Sovereign Debt, Default Risk and Fiscal Consolidation in the EZ Periphery

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Abstract

We consider non zero sovereign and private default probabilities in a monetary, open economy, search and matching model to allow the emergence of a sovereign risk channel and to test whether this financial wedge can reverse the sign of the Keynesian fiscal multiplier, conditional to alternative fiscal consolidation measures. The model is estimated with Bayesian techniques using data of EZ peripheral countries (Greece, Ireland, Italy, Portugal and Spain). From posterior simulations we show that the relation between sovereign risk and macroeconomic fundamentals is weak. Moreover, fiscal contractions are self-defeating, such that the sovereign risk channel amplifies the effects of the fiscal contraction. The consideration of a liquidity trap environment does not reverse, and reinforces, these results.

JEL classification: E62, H25, H30, J20, C11

Keywords: wage and hiring subsidies, search and matching, fiscal multiplier, zero lower bound, Bayesian estimation.

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Introduction

A number of advanced economies, following the global financial crisis, exprienced increases in sovereign debt that were unrecedented during peacetime. Such an evolution, which is still ongoing, has been particularly worrying in the periphery of the euro-zone. Even if different factors are likely to have played a role, the early stages of the sovereign debt surge were characterized by strong uncertainty about sovereign debt sustainability in all the peripheral countries, leading to rising bond and credit rates that worsened the stressed public and private finances. As an immediate result, debt reduction and fiscal consolidation became the major declared goals of European policymakers. Concerns about the risks of contagion (Guerrieri *et al.* 2012) led governments and European institutions to to set-up coordinated measures targeted to gain control over strained public budgets, i.e. to debt reduction and fiscal consolidation.

Despite the general acknowledgement of the fact that, historically, a number of alternative and not mutually exclusive factors played a role in successful debt reductions (Reinart and Sbrancia, 2011)[2], the recently signed Treaty on Stability, Coordination and Governance ("fiscal compact"), to be ratified in national parliaments by the end of 2013, establishes a set of policy measures that are - to a large extent - rooted in the automatic implementation of austerity plans in the case of structural deficits. These fiscal arrangements, backed by the hypothesis of expansionary fiscal contractions (Giavazzi and Pagano, 1990; 1996; Alesina and Perotti, 1997; Alesina and Ardagna, 2010), miss a widespread scientific consensus about their actual effectiveness (Romer and Romer, 2010; Guajardo *et. al.,* 2011; Ramey, 2011), but continue to receive large interest in macroeconomic research.

The hypothesis of a sovereign risk channel, suggested by the observation of a strong correlation between government bond and private sector spreads (Harjes, 2011), has recently provided further support to the idea of expansionary austerity, aside from the concepts of Ricardian equivalence and crowding-out effects of private expenditure.

From the theoretical perspective, Corsetti *et al.* (2013) show that, by modelling the sovereign default risk as an increasing function of the debt *level*, and considering a spillover effect from government bond rates to the private sector's credit conditions, fiscal contractions lead to a reduction of the government expenditure fiscal multiplier. For high levels of public debt and when the economy operates in a liquidity-trap environment, the sign of the multiplier can even be reversed, giving rise to expansionary fiscal contractions. A fiscal contraction, by reducing the level of debt, leads to a reduction in the sovereign default risk, which is translated into reduced bond and lending rates to the private sector. The improved credit conditions, i.e. reduced real interest rates, may stimulate an economic expansion.

The consideration of a sovereign risk channel can thus overturn the results of a recent literature showing that the interaction between fiscal and monetary policy regimes is crucial for the efficacy of the fiscal stimulus, particularly in a liquidity-trap (Christiano *et al.* 2011; Eggertsson, 2011; Eggertsson and Krugman, 2012). According to this literature, if the monetary authority is constrained by a binding zero-lower-bound (ZLB),

[2] These range from sustained economic growth to financial repression with inflationary commitment, from default or restructuring of debt to the implementation of austerity plans.

fiscal contractions - because of their deflationary implications - induce a rise in the real interest rate of the same size of the deflation, leading to a strong economic contraction.

