross-country spillover effects rather than informational market efficiency. The first half of a European trading day then affects US opening prices, the second half lies parallel to US trading. The other way round, the second half of the lagged US trading day can be taken into account by European opening prices. Consequently, the total (long-run) effects in a structural VAR mirror the correct "balance of power", since they contain all mutual impacts from both parallel and past trading as well as on opening prices.

The sample was set to 01/03/1994 - 12/29/2006, excluding weekends. The starting point guarantees a relatively homogenous estimation period, because events like the German reunification, the crises of the European Monetary System or the foundation of the EU have already taken place in the early 1990s. Later incidents like the euro introduction or the 9/11 terrorist attacks did not prove crucial for the results. In addition, the sample length should be sufficient for picking up signals of time-varying volatility in the EGARCH model (10) and of long-run adjustment in the cointegrating system (4).

Figure 1 provides a graphical overview of the data. Among the interest rates, the bond yields seem to follow a fairly symmetric course. The sample trend is slightly downward sloping, and the troughs and peaks relate to the business cycle course, for example in the economic boom at the turn of the millennium. In the money market rates, the idiosyncratic behaviour of the central banks is more apparent. A striking reduction can be observed during the recession in the first years of the new decade. This same recession also caused the drop in both the Dow Jones and the Euro Stoxx 50 indices. Apart from that period though, the stock prices follow an overall positive trend. The euro graph clearly illustrates the quick devaluation and recovery of the new European currency immediately after its introduction in 1999. In contrast, the dollar depreciated considerably in the last sample years.

For the empirical analysis, the exchange rates and the stock indices are transformed to logarithms multiplied by 100. Taking first differences thus generates continuously compounded asset returns or growth rates in percentage points. On the interest rates, already measured in percentage points and remaining rather low, no such transformation is applied. Tackling the degree of persistence in the data, Table 1 displays the p-values of ADF-tests including constants and centred seasonal dummies. As the hypotheses of non-stationarity can be confirmed, and additionally, the first differences are clearly I(0), all series can be assumed integrated of order one. The calculations in this paper have been done in Gauss 8.0, JMulti 4.14 and EViews 5.0.

