

Fiscal Spillovers in the Euro Area

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CESIFO WORKING PAPER NO. 3693

CATEGORY 6: FISCAL POLICY, MACROECONOMICS AND GROWTH

DECEMBER 2011

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Abstract

This paper analyses the dynamic effects of fiscal imbalances in a given EMU member state on the borrowing costs of other countries in the euro area. The estimation of a multivariate, multi-country time series model (specifically a Global VAR, or GVAR) using quarterly data for the EMU period suggests that euro-denominated government yields are strongly linked with each other. However, financial markets seem to be able to discriminate among different issuers. Consequently, fiscal imbalances in Italy and in other peripheral countries should be closely monitored by their EMU partners and the European institutions.

JEL-Code: C320, E620, F420, H630.

Keywords: global VAR methodology, fiscal spillovers, euro area, public debt.

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October 2011

1. Introduction

With the introduction of the euro in Stage III of the European Monetary Union (EMU), members of the euro area have redenominated their outstanding debt in the new common currency and started to issue euro-denominated government securities to finance their national debt. The degree of substitution between government bonds of different euro member states rose immediately following the elimination of the exchange rate risk. However, in the aftermath of the financial crisis of 2007-2008 the soaring costs of financial support schemes as well as recession-induced falls in tax revenues have caused in most advanced economies (including those belonging to the EMU) a dramatic increase in the supply of government paper and raised concerns about the impact of fiscal imbalances on long-term debt financing costs. As a result markets have become much more careful to discriminate issuers on the basis of their fiscal performance and their macroeconomic fundamentals (Blommestein, 2009; Schuknecht et al. 2011).

Understanding the ultimate effects of growing levels of public indebtedness on long-term yields is a difficult task. Economic theory suggests that worsening fiscal positions will lead to higher (real) interest rates with a detrimental effect on private investment and consumption plans (Buiter, 1977; Friedman, 1978; Elmendorf and Mankiw, 1999). However, this effect can be tempered, or even reversed, if the decrease in public saving is offset by an increase in private saving and/or international capital flows or perfect substitutability between private and public spending (Faini, 2006; Linneman and Schabert, 2004). Financial markets also play a key role. On the one hand, investors may see high debt levels as beneficial since they imply more liquid markets and lower liquidity premia for actively traded government debt securities. On the other hand, by issuing more debt, countries might create excess supply and therefore drive bond prices down (and thus interest rates up) to persuade markets to absorb the higher amount of debt, leading to the perception of an increasing default risk.

Consequently, determining the effects of a deterioration of a country's fiscal position on

government bond yields is largely an empirical issue. The existing literature provides mixed evidence, although the majority of papers report a significant positive effect (see the surveys of Gale and Orszag, 2002 and Engen and Hubbard, 2004). In general, higher increases in government yields have been observed in the European countries (Caporale and Williams, 2002; Chinn and Frankel, 2007) compared to the US (Engen and Hubbard 2004; Evans and Marshall, 2007; Laubach 2009).

The present paper aims to contribute to this literature by assessing the short- to medium-term effects of government debt accumulation on long-term interest rates in the case of euro area member states during the EMU period. It differs from previous studies in three ways. First, we adopt a dynamic multivariate time series approach instead of the static and/or univariate framework used, for instance, in Correia-Nunes and Stemitsiotis, 1995, Caporale and Williams, 2002, and Chinn and Frenkel, 2007. Second, we examine linkages between the yields of EMU countries in order to assess the degree of substitutability between bonds issued by different member states. In this respect, our analysis provides a macroeconomic perspective to the issue of bond market integration typically assessed in the financial literature using high frequency data (see Georgoutsos and Migiakis, 2010, among others). Third, our empirical framework is explicitly designed to identify shocks according to their geographical origin. Such a feature is particularly appealing for the study of fiscal spillovers across countries or regions within the euro area. In fact, we are able to identify the role played by fiscal imbalances in specific foreign economies in determining domestic bond yields in a multi-country environment including eleven EMU member states (only a subset of which has been considered in earlier papers such as Chinn and Frenkel, 2003, and Paesani et al., 2006).

More specifically, our econometric analysis is based on aggregating a number of Vector Error Correction (VEC) systems (one for each EMU country) in a Global Vector Auto Regressive (GVAR) model describing the euro area economy (Pesaran et al., 2004) in order to perform dynamic simulation exercises. Using quarterly data over the period 1998:1-2010:4,

we document that long-term yields dynamics for euro-denominated government securities are strongly linked to each other (except for the case of Greece): foreign factors appear to account for the largest percentage of variability in long-term interest rates over all simulation horizons in all EMU countries. However, the quality of government debt also plays an important role. In particular, increasing debt tends to reduce the yields on euro-denominated government securities if issuances are of high quality (such as the triple-A French or German government securities). By contrast, if the fiscal position worsens in a highly indebted issuer (such as Italy) or in a peripheral economy, real interest rates increase since country risk factors dominate liquidity effects. We also show that unanticipated flight-to-quality and/or flight-from-risk phenomena lead to widening spreads for Italy and the peripheral countries vis-à-vis the core EMU economies, implying higher debt financing costs for the economies with the weakest macroeconomic fundamentals.

The remainder of the paper is structured as follows. Section 2 outlines the inter-regional macro-econometric framework used for the empirical analysis. Section 3 presents the main estimation results from the country VEC models as well as the GVAR system. Section 4 discusses the dynamic simulation results. Section 6 offers some concluding remarks.

2. The econometric framework

In order to analyse the impact of domestic and foreign developments affecting the government debt/GDP ratio on long-term interest rates, we adopt the Global Vector Auto Regressive (GVAR) methodology proposed by Pesaran et al. (2004). This procedure stacks in a single coherent model of the global economy a number of country-specific Vector Error Correction (VEC) systems (Johansen, 1992) and explicitly allows for interdependencies across economies in a multi-country setting. Its main advantage is that the geographical origin (though not the economic nature) of the shocks hitting the variables of the global system can be identified. This is because each country-specific system in the multi-country model is

estimated conditionally on foreign variables, with only small remaining correlations between cross-country shocks and endogenous factors.