In this paper we develop a monetary model to evaluate the empirical validity of the sovereign risk channel hypothesis and of the related hypotesis of expansionary fiscal contractions. We calculate and compare the country-specific dynamic multipliers of financially equivalent fiscal policies affecting government consumption, transfers and investments on the expenditure side, and direct and indirect consumption taxes on the revenue side. The monetary model is estimated with Bayesian techniques on a large set of data for five major EZ peripheral countries, i.e. Greece, Ireland, Italy, Portugal and Spain (the PIIGS). Policy simulations consider both a standard environment in which the domestic economies operate at their full potential and a non standard liquidity-trap environment, with a binding ZLB.

The model is characterized by the joint consideration, in an otherwise standard closed-economy monetary model with nominal and real imperfections (Christiano et al., 2005; Smets and Wouters, 2007), of some theoretical extensions that are functional to the analysis.

The design of the monopolistically competitive financial sector (Curdia and Woodford, 2010), in which we assume costly Rotemberg pricing and non zero default probabilities on the side of both private and public borrowers, is key for the emergence of the sovereign risk channel. On this respect, we follow the strategy adopted by Corsetti *et al.* (2013) by formalizing a relation between sovereign default probability and interest rate spreads without providing an explicit model of the default event. However, we also substantially depart from their formal setting by assuming a different shape of the cumulative distribution function for the sovereign default probability, partly different economic fundamentals, by considering both the debt and the net foreign asset to GDP *ratios* as arguments of the default probability function, and by explicitly formalizing a private sector default probability. The choice of considering the debt to GDP *ratio* in the place of the debt *level* has two major justifications: on the one hand, it ensures consistency with the empirical literature, addressing economic growth and the ability of the government to service its debt as fundamental triggers of the default risk (Levy-Yeyati and Panizza 2011; Mendoza and Yue, 2012; De Grauwe and Ji, 2013); on the other hand, it highlights the close link between the size of the fiscal multipliers and the sign of the sovereign risk channel effects. In fact, when the former are sufficiently high, the debt to GDP ratio can increase following a fiscal contraction, leading to further deflationary pressure through increased bond and lending rate spreads. The consideration of the net foreign assets position as an important trigger of sovereign default risk is common in the empirical literature (Edwards, 1986; De Grauwe and Ji, 2013). Default episodes are in fact often preceded by large imbalances in the foreign asset position. A fiscal retrenchment, by improving the foreign position through reduced imports, is likely to mitigate the financial pressure of international lenders.

The consideration of a search and matching labor market framework (Diamond, 1982; Mortensen and Pissarides, 1994; Pissarides, 2000) with staggered Nash-wage bargaining (Gertler *et al.*, 2008; Gertler and Trigari, 2009), allows the evaluation of the unemployment implications of the alternative fiscal policies. The formalization of an optimal rule for government investment expenditure decisions ensures that the production potential is optimized. The small open economy framework, developed along the lines of Adolfson *et al.* (2007;

2008) and Christiano *et al.* (2011), in which the foreign sector is described by a structural vector auto-regressive system (SVAR), allows the evaluation of the effects of the policies on the net foreign position.

Results show that the default risk channel is only marginally effective, since the estimated relation between fundamentals and spreads is weak and, most importantly, it operates in the opposite direction than predicted. The reason for the latter result is that, irrespective of the fiscal instrument being considered, the fiscal contraction leads to an increase of the debt to GDP ratio, triggering a rise in default probabilities and interest rate spreads, whilst the improvement in the NFA position to GDP ratio, stimulating a reduction in default probabilities and spreads, is not sufficient to counterbalance the former effect.

The analysis also shows that two key factors are responsible for such result: *First* the low estimated elasticities of the default probability to the debt to GDP and NFA position ratios lead to very small variations in bond and lending rates; *ii)* the relatively high size of the fiscal multipliers implies that a decrease in the debt to GDP ratio is never observed following a fiscal contraction, ruling out even trascurable reductions in the interest rates.