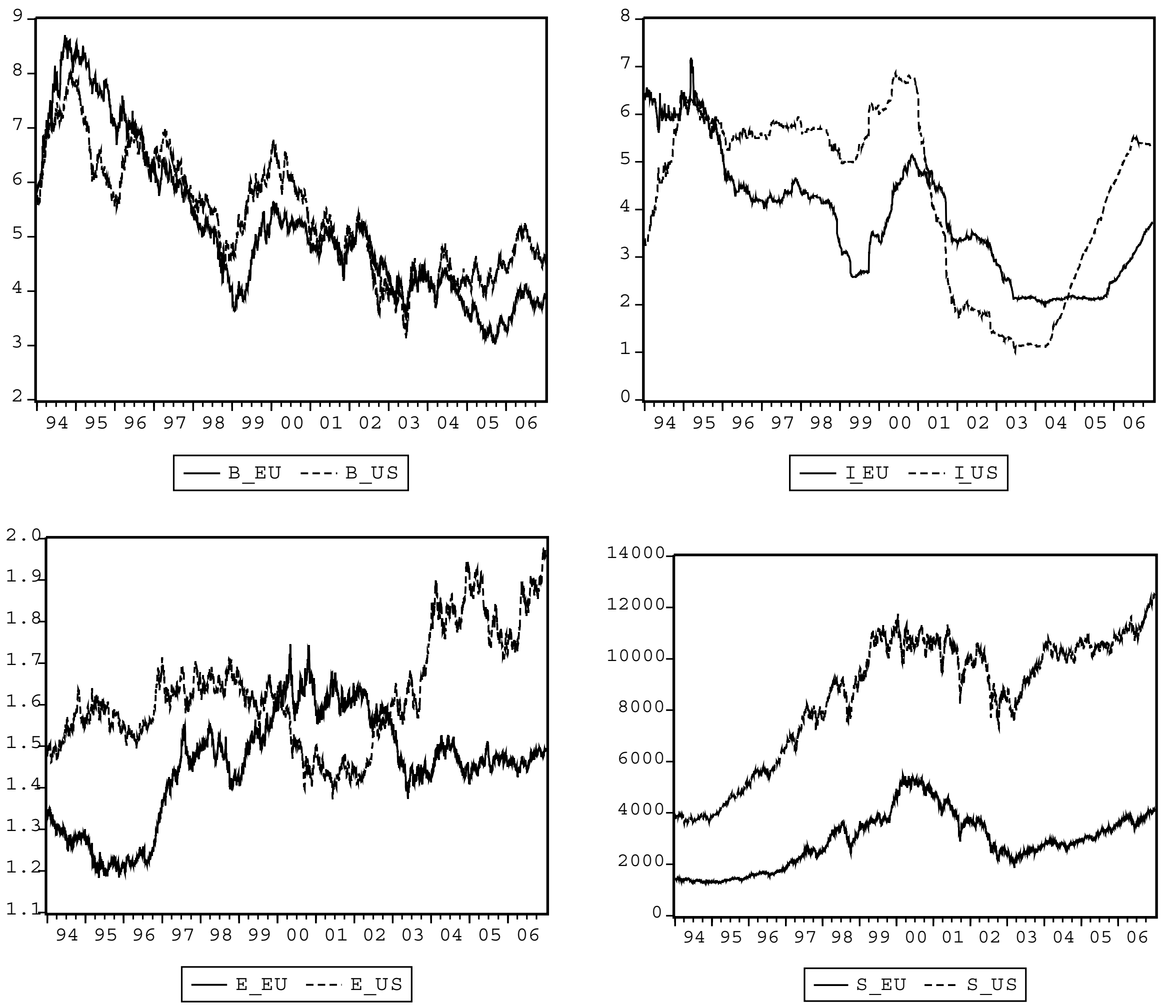

Figure 1: Bond yields (B), LIBORs (I), exchange rates to GBP (E), stock indices (S)

4.2 Specification of the Reduced-Form Models

As a first step, I specify four bivariate reduced-form models as in (5), which include the respective US and European financial variables belonging to the same market. In this, the numbers of lags and of cointegrating relations have to be determined. In order to check the lag lengths chosen by the information criteria, Table 2 presents LM tests with the null hypothesis of no autocorrelation. Since the overall impression is quite satisfying, the serial correlation should be completely captured. To avoid the low p-value for the stock model, the information criteria would have allowed the inclusion of additional lags, but later on these proved insignificant.

Concerning the cointegration properties, Table 3 shows the p-values of the respective trace tests. As could be expected, the exchange rates and stock indices are not cointe-

	b_{eu}	b_{us}	e_{eu}	e_{us}	i_{eu}	i_{us}	s_{eu}	s_{us}
ADF p-values	0.80	0.49	0.54	0.70	0.23	0.84	0.66	0.34
lags	3	0	2	1	1	3	6	0
Deterministics: constant, daily dummies								

Table 1: ADF-tests

	b	e	i	s
LM(1)	0.47	0.67	0.76	0.55
LM(5)	0.11	0.75	0.26	0.01
lags	3	2	5	3

Table 2: p-values of LM-tests for no residual autocorrelation in bivariate models

grated. While the corresponding models are thus specified as VARs in growth rates, I adopt the cointegration assumption for both the short- and the long-term interest rates. For the latter, the p-value of 0.12 does not indicate overwhelming significance, but the cointegration constraints led to very robust and sensible results. The further analysis can logically follow the theoretical UIP implications of equilibrium adjustment and transitory dynamics.

	b	e	i	s
$H_0 : r = 0$	0.12	0.90	0.02	0.80

Table 3: p-values of bivariate trace tests for cointegration

4.3 Financial Volatility Transmission

While the reduced-form models have already been specified, the structural VARs and VECMs can only be estimated in the last step after the identification of the contemporaneous impact matrices. Since this takes place in the EGARCH-procedure, at first the results for the generating processes of the conditional variances shall be presented. While insignificant parameters are sequentially deleted, naturally, there are no such cases on the diagonals of the first two matrices, what shows the pure presence of ARCH effects in the data. The QML standard errors are put in parentheses below the coefficients.