2.1. Individual country VEC models

There are $N + 1$ countries indexed by $i = 0, 1, \dots, N$.¹ For each country the following VEC model is estimated:²

$$\Delta \mathbf{x}_{it} = \mathbf{a}_i + \mathbf{\Pi}_i \boldsymbol{\kappa}_i - \mathbf{\Pi}_i [\mathbf{z}_{i,t-1} - \boldsymbol{\kappa}_i (t-1)] + \boldsymbol{\Lambda}_{i0} \Delta \mathbf{x}_{it}^* + \boldsymbol{\varepsilon}_{it} \quad (1)$$

where \mathbf{x}_{it} is a $(k_i \times 1)$ vector of country i domestic variables, \mathbf{x}_{it}^* is a $(k_i^* \times 1)$ vector of foreign variables specific to country i (to be defined below), $\mathbf{z}_{it} \equiv (\mathbf{x}'_{it}, \mathbf{x}^*_{it})'$, \mathbf{a}_i is a $(k_i \times 1)$ vector of fixed intercepts, $\boldsymbol{\Lambda}_{i0}$ is a $(k_i \times k_i^*)$ matrix of coefficients associated to the foreign variables, $\boldsymbol{\varepsilon}_{it}$ is a $(k_i \times 1)$ vector of country-specific shocks, with $\boldsymbol{\varepsilon}_{it} \sim N(\mathbf{0}, \boldsymbol{\Sigma}_{ii})$, where $\boldsymbol{\Sigma}_{ii}$ is a non-singular variance-covariance matrix, and where $t = 1, 2, \dots, T$ indexes time. Moreover, $E(\boldsymbol{\varepsilon}_{it}, \boldsymbol{\varepsilon}'_{js}) = \boldsymbol{\Sigma}_{ij}$, for $t = s$, and $E(\boldsymbol{\varepsilon}_{it}, \boldsymbol{\varepsilon}'_{js}) = \mathbf{0}$, for $t \neq s$, so that cross-country correlations among the idiosyncratic shocks are allowed. The number of long-run relations is given by the rank $r_i \leq k_i$ of the $k_i \times (k_i + k_i^*)$ matrix $\mathbf{\Pi}_i$. Finally, in order to avoid introducing quadratic trends in the levels of the variables when $\mathbf{\Pi}_i$ is rank-deficient, $k_i - r_i$ restrictions $\mathbf{a}_i = \mathbf{\Pi}_i \boldsymbol{\kappa}_i$ are imposed on the trend coefficients, where \mathbf{a}_i is the coefficient on the time trend in the isomorphic level Vector AutoRegression (VAR) form of (1) and $\boldsymbol{\kappa}_i$ is a $(k_i + k_i^*) \times 1$ vector of constants.

¹ $N = 10$ in this paper. $i = 0$ is the reference country (Germany). The GVAR approach is applied here to investigate cross-country linkages within the EMU region, although it would clearly be of interest also to model the global linkages between the euro area and other economies (as in Dees et al., 2007, and Boschi and Girardi, 2011, among others).

² The exposition refers to a model of autoregressive order one, as suggested by the standard information criteria and by the diagnostic tests discussed in Section 3.1 below.

The foreign variables \mathbf{x}_{it}^* are weighted averages of the variables of the rest of the euro area with country-specific weights, w_{ij} , given by trade shares, i.e. the share of country j in the total trade of country i over the years 2004-2006, measured in US dollars. Thus a generic foreign variable x_{it}^* is given by:

$$x_{it}^* = \sum_{j=0}^N w_{ij} x_{jt} \quad (2)$$

where $w_{ii} = 0$, $\forall i = 0, 1, \dots, N$ and $\sum_{j=0}^N w_{ij} = 1$, $\forall i, j = 0, 1, \dots, N$. In our set-up, all foreign variables included in the vector \mathbf{x}_{it}^* are treated as long-run forcing variables.

For each country i , with $i = 1, \dots, 10$, a VEC model (1) is estimated, where the vector of endogenous variables, \mathbf{x}_{it} , includes lr_t , ry_t , dp_t and db_t , denoting nominal long-term rates, real output, the (expected) inflation rate and the debt/GDP ratio respectively; the vector of country-specific foreign variables, \mathbf{x}_{it}^* , includes lr_t^* , ry_t^* , dp_t^* and db_t^* representing the long-term interest rate, real output, the (expected) inflation rate and the debt/GDP ratio, respectively, in the rest of the euro area as well as the 3-month Euribor rate, sr_t , which is treated as a global variable in the GVAR. For the reference country, the vector \mathbf{x}_{0t} , includes lr_t , ry_t , dp_t , db_t and sr_t , whilst the vector of country-specific foreign variables, \mathbf{x}_{0t}^* , includes only lr_t^* , ry_t^* , dp_t^* and db_t^* . The chosen variables are similar to those employed in previous works (Bandholz et al., 2009; Chinn and Frankel, 2007; Paesani et al., 2006) and are consistent with a theoretical framework based on the loanable funds approach to the determination of interest rates. The inclusion of the Euribor rate as a global variable aims at capturing the common monetary policy shared by the EMU countries. In particular, our specification implies that the Euribor rate is exogenous in all models but the one for Germany,

in order to emphasise the polar role of the German economy within the euro area.³

2.2. The GVAR system

Rather than estimating directly the complete system comprising the $N+1$ country-specific models (1) together with (2), we estimate the parameters of each country-specific model separately and then stack the coefficient estimates in a GVAR model. All country/region-specific endogenous variables are collected in the $(k \times 1)$ global vector $\mathbf{x}_t = (\mathbf{x}'_{0t}, \mathbf{x}'_{1t}, \dots, \mathbf{x}'_{Nt})'$ where $k = \sum_{i=0}^N k_i$. Then we have that $\mathbf{z}_{it} = \mathbf{W}_i \mathbf{x}_t$, where \mathbf{W}_i is the $(k_i + k_i^*) \times k$ matrix collecting the trade weights w_{ij} , $\forall i, j = 0, 1, \dots, N$.

Therefore, for each country the following VAR representation of model (1) is obtained:

$$\mathbf{A}_i \mathbf{W}_i \mathbf{x}_t = \mathbf{a}_{i0} + \mathbf{a}_{i1}t + \mathbf{B}_i \mathbf{W}_i \mathbf{x}_{t-1} + \boldsymbol{\varepsilon}_{it} \quad (3)$$

where \mathbf{A}_i and \mathbf{B}_i are matrices of dimension $k_i \times (k_i + k_i^*)$ and matrix \mathbf{A}_i has full row rank.

Stacking the $N+1$ systems (3) yields the following GVAR in level form:

$$\mathbf{G} \mathbf{x}_t = \mathbf{a}_0 + \mathbf{a}_1 t + \mathbf{H} \mathbf{x}_{t-1} + \boldsymbol{\varepsilon}_t \quad (4)$$

where \mathbf{G} is a $k \times k$ full rank matrix, $\mathbf{a}_h = (\mathbf{a}_{0h}, \dots, \mathbf{a}_{Nh})'$ for $h = 0, 1, 2$,

$\mathbf{G} = (\mathbf{A}_0 \mathbf{W}_0, \dots, \mathbf{A}_N \mathbf{W}_N)'$, $\mathbf{H} = (\mathbf{B}_0 \mathbf{W}_0, \dots, \mathbf{B}_N \mathbf{W}_N)'$, for $h = 0, 1$. The GVAR has the

reduced form:

$$\mathbf{x}_t = \mathbf{b}_0 + \mathbf{b}_1 t + \mathbf{F} \mathbf{x}_{t-1} + \mathbf{u}_t \quad (5)$$

where $\mathbf{b}_h = \mathbf{G}^{-1} \mathbf{a}_h$ for $h = 0, 1, 2$, $\mathbf{F} = \mathbf{G}^{-1} \mathbf{H}$, and $\mathbf{u}_t = \mathbf{G}^{-1} \boldsymbol{\varepsilon}_t$.