The consideration of a deep recession characterized by a binding ZLB highlights the role played by the monetary policy regime. Results show that, in this situation, the effectiveness of policies based on reduced marginal costs and internal deflations is weakened and delayed, because of the impossibility of accommodating the deflation with a relevant nominal interest rate drop. Such a result holds both for the labor market targeted fiscal policies (hiring and wage subsidies for new hires of labor) and for fiscal expansions based on tax cuts. On the contrary, and in line with the results of a recent literature (Christiano et al. 2011; Eggertsson, 2011; Eggertsson and Krugman, 2012), the efficacy of standard inflationary fiscal measures, as are the policies based on increased government expenditure, is increased by the reduced counteracting response of the monetary policy.

The paper is organized as follows: Section one describes the model, focusing in particular on the theoretical extensions implemented in the design of the labor market. Section two provides the details of the Bayesian estimation of the country-specific models. Here we describe the data and their transformations, we address issues of empirical identification, the calibration and the elicidation of priors for the structural model and the Bayesian SVAR parameters, and discuss the posterior estimates. Section three provides a discussion of simulation results, explaining the propagation mechanics in the standard time and binding ZLB environments. Section four concludes.

1 The model

We jointly consider a number of extensions to the now standard set-up of the new-Keynesian monetary model, characterized by the presence of nominal and real frictions in both goods and labor markets (Christiano *et al.*, 2005; Smets and Wouters, 2007). *First*, we introduce a monopolistically competitive financial sector (Curdia and Woodford, 2010) which is subject to non zero default probabilities on the side of both public and private borrowers, such that a sovereign default risk channel emerges (Corsetti *et al.*„ 2013). *Second*, we consider a small open economy framework, developed along the lines of Adolfson *et al.* (2007; 2008) and

Christiano *et al.* (2011), in which the foreign sector is exogenous with respect the domestic economy and its evolution is described by a structural vector auto-regressive system (SVAR). *Third*, we adopt a reasonably detailed specification of the fiscal sector, whose relevance for macroeconomic dynamics is recuperated by considering that a fraction of households are liquidity constrained. The design of the fiscal sector marginally resembles that proposed in Drautzburg and Uhlig (2011), and considers unemployment benefits in addition to the standard fiscal instruments characterizing the expenditure and revenues sides of fiscal models, and an optimal definition of the public investment and capital decisions. *Fourth*, we develop a detailed representation of the non Walrasian labor market, basically following Diamond (1982), Mortensen and Pissarides (1994), and Pissarides (2000) for the introduction of hiring costs and matching frictions, and Gertler *et al.* (2008) and Gertler and Trigari (2009) for the representation of the staggered Nash-wage bargaining between unions and firms.

The major novelty in the design of the monopolistically competitive financial sector is the consideration of a non zero default probability for both private sector and public sector borrowers, obtained by formalizing a cumulative distribution function relating the sovereign default probability to the debt and the NFA position to GDP ratios, and the private sector default probability to the sovereign default probability. Default risks are traduced in bond and lending rate spreads through the consideration of a no arbitrage condition between deposits and domestic bond holdings, and an optimality condition for credit institutions including the Loss Given Default of the bank in the case of counterparty default, respectively.

1.1 Households

1.1.1 Optimizers

A continuum of liquidity unconstrained households indexed by $j \in [0,1]$ have access to a complete set of contingent claims[3]. The representative household is assumed to maximize the following lifetime utility function:

$$\max_{C_t^r, B_t^r, B_t^{*r}, K_t^{p,r}, I_t^r, u_t^k} E_0 \sum_{t=0}^{\infty} \beta^t \left[\xi_t \frac{(C_t^r - h\widetilde{C}_{t-1})^{1-\sigma_c}}{1-\sigma_c} - \chi_t n_t \right] \tag{1}$$

where C_t^r is a composite consumption index, $h\widetilde{C}_{t-1}$ denotes external habits σ_c is the consumption curvature parameter and $0 \leq n_t \leq 1$ denotes the fraction of household members who are employed. ξ_t and χ_t are two preference shocks which are assumed to follow the i.i.d. processes $\xi_t = e^{\varepsilon_{\xi,t}}$ and $\chi_t = \chi\mu^{(1-\sigma_c)t}\xi_t^n$, respectively, where $\xi_t^n = e^{\varepsilon_{\xi^n,t}}$[4].

