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Interactions of U.S., German and Greek Bond Markets in Times of Financial Crisis. A Bayesian Analysis of Exogeneity in the VAR-SV Model

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Abstract

This paper examines interactions of the U.S., German and Greek bond markets in times of financial crisis. Specifically, the connections among daily and weekly growth rates of the 10-year government bond yields of the U.S., Germany and Greece from July 13, 2006 to June 28, 2012 are considered and an empirical illustration of those, based on the vector autoregressive (VAR) model with stochastic volatility (SV) disturbances, is provided. Finally, sufficient weak and strong exogeneity conditions in the VAR-SV models are tested. Since the strong exogeneity hypothesis of the 10-year US bond yields' growth rates has not been rejected by the data, they can be predicted from the marginal model only.

JEL classification: G15, G01, C11, C32.

Keywords: international bond markets, financial crisis, exogeneity, stochastic volatility

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Anita Modrzejewska, Anna Pajor

1. Introduction

The financial crisis that originated in the crash of the U.S. housing market in mid-2007 followed by the Lehman Brothers bankruptcy in 2008 has reversed the convergence trend in the Eurozone government bond market, which was clearly recognizable after the introduction of the common currency in 1999. The onset of the sovereign debt crisis in Europe at the end of 2009, closely related to the worsening economic situation in Greece, has reinforced the tendency towards a cross-country bond yields differentiation leading to a greater market fragmentation. As a result, European government bond spreads remained at higher levels in 2012 in comparison to the pre-crisis years resembling those of the pre-Euro introduction period.

One of the reasons for differentiated pricing of sovereign debt securities across European countries was a stronger impact of domestic factors, highly neglected by the bond market participants in the years 1999-2007. The influence of country-specific factors, such as national fiscal conditions and macroeconomic imbalances (credit risk) together with market liquidity (liquidity risk) on government bond yields, was confirmed by Barrios et al. (2009). Nevertheless, the major role in explaining the differentials of Euro area sovereign bond yields was played by international factors, especially investors' risk aversion (Barrios et al., 2009). In times of financial instability, investors typically engage in a 'flight to quality' (and a 'flight to liquidity'[1]) i.e. they substitute safe assets for the more risky ones. Since government bonds are perceived as 'safe havens', investors substitute bonds for stocks, and in consequence, bond and stock market returns become negatively correlated (Kim et al., 2006). Similarly, as in terms of credit quality and liquidity, German Bunds have been considered the 'safest haven' among European sovereign bonds (Barrios et al., 2009) and the Greek ones the least safe, yields on both should be negatively correlated during a financial crisis.

However, the developments in the Euro area government bond markets cannot merely be assessed from the regional perspective. Discernible cross-border co-movements of bond yields signal the presence of common driving factors at the global level as well. In particular, there is

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[1] Risky assets are usually less liquid.

much empirical evidence supporting a substantial interdependence between the U.S. and European sovereign bond markets (Goldberg and Leonard, 2003; Engsted and Tanggaard, 2005, Ehrmann et al., 2005; Andersson et al., 2006; Laopodis, 2008). In times of financial stability, government bond yields in these markets tend to move together displaying a high degree of positive correlation. This holds true especially for the U.S. Treasuries and the German Bunds (Goldberg and Leonard, 2003; Engsted and Tanggaard, 2005), which in the global sovereign debt market are deemed to be benchmark securities (U.S. Treasuries serve as international benchmarks, and German Bunds – as regional ones). Considering the facts that the current financial crisis began in the U.S. and the U.S. Treasuries as global benchmark securities assign prices to other assets, the question arises whether, and if so, to what extent the changes in Euro area government bond yields were affected by the changes in yields of U.S. Treasuries.

On the other hand, downgrading of the Greek public debt to junk level (BBB-) by rating agency Fitch in April 2010 has threatened the stability of the whole Euro area, bringing the bond yields of larger European economies under pressure in 2011 (ECB, 2012). On November 23, 2011, Germany, considered a stalwart pillar of the Eurozone economy, failed to get bids for 35% of the 10-year government bonds offered for the auction sale, which made it clear that even German Bunds are not immune to the ongoing sovereign debt crisis in Europe (http://www.bloomberg.com/news/2011-11-23/germany-fails-to-receive-bids-for-35-of-10-year-bunds-offered-at-auction.html). As the German bond market seems to have been affected by the dysfunction of the Greek bond market since the end of April 2010, another question addressed in this paper is about the potential impact of the European bond markets on that of the U.S. during a financial turmoil. It's worth noting that earlier the same year, on August 5, 2011, Standard & Poor's downgraded the U.S. sovereign debt from AAA to AA+, which increased speculation about a potential loss of the benchmark status of U.S. Treasuries (IMF, 2012). Hence, an analysis of connections between the U.S. and German government bonds, acting as close substitutes in the global bond market, seems to be especially interesting.

Thus the aim of this paper is to study interactions among selected government bond markets, namely that of the U.S. (the global benchmark), Germany (the regional benchmark) and Greece (the country most severely affected by the sovereign-debt crisis) in the presence of high levels of market uncertainty. In particular, we examine the strength and direction (the sign) of bilateral linkages among yields in the sovereign debt markets of those countries, i.e. we consider how the past changes in 10-year government bond yields in one country have determined the present changes in 10-year government bond yields in another country. Additionally, we test the sufficient weak and strong exogeneity conditions in the vector autoregressive (VAR) model whose disturbances follow the stochastic volatility (SV) process. The Bayesian concept of exogeneity adopted in this paper

was proposed by Florens and Mouchart (Florens and Mouchart, 1977, 1980, 1982; Florens et al., 1990; Mouchart et al,. 2007; Osiewalski and Steel, 1996), and then extended by Pajor (2011) to the class of models with latent variables (in which a subset of parameters and latent variables is of interest).

The remainder of this paper is as follows: section 2 briefly reviews empirical studies which fit the line of our research, section 3 introduces the Bayesian econometric VAR-SV model together with the formal definition of exogeneity, section 4 discusses the dataset and empirical results, and finally, section 5 concludes by providing the key findings.

2. Literature review

The sovereign debt crisis in Europe shifted interest to government bond markets again, thus raising the problem of their international interdependence in the context of transmission of financial shocks. Previous analyses of interactions among European sovereign bond markets have seldom been carried out in isolation from the global market. For instance, Sosvilla-Rivero and Morales-Zumaquero (2011) studied the volatility of the daily 10-year sovereign bond yields for 11 EMU countries during the 2001-2010 decade, using a C-GARCH model and cluster analysis. Their results suggest that for all countries and periods, the transitory (short-run) component of bond-yield volatility was less important than the permanent (long-run) one. In other words, shocks to underlying economic fundamentals had a greater impact on volatility of government bond yields than temporary shifts in bond markets. Moreover, the correlation analyses and Granger causality tests between transitory and permanent volatilities of sovereign bond yields confirmed the existence of two groups of EMU countries, characterized by different degree of credibility ascribed to the announcements made by policymakers and by various positions regarding the stability of public finance. It came as no surprise that Germany and Greece belonged to the distinct groups, i.e. to the EMU-core and -periphery respectively, which means that there was a rather weak relation between volatilities of their bond yields in the years 2001-2010.

Nevertheless, the empirical study of Sosvilla-Rivero and Morales-Zumaquero (2011) did not cover the actual period of sovereign debt crisis in Europe, and in times of financial instability, markets behave differently. They are more volatile and become more highly correlated (Chordia et al., 2001). Therefore, during periods of financial stress the external factor plays a more important role than the domestic one (Clare and Lekkos, 2000). Claeys and Vašiček (2012), who measured the bilateral spillovers between 16 EU government bond markets since May 2000 up to February 2012, using the forecast-error variance decompositions from a VAR model on sovereign bond yield spreads relative to the German 10-year bond yield, indicated that 59% of the variations in sovereign bond spreads could be explained by shocks to other European countries, i.e. the regional factor, and

only the remaining 41% of all movements were determined by a purely domestic factor. The authors also suggest that spillover has substantially increased since 2007, and it was of greater significance than domestic factors to all EMU countries due to the influence of the common factor as well as bilateral linkages. In case of Greece, however, domestic dynamics were slightly more important than the international ones.

Furthermore, Claeys and Vašiček (2012) examined the spillover of fiscal problems in Greece to other European bond markets. They observed that the influence of changes in Greek sovereign bond spreads on other markets was varying significantly over time, and it was quite different across four identified groups of countries. The most affected were the bond markets of the PIIS (Portugal, Ireland, Italy and Spain) and those of the core EMU countries (Austria, Belgium, France, Finland and the Netherlands), whereas the bond markets of the CEE (the Czech Republic, Hungary and Poland) and those of the non-EMU countries (Denmark, Sweden and the UK) remained barely influenced. Since May 2010 the spillover to the European bond markets was magnified by the crisis. The structure of the spillover, however, did not change. The time-varying effect of shocks was also measured in the opposite direction, i.e. the influence of other European bond markets on the Greek one was calculated. The overall effect was stable, and again, there were stronger links between the core EMU, the PIIS countries, and Greece. It is worth noting that in both cases, Greece contributed more to spread movements in other markets than it received from, so the influence of the Greek sovereign bond market on the other markets was more powerful than the reverse effect.

As mentioned in the other paragraph, analyses of interactions among European bond markets have usually been done within the global context, i.e. next to the influence of the domestic and regional factor also the global factor has been taken into consideration. Such an approach was adopted by Ch. Christiansen (2003) who studied volatility spillovers from the US and aggregate European bond markets into 6 EMU (Belgium, France, Germany, Italy, the Netherlands, and Spain) and 3 non-EMU countries (Denmark, Sweden, and the UK), using a GARCH volatility-spillover model with the weekly data of total return government bond indices for the period from January 6, 1988 to November 27, 2002. She measured three volatility effects: the global, regional and local one. The regional effect followed by the local one appeared to be the most important for the EMU countries and Denmark. Additionally, the introduction of common currency has strengthened the European and weakened the US volatility-spillover effect for EMU bond markets. In case of non-EMU countries the local and global effects were of greater importance than the regional one which turned out to be rather weak.

J. M. C. da Costa et al. (2004), who followed the research method applied by Christiansen, examined the volatility-spillover effects from the U.S. and Germany to EMU countries with the weekly data of return indices for 7 and 10-year government bond markets from October 1993 to

May 2004. They proved, however, that most European bond markets were strongly affected by volatility-spillover effects, coming both from the global (the U.S.) and regional (German) markets. Moreover, the spillover effects have increased significantly after the introduction of the common currency, whereas the local effect has declined in importance after 1999.

On the other hand, the findings of P. Abad et al. (2009) are consistent with the results of Christiansen, although a different research method was used. P. Abad et al. adopted Bekaert and Harvey's CAPM-based model (1995) to analyse the behaviour of 10-year sovereign bond returns of 13 European countries (Greece and Luxembourg were not included into the sample) with a weekly dataset covering the period from January 1999 to June 2008. The adoption of this model also allowed them to separate each individual country's government bond return into three effects: a local (own country), a regional (Eurozone), and a global (the U.S.) effect. Significant distinctions between EMU and non-EMU countries were recognized. The degree of integration with the US and German bond markets varied clearly between these two groups of countries. The authors indicated that the EMU countries, sharing a single monetary policy, were more vulnerable to the influence of regional risk factors and less vulnerable to the global risk factors. Their results suggest that the EMU bond markets were only partially integrated with the German market because of differences in their market liquidity or default risk. In contrast, the non-EMU countries presented higher vulnerability to external risk factors i.e. their bond markets were more affected by the U.S. market.