As pointed out by Pesaran et al. (2004), three conditions need to be fulfilled to ensure

³ Views differ on what fiscal variable (actual/projected public consumption/deficit/debt) should be employed to proxy fiscal positions (Laubach, 2009). As in Paesani et al. (2006) and Baldacci and Kumar (2010), we use the debt/GDP ratio since we would argue that the stock of debt is more informative than the flow of fiscal imbalances in the present context. For instance, in the most recent years, Italy has recorded a relatively better budgetary position with respect to its EMU partners. Notwithstanding this, its very high debt/GDP ratio remains a weakness of that economy.

that the GVAR estimation procedure is equivalent to the simultaneous estimation of the VAR model of the global economy. First, model (5) must be dynamically stable, i.e. the eigenvalues of matrix F must lie either on or inside the unit circle. Second, the trade weights must be such that $\sum_{j=0}^N w_{ij}^2 \rightarrow 0$ as $N \rightarrow \infty$, $\forall i$. Third, the cross-dependence of the idiosyncratic shocks must be sufficiently small, so that $\frac{1}{N} \sum_{j=0}^N \sigma_{ij,ls} \rightarrow \infty$, $\forall i, l, s$, where $\sigma_{ij,ls} = \text{cov}(\epsilon_{ilt}, \epsilon_{jst})$ is the covariance of the l^{th} variable in country i with the s^{th} variable in country j .

The dynamic properties of the GVAR model in (5) are then analysed by using the Generalized Forecast Error Variance Decomposition (GFEVD) and Generalized Impulse Response Functions (GIRFs) developed by Pesaran and Shin (1998). Although these methods do not allow a structural interpretation of the shocks, they overcome the identification problem by providing a meaningful characterisation of the dynamic responses to observable shocks.

3. Estimation results

3.1. Preliminary analysis and country model estimates

Quarterly seasonally adjusted series are used over the sample period 1999:1-2010:4 and for the following countries: Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, the Netherlands, Portugal and Spain.⁴ The Appendix provides details of the data sources and variable definitions.

The matrix of trade weights used to construct the country-specific foreign variables is reported in Table 1, where the 2004-2006 trade shares are displayed by country in each

⁴ The year 1999 marks the beginning of the euro area for all countries but Greece, who joined in 2001. To simplify the analysis, we assume euro membership for Greece as well since 1999. Given the small size of its economy relative to the euro area as a whole, this should not affect the results significantly. The same applies to the omission of later entrants (namely Slovenia, Cyprus, Luxembourg and Malta).

column. The vast majority of the weights are small. The biggest values are those for Austria vis-à-vis Germany and for Portugal vis-à-vis Spain (0.656 and 0.424, respectively).

[Table 1]

As a preliminary step, standard ADF unit root tests are performed. The ADF results based on the Schwarz Bayesian criterion (SBC) for lag selection (with the maximum lag length set equal to four) as well as those from the DF-GLS tests (Elliot et al., 1996) indicate that the unit root null hypothesis cannot be rejected at conventional significance levels. The KPSS (Kwiatkowski et al., 1992) stationarity tests corroborate these conclusions. On the other hand, differencing the series appears to induce stationarity.⁵

We set the lag length of the endogenous and exogenous variables, p_i , by combining standard selection criteria, namely the Akaike information criterion (AIC), and the SBC. The latter suggests order one for all models except Italy, whilst the former selects order one for Austria, Portugal and Germany and two for the remaining models. Therefore, taking into account the limited sample size compared to the number of unknown parameters in each VARX* model, where X* indicates foreign exogenous variables, the lag length p_i is set equal to 1. In order to choose the lag length of the foreign specific variables, q_i , an unrestricted VAR is estimated for each country in which the foreign variables are treated as endogenous, obtaining similar results. Given this evidence, we set q_i equal to one in all models. The selected lag length produces data congruent models, with the null hypotheses of no serial correlation and normality being rejected, at the 1 percent confidence level, only in 6 and 7 out of 45 cases, respectively.

Given the $I(1)$ -ness of the series, testing for cointegration is the next logical step. In particular, the VAR specification considered here is model IV according to the notation in

⁵ Results for unit root and stationarity tests as well as for univariate mis-specification tests (including stability tests) are not reported for reasons of space but are available on request.

Pesaran et al. (2000), where a linear deterministic trend is implicitly allowed for in the cointegration space, but can be eliminated in the dynamic part of the VEC model. Given the well-known limitations of the trace and (especially) the maximum eigenvalue test statistics in small samples, the number of cointegration relationships in the matrix Π_i in (1) has been determined by bootstrapping techniques. Since testing for cointegration basically implies imposing restrictions on an otherwise unrestricted VAR, we performed for each country model a battery of likelihood-ratio tests for the hypotheses $r_i = \varpi_i - 1$ against the alternative $r_i = \varpi_i$, starting from $\varpi_i = k_i$ to $\varpi_i = 1$. Table 2 reports the p-values computed with 1000 replications, with the selected rank in bold.

[Table 2]

The testing procedure indicates a cointegration rank of 2 for Belgium, Finland and Greece, of 1 for Austria, France, Portugal and of 0 for Italy and the Netherlands at the 10 percent significance level. For the other models (Spain, Finland and Germany), the results are less clear-cut. After considerable experimentation, a rank of 1 was chosen for Spain and Finland, whilst we set $r = 0$ in the case of Germany in order to have a more stable GVAR.⁶

3.2. *Properties of the GVAR*

Since in the GVAR the total number of endogenous variables is 45 and that of the cointegrating relations is at most 9,⁷ it follows that matrix F in equation (5) must have at least $45-9=36$ eigenvalues that fall on the unit circle in order to ensure stability of the global model. Our results confirm this; the matrix F estimated from the country-specific models has exactly 36 eigenvalues falling on the unit circle, while the remaining 9 are all less than one (in

⁶ Although the GVAR model is an a-theoretical framework, it can incorporate structural restrictions (e.g. Dees et al., 2007), but given the focus of this work on cross-country short- to medium-run linkages, we consider only unrestricted specifications. The finding of a zero rank is not a concern, since our aim is not to identify long-run relationships, but to analyse the dynamic effects of public debt accumulation on long-term interest rates.

⁷ That is the sum of the ranks of matrix Π_i in equation (1) for each country $i = 0, \dots, N + 1$ (Pesaran et al., 2004).

modulus). The long-run properties of the GVAR are also analysed by looking at their persistence profiles (Pesaran and Shin, 1996), which makes it possible to assess how long the system takes to revert to its steady-state path, after being hit by a system-wide shock.⁸ Figure 1 shows the absorption path of deviations from the steady-state, over a simulation horizon of 20 quarters (5 years). In all cases, we observe convergence towards the steady-state, with the adjustment process being completed within the fifth year of the simulation.

[Figure 1]

A second key assumption of the GVAR approach is that idiosyncratic shocks are cross-sectionally weakly correlated. The basic idea is that conditioning the estimation of country-specific VEC models on foreign variables considered as proxies of “common” global factors will leave only a modest degree of correlation between the remaining shocks across countries. This is also important if one wants to interpret the disturbances in the simulation analysis as “geographically structural”: an external shock is truly external if its contemporaneous correlation with internal shocks is weak. In order to verify this, contemporaneous correlations of residuals across different country-specific models for each equation are computed. The hypothesis that these coefficients, computed as averages of the correlation coefficients between the residuals of each equation (variable) with all other countries’ equation residuals, are significantly different from zero at the 5 percent level is rejected in all cases (Table 3). This suggests that the model is successful in capturing the effect of common factors driving the domestic variables.