Each household purchases consumption and investment goods by means of after tax labor and capital incomes, after tax unemployment benefits, dividends and government transfers. The budget constraint is thus given by:

[3] This standard hypothesis ensures that households are homogeneous with respect to consumption and asset holdings choices, such that the notation can be simplified by dropping the j-index.

[4] The peculiar specification of the stochastic scaling factor of labor disutility χ_t is chosen to ensure balanced growth.

$$
\begin{aligned}
& (1+\tau_t^c)C_t^r + I_t^r + \frac{B_t^r}{P_t R_t^g} + \frac{e_t B_t^{*r}}{P_t R_t^{g*} \Phi(\frac{A_t}{Y_t}, \frac{e_t}{e_{t-1}}, R_t^{e*} - R_t^e, \widetilde{\phi}_t)} + \frac{D_t^r}{P_t} \\
= \quad & Tr_t^r + \frac{R_{t-1} D_{t-1}^r}{P_t} + \left[\left(1 - p_t^{d,g}\right) + z p_t^{d,g}\right] \frac{B_{t-1}^r}{P_t} + \frac{e_t B_{t-1}^{*r}}{P_t} + \\
& (1-\tau_t^n)\left[\frac{W_t}{P_t} n_t + b_t^u (1-n_t)\right] + \left[(1-\tau_t^k)\left[\frac{R_t^k}{P_t} u_t - a(u_t)\right] + \delta \tau_t^k\right] k_{t-1}^{p,r} + \frac{\Pi_t^p \mu^t}{P_t}
\end{aligned} \tag{2}
$$

where I_t^r is private investment, $A_t = \frac{e_t B_{t+1}^*}{P_t}$ is the aggregate net foreign asset position of the domestic economy, e_t is the nominal effective exchange rate and $\frac{D_t^r}{P_t}$ denotes household's deposits to financial intermediaries in real terms. B_t^r and B_t^* are domestic and foreign bond holdings, respectively, P_t is the consumption price index and $R_t^g = R_t q_{b,t}$, $R_t^{g*} = R_t^* q_{b,t}^*$ are the domestic and foreign interest rates on government bonds, where R_t, R_t^* denote the respective policy rates and $q_{b,t}$, $q_{b,t}^*$ are the home and foreign spreads on government bonds, respectively, the latter defined within the SVAR system for the foreign variables. The variable $p_t^{d,g}$ and the parameter z denote the sovereign debt default probability and the recovery rate on defaulted bonds. $\frac{R_t^k}{P_t}$ is the real return on capital $K_t^{p,r}$, u_t^k and $a\left(u_t^k\right)$ denote the utilization rate and its adjustment cost[5], respectively, and δ is the private capital depreciation rate. $\frac{W_t}{P_t}$ is the real wage and $\frac{\Pi_t^p \mu^t}{P_t}$ define real dividends, where μ denotes the long-run trend growth of labor-augmenting productivity. Government transfers TR_t^r, unemployment benefits $b_t^u = b\mu^t$[6] and the tax rates on consumption τ_t^c, on labor income τ_t^n and on capital τ_t^k complete the budget constraint of the Ricardian household. The term $\Phi_t = \Phi(\frac{A_t}{Y_t}, \frac{e_t}{e_{t-1}}, R_t^{e*} - R_t^e, \widetilde{\phi}_t)$ in (2) denotes the risk premium on foreign bond holdings in the modified uncovered interest parity (UIP) equation $E_t\left(\frac{e_{t+1}}{e_t}\right) = \frac{R_t^e}{\Phi_t R_t^{e*}}$, i.e.:

$$
\Phi_t = \exp[-\widetilde{\phi}_a \left(\frac{A_t}{Y_t} - \frac{A}{Y}\right) - \widetilde{\phi}_r \left(R_t^{e*} - R_t^e\right) + \widetilde{\phi}_s \left(1 - \frac{e_t}{e_{t-1}}\right) + \widetilde{\phi}_t] \tag{3}
$$