Capital Market

$$\begin{pmatrix}\log h_{eu,t}\\ \log h_{us,t}\end{pmatrix}=\begin{pmatrix}\underset{(0.032)}{-0.099}\\ \underset{(0.022)}{-0.106}\end{pmatrix}+\begin{pmatrix}\underset{(0.002)}{0.995} & 0\\ 0 & \underset{(0.002)}{0.994}\end{pmatrix}\begin{pmatrix}\log h_{eu,t-1}\\ \log h_{us,t-1}\end{pmatrix}+\begin{pmatrix}\underset{(0.023)}{0.085} & 0\\ \underset{(0.013)}{0.045} & \underset{(0.014)}{0.047}\end{pmatrix}\begin{pmatrix}|\tilde{\varepsilon}_{eu,t-1}|\\ |\tilde{\varepsilon}_{us,t-1}|\end{pmatrix}+\begin{pmatrix}0 & 0\\ 0 & \underset{(0.006)}{-0.014}\end{pmatrix}\begin{pmatrix}\tilde{\varepsilon}_{eu,t-1}\\ \tilde{\varepsilon}_{us,t-1}\end{pmatrix}$$

In the capital market, shocks to the European bond yield cause a variance increase on the US side. This effect will as well be found in all remaining markets. It indicates at least a certain signalling function of European financial developments, since in the sense of Ross (1989), volatility spillovers can be given an interpretation as equivalent to information flows. The US variance is particularly boosted by *negative* innovations, which obviously provoke higher activity and insecurity than rising interest rates. One interpretation might be that expectations are driven by speculations on monetary policy loosening or downturns in the business cycle; indeed, in case of a fall in the long rate, the term structure would predict lower future short rates and weaker economic activity. Certain turbulences could then be the consequence of market resettling processes. Positive shocks however might promise a more foreseeable course of monetary policy and macroeconomic development.

Money Market

$$\begin{pmatrix}\log h_{eu,t}\\ \log h_{us,t}\end{pmatrix}=\begin{pmatrix}\underset{(0.115)}{-0.038}\\ \underset{(0.055)}{-0.130}\end{pmatrix}+\begin{pmatrix}\underset{(0.023)}{0.913} & \underset{(0.037)}{0.109}\\ 0 & \underset{(0.004)}{0.991}\end{pmatrix}\begin{pmatrix}\log h_{eu,t-1}\\ \log h_{us,t-1}\end{pmatrix}+\begin{pmatrix}\underset{(0.060)}{0.361} & 0\\ \underset{(0.031)}{0.061} & \underset{(0.031)}{0.060}\end{pmatrix}\begin{pmatrix}|\tilde{\varepsilon}_{eu,t-1}|\\ |\tilde{\varepsilon}_{us,t-1}|\end{pmatrix}+\begin{pmatrix}0 & \underset{(0.059)}{-0.121}\\ \underset{(0.033)}{-0.094} & 0\end{pmatrix}\begin{pmatrix}\tilde{\varepsilon}_{eu,t-1}\\ \tilde{\varepsilon}_{us,t-1}\end{pmatrix}$$

In the money market, two features from above, the variance spillover to the US and the more than proportional impacts of negative shocks reappear. Keeping with the line of argumentation, the asymmetric cross-country effects might be the consequence of lower foreign rates allowing reductions in the domestic ones and signalling the stance of foreign growth performance. The positive cross-GARCH parameter indicates a lasting influence of the US on the Euro rate; evidently, the European economy pays continual attention to any sort of US-side activities.

Stock Market

$$\begin{pmatrix}\log h_{eu,t}\\ \log h_{us,t}\end{pmatrix}=\begin{pmatrix}\underset{(0.019)}{-0.134}\\ \underset{(0.021)}{-0.146}\end{pmatrix}+\begin{pmatrix}\underset{(0.002)}{0.995} & \underset{(0.006)}{-0.013}\\ 0 & \underset{(0.006)}{0.977}\end{pmatrix}\begin{pmatrix}\log h_{eu,t-1}\\ \log h_{us,t-1}\end{pmatrix}+\begin{pmatrix}\underset{(0.018)}{0.112} & \underset{(0.018)}{0.056}\\ \underset{(0.023)}{0.082} & \underset{(0.020)}{0.097}\end{pmatrix}\begin{pmatrix}|\tilde{\varepsilon}_{eu,t-1}|\\ |\tilde{\varepsilon}_{us,t-1}|\end{pmatrix}+\begin{pmatrix}\underset{(0.010)}{-0.044} & \underset{(0.012)}{-0.073}\\ \underset{(0.014)}{-0.051} & \underset{(0.017)}{-0.092}\end{pmatrix}\begin{pmatrix}\tilde{\varepsilon}_{eu,t-1}\\ \tilde{\varepsilon}_{us,t-1}\end{pmatrix}$$

The stock market reveals strong causality-in-variance effects with shocks in both equity indices mutually affecting the respective variances. The asymmetry coefficients represent the well-known leverage-effect, where negative equity shocks have higher variance impacts than positive ones. This phenomenon reaches considerable magnitude especially for shocks on the US side and can as well be found in the cross-country relations. The negative non-diagonal autoregressive coefficient leads to a faster cushioning of the US impact in the Euro Stoxx variance: For instance, a negative shock on the Dow Jones initially drives up European volatility, but in the following periods this is partly compensated for by the opposite effect from the risen US variance.