[Table 3]

A third econometric issue concerns the weak exogeneity assumption. Following

⁸ Note that the persistence profiles are independent of the way in which the shocks are orthogonalised or the ordering of the variables and equations in the VAR model. By construction these profiles, normalised to unity on impact, should tend to zero as the number of simulation periods increases only if the cointegration vector is stationary, while in the case of $I(1)$ (or “near integrated”) series they can be different from zero for a long period of time.

Johansen (1992), we test the joint significance of the error correction terms in auxiliary equations of the country-specific foreign variables and the Euribor rate. Specifically, we run the following regression for each l -th element of country i 's vector of foreign variables, \mathbf{x}_{it}^* and for the Euribor rate:

$$\Delta \mathbf{x}_{il,t}^* = \mu_{il} + \sum_{j=1}^{r_i} \phi_{ijl} ECM_{i,t-1}^j + \theta_{il} \Delta \mathbf{x}_{i,t-1} + \gamma_{il} \Delta \mathbf{x}_{i,t-1}^* + \zeta_{il,t}$$

where $ECM_{i,t-1}^j$, $j = 1, \dots, r_i$ are the estimated error correction terms corresponding to the r_i cointegrating relations found in the i^{th} model. The results from an F -test of the joint hypothesis that $\phi_{ijl} = 0$, $j = 1, \dots, r_i$, indicate that test statistics are not significant at the 5 percent level. Given the overall statistical evidence, foreign variables and the Euribor rate are treated as weakly exogenous.⁹

[Table 4]

3.3. Impact elasticities

In order to provide a first assessment of the international linkages between domestic and foreign variables, Table 5 reports the estimates of the contemporaneous variation of a domestic variable due to changes in its foreign counterpart, with White's heteroscedasticity-consistent t -ratios in parentheses.

[Table 5]

For the main variable of interest, lr , it emerges that a 1 percent change in foreign long-term yields in a given quarter leads to a similar increase in domestic yields. This evidence holds for all countries but Greece. Our findings indicate strong (and proportional) comovements between each EMU country's yields and those of its EMU partners, suggesting a high degree of integration as previously documented by Claeys et al. (2008) and

⁹ The weak exogeneity assumption is also satisfied for the country models with no cointegration relationships, since no error correction terms appear.

Georgoutsos and Migiakis (2010), among others. Moreover, the estimated coefficients for Austria and the Netherlands are above one, suggesting an over-reaction to foreign bond market changes. The contemporaneous variations in the public debt/GDP ratio indicate a high degree of heterogeneity and weak synchronisation of fiscal positions across EMU countries: the strongest positive comovements are those for highly indebted countries (Ireland and Italy), whilst the elasticity for the best performing country in terms of public debt/GDP ratio (Finland) is negative and statistically significant. As for real growth, the estimated coefficients are all positive and statistically significant (except for Ireland), reflecting the high degree of trade openness within the EMU. Finally, we find positive and statistically significant elasticities for (expected) inflation. It is noteworthy that smaller economies (Finland, Greece, Ireland and Portugal) have higher elasticities. This is consistent with a transmission channel of inflation working mostly from large to small countries (Eun and Jeong, 1999).

4. Dynamic simulations

Although informative, elasticities do not shed light on the dynamic interlinkages between the equations of the GVAR model. In what follows we aim at: a) quantifying the main sources of interest rate fluctuations by distinguishing between domestic and foreign driving factors; b) analysing the dynamic trajectories of long-term interest rates after an unanticipated increase in the amount of paper issued (with the associated higher debt/GDP ratio) by the three largest economies (France, Germany and Italy) and intra-EMU flight-to-quality and/or flight-from-risk phenomena (represented by shocks to long-term interest rates). Specifically, the bulk of our dynamic simulations is conducted using the Generalized Forecast Error Variance Decomposition (GFEVD) and Generalized Impulse Response Functions (GIRFs), over a simulation horizon of 20 quarters.¹⁰

¹⁰ We are aware that most of the responses appear to be statistically insignificant even using 68 percent significance level confidence intervals. This is not surprising, however, given the fact that the model is estimated

Since we are interested in analysing possible asymmetries between peripheral and core EMU member states, we report the results based on the geographical transmission of shocks by considering the three largest economies and two regional aggregates: “Periphery” (which includes Greece, Ireland, Portugal, and Spain) and “Other” (which comprises Austria, Belgium, Finland and Netherlands).¹¹

4.1. Generalized Forecast Error Variance Decompositions

Table 6 reports the GFEVD for each EMU country’s long-term interest rate. Panel [A] and Panel [B] show the contribution of country/region-specific domestic and foreign shocks, respectively, whilst Panel [C] reports the overall contribution of domestic versus foreign contribution to each country’s long-term yields fluctuations.¹²

[Table 6]

A mixed picture of the domestic determinants of long-term interest rates variability emerges (Panel [A]). The role of macroeconomic and fiscal factors in the “Other” group increases substantially over the simulation span. Moreover, fiscal factors turn out to be the most important source of variability (around 8 percent) towards the end of the simulation horizon. This pattern is reversed for France and the “Periphery” aggregate: here real output and public debt are the main determinants of short-run government bond fluctuations, but their role tends to decrease over time (from 24 to 7 percent for France, and from 16 to 6 percent for the “Periphery” region). By contrast, inflation and government bond yields are the

using only a few dozens of observations. In order to compute more meaningful dynamic responses, we conduct simulations based on the median values of the statistics of interest computed with 1000 bootstrap replications of the GVAR model (5).

¹¹ In order to obtain a synthetic measure for regional GFEVD and GIRFs, we aggregate the country-specific results into regional ones by using GDP in PPP weights for country-specific shocks, such that their sum adds up to one standard error (see the Appendix).

¹² Even though GFEVD provides information on the role of domestic versus foreign shocks in determining the variability of long-term interest rates, the presence of contemporaneous correlations among innovations means that it has to be rescaled so that the sum of the variance decompositions is normalised to 100 as in Boschi and Girardi (2011).

main domestic driving factors over the whole forecast horizon for Germany and Italy, although their contribution is of a limited magnitude (from 5 to 8 percent).

As for foreign shocks (Panel [B]), common patterns for all country/region models can be detected. First, the contribution of real output and inflation expectations to total variability is around 12-14 and 18-19 percent, respectively. These percentages are more or less constant over the simulation span, except for the cases of real output in the case of Germany (where it decreases from 24 to 14 percent) and the “Periphery” region (where it increases from 10 to 14 percent). Second, foreign debt and interest rates are the main sources of total variability (except for Germany): the share for foreign debt ranges from 20 to 25 percent, whilst foreign interest rates explain roughly one-fifth of domestic long-term yield variability. For the reference economy (i.e. Germany), the role of foreign debt is still sizeable (around 18 percent) and increases over the simulation horizon. The contribution of foreign interest rates has instead a hump-shaped pattern, reaching its maximum in the first quarter (around 29 percent) and then converging to a value (22.5 percent) slightly lower than the one on impact (24.7 percent).