where $\widetilde{\phi}_t$ is a time varying shock to the risk premium, which is assumed to follow the AR(1) stochastic process $\widetilde{\phi}_t = \widetilde{\phi}_{t-1}^{\rho_{\widetilde{\phi}}} e^{\varepsilon_{\widetilde{\phi},t}}$ and $\widetilde{\phi}_a$, $\widetilde{\phi}_s$ and $\widetilde{\phi}_r$ are positive elasticities. Our specification ensures the satisfaction of the usual equilibrium requirements (Lundvik, 1992; Schmitt-Grohé and Uribe, 2001) and adds some flexibility to alternative modified UIP equations adopted in the literature (e.g. Adolfson *et al.* 2008 and Christiano *et al.* 2011). The log-linear representation of the modified UIP is the following:

$$
E_t\left(\Delta e_{t+1}\right) = \widetilde{\phi}_s \Delta e_t + \left(1 - \widetilde{\phi}_r\right)\left(R_t^e - R_t^{e*}\right) + \widetilde{\phi}_a \left(A_t - Y_t\right) - \widetilde{\phi}_t
$$

were the parameter $\widetilde{\phi}_s$ defines the autoregressive behavior of the expected change in the nominal exchange rate and $\widetilde{\phi}_r \geq 0$ denotes the elasticity to the interest rate differential on bond holdings, allowing for the emergence of the "forward premium puzzle" (for $\widetilde{\phi}_r > 1$), i.e. the negative correlation between interest rate differentials

[5]The function $a\left(u_t^k\right)$ is assumed to be strictly increasing and convex, with curvature parameter ψ^k. The utilization rate relates effective to physical capital in a standard fashion, i.e. $K_t^r(i) = K_{t-1}^{p,r}(i) u_t(i)$.

[6]In order to ensure long-run balanced growth, b_t^u is assumed to grow at the labor augmenting productivity growth rate μ.

and expected exchange rate variations often observed in empirical trials[7].

The law of motion of physical capital is described by the following equation:

$$K_t^{p,r} = (1-\delta)K_{t-1}^{p,r} + q_{i,t}\left[1 - S(\frac{I_t^r}{I_{t-1}^r})\right] I_t^r \tag{4}$$

where $S(\frac{I_t^r}{I_{t-1}^r})$ defines the private investment adjustment cost function, with curvature parameter ψ^i, and $q_{i,t}$ is an investment-specific shock, which is assumed to follow the i.i.d. stochastic process $q_{i,t} = e^{\varepsilon_{q_{i,t}}}$.

Aggregate demand for type X_t goods, $X_t = (C_t, I_t)$, is obtained as a CES index of domestically produced and imported goods, such that:

$$X_t = \left[(1-\nu)^{\frac{1}{\eta}} \left(X_t^d\right)^{\frac{\eta-1}{\eta}} + \nu^{\frac{1}{\eta}} \left(X_t^m\right)^{\frac{\eta-1}{\eta}}\right]^{\frac{\eta}{\eta-1}} \tag{5}$$

where, from households' cost minimization, $X_t^d\,(1-\nu)\left(\frac{P_t^d}{P_t}\right)^{-\eta} X_t$ and $X_t^m = \nu\left(\frac{P_t^m}{P_t}\right)^{-\eta} X_t$ are, respectively, the aggregate available domestic and foreign produced goods, ν denotes the import share parameter and η is the elasticity of substitution between domestic and imported goods. P_t^d and P_t^m denote the price indexes of domestic and imported goods, respectively, such that:

$$P_t = \left[(1-\nu)\left(P_t^d\right)^{1-\eta} + \nu\left(P_t^m\right)^{1-\eta}\right]^{\frac{1}{1-\eta}} \tag{6}$$