Another line of empirical research on interactions of European government bond markets puts a special emphasis on the relationship between the U.S. and Germany. The interdependence between these two markets has been usually investigated in the context of a high level of their co-movements. L. Goldberg and D. Leonard (2003) used 30 months of hourly yield data for the on-the-run U.S. and German 2- and 10-year notes from January 3, 2000, to June 28, 2002. The U.S. economic news was found to have a direct and large influence on German yields within an hour of its release. The effect in the opposite direction, however, appeared to be far less influential. According to T. Engsted and C. Tanggaard (2005), who used a VAR model for monthly data over the period of 1975-2003 to decompose the U.S. and German unexpected bond returns into 3 news components, news about: future inflation, future real interest rates and future excess bond returns, the main driving force behind the co-movement of the U.S. and German bond markets was inflation news coming from the U.S. Furthermore, they proved that there were important spillover effects from the U.S. to the German bond market, but only limited ones from Germany to the U.S., and that these two markets were closely linked together, since one-month excess bond returns in the U.S. and Germany showed a contemporaneous positive correlation of 0,54. The significant impact of the U.S. announcements on German government bond returns was also confirmed by M. Andersson et al. (2006), who examined the effects of macroeconomic data releases and the ECB's monetary

policy statements on the German long-term bond market segment of the yield curve in the period from January 1999 to December 2005. Their results suggest that the U.S. macroeconomic releases exert a stronger influence on German bond markets than the Euro area and domestic news, and the strength of those reactions has increased over the period considered.

Interactions among international government bond markets have also been studied across various classes of financial assets (currencies, stocks, and bonds etc.). Despite the fact that this kind of approach, similarly to the previous paragraph, goes beyond the scope of this research, it is worth noting that in times of financial crisis, shocks emanating from the U.S. asset markets are transmitted worldwide and in a negative direction within-market (as in case of equity and bond markets) as well as cross-market from the U.S. equity to bond market, whereas the Euro area shocks mainly have a negative effect on bond markets and are largely confined to spillovers to the bond markets of advanced economies. J. Beirne and J. Gieck (2012) have come to such conclusions, while investigating over 60 economies in periods of financial stress between 1998 and 2011 with the use of a GVAR model. The dominance of the U.S. as the main driver of global financial markets was also proved by M. Ehrmann et al. (2005). They applied a multifactor model to analyse daily returns of a 16-year period of 1998-2004 for 7 asset prices: short-term interest rates, bond yields, equity market returns and the exchange rate in the U.S. and the Euro area. Their results suggest that the U.S. financial markets explained, on average, more than 25% of movements in the Eurozone financial markets, while the Euro area markets were responsible only for 8% of the U.S. price changes. The international transmission of shocks was strengthened in times of recession.

The investigations hitherto conducted on interactions among international bond markets allow us to assume a stronger relationship between the U.S. and Germany and between Germany and Greece than between the U.S. and Greece. The direction (the sign) of bilateral linkages between each pair of these countries will determine whether the crisis stimulated a 'flight to quality' phenomenon.

This paper contributes to the existing literature in a number of ways. Firstly, it examines connections among the selected government bond markets in times of financial crisis. Secondly, it employs a sophisticated econometric model to investigate the relationship of exogeneity between international bond markets. Thirdly, it uses the Bayesian framework to analyse these linkages, and to our knowledge, none of the researchers has so far investigated the interactions of international bond markets with use of a Bayesian analysis.

3. Methodological framework
3.1. Bayesian econometric VAR-SV model

Let $x_{t,j}$ denote the 10-year bond yields of country j at time t for $j = 1, 2, 3$ and $t = 1, 2, ..., T$. The vector of growth rates $y_t = (y_{t,1}, y_{t,2}, y_{t,3})'$, each defined by the formula $y_{t,j} = 100 \ln (x_{t,j}/x_{t-1,j})$, is modelled using the basic VAR(1) framework:

$$y_t - \delta = R(y_{t-1} - \delta) + \xi_t, \tag{1}$$

where R is a 3×3 matrix of the autoregressive coefficients, δ is a 3×1 mean vector, $\{\xi_t\}$ is a trivariate SV process. More specifically:

$$\begin{bmatrix} y_{t,1} \\ y_{t,2} \\ y_{t,3} \end{bmatrix} - \begin{bmatrix} \delta_1 \\ \delta_2 \\ \delta_3 \end{bmatrix} = \begin{bmatrix} r_{11} & r_{12} & r_{13} \\ r_{21} & r_{22} & r_{23} \\ r_{31} & r_{32} & r_{33} \end{bmatrix} \left(\begin{bmatrix} y_{t-1,1} \\ y_{t-1,2} \\ y_{t-1,3} \end{bmatrix} - \begin{bmatrix} \delta_1 \\ \delta_2 \\ \delta_3 \end{bmatrix} \right) + \begin{bmatrix} \xi_{t,1} \\ \xi_{t,2} \\ \xi_{t,3} \end{bmatrix}, \quad t = 1, ..., T, \tag{2}$$

We assume that conditionally on the latent variable vector, (h_t, introduced later) ξ_t follows a trivariate Gaussian distribution with mean vector $0_{[3\times1]}$ and covariance matrix Σ_t, i.e. $\xi_t | h_t \sim N(0_{[3\times1]}, \Sigma_t)$, $t = 1, 2, ..., T$. We employ the specification of the conditional covariance matrix as in Tsay (2002). Thus the Cholesky decomposition is used:

$$\Sigma_t = L_t G_t L_t', \tag{3}$$

where L_t is a lower triangular matrix with unitary diagonal elements, and G_t is a diagonal matrix with positive diagonal elements:

$$L_t = \begin{bmatrix} 1 & 0 & 0 \\ q_{t,21} & 1 & 0 \\ q_{t,31} & q_{t,32} & 1 \end{bmatrix}, \quad G_t = \begin{bmatrix} q_{t,11} & 0 & 0 \\ 0 & q_{t,22} & 0 \\ 0 & 0 & q_{t,33} \end{bmatrix}.$$

Hence, we have

$$\Sigma_t = \begin{bmatrix} q_{t,11} & q_{t,11}q_{t,21} & q_{t,11}q_{t,31} \\ q_{t,11}q_{t,21} & q_{t,11}q_{t,21}^2 + q_{t,22} & q_{t,11}q_{t,21}q_{t,31} + q_{t,22}q_{t,32} \\ q_{t,11}q_{t,31} & q_{t,11}q_{t,21}q_{t,31} + q_{t,22}q_{t,32} & q_{t,11}q_{t,31}^2 + q_{t,22}q_{t,32}^2 + q_{t,33} \end{bmatrix}, \tag{4}$$

where $\{q_{t,ij}\}$, and $\{\ln q_{t,jj}\}$ ($i, j = 1, 2, 3, i > j$), as in the univariate SV, are standard univariate autoregressive processes of order one, namely:

$$\ln q_{t,jj} - \gamma_{jj} = \phi_{jj}(\ln q_{t-1,jj} - \gamma_{jj}) + \sigma_{jj}\eta_{t,jj}, j = 1, 2, 3,$$

$$q_{t,ij} - \gamma_{ij} = \phi_{ij}(q_{t-1,ij} - \gamma_{ij}) + \sigma_{ij}\eta_{t,ij}, j, i \in \{1, 2, 3\}, i > j,$$

$\eta_t = (\eta_{t,11} \ \eta_{t,22}, \ \eta_{t,33}, \ \eta_{t,21}, \ \eta_{t,31}, \ \eta_{t,32})'$, $\{\eta_t\} \sim iiN_6(0_{[6\times1]}, I_6)$, $t \in \{1, ..., T\}$,

$h_t = (q_{t,11}, q_{t,22}, q_{t,33}, q_{t,21}, q_{t,31}, q_{t,32})'$.

It can easily be shown that if the absolute values of ϕ_{ij} are less than one, the SV process is a vector white noise (see Pajor 2005a). The conditional distribution of y_t (given the past of the process Y_{t-1}, the parameters and the latent variable vector, h_t) is a trivariate Normal with mean vector $\mu_t = \delta + R(y_{t-1} - \delta)$ and covariance matrix Σ_t: $(y_{t,1}, y_{t,2}, y_{t,3})' | \delta, R, h_t, Y_{t-1} \sim N_3(\mu_t, \Sigma_t)$.

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3.2. Exogeneity in the VAR-SV model

To formally test the sufficient conditions for the strong exogeneity of the growth rates of 10-year U.S. bond yields we use the Lindley type test in the VAR-SV model for selected countries: the U.S., Germany and Greece.

Following Pajor (2011), we partition y_t and Σ_t, conformably, into:

$$y_t = \begin{bmatrix} z_t \\ w_t \end{bmatrix} \text{ and } \Sigma_t = \begin{bmatrix} \Sigma_{t,11} & \Sigma_{t,12} \\ \Sigma_{t,21} & \Sigma_{t,22} \end{bmatrix}, \tag{5}$$

where z_t is a scalar: $z_t = y_{t,1}$, and w_t has two elements: $w_t = (y_{t,2}, y_{t,3})'$. After partitioning, equation (2) becomes:

$$\begin{bmatrix} z_t \\ w_t \end{bmatrix} = \begin{bmatrix} \delta_1 \\ \delta_{-1} \end{bmatrix} + \begin{bmatrix} R_{11} & R_{12} \\ R_{21} & R_{22} \end{bmatrix} \left(\begin{bmatrix} z_{t-1} \\ w_{t-1} \end{bmatrix} - \begin{bmatrix} \delta_1 \\ \delta_{-1} \end{bmatrix} \right) + \begin{bmatrix} \varepsilon_{t,1} \\ \varepsilon_{t,2} \end{bmatrix}, \tag{6}$$

where $\delta_{-1} = (\delta_2, \delta_3)'$, $\varepsilon_{2,t} = (\xi_{t,2}, \xi_{t,3})'$.

The VAR form in (6) can be reparameterized as follows:

$$z_t - \delta_1 = R_{11}(z_{t-1} - \delta_1) + R_{12}(w_{t-1} - \delta_{-1}) + \varepsilon_{t,1} \tag{7}$$

$$w_t - \delta_{-1} = (R_{22} - D_t R_{12})(w_{t-1} - \delta_{-1}) + (R_{21} - D_t R_{11})(z_{t-1} - \delta_1) + D_t(z_t - \delta_1) + (\varepsilon_{2,t} - D_t \varepsilon_{t,1}), \tag{8}$$

where

$$\begin{bmatrix} \varepsilon_{t,1} \\ \varepsilon_{t,2} - D_t \varepsilon_{t,1} \end{bmatrix} | h_t \sim N_3 \left(0, \begin{bmatrix} \Sigma_{t,11} & 0 \\ 0 & \Sigma_{t,22.1} \end{bmatrix} \right), \quad \Sigma_{t,22.1} = \Sigma_{t,22} - \Sigma_{t,21}\Sigma_{t,11}^{-1}\Sigma_{t,12},$$

$$D_t = \Sigma_{t,21}\Sigma_{t,11}^{-1} = \begin{bmatrix} q_{t,21} \\ q_{t,31} \end{bmatrix}, \quad \Sigma_{11,t} = [q_{t,11}], \quad \Sigma_{t,22.1} = \begin{bmatrix} q_{t,22} & q_{t,22}q_{t,32} \\ q_{t,22}q_{t,32} & q_{t,33} + q_{t,22}q_{t,32}^2 \end{bmatrix}.$$

The latent variables and parameters of the conditional model (8) are

$$(vec(H_T^{1,w})', \theta_1')' = ((R_{21} - D_1 R_{11})', ..., (R_{21} - D_T R_{11})', vec(R_{22} - D_1 R_{12})', ..., vec(R_{22} - D_T R_{12}])',$$
$$, D_1', ..., D_T', vec(\Sigma_{1,22.1})', ..., vec(\Sigma_{T,22.1})', \delta_1, \delta_{-1}', \theta_1^{h,w})',$$

and those of the marginal model (7) are

$$(vec(H_T^{1,z})', \theta_2')' = (\Sigma_{1,11}, ..., \Sigma_{T,11}, \delta_1, \delta_{-1}', R_{11}, R_{12}, \theta_2^{h,z})',$$

where $\theta_1^{h,w}$ and $\theta_2^{h,z}$ are vectors of parameters in the equations for $h_t^{1,w} = (q_{t,21}, q_{t,31}, q_{t,22}, q_{t,32}, q_{t,33})'$, and $h_t^{1,z} = q_{t,11}$,, respectively, $H_T^{1,w} = [h_1^{1,w} h_2^{1,w} ... h_T^{1,w}]$, $H_T^{1,z} = [h_1^{1,z} h_2^{1,z} ... h_T^{1,z}]$.