Concerning the relative importance of foreign and domestic factors (Panel [C]), it appears that the former contribute far more than the latter to domestic long-term interest rate variability. This is true for all countries/regions at all horizons, the difference between the percentage contribution of foreign shocks and that of domestic ones ranging from 46 percent for Germany to 67 percent for Italy towards the end of the simulation horizon. Interestingly, Italian bond yields are more affected by external developments than those in the “Periphery”. Moreover, except for the case of the “Other” aggregate, the relevance of foreign factors tends to increase over time. This finding is in contrast to previous studies reporting a dominant role for domestic factors relative to foreign ones (for instance, Breedon et al., 1999, for the G-3 bloc; Ardagna et al., 2007, for a number OECD economies; Chinn and Frenkel, 2003, for selected European economies). It supports instead the argument of Ford and Laxton (1999)

according to whom with integrated capital markets it is aggregate debt (of the EMU region as a whole in the present context) that matters for the determination of country/region-specific interest rates.

4.2. Generalised Impulse Response Functions

To investigate further the role of fiscal developments in determining interest rates dynamics, we show in Figure 2 the responses of (real) long-term yields to a (one positive standard error) shock to the debt/GDP ratio in the three largest economies in the EMU region.¹³ A negative response is estimated for all countries/regions after a shock originating in France and Germany. These results are consistent with a “liquidity effect” (Caporale and Williams, 2002; Ardagna et al., 2007) according to which if sovereign debt in some countries (France and Germany) is viewed as less risky than issuances of other EMU countries, the overall debt stock of euro-denominated government securities reduces the aggregate risk premium and thus the interest rate for their EMU partners. At the end of the simulation horizon, the decrease in yields is in the range of 25-35 basis points after a shock from Germany, whilst the liquidity effect for France is less pronounced (around 5-10 basis points). By contrast, when the same shock originates from the “Periphery” and Italy, country risk considerations dominate liquidity effects: a worsening of the fiscal position in Italy increases debt financing costs for its EMU partners by up to 15 basis points in the first year of the simulation and about 10 basis points at the end of the simulation span. An unexpected increase in public debt in the “Periphery” has similar effects, although of a smaller size. Moreover, there is a decrease in German real interest rates, indicating flight-to-quality. The dynamic responses of interest rates thus suggest that financial markets are able to discriminate among different issuers, and that the same policy action will have different effects depending on the identity of the debtor. This is consistent with the results by Ardagna et al. (2007) and Baldacci and

¹³ Dynamic responses for real interest rates are computed as differences between the GIRFs for nominal yields and (annualised) expected inflation.

Kumar (2010) according to which financial markets behave in nonlinear ways, requiring fiscal discipline for borrowers when fiscal imbalances exceed a given threshold.

[Figure 2]

Further evidence on possible intra-EMU flight-to-quality and/or flight-from-risk phenomena is provided by Figure 3, where the effects of a (one negative standard error) shock to interest rates in Germany and France and of a (one positive standard error) shock to Italian long-term yields on the (real) borrowing costs faced by EMU members are shown. An unanticipated fall in bond yields in France and (especially) in Germany tends to lower real debt financing costs (up to 40 basis points) in all countries/regions in the first year of the simulation; subsequently, real yields for Germany, France and the Other aggregate remain below the baseline, whilst those for Italy and the “Periphery” region revert to their pre-shock values. A similar picture emerges when a positive shock to government yields in Italy is considered. The medium-run effect of an unexpected fall in the demand for Italian Treasury securities translates into higher real borrowing costs for Italy and the peripheral countries, with no effects on the other countries. Finally, a flight-from-risk effect for the “Periphery” region translates into higher real yields for all countries/regions. This is especially true for the peripheral countries and for Italy. Therefore crowding out effects can also arise if investors suddenly change the weights of high- and low-quality euro-denominated Treasury securities held in their portfolios.

[Figure 3]

5. Conclusions

This paper examines how fiscal deterioration affects long-term interest rates; specifically, it focuses on the dynamics effects of fiscal imbalances in a given EMU country on the borrowing costs faced by its EMU partners. To investigate this issue, we estimate a dynamic multi-country model (specifically, a Global VAR, or GVAR) using quarterly data for the EMU period.

We document strong linkages between long-term interest rates in the presence of a single pool of funds (the euro-denominated market for Treasury securities) for which all EMU countries compete. However, in contrast to a widely held view, our estimates show that the percentage of variability in long-term interest rates of EMU countries explained by domestic factors is modest relative to that accounted for by foreign (that is, EMU aggregate) shocks. Moreover, financial markets discriminate among different issuers: increasing demand for new issues from France and Germany tends to reduce interest rate pressures in the other country/regions in the euro area; the opposite holds when new paper is issued in Italy or in peripheral countries. We also show that unanticipated flight-to-quality and/or flight-from-risk phenomena lead to widening spreads relative to the core Europe economies for those with the weakest macroeconomic fundamentals.

Overall our empirical evidence gives support to the need for fiscal discipline as envisaged in the Stability Growth Pact since negative externalities imposed by fiscal imbalances in Italy and in the peripheral countries can result in possible crowding out effects in all other countries/regions where the real interest rates would otherwise be lower. We also show that these effects can occur when investors suddenly rebalance their portfolios by increasing the share of high-quality euro-denominated Treasury securities and lowering their exposure in highly risky securities issued by peripheral countries. In this respect, the unconventional policies recently undertaken by the European Central Bank might be viewed as (temporary) useful measures aimed at reducing these adverse effects on real debt financing costs for EMU member states. From a more general perspective, the evidence of close linkages between EMU economies suggests that fiscal developments in a given country are likely to reverberate throughout the euro area. Therefore, greater co-ordination among EMU member states aimed at tackling worsened issuance conditions and liquidity evaporation in the secondary markets for euro-denominated government securities should be pursued, especially in the present period of financial turmoil.

Our analysis could be extended to include economies outside the euro area in order to assess whether core and peripheral EMU countries react symmetrically to external fiscal shocks. Moreover, our findings are based on a non-structural model. Future research should focus on the structural identification of shocks within a global model of the world economy to gain a deeper economic understanding of fiscal linkages.

Appendix

The source for real output (YCC), consumer price index (CPI) and long-term interest rates (LTR) data is the OECD Main Economic Indicators database. The Government debt/GDP ratio (GDY) and the 3-month Euribor rate (EUR) series are taken from Eurostat Quarterly Government Finance Statistics and ECB Statistical Data Warehouse, respectively. The variables used in the estimation of each country-specific VEC model are constructed from the series above as follows: $ry_t = \ln[100 \cdot (YCC_t / YCC_{2005})]$, $lr_t = \ln(1 + LTR_t / 100)$, $db_t = GDY_t / 100$, $dp_t = f[\ln(CPI_t / CPI_{t-1})]$, where YCC_{2005} denotes the 2005 average for real output and $f[.]$ indicates the permanent component computed with the Holt-Winters' method.