Foreign Exchange Market

$$\begin{pmatrix}\log h_{eu,t}\\ \log h_{us,t}\end{pmatrix} = \begin{pmatrix}\underset{(0.020)}{-0.067}\\ \underset{(0.021)}{-0.083}\end{pmatrix} + \begin{pmatrix}\underset{(0.002)}{0.995} & 0\\ 0 & \underset{(0.006)}{0.987}\end{pmatrix}\begin{pmatrix}\log h_{eu,t-1}\\ \log h_{us,t-1}\end{pmatrix} + \begin{pmatrix}\underset{(0.021)}{0.076} & 0\\ \underset{(0.014)}{0.033} & \underset{(0.014)}{0.048}\end{pmatrix}\begin{pmatrix}|\tilde{\varepsilon}_{eu,t-1}|\\ |\tilde{\varepsilon}_{us,t-1}|\end{pmatrix} + \begin{pmatrix}0 & 0\\ 0 & 0\end{pmatrix}\begin{pmatrix}\tilde{\varepsilon}_{eu,t-1}\\ \tilde{\varepsilon}_{us,t-1}\end{pmatrix}$$

The foreign exchange model happens to reduce to a diagonal symmetric EGARCH, apart from a small spillover to the US variance.

On balance, the EGARCH results can be summarised as follows: The capital and foreign exchange market variances are mostly independent, the money market reveals asymmetric spillovers of negative shocks, and in the equity market exist the most extensive interactions. All in all, this supports an impression of a calm long-term bond development, interdependent central banks and highly reactive stock exchanges.

Finally, the multivariate EGARCH models shall be checked for sufficiently catching up the heteroscedasticity in the data. The p-values for the ARCH-LM null hypothesis of no remaining ARCH in the standardised disturbances $\tilde{\varepsilon}_{jt}$ in Table 4 confirm the common literature result that GARCH models of orders 1,1 are fairly appropriate for financial markets data. Solely for the European bond and exchange rate equations, several residual outliers evoke the impression that the model might fail in absorbing the time-variation in volatility. At last, all eigenvalues of the autoregressive matrices are smaller than one and therefore meet the stability criterion; the still high persistence is a common feature throughout the GARCH literature.

	b_{eu}	b_{us}	e_{eu}	e_{us}	i_{eu}	i_{us}	s_{eu}	s_{us}
LM(1)	0.00	0.17	0.04	0.96	0.83	0.52	0.51	0.65
LM(5)	0.00	0.12	0.00	0.92	1.00	0.98	0.01	0.62

Table 4: p-values of LM-tests for no residual ARCH

4.4 Financial Markets Leadership

Now I proceed to the core results on the main research topic of financial markets leadership. By maximising the likelihood function (11), besides the EGARCH parameters, estimates of the contemporaneous impacts are obtained. In the following SVAR and SVECM systems, the corresponding A matrices are treated as given, what therefore allows the estimation of the right-hand sides of (4), including the cointegrating relations. The dots at the end of the equations serve as placeholders for the deterministics and residuals. As summary measures, I provide impulse response functions and variance decompositions in the different models. Whilst the former stand for the actual reaction to foreign impulses measured in percentage points (interest rates) or percent (stock indices and exchange rates), the latter provide the proportion of total variance governed by foreign shocks. For the cointegrating models, the calculation of these measures is based on the restricted SVAR representations of the SVECMs; in case of specification in first differences, *accumulated* impulse responses are displayed.