The conditional and marginal models contain two completely separate sets of latent variables. Thus under some additional assumptions, we can make efficient inference about the parameters and latent variables in the conditional model, and the marginal model can be neglected due to no information loss. Folowing Pajor (2011) we have:

Lemma 1. If

(i) $R_{11} = 0$, $\mathrm{R}_{12} = 0$, $\delta_1 = 0$,

(ii) $(\delta_{-1}, \mathrm{R}_{21}{}', vec(\mathrm{R}_{22})', \theta_1^{h,w}{}')'$ and $\theta_2^{h,z}$ are a priori independent,

or

(iii) $\delta_1 = 0$, $\forall t \in \{1, ..., T\}$ $\mathrm{D}_t = 0$,

(iv) $(\delta_{-1}{}', \mathrm{R}_{21}{}', vec(\mathrm{R}_{22})', \theta_1^{h,w}{}')'$ and $(R_{11}, R_{12}, \theta_2^{h,z}{}')'$ are a priori independent,

then z_t is weakly exogenous for $f_0(\mathrm{H}_T^{1,w}, \theta_1)$.

Hence, assumptions (i) and (ii) in Lemma 1 are sufficient for the Bayesian sequential cut and, consequently, for weak exogeneity of z_t for $(vec(\mathrm{H}_t^{1,w})', \theta_1{}')'$. It means that the inference about the parameters and latent variables of interest (i.e. $(vec(\mathrm{H}_t^{1,w})', \theta_1{}')')$ based on the complete process $\{y_t\}$ is the same as the one based on the conditional process $\{w_t|z_t\}$. Conditions (i) and (ii) in Lemma 1 are sufficient but not necessary for weak exogeneity. In Lemma 1 an alternative set of sufficient conditions that guarantee weak exogeneity of z_t for $(vec(\mathrm{H}_T^{1,w})', \theta_1{}')'$ is presented: $\mathrm{D}_t = 0$ ($\mathrm{D}_t \neq 0$ implying contemporaneous conditional correlation between w_t and z_t), $t = 1, ..., T$, $\delta_1 = 0$, and the prior independence of $(\delta_{-1}{}', \mathrm{R}_{21}{}', vec(\mathrm{R}_{22})', \theta_1^{h,w}{}')'$ and $(R_{11}, \mathrm{R}_{12}, \theta_2^{h,z}{}')'$.

Once conditions (i) and (ii) in Lemma 1 are satisfied, the conditional model of w_t given z_t assumes the form:

$$\begin{bmatrix} y_{t,2} \\ y_{t,3} \end{bmatrix} - \begin{bmatrix} \delta_2 \\ \delta_3 \end{bmatrix} = \begin{bmatrix} r_{22} & r_{23} \\ r_{32} & r_{33} \end{bmatrix} \left(\begin{bmatrix} y_{t-1,2} \\ y_{t-1,3} \end{bmatrix} - \begin{bmatrix} \delta_2 \\ \delta_3 \end{bmatrix} \right) + \begin{bmatrix} q_{t,21} \\ q_{t,31} \end{bmatrix} y_{t,1} + \begin{bmatrix} \beta_2 \\ \beta_3 \end{bmatrix} y_{t-1,1} + \begin{bmatrix} \xi_{t,2} \\ \xi_{t,3} \end{bmatrix}, \tag{9}$$

where $\{(\xi_{t,2}, \xi_{t,3})'\}$ is the bivariate SV process.

When conditions (iii) and (iv) in Lemma 1 hold, equations (7) and (8) can be written as:

$$z_t = R_{11} z_{t-1} + \mathrm{R}_{12}(\mathrm{w}_{t-1} - \delta_{-1}) + \varepsilon_{t,1}, \tag{10}$$

$$\mathrm{w}_t - \delta_{-1} = \mathrm{R}_{22}(\mathrm{w}_{t-1} - \delta_{-1}) + \mathrm{R}_{21} z_{t-1} + \varepsilon_{t,2}, \tag{11}$$

Note that conditions (i) and (ii) in Lemma 1 remain the same for strong exogeneity of z_t for $(vec(\mathrm{H}_T^{1,w})', \theta_1{}')'$. However, conditions (iii) and (iv) are not sufficient for strong exogeneity of z_t for $(vec(\mathrm{H}_T^{1,w})', \theta_1{}')'$. To ensure it, we must assume, in addition, that $R_{12} = 0$.

The strong exogeneity of z_t for $(vec(\mathrm{H}_T^{1,w})', \theta_1{}')'$ implies that the marginal model suffices (without a loss of information) for predictive inference on z_t and h_t^z. On the other hand, the strong exogeneity of z_t permits to forecast w_t from the conditional model, given the forecast of z_t from the marginal model.

In this paper, the conditions for exogeneity need to be tested. Unfortunately, tests of weak or strong exogeneity hypothesis require the joint model to be specified. To test the sufficient conditions (presented in Lemma 1, i-ii) for the strong exogeneity of z_t for a function of $(vec(H_t^{1,w})', \theta_1')'$ we use the Lindley type test (based on the highest posterior density region, see Box and Tiao, 1973; Osiewalski and Steel, 1993). We set the null and alternative hypotheses as: H_0: $\theta_{01} = 0$ and H_1: $\theta_{01} \neq 0$, where θ_{01} is the $k \times 1$ vector $(\theta_{01} = (\delta_1, r_{11}, r_{12}, r_{13})'$, $k = 4)$. Then, a posterior, the quadratic form: $F(\theta_{01}) = [\theta_{01} - E(\theta_{01} | y)]' V^{-1}(\theta_{01} | y)[\theta_{01} - E(\theta_{01} | y)]$, where $E(\theta_{01} | y)$ and $V(\theta_{01} | y)$ are the vector of posterior means of θ_{01} and the posterior covariance matrix respectively (in the notation we omit the initial conditions), is approximately $(T \to \infty)$ χ^2 distributed with k degrees of freedom. We assume that if $p(F(\theta_{01}) > F(0)|y) \geq 0.01$, the sufficient conditions of the strong exogeneity are not rejected by the data. In the same way, we examine whether $\theta_{02} = 0$, where $\theta_{02} = (\delta_1, r_{11}, r_{12}, r_{13}, \gamma_{21}, \phi_{21}, \sigma_{21}, \gamma_{31}, \phi_{31}, \sigma_{31})'$ (see conditions (iii) and (iv) in Lemma 1).

4. Dataset and empirical findings

Let us consider the daily and weekly 10-year bond yields of the U.S. ($y_{t,1}$), Germany ($y_{t,2}$), and Greece ($y_{t,3}$), over the period from July 13, 2006 to June 28, 2012 (downloaded from the website www.stooq.pl). The dataset of the daily (weekly) logarithmic growth rates (expressed in percentage points), consists of 1550 (311) observations (for each series). The first observation (in each time series) is used to construct the initial conditions. We consider both the daily and weekly data to examine whether differences in closing times of markets as well as differences in time zones affect the conclusions regarding the exogeneity of the growth rates of 10-year US bond yields.

The selected data are plotted in Figure 1 and 2. It can be seen from the graphs that the growth rates seem to be centered around zero, with time-varying volatility and the presence of outliers. The daily and weekly growth rates are more volatile in the periods of the financial crisis 2008-2009 and 2011-2012. However, the volatility of the growth rates of the 10-year bond yields of Greece is higher than that of the U.S. and Germany. Summary statistics for the time series are shown in Table 1. The arithmetic mean of the growth rates of the 10-year Greek bond yields is positive (with a relatively high standard deviation), while means of the growth rates of the 10-year German and U.S. bond yields are negative, which signals increased risk aversion among investors and can be interpreted as a confirmation of the assumption about the 'flight-to-safety' effect. The kurtosis (much larger than 3) suggests that distributions of the growth series are leptokurtic. As expected, the growth rates of the 10-year bond yields of the U.S. and Germany are positively correlated.

To obtain the Bayesian model, we specify a prior distribution of the parameters in the same way as in Pajor (2011). The prior distributions used are relatively non-informative. For daily data we start with T=550 initial observations (from the period July 13, 2006 – August, 24, 2008, selected ad hoc) and calculate p=1000 posterior distributions of parameters and latent variables based on the dataset available at time $T+k$ for each $k = 0, 1, …, p$-1 (up to $T+p$-1 = 1549). Thus we receive 1000 posterior distributions for the quadratic form $F(\theta_0)$ and posterior characteristics for each parameter and a latent variable. For weekly data we start with T=98 initial observations (from the period July 14, 2006 – June, 13, 2008) and calculate p=213 posterior distributions of parameters and latent variables based on the data set available at time $T+k$ for each $k = 0, 1, …, p$-1 (up to $T+p$-1 = 310).

All presented results were obtained with the use of the Metropolis and Hastings algorithm within the Gibbs sampler using 10^6 iterations after $5 \cdot 10^4$ burn-in Gibbs steps (see for details, Gamerman 1997, Tsay 2002 and Pajor 2005a, 2007).

Table 2 reports the posterior means and standard deviations (in parentheses) of all parameters of the VAR-SV model for the daily data. The main posterior characteristics of matrix R indicate that the daily growth rates of German and Greek 10-year bond yields 'significantly' (and positively) depend on the past daily growth rates of 10-year U.S. bond yields (the posterior standard deviations of r_{21} and r_{31} are relatively small compared to the posterior means). In other words, the daily growth rates of German and Greek 10-year bond yields are affected by the past movements of the U.S. sovereign bond yields. However, the daily growth rates of Greek 10-year bond yields 'significantly' (and negatively) depend on the past daily growth rates of German 10-year bond yields. On the other hand, the relatively large standard deviations of r_{12} and r_{13} indicate that the U.S. government bond yields are 'insignificantly' affected by the past movements in the German and Greek bond markets.

Table 3 reports the posterior means and standard deviations (in parentheses) of all parameters of the VAR-SV model for the weekly data. For this set of data, the main posterior characteristics of matrix R indicate that the weekly growth rates of 10-year German and Greek bond yields do not 'significantly' depend on the past weekly growth rates of 10-year U.S. bond yields. Similarly, the weekly growth rates of 10-year Greek bond yields do not 'significantly' depend on the past weekly growth rates of 10-year German bond yields. These results may be caused by using more aggregate data, and aggregation usually implies a loss of information. Finally, we can see that, similar to the results obtained for daily data, the U.S. bond market is not affected by the past movements in the German and Greek bond markets.

In Figure 3 and 4, the dashed lines represent the posterior means plus (or minus) the standard deviation, and the solid ones - the posterior means. These lines illustrate the alteration of characteristics of the posterior distribution of r_{ij} when the dataset was being increased one by one

observation, from the initial 550 vectors of daily data observations to 1550, and from the initial 98 vectors of weekly data observations to 311. The more observations are used, the more concentrated around their means the posterior distributions are. This effect can be seen both in Figure 3 and 4, where we observe the decrease of the posterior standard deviations of r_{ij} (the diminishing distance between the solid and the dashed lines). The location of the posterior distributions of r_{12} changes significantly, but only for the data from the beginning of the analyzed period, August 2008 to January 2010. In the middle of November 2008 particularly, the growth rates of 10-year U.S. Treasuries bond yields seem to be negatively affected by the growth rates of 10-year German bonds.