The regional aggregates are constructed as weighted averages of the corresponding series of individual countries. The “Periphery” region comprises Greece, Ireland, Portugal, and Spain, whilst the “Other” group includes Austria, Belgium, Finland and Netherlands. In both cases, the weights are each country's average shares of the regional aggregate's real GDP in PPP over the period 1999-2009. The real GDP in PPP series are obtained from OECD National Accounts. The country weights for the “Periphery” aggregate are 0.602 for Spain, 0.135 for Greece, 0.131 for Ireland and 0.132 for Portugal. Those for the “Other” region are 0.197 for Austria, 0.281 for Belgium, 0.129 for Finland and 0.393 for the Netherlands.

The computations are carried out using R routines (unit root and stationarity tests; structural stability tests) and customised versions of the Matlab codes provided by A. Warne (bootstrapping procedures to the determination of the cointegration ranks) and by A. Galesi and V. Smith (GVAR estimation and simulations).

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Table 1. Trade weights

	Aut	Bel	Spa	Fin	Fra	Gre	Ire	Ita	Net	Por	Ger
Aut	0	0.013	0.016	0.028	0.019	0.023	0.012	0.055	0.020	0.009	0.120
Bel	0.028	0	0.056	0.074	0.156	0.069	0.280	0.077	0.215	0.047	0.126
Spa	0.036	0.047	0	0.066	0.171	0.083	0.076	0.128	0.058	0.424	0.099
Fin	0.012	0.010	0.010	0	0.012	0.021	0.013	0.012	0.022	0.009	0.027
Fra	0.071	0.238	0.299	0.116	0	0.125	0.151	0.242	0.141	0.157	0.231
Gre	0.006	0.006	0.011	0.013	0.011	0	0.008	0.029	0.010	0.004	0.014
Ire	0.010	0.057	0.021	0.024	0.026	0.017	0	0.020	0.023	0.011	0.035
Ita	0.133	0.070	0.161	0.108	0.180	0.264	0.102	0	0.080	0.076	0.157
Net	0.041	0.241	0.069	0.150	0.083	0.104	0.123	0.086	0	0.061	0.171
Por	0.007	0.009	0.107	0.015	0.023	0.007	0.012	0.017	0.012	0	0.019
Ger	0.656	0.308	0.251	0.405	0.319	0.286	0.225	0.334	0.419	0.201	0

Notes. Trade weights, computed as shares of exports and imports in 2004-2006, are displayed in each column by country. Columns, but not rows, sum up to one. The countries considered in the analysis are: Austria (Aut), Belgium (Bel), Spain (Spa), Finland (Fin), France (Fra), Greece (Gre), Ireland (Ire), Italy (Ita), Netherlands (Net), Portugal (Por), and Germany (Ger).

Table 2. Determination of the cointegration rank

		Hypotheses to test				
		$r_i = 4$ vs $r_i = 5$	$r_i = 3$ vs $r_i = 4$	$r_i = 2$ vs $r_i = 3$	$r_i = 1$ vs $r_i = 2$	$r_i = 0$ vs $r_i = 1$
Aut	.	[0.798]	[0.828]	[0.566]	[0.004]	[0.004]
Bel	.	[0.471]	[0.233]	[0.009]	[0.000]	[0.000]
Spa	.	[0.697]	[0.707]	[0.424]	[0.062]	[0.000]
Fin	.	[0.253]	[0.151]	[0.019]	[0.010]	[0.010]
Fra	.	[0.929]	[0.444]	[0.374]	[0.011]	[0.011]
Gre	.	[0.556]	[0.263]	[0.002]	[0.000]	[0.000]
Ire	.	[0.626]	[0.438]	[0.179]	[0.087]	[0.000]
Ita	.	[0.828]	[0.364]	[0.242]	[0.131]	[0.000]
Net	.	[0.828]	[0.657]	[0.404]	[0.242]	[0.000]
Por	.	[0.879]	[0.838]	[0.667]	[0.014]	[0.014]
Ger	[0.495]	[0.250]	[0.181]	[0.117]	[0.042]	[0.042]

Notes. p -values from a χ^2 -test for testing sequentially the null of reduced rank regression in square brackets.

r_i is the rank of the long-run matrix Π_i in (1). Values in bold indicate the chosen cointegration rank.

Table 3. Contemporaneous Effects of Foreign Variables on Domestic Counterparts

	<i>lr</i>	<i>gd</i>	<i>yr</i>	<i>cp</i>
Aut	1.044 [0.020]	1.218 [0.864]	0.610 [0.187]	0.585 [0.101]
Bel	0.995 [0.021]	0.706 [0.285]	0.584 [0.170]	0.375 [0.055]
Spa	0.983 [0.046]	0.743 [0.168]	0.628 [0.153]	1.127 [0.128]
Fin	1.085 [0.023]	-1.137 [0.201]	1.216 [0.261]	2.032 [0.273]
Fra	1.027 [0.017]	0.349 [0.119]	0.646 [0.135]	1.151 [0.112]
Gre	0.280 [0.331]	0.458 [0.116]	0.657 [0.278]	2.057 [0.161]
Ire	1.129 [0.207]	1.426 [0.408]	1.072 [0.648]	4.765 [0.591]
Ita	0.925 [0.051]	1.224 [0.384]	1.022 [0.108]	0.881 [0.115]
Net	1.056 [0.024]	0.505 [0.214]	0.678 [0.101]	1.104 [0.166]
Por	1.050 [0.060]	-0.207 [0.366]	0.776 [0.240]	2.096 [0.161]
Ger	0.949 [0.048]	0.216 [0.090]	1.250 [0.225]	0.398 [0.077]

Notes. White's heteroscedasticity-consistent *t*-ratios in square brackets.

Table 4. Average Pairwise Cross-Section Correlations

	<i>lr</i>	<i>gd</i>	<i>yr</i>	<i>cp</i>
Aut	[0.450]	[0.799]	[0.615]	[0.664]
Bel	[0.479]	[0.940]	[0.852]	[0.799]
Spa	[0.545]	[0.793]	[0.741]	[0.635]
Fin	[0.705]	[0.923]	[0.874]	[0.783]
Fra	[0.187]	[0.929]	[0.705]	[0.690]
Gre	[0.505]	[0.259]	[0.831]	[0.559]
Ire	[0.978]	[0.934]	[0.885]	[0.940]
Ita	[0.383]	[0.577]	[0.973]	[0.901]
Net	[0.438]	[0.762]	[0.842]	[0.825]
Por	[0.550]	[0.852]	[0.995]	[0.751]
Ger	[0.159]	[0.625]	[0.336]	[0.285]

Notes. *p*-values from a two-tailed *t*-test statistic with 48 d.o.f. in square brackets. The null hypothesis is no correlation.

Table 5. Weak exogeneity tests

	F test	lr^*	gd^*	yr^*	cp^*	er
Aut	F(1,34)	[0.883]	[0.567]	[0.528]	[0.598]	[0.904]
Bel	F(2,33)	[0.726]	[0.753]	[0.296]	[0.691]	[0.549]
Spa	.			.		
Fin	F(2,33)	[0.781]	[0.643]	[0.508]	[0.783]	[0.963]
Fra	F(1,34)	[0.205]	[0.634]	[0.209]	[0.779]	[0.834]
Gre	F(2,33)	[0.571]	[0.406]	[0.719]	[0.801]	[0.507]
Ire	.			.		
Ita	.			.		
Net	.			.		
Por	F(1,34)	[0.360]	[0.187]	[0.495]	[0.110]	[0.593]
Ger	.			.		