Money Market

$$\begin{pmatrix} 1 & \underset{(0.046)}{-0.119} \\ \underset{(0.035)}{-0.153} & 1 \end{pmatrix} \begin{pmatrix} \Delta i_{eu,t} \\ \Delta i_{us,t} \end{pmatrix} = \begin{pmatrix} \underset{(0.001)}{-0.002} \\ 0 \end{pmatrix} \left(i_{eu,t-1} - \underset{(0.077)}{0.769} i_{us,t-1} \right) + \begin{pmatrix} \underset{(0.007)}{0.075} & 0 \\ 0 & \underset{(0.009)}{0.123} \end{pmatrix} \begin{pmatrix} \Delta i_{eu,t-1} \\ \Delta i_{us,t-1} \end{pmatrix} + \begin{pmatrix} \underset{(0.010)}{-0.040} & 0 \\ 0 & \underset{(0.011)}{0.081} \end{pmatrix} \begin{pmatrix} \Delta i_{eu,t-2} \\ \Delta i_{us,t-2} \end{pmatrix}$$

$$+ \begin{pmatrix} \underset{(0.010)}{0.058} & 0 \\ 0 & \underset{(0.014)}{0.046} \end{pmatrix} \begin{pmatrix} \Delta i_{eu,t-3} \\ \Delta i_{us,t-3} \end{pmatrix} + \begin{pmatrix} \underset{(0.008)}{0.034} & 0 \\ 0 & 0 \end{pmatrix} \begin{pmatrix} \Delta i_{eu,t-4} \\ \Delta i_{us,t-4} \end{pmatrix} + \begin{pmatrix} \underset{(0.006)}{0.023} & 0 \\ \underset{(0.011)}{-0.040} & \underset{(0.012)}{0.029} \end{pmatrix} \begin{pmatrix} \Delta i_{eu,t-5} \\ \Delta i_{us,t-5} \end{pmatrix} + \begin{pmatrix} \underset{(0.009)}{-0.031} & \underset{(0.014)}{0.031} \\ 0 & \underset{(0.012)}{0.035} \end{pmatrix} \begin{pmatrix} \Delta i_{eu,t-6} \\ \Delta i_{us,t-6} \end{pmatrix} + \dots$$

The money market model is characterised by the presence of a cointegrating relation, based on the test in Table 3. Restricting β' to $(1, -1)$, as implied by UIP theory, is rejected by an LR test (p-value=0.003), but the coefficient $\beta_2 = -0.769$ is still in a sensible range. The constant is left out due to insignificance, signalling the absence of permanent risk premia. The influence from the US on Euroland is clearly stronger than in the reverse direction; the situation might still be balanced in the short-run, but due to insignificance of the cointegrating term in the US equation, the equilibrium adjustment works exclusively through the European rate, as it is visualised by the impulse responses and variance decompositions in Figure 2.

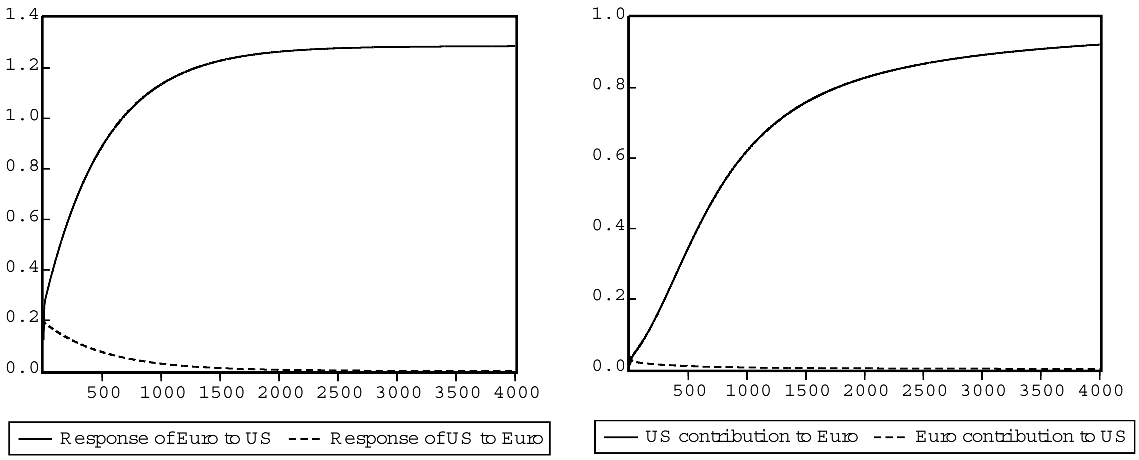

Figure 2: Impulse responses and variance decomposition in the money market

Capital Market

$$\begin{pmatrix} 1 & \underset{(0.006)}{-0.200} \\ \underset{(0.009)}{-0.238} & 1 \end{pmatrix} \begin{pmatrix} \Delta b_{eu,t} \\ \Delta b_{us,t} \end{pmatrix} = \begin{pmatrix} \underset{(0.001)}{-0.002} \\ \underset{(0.001)}{0.002} \end{pmatrix} \begin{pmatrix} b_{eu,t-1} - \underset{(0.295)}{1.595} b_{us,t-1} + \underset{(1.744)}{3.518} \end{pmatrix} + \begin{pmatrix} \underset{(0.013)}{-0.119} & \underset{(0.012)}{0.200} \\ \underset{(0.019)}{0.064} & \underset{(0.012)}{-0.045} \end{pmatrix} \begin{pmatrix} \Delta b_{eu,t-1} \\ \Delta b_{us,t-1} \end{pmatrix}$$