For both datasets the posterior means and standard deviations of δ_i ($i = 1, 2, 3$) indicate that vector δ is 'insignificantly' different from 0. The formal Bayesian testing (not presented here) would not lead to rejection of the null hypothesis that $\delta_i = 0$ for $i = 1, 2, 3$.

It is interesting to analyse the main characteristics of the posterior distributions of the conditional correlation coefficients. The time plots of conditional correlation between the growth rates of the selected countries' sovereign bond yields (for each $t=1,2, ..., T+p-1$) are shown in Figure 5a-b and 6a-b, where the upper line represents the posterior mean plus the standard deviation, and the lower one – the posterior mean minus the standard deviation. It can easily be seen that the conditional correlation coefficients vary over time. Generally, the correlation between the U.S. and Germany is positive and rather strong - the average posterior mean of $\rho_{12,t}$ ($t = 1, ..., T+p-1$) is equal to 0.636 (with average standard deviation equal to 0.095) for the daily data and 0.786 (with average standard deviation 0.084) for the weekly growth rate respectively. Different results were obtained for conditional correlations between the U.S. and Greece as well as between Germany and Greece. From July 2006 to October 2008 all conditional correlations are strongly positive, but from October 2008 (after the collapse of Lehman Brothers) the downward trend related to the posterior mean of $\rho_{13,t}$ and $\rho_{23,t}$ ($t = 1, ..., T+p-1$) can be clearly seen. From the beginning of 2010 these correlations are weakly negative – at the end of analyzed period the posterior mean of $\rho_{13,T+p-1}$ is equal to -0.058 for the daily data and -0.089 for the weekly data, while the posterior mean of $\rho_{23,T+p-1}$ is equal to -0.174 for the daily data and -0.385 for the weekly data. Among other things, this result is connected with the onset of the sovereign debt crisis in Europe at the end of 2009 (in late 2009 a new government in Greece revealed that its predecessors had concealed enormous budget deficit of 12,7% of GDP in contrast to the projected 3,7% of GDP, by the spring of 2010 Greece was unable to borrow on the open markets at an affordable interest rate, and was devised a series of bailout package by the so-called troika, the EU, the IMF and ECB, http://topics.nytimes.com/top/reference/timestopics/subjects/e/european_sovereign_debt_crisis/index.html).

Additionally, the conditional correlations presented in Figure 5a, 5b, 6a, and 6b, provide information about possible contagion effects. If we define contagion between bond markets as a significant increase of the conditional correlation coefficient in a crisis period compared to a pre-crisis period, then no contagion between all 10-year government bond yield pairs considered can be observed[2], whereas a significant decrease in the conditional correlation coefficient between Germany and Greece as well as between the U.S. and Greece may be explained by the effect known as 'flight to quality' (see Baur and Lucey 2009).

Individual volatility of each time series is measured by the conditional standard deviation. The time plots of conditional standard deviations of the growth rates are presented in Figure 7 and 8, where the upper line represents the posterior mean plus the standard deviation, and the lower one – the posterior mean minus the standard deviation. It can be seen from the graph that volatility of the growth rates of 10-year government bonds varies over time. Surprisingly, in case of the daily data the average volatility of the 10-year U.S. bond yields (which is equal to 2.04) is higher than that of the 10-year German and Greece bond yields (which are equal to 1.64 and 1.51 respectively). In case of the weekly data the average volatility of Greek bonds (which is equal to 4.45) is the highest one (for the U.S. and German bonds the averages of the conditional standard deviation are equal to 4.32 and 3.80 respectively). For both datasets the volatility plots for the U.S. and German bonds are relatively smooth. The dynamics of volatility of the 10-year Greek bond yields is different. The posterior means of the conditional standard deviation are most volatile, with the highest peaks. The first high peaks appeared in-between April and May 2010 and were connected with the economic situation in Greece (On April 27, 2010 Greece's debt was downgraded to junk level by rating agencies, in May 2010 a series of general strikes and demonstrations were taking place across Greece), the second one (in March 2012 when Greek 10-year bonds topped 38.5%) with a sudden decrease of the 10-year Greece bond yields from 36.562 on March 12, 2012 to 19.018 on March 13, 2012 (on May 13, 2012, credit rating agency Fitch upgraded Greece's long-term foreign and local currency issuer default ratings to 'B-' with stable outlooks from 'restricted default', http://news.xinhuanet.com/english/business/2012-03/14/c_122830872.htm).

Suppose that some function of $(vec(\mathrm{H}_T^{1,w})', \theta_1')'$ is of interest. As previously mentioned, to test the sufficient conditions (presented in Lemma 1) for the strong exogeneity of the growth rates of the 10-year U.S. bond yields we use the Lindley type test. For the complete daily dataset we obtain $F(0) = 12.412$, and $p(F(\theta_{01}) > F(0)|y) \approx 0.03$. The 99% highest posterior density interval (HPDI) for $F(\theta_{01})$ does include $F(0)$, but the value $F(0)$ lies outside the 95% HPDI (see Figure 9). Hence, our result precisely depends on that which HPDI we chose. The 99% HPDI implies that $\theta_{01} = 0$. Thus

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[2] A formal Bayesian test whether changes in excess correlation are statistically significant has not been presented here.

the sufficient conditions of the strong exogeneity are not rejected by the data. The result obtained for all weekly data provides similar conclusion, because of $F(0) = 9.625$, and $p(F(\theta_{01}) > F(0)|y) \approx 0.09$. The 95% HPDI implies that the sufficient conditions of the strong exogeneity are supported by the data.

A different result is obtained when we examine whether $\theta_{02} = 0$ (see conditions (iii) and (iv) in Lemma 1). For the daily data $F(0) = 150445.2$, that is $p(F(\theta_{02}) > F(0)|y) \approx 0.0000$; and for the weekly data $F(0) = 25060.1$, that is $p(F(\theta_{02}) > F(0)|y) \approx 0.0000$. None of the reasonable HPDIs contains $F(0)$. Thus the hypothesis that $\theta_{02} = 0$ is strongly rejected by the data.

As a byproduct of the strong exogeneity testing, it has been formally shown that $w_t = (y_{t,2}, y_{t,3})'$ does not cause $z_t = y_{t,1}$, given $H_t^{1,z}$ and θ_2 (see Pajor 2011). Thus the past of w_t does not influence the conditional distribution of z_t given the past of z_t, $H_t^{1,z}$ and θ_2. Roughly speaking, the past values of the growth rates of 10-year German and Greece bond yields do not influence the growth rates of 10-year U.S. bonds. Therefore, the forecast of the growth rates of 10-year U.S. bonds can be constructed from the marginal model only (see equation (7) with $R_{12} = 0$) and then forecasts of the growth rates of 10-year German and Greek bond yields can be obtained from the conditional model (see equation (9)) - without loss of relevant sample information. The result is consistent with our intuition.

In Figure 10 we present the posterior probabilities that $F(\theta_{01}) > F(0)$ (and respectively that $F(\theta_{02}) > F(0)$) at each time $t = T+i$-1 based on updated data set up to time $t = T+i$-1 (for $i=1, 2, ..., p$) The dataset is updated every day from August, 24, 2008 for the daily data and from June, 13, 2008 for the weekly data to June, 28, 2012. Each time when single vector observation was updated into a dataset, we computed these posterior probabilities. It can easily be seen that these probabilities vary with updating of the dataset. For the daily data, the lowest values are obtained in July-October 2009 (but $p(F(\theta_{01}) > F(0)|y) \geq 0.1$) and in May-June 2012 (but $p(F(\theta_{01}) > F(0)|y) \geq 0.01$). In the whole analyzed period the 99% HPDI implies that $\theta_{01} = 0$, that is the sufficient conditions of the strong exogeneity of the growth rates of 10-years U.S. bond yields are not rejected by the data. The value $F(0)$ lies outside the 95% HPDI from May 22, 2012 to the end of the observation period, i.e. June 28, 2012 (see Figure 10). For the weekly data, relatively low (but above 0.05) values of the posterior probability that $F(\theta_{01}) > F(0)$ are in May-November 2010 and from February 2012 to June 2012. Thus the 95% HPDI implies that the hypothesis that $\theta_{01} = 0$ is not rejected by each of the datasets.

As regards the posterior probability that $F(\theta_{02}) > F(0)$, it is from January 2009 (for the daily data) and December 2010 (for the weekly data) to June 2012 close to zero, which means that the

second set of sufficient conditions of the strong exogeneity of the growth rates of 10-year U.S. bond yields is strongly rejected by the data.

5. Conclusions

The examination of interactions among the U.S., German and Greek bond markets confirmed the asymmetric nature of their relationships. Over the whole period of our analysis, except for a brief interval in November 2008, the status of U.S. Treasuries as a global 'safe haven' remained unchanged. During the financial crisis, the growth rates of 10-year U.S. bond yields were not affected by the past growth rates of 10-year German and Greek bond yields. Our results indicate that contagion effects were absent among all 10-year bond markets considered. However, a 'flight to quality' effect between Germany and Greece occurred in the crisis period.

The econometric research of time-varying interdependence, based on the joint (complete) model for all variables, shows that the growth rates of 10-year U.S. bond yields can be forecast from the marginal model only, i.e. without taking German and Greek bond yields into consideration (the strong exogeneity hypothesis has not been rejected by the data), and the growth rates of 10-year German and Greek bond yields can be predicted, without loss of information, from the conditional model, i.e. the predicted values of 10-year U.S. bond yields obtained from the marginal model may be treated as fixed.

References

Abad P., Chuliá H., Gómez-Puig M., (2009), EMU and European Government Bond Market Integration, EBC Working Paper Series No. 1079/August.

Andersson M., Hansen L. J., Sebestyén S., (2006), Which News Moves the Euro Are Bond Market?, ECB Working Paper Series No. 631/May.

Barrios S., Iversen P., Lewandowska M., Setzer R., (2009), Determinants of intra-euro area government bond spreads during the financial crisis, European Ecocnomy – Economic Papers 388, Directorate General Economic and Monetary Affairs (DG ECFIN), European Commission.

Baur, D.G., Lucey, B.M., (2009), Flights and contagion. An empirical analysis of stock–bond correlations. Journal of Financial Stability 5, 339-352.

Beirne J., Gieck J., (2012), Interdependence and Contagion in Global Asset Markets, ECB Working Paper Series No. 1480/October.

Box G.E.P., Tiao G.C., (1973), *Bayesian Inference in Statistical Analysis*, Addison-Wesley Publishing Company, Massachusetts.

Christiansen Ch., (2003), Volatility-Spillover Effects in European Bond Markets, available at: http://pure.au.dk/portal/files/12707/wp_03-8.

Claeys P., Vašiček B., (2012), Measuring Sovereign Bond Spillover in Europe and the Impact of Rating News, Working Papers 2012/07, Czech National Bank, Research Department, available at: http://www.cnb.cz/en/research/research_publications/cnb_wp/download/cnbwp_2012_07.pdf.

Clare A., Lekkos I., (2000), An analysis of the relationship between international bond markets, Bank of England's Working Paper, available at: http://www.bankofengland.co.uk/publications/Documents/workingpapers/wp123.pdf.

Da Costa J. N. C., Cruz L. V., Leiria P. S., (2004), Government bond markets: what is the magnitude of volatility-spillover effects in the euro area?, available at: http://docentes.fe.unl.pt/~jmccosta/JMCCCruzLeiria_04_Volatility-spillovers.pdf

De Luna X., P. Johansson P., (2006), Exogeneity in structural equation models, *Journal of Econometrics* 132 (2), 527-543.

Ehrmann M., Fratzscher M., Rigobon R., (2005), Stocks, Bonds, Money Markets and Exchange Rates. Measuring International Financial Transmission, ECB Working Paper Series No. 452/March.

Engsted T., Tanggaard C., (2005), The comovement of US and German bond markets, available at: http://www.ucd.ie/t4cms/s51124.pdf.

Engle R.F., Granger C.W.J., (1987), Co-integration and error correction: representation, estimation and testing, *Econometrica* 55, 251-276.