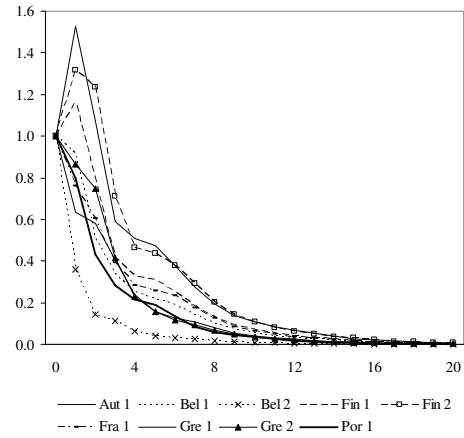
Notes. p -values from a F -test statistics with degrees of freedom reported in the first column in square brackets.

Table 6. Generalised Forecast Error Variance Decomposition

Horizon	Panel [A]					Panel [B]					Panel [C]	
	<i>lr</i>	<i>gd</i>	<i>yr</i>	<i>cp</i>	<i>er</i>	<i>lr</i> *	<i>gd</i> *	<i>yr</i> *	<i>cp</i> *	<i>er</i>	All domestic factors	All foreign factors
Fra												
0	7.73	14.63	9.53	4.90		24.18	13.09	13.34	9.77	2.83	36.79	63.21
1	8.34	5.97	4.28	9.50		27.62	16.73	6.98	18.76	1.82	28.08	71.92
2	7.15	4.94	3.83	10.39		24.42	19.22	7.30	20.76	1.98	26.31	73.69
4	6.37	4.46	3.94	9.34		22.99	20.27	10.09	19.76	2.78	24.11	75.89
8	6.26	3.82	4.28	8.04		24.07	18.97	12.03	18.87	3.65	22.41	77.59
12	5.94	3.65	4.09	8.34		22.92	19.68	12.11	19.56	3.71	22.02	77.98
20	5.87	3.51	3.94	8.27		23.07	19.93	12.11	19.54	3.77	21.58	78.42
Ger												
0	16.50	3.28	3.27	3.44	3.49	24.71	12.08	24.52	8.71		29.98	70.02
1	13.18	5.42	2.05	4.86	2.42	29.34	12.71	12.10	17.92		27.94	72.06
2	10.39	6.74	2.36	6.04	2.71	24.59	14.93	11.93	20.32		28.23	71.77
4	8.46	6.00	2.89	6.13	3.66	23.53	16.06	13.46	19.82		27.13	72.87
8	7.03	4.86	2.91	6.70	4.78	23.58	16.29	15.11	18.74		26.28	73.72
12	6.63	4.86	2.74	7.47	4.83	22.78	17.52	14.46	18.71		26.53	73.47
20	6.49	4.69	2.62	7.92	5.08	22.55	17.77	14.46	18.41		26.81	73.19
Ita												
0	4.65	4.45	4.94	6.53		21.76	28.05	13.64	13.36	2.62	20.57	79.43
1	4.30	3.13	2.90	4.94		23.01	29.37	10.41	19.89	2.05	15.27	84.73
2	3.69	3.35	2.76	4.29		19.20	32.37	10.93	21.14	2.26	14.09	85.91
4	4.60	3.25	3.01	4.70		19.61	29.26	12.66	19.95	2.96	15.56	84.44
8	5.59	3.07	2.84	5.22		20.99	25.24	13.99	19.33	3.71	16.73	83.27
12	5.55	3.03	2.78	5.18		20.69	25.39	14.10	19.45	3.84	16.53	83.47
20	5.71	2.94	2.64	5.14		20.61	25.46	14.31	19.37	3.83	16.43	83.57
Other												
0	5.49	5.40	3.85	2.83		25.27	24.73	12.69	17.02	2.73	17.57	82.43
1	5.30	4.60	3.46	3.56		21.89	24.09	11.55	22.95	2.61	16.92	83.08
2	4.65	5.70	3.27	3.47		19.10	25.92	10.97	24.24	2.68	17.09	82.91
4	4.77	6.23	3.98	3.73		20.26	22.99	12.25	22.48	3.31	18.71	81.29
8	4.94	6.77	4.83	4.32		21.41	20.33	13.31	20.02	4.07	20.86	79.14
12	4.77	7.43	5.02	4.38		20.79	20.50	13.39	19.59	4.12	21.61	78.39
20	4.70	7.79	5.16	4.53		21.06	20.02	13.42	18.98	4.35	22.18	77.82
Periphery												
0	7.27	7.35	8.25	5.41		17.57	28.82	10.24	12.75	2.34	28.27	71.73
1	5.80	4.59	4.83	5.74		24.04	23.92	10.40	18.08	2.61	20.96	79.04
2	4.92	4.42	4.33	5.76		21.50	25.63	11.04	19.62	2.77	19.43	80.57
4	6.44	4.11	3.78	5.01		21.48	23.96	13.08	18.58	3.55	19.34	80.66
8	8.66	3.32	3.26	4.12		22.67	21.25	14.68	17.75	4.29	19.36	80.64
12	8.95	3.23	3.09	4.05		21.82	21.85	14.52	18.28	4.21	19.32	80.68
20	10.05	2.96	2.86	3.84		22.04	21.43	14.40	18.14	4.29	19.70	80.30