$$+ \begin{pmatrix} 0 & \underset{(0.011)}{0.046} \\ 0 & 0 \end{pmatrix} \begin{pmatrix} \Delta b_{eu,t-2} \\ \Delta b_{us,t-2} \end{pmatrix} + \begin{pmatrix} \underset{(0.012)}{-0.026} & 0 \\ \underset{(0.019)}{-0.043} & 0 \end{pmatrix} \begin{pmatrix} \Delta b_{eu,t-3} \\ \Delta b_{us,t-3} \end{pmatrix} + \dots$$

In contrast, the impulse responses and variance contributions are far more balanced in the capital market, see Figure 3. Evidently, the long end of the yield curve is characterised by European feedback effects accounting for more than a half of the US influence, while at the short end, the Federal Reserve clearly held on to the leading position. The constant in the cointegrating relation suggests a permanent risk premium burdened on the US bonds (even though the cointegrating parameter different from -1 has to be taken into regard), and the UIP restriction $\beta' = (1, -1)$ is borderline rejected with a p-value of 0.04.

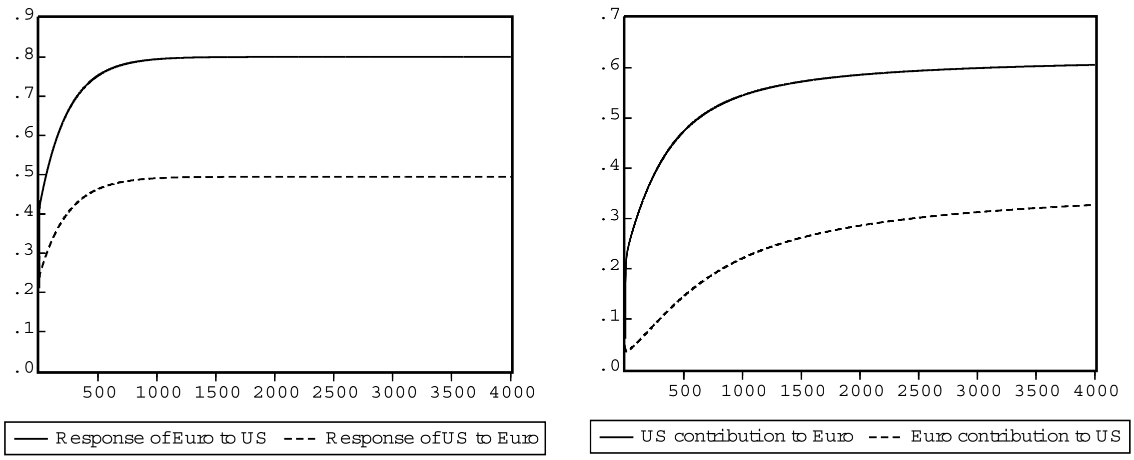

Figure 3: Impulse responses and variance decomposition in the bond market

Stock Market

$$\begin{pmatrix} 1 & \underset{(0.006)}{-0.284} \\ \underset{(0.004)}{-0.230} & 1 \end{pmatrix}\begin{pmatrix} \Delta s_{eu,t} \\ \Delta s_{us,t} \end{pmatrix} = \begin{pmatrix} \underset{(0.013)}{-0.181} & \underset{(0.016)}{0.465} \\ \underset{(0.009)}{0.079} & \underset{(0.012)}{-0.134} \end{pmatrix}\begin{pmatrix} \Delta s_{eu,t-1} \\ \Delta s_{us,t-1} \end{pmatrix} + \begin{pmatrix} \underset{(0.012)}{-0.076} & \underset{(0.018)}{0.089} \\ \underset{(0.011)}{0.024} & \underset{(0.014)}{-0.063} \end{pmatrix}\begin{pmatrix} \Delta s_{eu,t-2} \\ \Delta s_{us,t-2} \end{pmatrix} + \begin{pmatrix} \underset{(0.011)}{-0.068} & \underset{(0.017)}{0.066} \\ 0 & \underset{(0.012)}{-0.027} \end{pmatrix}\begin{pmatrix} \Delta s_{eu,t-3} \\ \Delta s_{us,t-3} \end{pmatrix} + \dots$$