Engle R.F., Hendry D.F., Richard J.F., (1983), Exogeneity, *Econometrica* 51, 277-304.

Ericsson N.R., Hendry D.F., Mizon G.E., (1998), Exogeneity, Cointegration, and Economic Policy Analysis, *Journal of Business and Economic Statistic* 6 (4), 370-387.

European Debt Crisis, available at: http://topics.nytimes.com/top/reference/timestopics/subjects/e/european_sovereign_debt_crisis/index.html.

Financial Integration in Europe, (2012), EBC, April, available at: http://www.ecb.int/pub/pdf/other/financialintegrationineurope201204en.pdf.

Florens J.P., Mouchart M., (1977), Reduction of Bayesian experiments, *CORE disscusion paper* no. 7737, Université Catholique de Louvain, Louvain-la-Neuve, Belgium.

Florens J.P., Mouchart M., (1980), Initial and Sequential Reduction of Bayesian Experiments, *CORE disscusion paper* no. 8015, Université Catholique de Louvain, Louvain-la-Neuve, Belgium.

Florens J.P., Mouchart M., (1982), A Note on Noncausality, *Econometrica* 50 (3), 583-591.

Florens J.P., Mouchart M., (1985), Conditioning in dynamic models, *Journal of Time Series Analysis* 53(1), 15-35.

Florens J.P., Mouchart M., Rolin J.M., (1990), *Elements of Bayesian statistic*, Marcel Dekker, Inc, New York and Basel.

Gamerman D., (1997), *Markov Chain Monte Carlo. Statistic simulation for Bayesian inference*, Chapman and Hall, London.

German Auction 'Disaster' Stirs Crisis Concern, available at: http://www.bloomberg.com/news/2011-11-23/germany-fails-to-receive-bids-for-35-of-10-year-bunds-offered-at-auction.html.

Goldberg L., Leonard D., (2003), What Moves Sovereign Bond Markets? The Effects of Economic News on U.S. and German Yields, Federal Reserve Bank of New York Vol. 9, No. 9, September, available at: http://www.newyorkfed.org/research/current_issues/ci9-9.pdf.

Granger C.W.J., (1969), Investigating causal relation by econometric and cross-sectional method, *Econometrica* 37, 424–438.

Hendry D.F., Richard J.F., (1982), On the Formulation of Empirical Models in Dynamic Econometrics, *Journal of Econometrics* 20, 3-33.

Hendry D.F., Richard J.F., (1983), The Econometric Analysis of Economic Time Series, *International Statistical Review* 51, 111-163.

Kim S.-J., Moshirian F., Wu E., (2006), Evolution of International Stock and Bond Market Integration: Influence of the European Monetary Union, Journal of Banking and Finance, 30:5, 1507-1534.

Laopodis N. T., (2008), Government Bond Market Integration within European Union, International Research Journal of Finance and Economics, Issue 19.

Mouchart M., Russo F., Wunsch G., (2007), Causality, structural modelling, and exogeneity, *IAP Statistics Network*, 1-26.

Newton M.A., Raftery A.E., (1994), Approximate Bayesian inference by the weighted likelihood bootstrap [with discussion], *Journal of the Royal Statistical Society B*, 56(1), 3-48

O'Hagan A., (1994), *Bayesian Inference*, Halsted Press, New York.

Osiewalski J., Pipień M., (2004), Bayesian comparison of bivariate GARCH processes. The role of the conditional mean specification, Chapter 7, [in:] *Contributions to Economic Analysis 269, New Directions in Macromodelling*, [ed.:] A. Welfe,Elsevier, Amsterdam, 173-196.

Osiewalski J., Steel M.F.J., (1993), Una perspectiva bayesiana en selección de modelos, *Cuadernos Economicos* 55/3, 327-351, (typescript: A Bayesian perspective on model selection).

Osiewalski J., Steel M.F.J., (1996), A Bayesian analysis of exogeneity in models pooling time-series and cross-section data, *Journal of Statistical Planning and Inference* 50, 187-206.

Pajor A., (2005a), Bayesian Analysis of Stochastic Volatility Model and Portfolio Allocation*, Acta Universitatis Lodzensis – Folia Oeconomica 1*92, 229-249.

Pajor A., (2005b), Bayesian comparison of bivariate SV models for two related time series, *Acta Universitatis Lodziensis – Folia Oeconomica* 190, 177-196.

Pajor A., (2006), Bayesian VECM-SV Models for the Main Polish Exchange Rates, [in:], *32-nd International Conference, MACROMODELS'05*, [ed.:] W. Welfe, A. Welfe, Łódź, 135-154.

Pajor A., (2007), VECM-TSV Models for Two Polish Official Exchange Rates, [in:] *Financial Markets: Principles of Modeling, Forecasting and Decision-Making, FindEcon Monograph Series: Advance in Financial Market Analysis*, No. 2, [ed.:] W. Milo, P. Wdowiński, Łódź University Press, Łódź, 49-66.

Pajor A., (2008), A Bayesian Analysis of Weak Exogeneity in VECM-SV Models, [in:] *Metody ilościowe w naukach ekonomicznych (Quantitative Methods in Economic Sciences),* [ed.:] A. Welfe, Warsaw School of Economics, Warszawa, 235-249.

Pajor A., (2010), *Wielowymiarowe procesy wariancji stochastycznej w ekonometrii finansowej. Ujecie bayesowskie (Multivariate Stochastic Variance Processes in Financial Econometrics. Bayesian approach, in Polish)*, Cracow University of Economics, Kraków.

Pajor A., (2011), A Bayesian Analysis of Exogeneity in Models with Latent Variables, *Central European Journal of Economic Modelling and Econometrics, vol. 3 (issue 2)*, 49-73, http://www.cejeme.com.

Rault Ch., (2011), Long-run strong-exogeneity, *Economics Bulletin* 31 (1), 1-8.

Safe Assets: Financial System Cornerstone?, (2012), IMF, April, available at: https://www.imf.org/external/pubs/ft/gfsr/2012/01/pdf/c3.pdf.

Sosvilla-Rivero S., Morales-Zumaquero A., (2011), Volatility in EMU sovereign bond yields: permanent and transitory components, ICEI Working Papers, WP06/11, available at: http://eprints.ucm.es/12896/1/WP_06-11.pdf.

Tsay R.S., (2002), *Analysis of Financial Time Series. Financial Econometrics*, A Wiley-Interscience Publication, John Wiley & Sons, INC, New York.

Table 1.

Summary statistics of $y_{i,t}$

	daily data			weekly data		
	U.S.	Germany	Greece	U.S.	Germany	Greece
Mean	-0.075	-0.063	0.117	-0.372	-0.309	0.574
std dev.	2.258	1.936	2.738	4.691	4.190	6.683
skewness	-0.278	-0.009	-11.907	-0.211	-0.233	-4.390
Kurtosis	6.697	7.668	271.697	4.874	5.171	51.973
minimum	-17.021	-13.572	-65.362	-19.468	-17.746	-69.839
maximum	9.221	11.273	13.577	12.929	14.018	32.865
correlation coefficient						
U.S.	1	0.585	-0.011	1	0.757	-0.097
Germany		1	-0.046		1	-0.183
Greece			1			1

Table 2. The posterior means and standard deviations (in parentheses) of the parameters of the trivariate VAR-SV model (results based on all available daily data)

δ_1	δ_2	δ_3	r_{11}	r_{12}	r_{13}	r_{21}	r_{22}	r_{23}	r_{31}
-0.0117	0.0324	0.0211	-0.0748	0.0910	-0.0387	0.1204	-0.0168	-0.0232	0.1009
(0.0318)	(0.0253)	(0.0229)	(0.029)	(0.0349)	(0.0195)	(0.0179)	(0.0257)	(0.0169)	(0.0155)
r_{32}	r_{33}	γ_{11}	ϕ_{11}	σ_{11}^2	γ_{22}	ϕ_{22}	σ_{22}^2	γ_{33}	ϕ_{33}
-0.1735	0.1473	1.8820	0.9953	0.0107	1.9118	0.9974	0.0115	-0.7366	0.9793
(0.0221)	(0.0233)	(2.324)	(0.003)	(0.0031)	(3.9849)	(0.0024)	(0.0039)	(0.8209)	(0.0063)
σ_{33}^2	γ_{21}	ϕ_{21}	σ_{21}^2	γ_{31}	ϕ_{31}	σ_{31}^2	γ_{32}	ϕ_{32}	σ_{32}^2
0.2753	0.5331	0.9869	0.0007	-0.2689	0.9960	0.0005	-0.8932	0.9984	0.0009
(0.0452)	(1.1028)	(0.0084)	(0.0002)	(2.2384)	(0.003)	(0.0001)	(3.5398)	(0.0016)	(0.0003)
ln $q_{11,0}$	ln $q_{22,0}$	ln$q_{33,0}$	$q_{21,0}$	$q_{31,0}$	$q_{32,0}$				
-0.8775	-1.2962	-6.7543	0.6423	0.5827	0.9382				
(0.3741)	(0.4386)	(1.3319)	(0.1521)	(0.138)	(0.0809)				

Source: own calculations

Table 3. The posterior means and standard deviations (in parentheses) of the parameters of the trivariate VAR-SV model (results based on all available weekly data)

δ_1	δ_2	δ_3	r_{11}	r_{12}	r_{13}	r_{21}	r_{22}	r_{23}	r_{31}
-0.1460	0.1233	0.0755	-0.1202	-0.0599	0.0045	0.0008	-0.0735	0.0386	0.0262
(0.1403)	(0.1247)	(0.1147)	(0.0836)	(0.0942)	(0.0383)	(0.0668)	(0.0753)	(0.0339)	(0.0614)
r_{32}	r_{33}	γ_{11}	ϕ_{11}	σ_{11}^2	γ_{22}	ϕ_{22}	σ_{22}^2	γ_{33}	ϕ_{33}
-0.0022	-0.0327	3.7574	0.9832	0.0244	3.3795	0.9899	0.0378	2.1575	0.9636
(0.07)	(0.0508)	(2.5727)	(0.0134)	(0.0136)	(3.8537)	(0.0093)	(0.0221)	(2.1464)	(0.0181)
σ_{33}^2	γ_{21}	ϕ_{21}	σ_{21}^2	γ_{31}	ϕ_{31}	σ_{31}^2	γ_{32}	ϕ_{32}	σ_{32}^2
0.5398	0.6766	-0.2690	0.0086	-1.3111	0.9950	0.0015	-2.5304	0.9963	0.0036
(0.1907)	(0.0293)	(0.2457)	(0.0043)	(3.6984)	(0.0056)	(0.0006)	(4.4741)	(0.0044)	(0.002)
ln $q_{11,0}$	ln $q_{22,0}$	ln$q_{33,0}$	$q_{21,0}$	$q_{31,0}$	$q_{32,0}$				
0.8110	-0.5809	-9.3019	0.0898	0.5635	0.8822				
(0.5683)	(0.5888)	(3.9547)	(3.7334)	(0.0804)	(0.1475)				