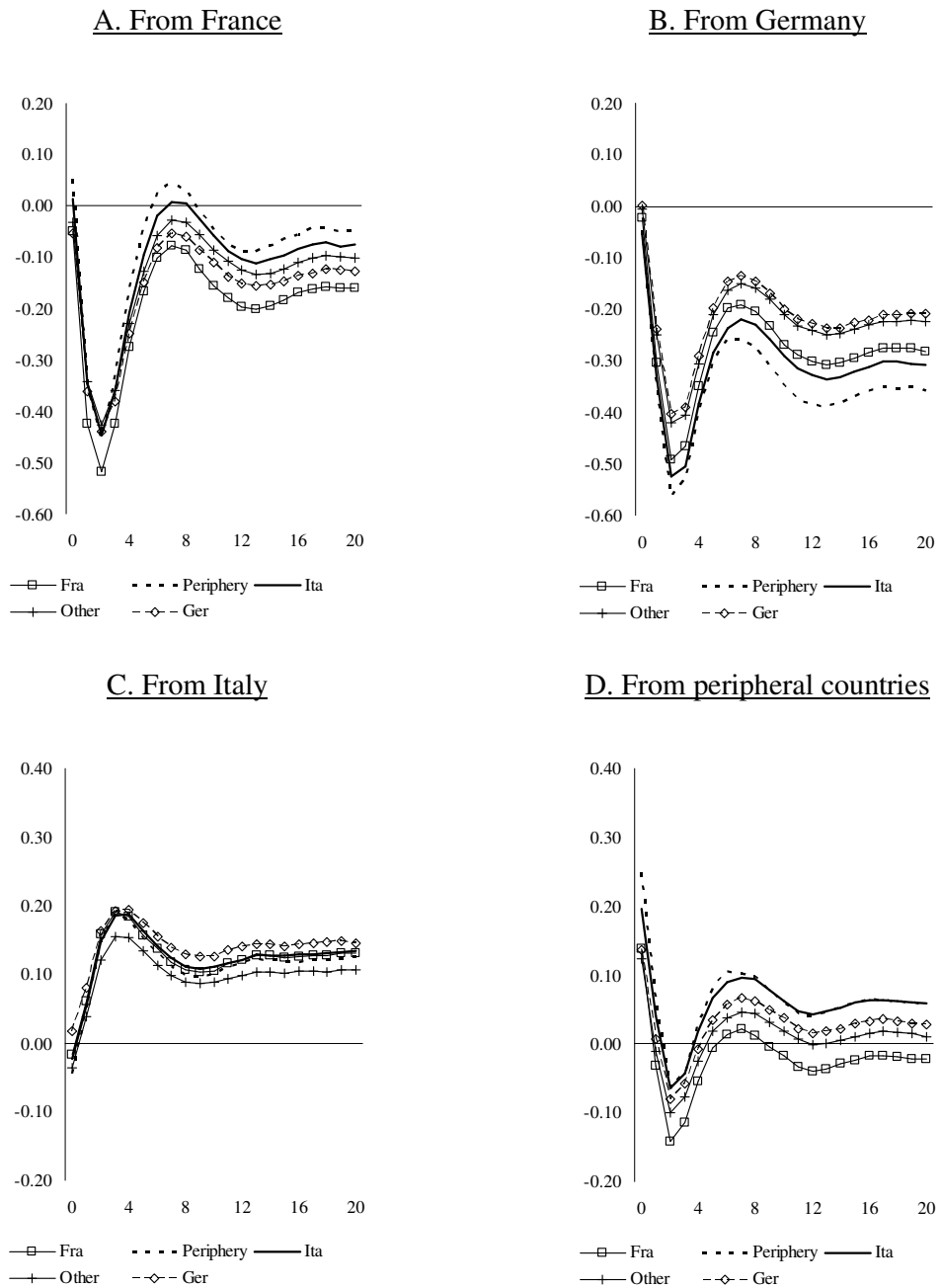
Notes. Share of the k -step ahead GFEVD of domestic long-term yields explained by the shocks in the corresponding column. Entries are normalised so as to sum up to 100. Each entry in columns “All domestic factors” and “All foreign factors” is the row sum of the columns of Panel [A] and Panel [B], respectively.

Figure 1. Persistence profiles



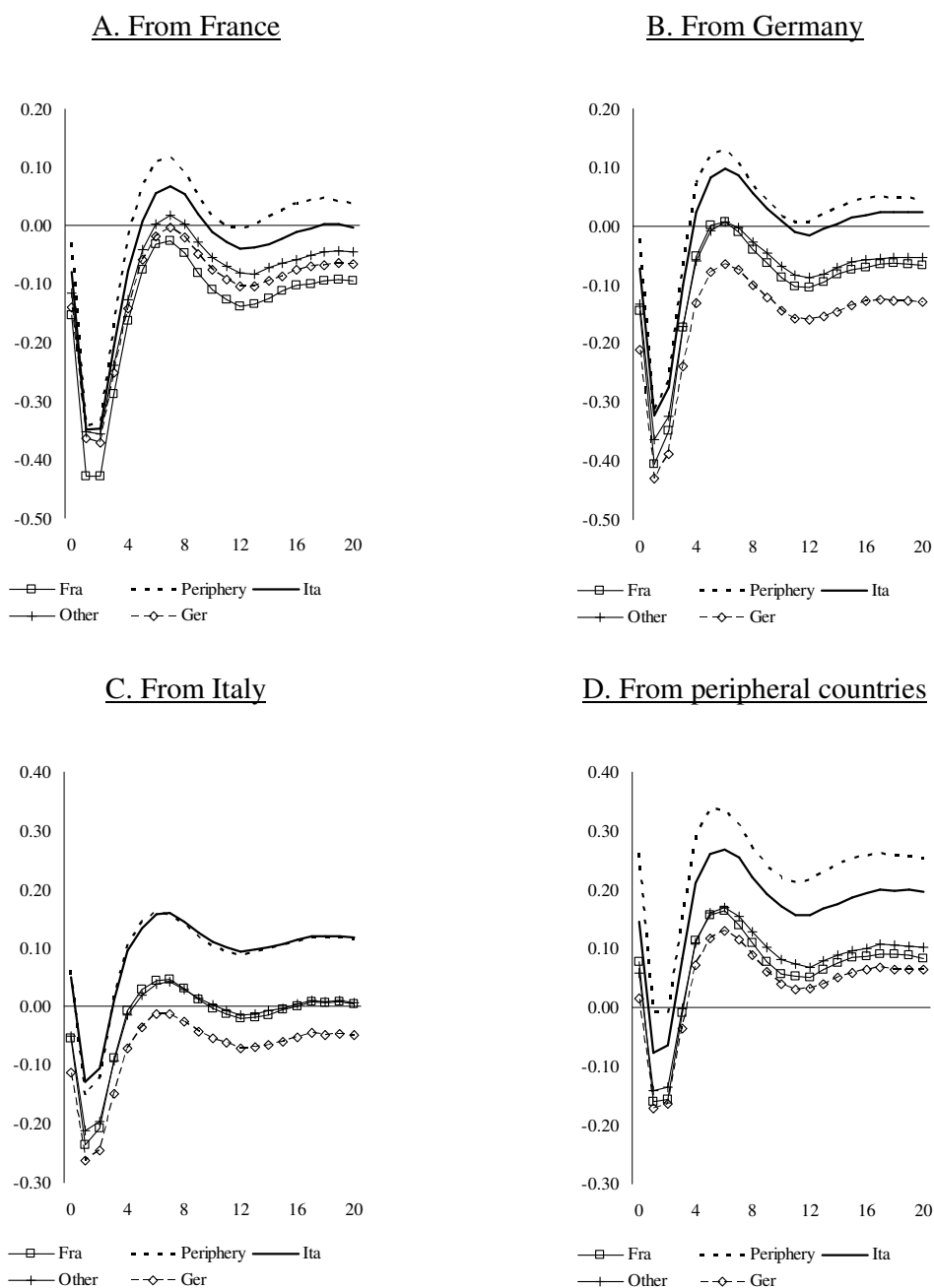
Note. Persistence profiles of the effect of system-wide shocks (normalised to unity on impact) on the cointegrating relationships of the GVAR model. Bootstrap median estimates computed on the basis of 1000 replications.

Figure 2. Shocks to debt



Note. GIRFs of a positive (one standard error) shock to the debt/GDP ratio on real long-term yields, computed as differences between the GIRFs for nominal yields and for (annualised) expected inflation. Bootstrap median estimates computed on the basis of 1000 replications.

Figure 3. Shocks to long-term interest rate



Note. GIRFs of a (one standard error) shock to nominal long-term interest rates on real long-term yields, computed as differences between the GIRFs for nominal yields and for (annualised) expected inflation. The graphs refer to positive shocks in Panel A. and Panel B. and negative shocks for Panel C. and Panel D. Bootstrap median estimates computed on 1000 replications.