In the stock market, the European accumulated impulse response (Figure 4 left panel) more than doubles the reverse effect on US returns. However, the gap between the variance contributions (right panel) is far smaller, showing relatively comparable strength in governing the total stock variability. While the bulk of the Euro Stoxx impacts is limited to the contemporaneous period, at least one further day contributes substantially to the total Dow Jones effect; such a constellation can be seen as typical for non-synchronous trading as discussed in section 4.1. As another feature, note that in the stock market, the impulse adjustment is completed within few days, while the long-run relations between the levels of the interest rates generated far more sluggish processes.

Foreign Exchange Market

$$\begin{pmatrix} 1 & \underset{(0.005)}{-0.247} \\ \underset{(0.004)}{-0.223} & 1 \end{pmatrix}\begin{pmatrix} \Delta e_{eu,t} \\ \Delta e_{us,t} \end{pmatrix} = \begin{pmatrix} \underset{(0.016)}{-0.170} & \underset{(0.017)}{-0.213} \\ 0 & \underset{(0.015)}{0.035} \end{pmatrix}\begin{pmatrix} \Delta e_{eu,t-1} \\ \Delta e_{us,t-1} \end{pmatrix} + \begin{pmatrix} \underset{(0.014)}{-0.054} & 0 \\ 0 & 0 \end{pmatrix}\begin{pmatrix} \Delta e_{eu,t-2} \\ \Delta e_{us,t-2} \end{pmatrix} + \dots$$

Similar to the stocks, the fluctuations in the foreign exchange market are limited to a duration of one, or at most two days. The variance decomposition in Figure 5 (right

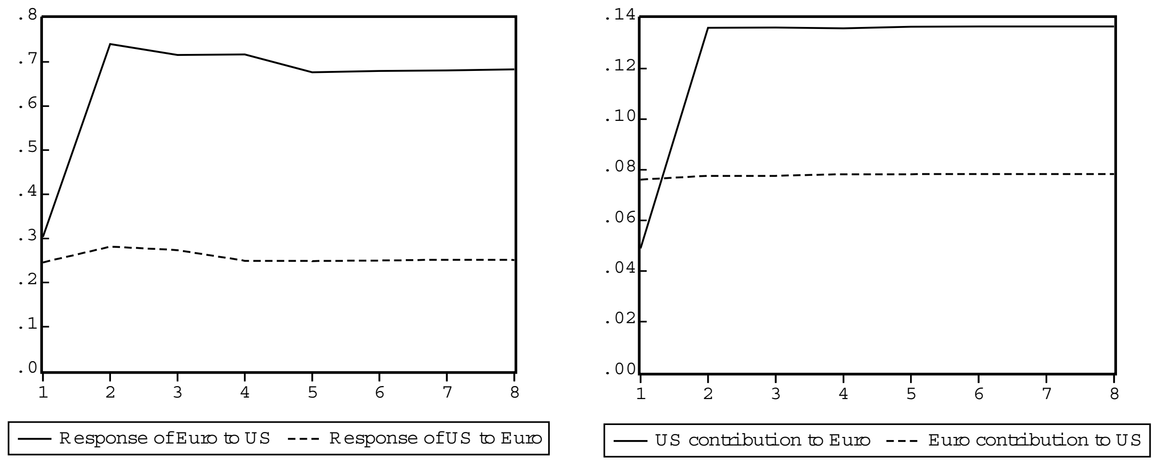

Figure 4: Impulse responses and variance decomposition in the stock market

panel) resembles the one from the stock market. The drop in the European accumulated impulse response function (left panel) is triggered by the one-day lagged appreciating effect on the euro stemming from the unit depreciating shock in the US dollar.