Source: own calculations
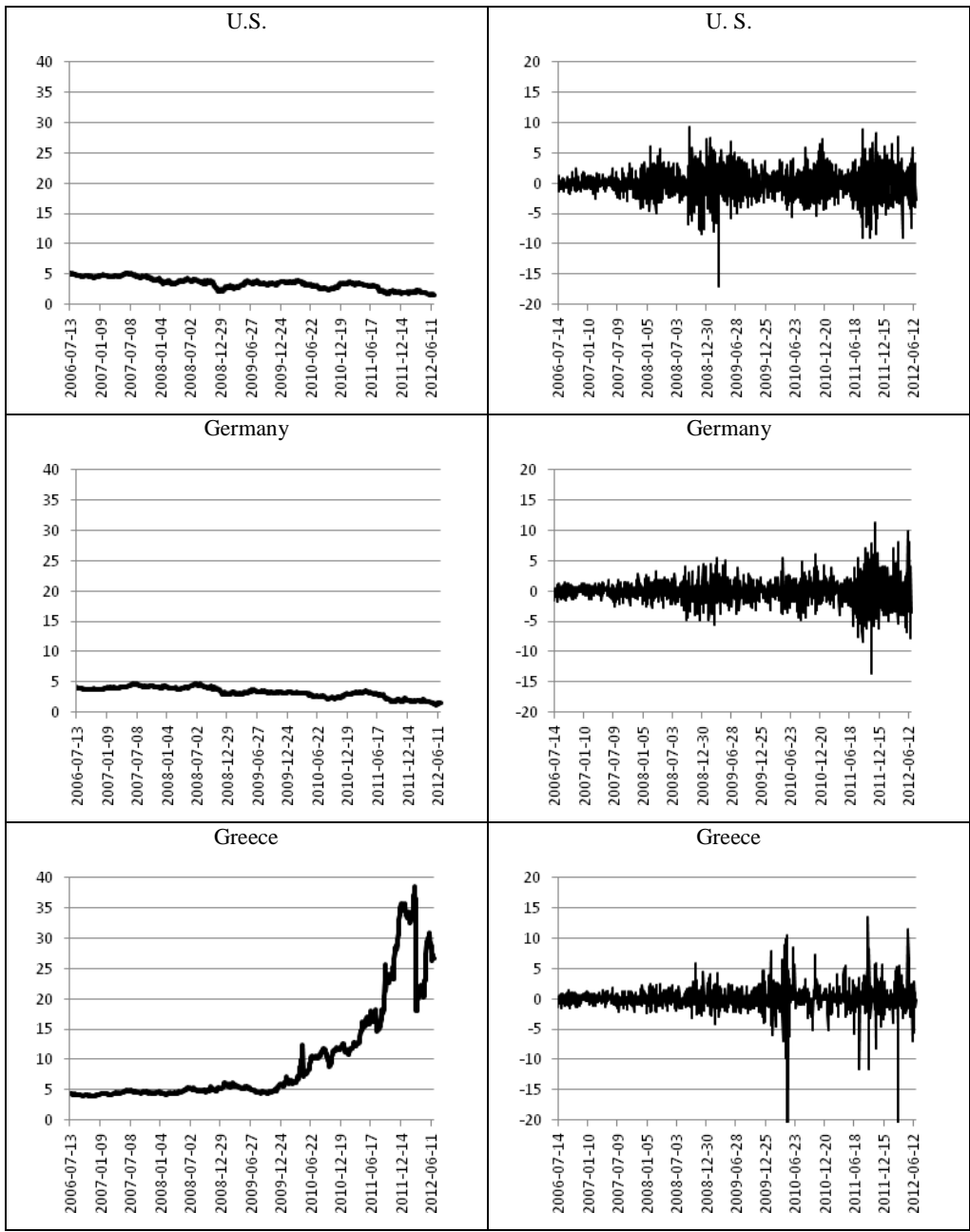
Figure 1. Daily data: 10-year bond yields and their growth rates (July 14, 2006 – June 28, 2012)
Source: own calculations
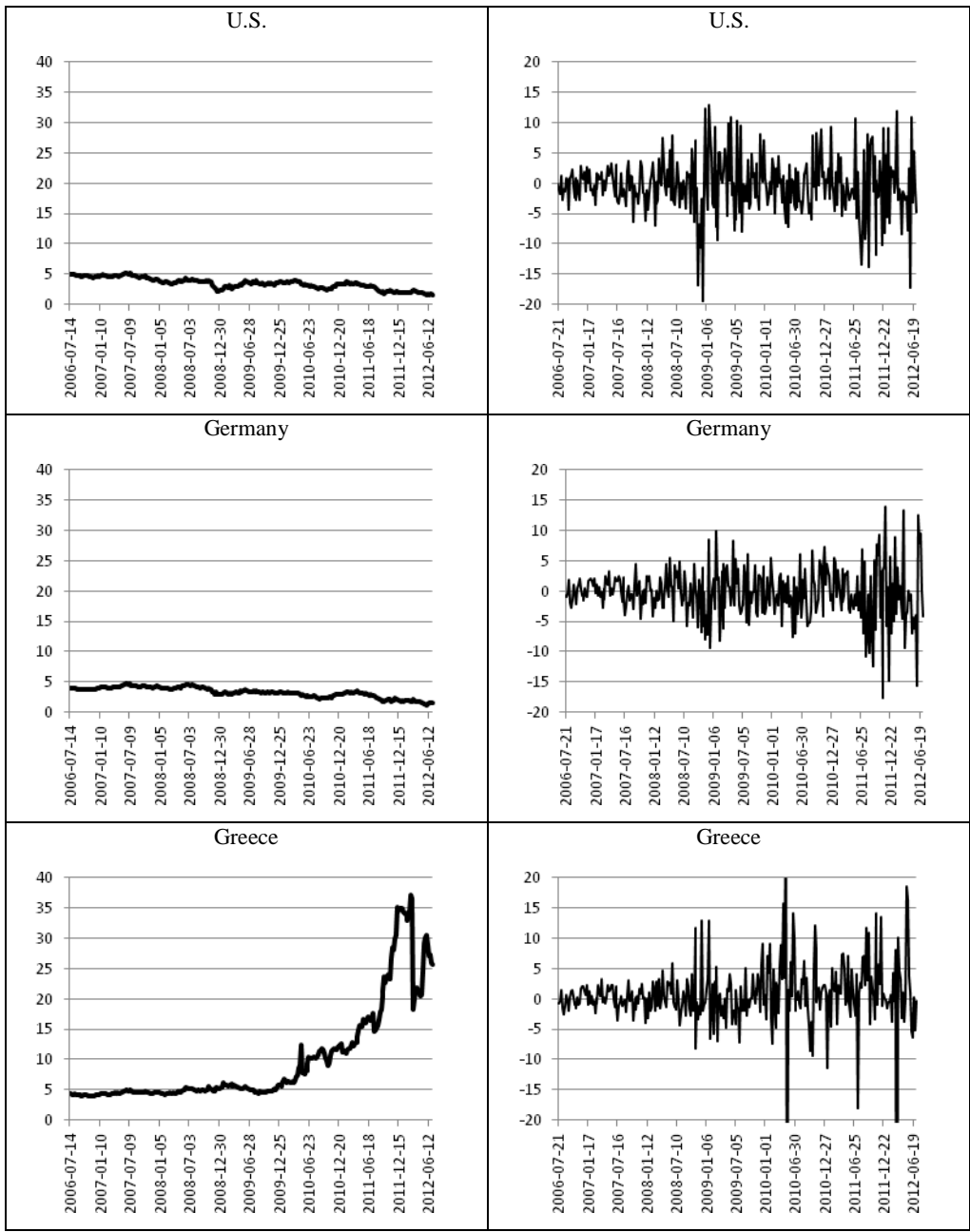
Figure 2. Weekly data: 10-year bond yields and their growth rates (July 14, 2006 – June 28, 2012)
Source: own calculations
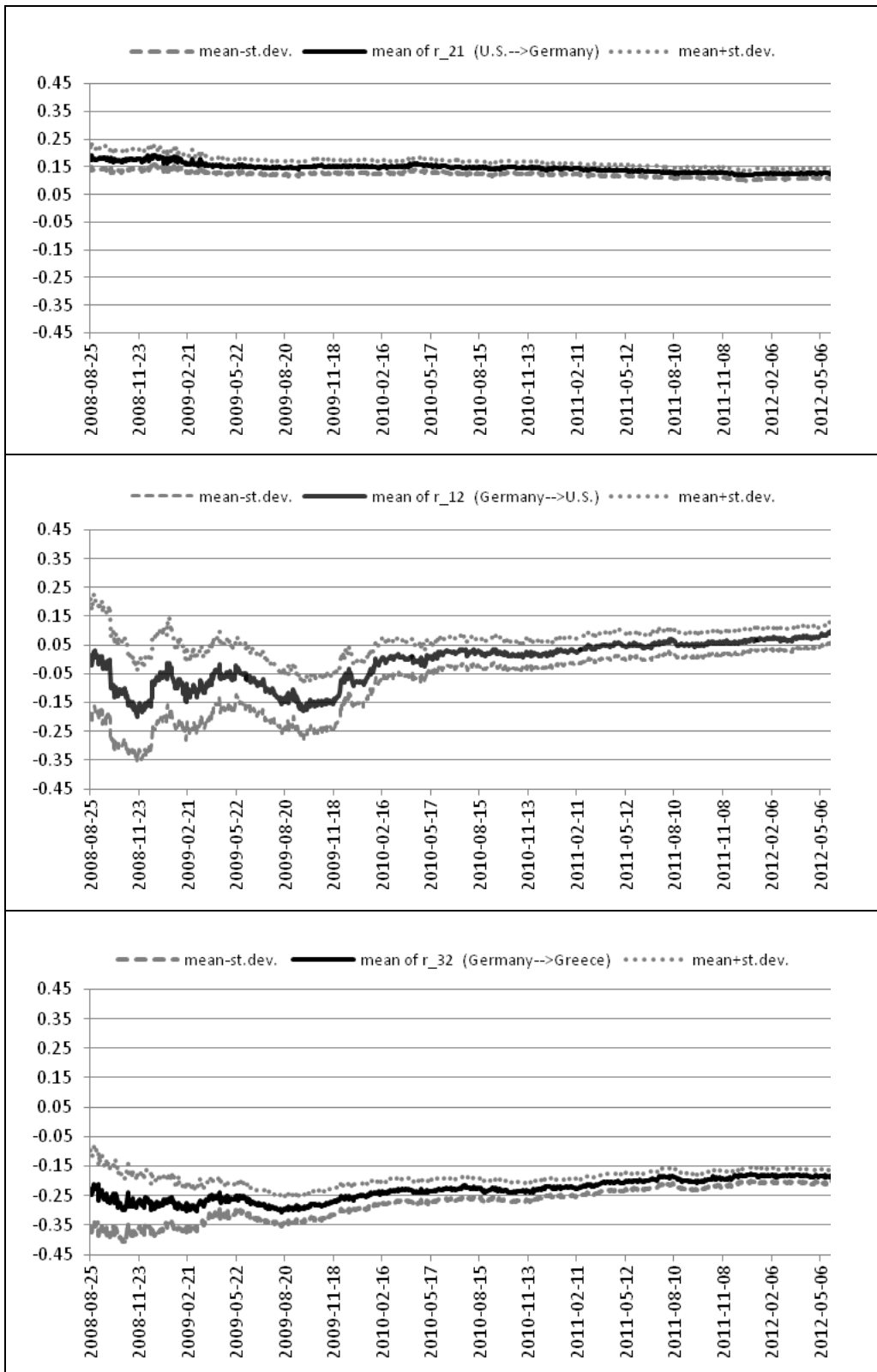
Figure 3. Selected entries of matrix R (posterior mean ± standard deviation, daily data). Results based on the data at time $T+k$, $k = 0$, ..., p -1
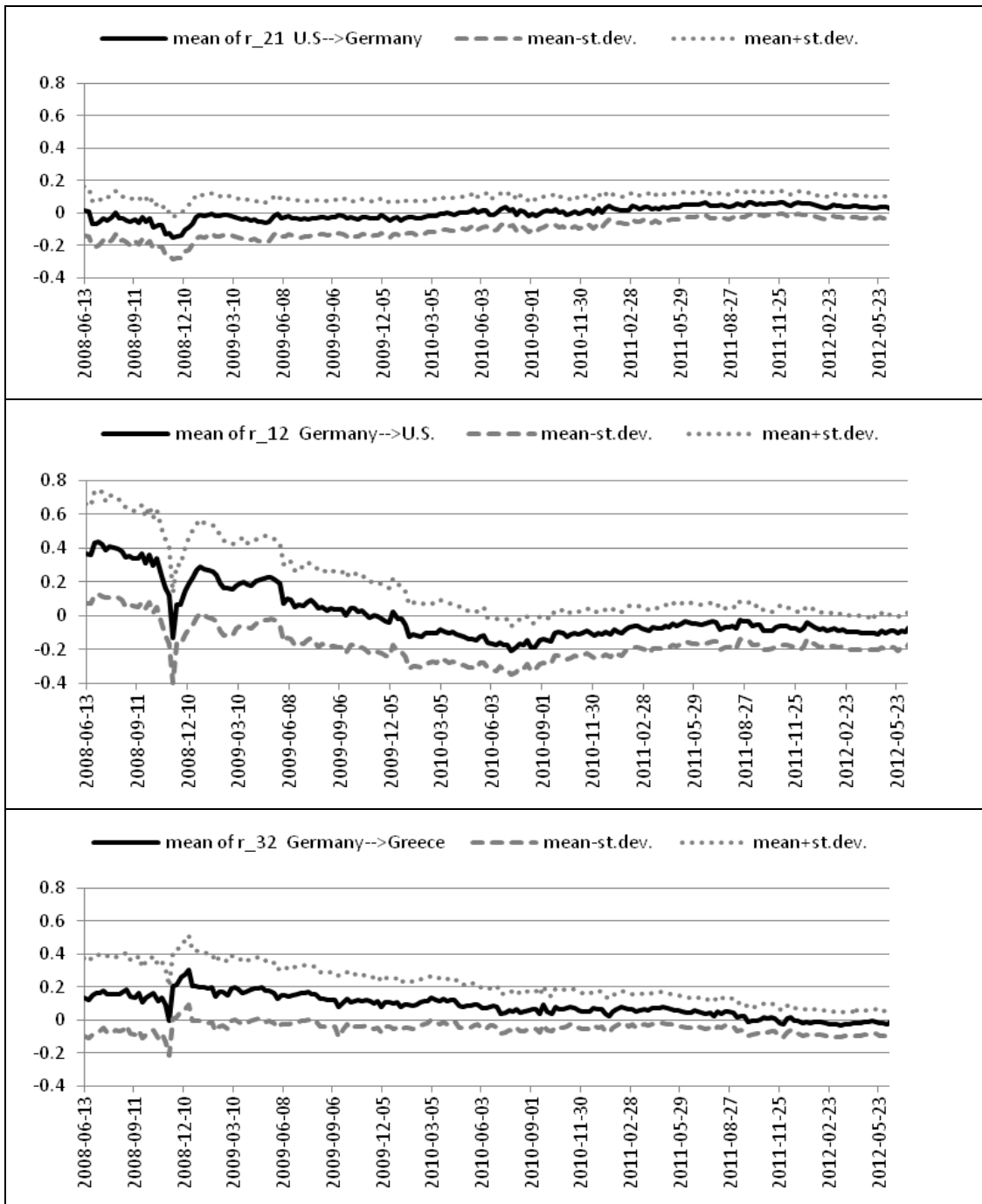
Figure 4. Selected entries of matrix R (posterior mean ± standard deviation, weekly data). Results based on the data at time $T+k$, $k = 0, ...,p$ -1.
Source: own calculations
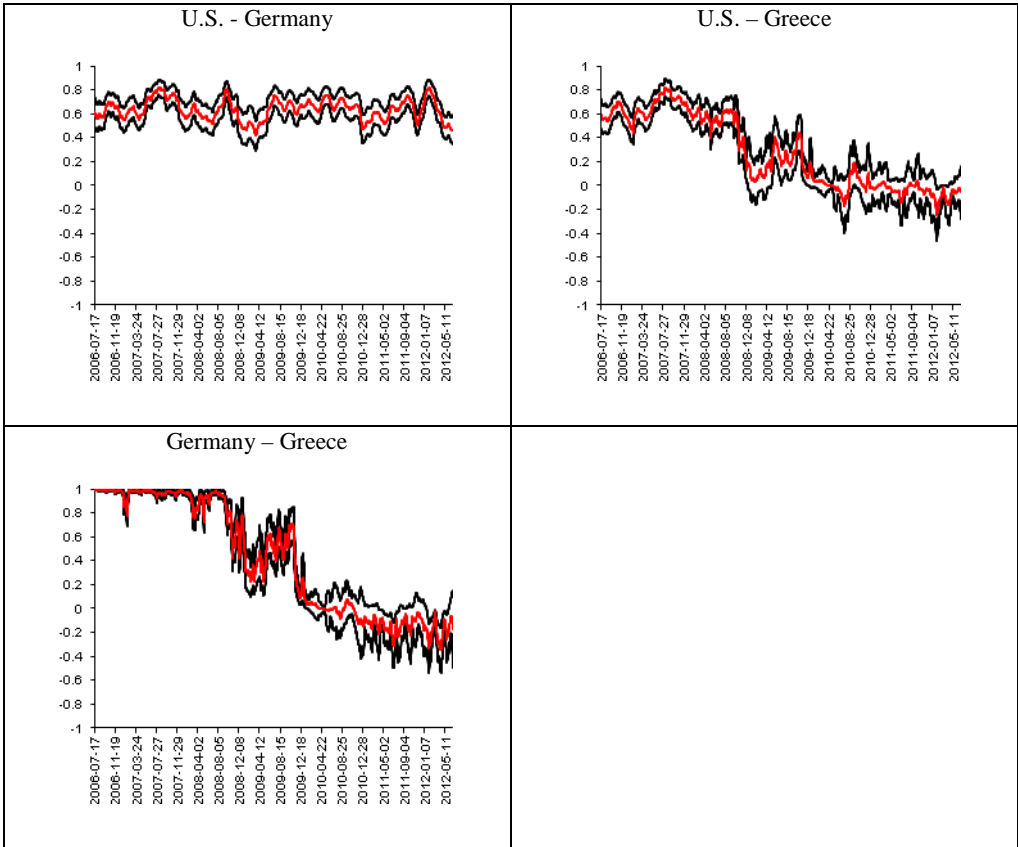
Figure 5a. Conditional correlation coefficients between <u>daily</u> growth rates of the 10-yeard bond yields (posterior mean ± standard deviation, results based on all available data)
Source: own calculations
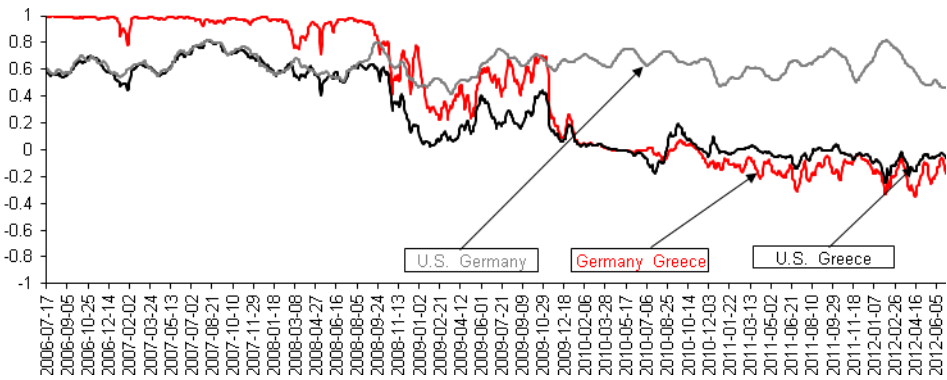
Figure 5b. Conditional correlation coefficients between <u>daily</u> growth rates of the 10-yeard bond yields (results based on all available data)
Source: own calculations
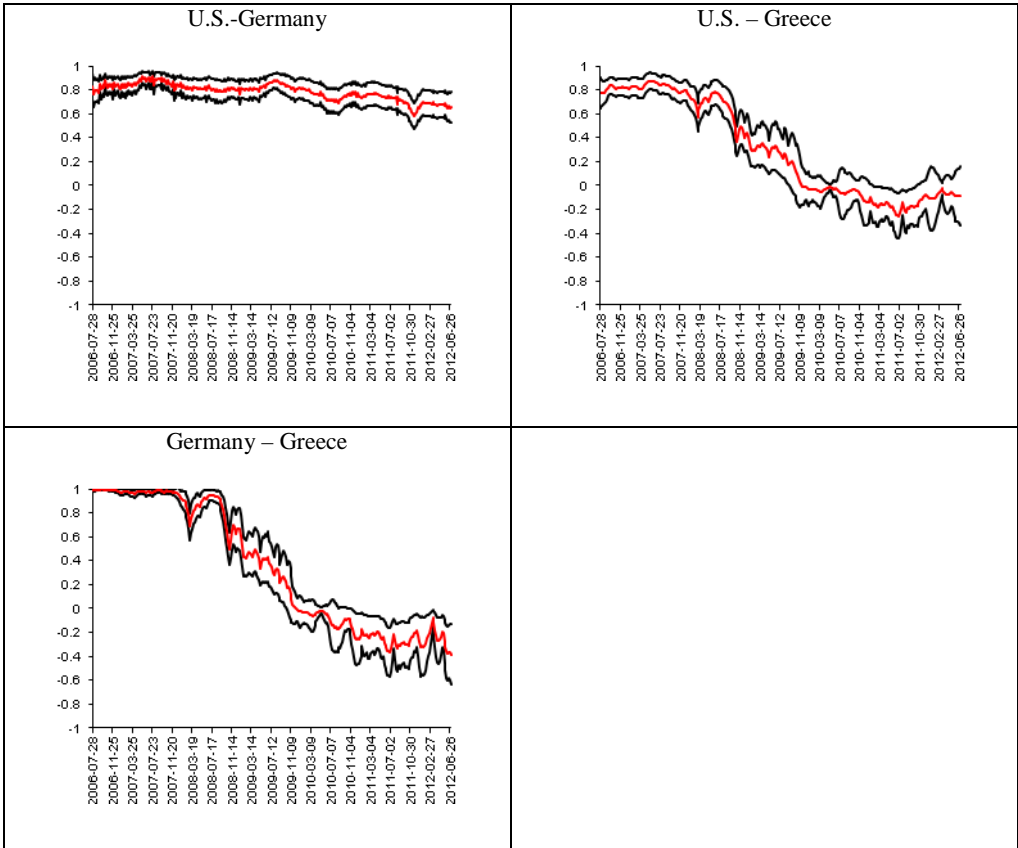
Figure 6a. Conditional correlation coefficients between <u>weekly</u> growth rates of the 10-yeard bond yields (posterior mean ± standard deviation, results based on all available data)
Source: own calculations
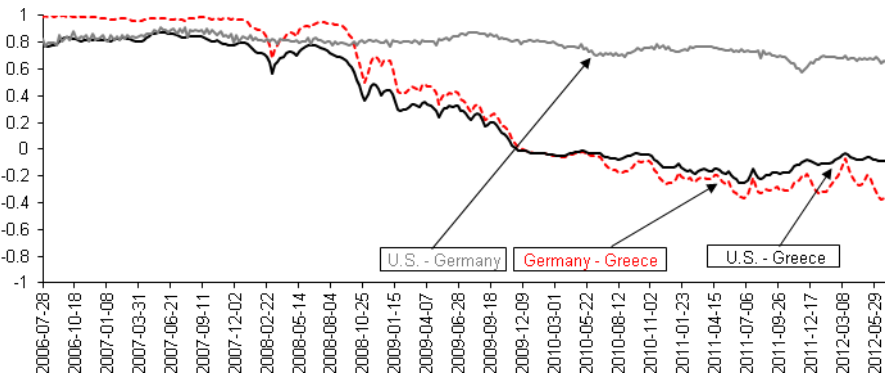
Figure 6b. Conditional correlation coefficients between <u>weekly</u> growth rates of the 10-yeard bond yields (results based on all available data)
Source: own calculations

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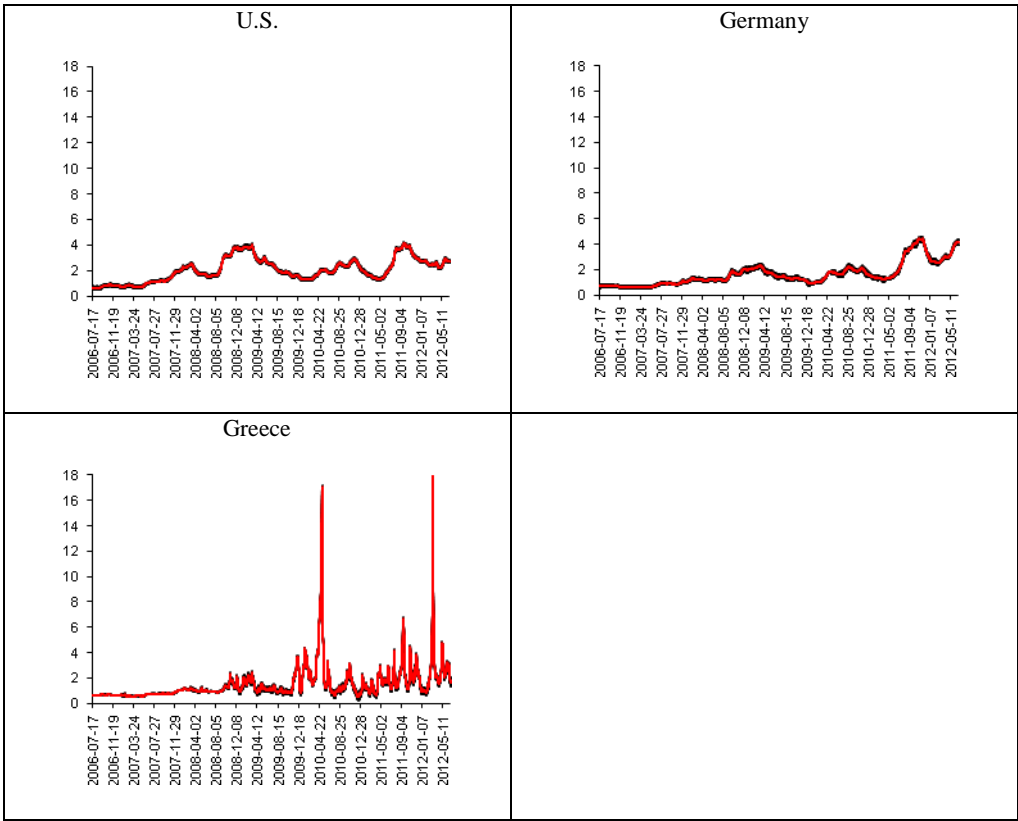
Figure 7. Conditional standard deviations for <u>daily</u> data (posterior mean ± standard deviation, results based on all available data)
Source: own calculations
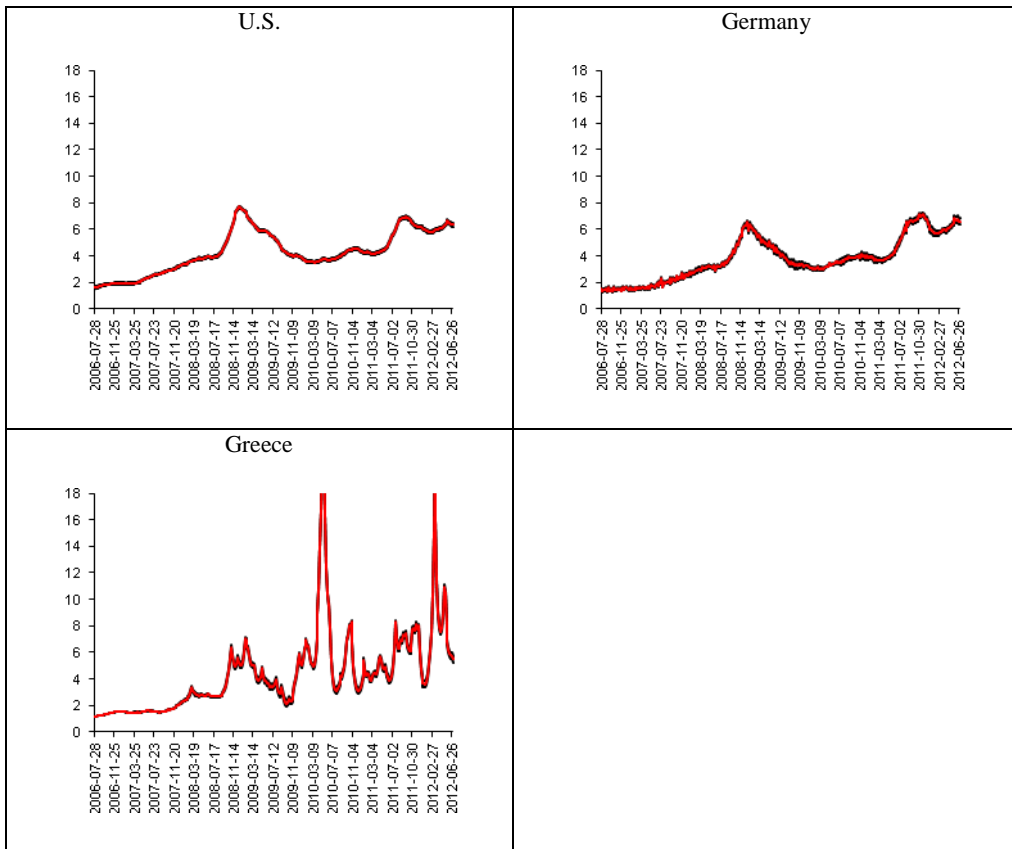
Figure 8. Conditional standard deviations for <u>weekly</u> data (posterior mean ± standard deviation, results based on all available data)
Source: own calculations