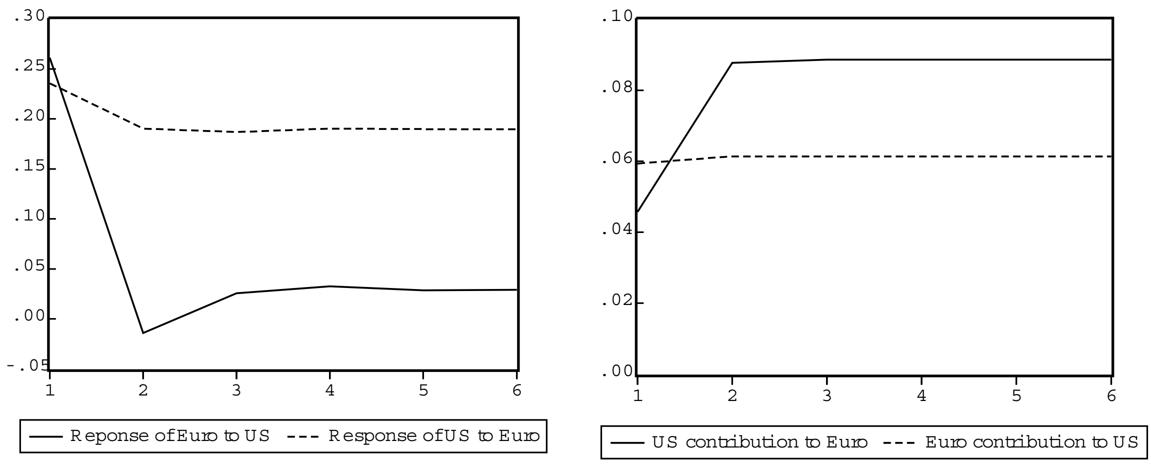

Figure 5: Impulse responses and variance decomposition in the foreign exchange market

One important fact should be noted in conclusion: Evidently, the contemporaneous reactions already deliver substantial parts of the total effects. The employed methodology is therefore crucial for uncovering the real dimensions of the international interdependence. Even more, this has to be pronounced for the bond market: Not considering the contemporaneous impacts in this case, that is specifying the VECM in reduced form, leaves the US adjustment coefficient (α_2) clearly insignificant. In such a situation, the European interest rate would be governed by the need to re-equilibrate any deviations from the UIP condition, and causalities on the US economy would shrink to transitory influences in the short-run dynamics. In the end, the overall impression would be that of totally dependent European financial markets without any feedback and degrees of freedom.

Collecting the important facts from the analysis produces the following outcome: First of all, the initial hypothesis of US predominance cannot be rejected. Second, in sum, the European influence is not negligible. Third, this second finding depends crucially on an adequate modelling in presence of extremely short time spans within which impulses are processed in highly developed financial markets.

5 Concluding Summary

Analysing the international balance of economic power attracts the interest of many different kinds of econometric research. This special paper examines the interdependence between financial markets of Euroland and the USA, the two superpowers of the industrialised world. Thereby, the focus lies on the markets for capital, currencies, money and stocks.

The first moments of the respective variables are modelled in the VAR form, or VECM in case of cointegration. These systems are complemented by bivariate EGARCH models for the structural residuals, adding two features to the analysis: First, causality-in-variance effects are assessed, which play an important role in a financial markets context, and second, the simultaneous impacts in conditional mean can be estimated in the fashion of identification through heteroscedasticity.

The empirical results confirm on the one hand the generally accepted leading role of the USA. Though, on the other hand, non-trivial repercussions originating from the European side can be quantified. While the market for short-term money however is largely dominated by US influence, the bond, equity and foreign exchange markets tend to more symmetric behaviour. Interestingly, without applying the special econometric identification procedure, according results would erroneously indicate nearly complete dependence of the euro zone on US developments. The conditional variance analysis reveals strong volatility spillovers between the Dow Jones and Euro Stoxx indices as well as asymmetric impacts of interest rate reductions in the money market. Contrarily, shocks in the long-term bond yields and the exchange rates scarcely translate into significant cross-country variability effects.

At the final count, there lasts an ambivalent picture of political reality: The US economic and monetary policy *can* act in the knowledge of being the measure of all things, but not without limitations. Despite the US predominance, the EMU has the potential to maintain a certain, even if by far not complete autonomy in the international economy. Put it the other way round, the unified Europe has to take into account foreign consequences of its own actions as well as substantial impacts from abroad, as it is the typical situation of one of the "large countries".

In future research, one methodological extension might be found in allowing for contemporaneous correlation in the structural equations. Whereas the presently considered superpowers US and Europe might still cover the vast majority of financial information, accounting for common factors unrelated to direct spillovers between the included endogenous variables would contribute to generalising the analysis. In the same vein, intraday data could be used to shed light on the contributions of truly contemporaneous spillovers and overnight effects to overall shock transmission.

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