daily data	weekly data
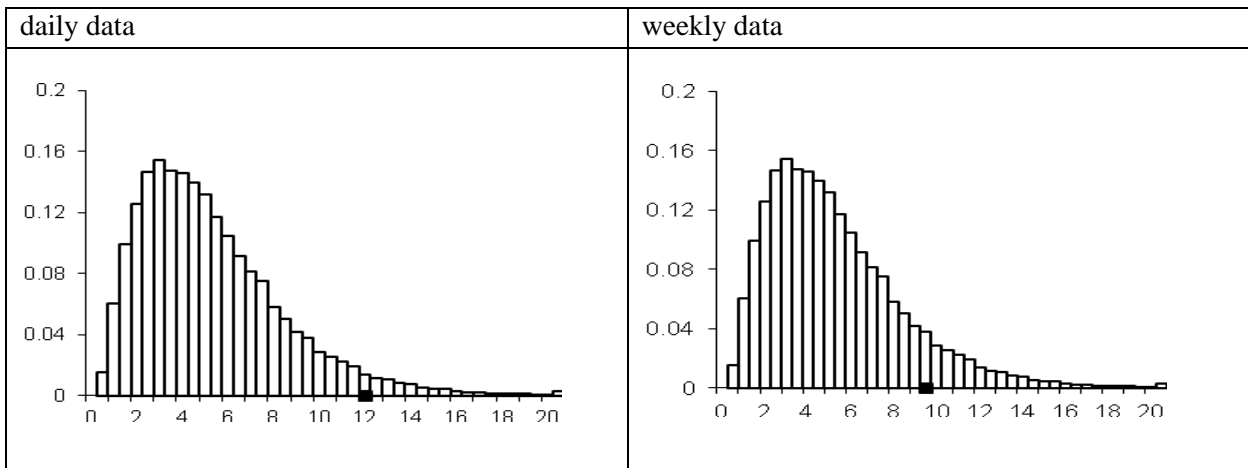
Figure 9. Histogram of the posterior distribution for $F(\theta_0)$. The black square represents F(0), obtained based on all available data (T = 1549 and T=310, respectively)

Source: own calculations

conditions i-ii in Lemma 1	conditions iii-iv in Lemma 1
daily data	
weekly data	
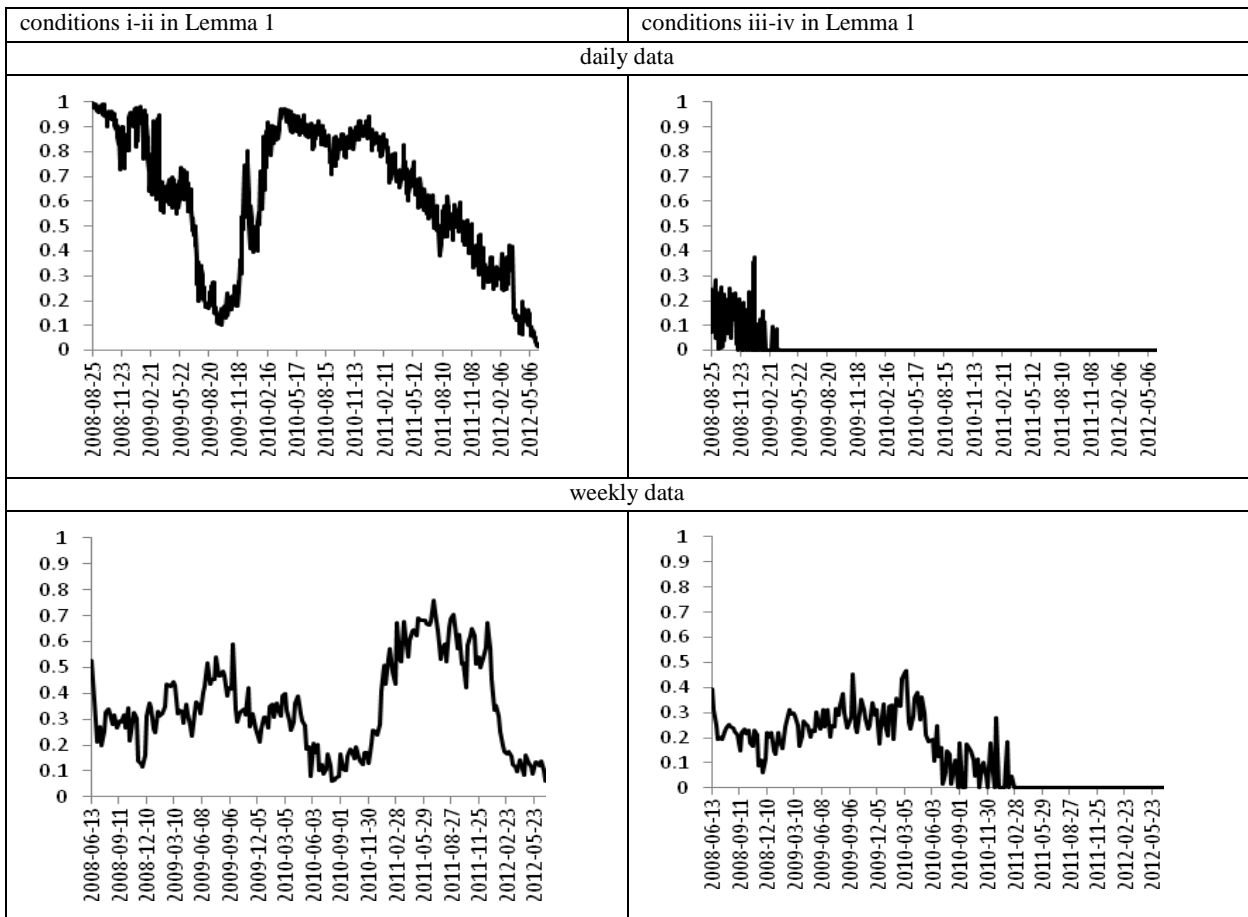
Figure 10. Results for $p(F(\theta_0) > F(0)|y)$ and $p(F(\theta_{01}) > F(0)|y)$, based on the data at time $T+k$, $k = 0, ...,p$ -1

Source: own calculations

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