From the first order condition (F.O.C.) for consumption, the following consumption Euler equation is obtained:

$$C_t^r - hC_{t-1}^r = \left[\beta R_t^e \frac{P_t}{P_{t+1}} \frac{(1+\tau_t^c)}{(1+\tau_{t+1}^c)} \frac{\xi_{t+1}}{\xi_t}\right]^{-\frac{1}{\sigma_c}} \left(C_{t+1}^r - hC_t^r\right) \tag{7}$$

1.1.2 The rule-of-thumb household

Liquidity constrained and unconstrained households have the same number of workers:

$$n_t = n_t^r = n_t^{nr} \tag{8}$$

From the budget constraint of the liquidity constrained household the following consumption equation is obtained:

$$C_t^{nr} = \frac{1}{(1+\tau_t^c)}\left[Tr_t^{nr} + (1-\tau_t^n)\frac{W_t}{P_t}n_t + (1-\tau_t^n)b^u(1-n_t)\right] \tag{9}$$

where it is evident that rule-of-thumbers spend all their net income (from labor, government transfers and unemployment benefits) in consumption goods.

[7] In the modified UIP adopted in Adolfson et al. (2008) the autoregressive component is not independent on the elasticity to the interest rate differential, and the chosen prior does not allow for a direct emergence of the forward premium puzzle. Compared to the specification adopted in Christiano et al. (2011), our modified UIP adds the autoregressive component.

1.2 Firms

1.2.1 Intermediate sector

Each intermediate firm (i) operates in a perfectly competitive environment combining private capital public infrastructures and labor. The production technology is as follows:

$$Y_t^i(i) = \xi_t^a \left[\frac{K_{t-1}^g}{\int_0^1 Y_t^i(j)dj} \right]^{\frac{\xi}{1-\xi}} [K_t(i)]^{\alpha} \left[\mu^t n_t(i)\right]^{(1-\alpha)} \tag{10}$$

where K_t^g is public capital, α and ξ are the private and public capital shares in production, respectively, and $\xi_t^a = \xi_{t-1}^{a^{\rho_{\xi^a}}} e^{\varepsilon_{\xi^a,t}}$ is an AR(1) process defining the evolution of total factor productivity.

The optimizing firm chooses the optimal quantity of capital by solving the following maximization problem:

$$\max_{K_t(i)} P_t^i Y_t^i(i) - R_t^k K_t(i) \quad \text{s.t. } (10)$$

whose re-arranged F.O.C. yields:

$$R_t^k(i) = \alpha P_t^i(i) \frac{Y_t^i(i)}{K_t(i)} \tag{11}$$

where $P_t^i(i)$ is the intermediate sector price index.

Since a fraction ϑ^b of the wage bill $W_t n_t$ is anticipated by borrowing from financial intermediaries, the cost of one unit of labor is $R_t^t W_t$, where:

$$R_t^t(i) = \vartheta^b \left[1 - p_t^{dp}(i)\right] R_t^l(i) + \left(1 - \vartheta^b\right) + d_t^{cp}(i) \tag{12}$$

is the effective interest rate. $p_t^{dp}(i)$ denotes the firm's default probability and $d_t^{cp}(i) = \vartheta^b p_t^{dp}(i) R_t^l(i)$ is the cost of default per unit of borrowed cost of labor.

1.2.2 Final sector: wholesalers and retailers in the domestic, import and export sectors

For expositional convenience, a joint description of the structure of the final good sector, composed of domestic, import and export wholesalers and retailers, is provided.

Domestic wholesale firms buy the homogenous good $Y_t^i(i)$ from domestic intermediate good producers at the price $P_t^i(i)$, and differentiate the homogeneous product into $Y_t^d(i)$ using a linear technology. Wholesalers sell their goods under monopolistic competition to domestic retailers, who use the differentiated goods $Y_t^d(i)$ to produce the composite final good Y_t^d.

Wholesale firms in the import sector buy the homogenous good Y_t^* from foreign retailers at the foreign price P_t^*, and obtain a differentiated good $Y_t^m(i)$. Wholesale importing firms sell their goods under monopolistic competition to import retailers who use the differentiated goods $Y_t^m(i)$ to produce the composite final good Y_t^m.

Finally, wholesale export firms buy the homogenous good Y_t^d from domestic retailers at the price P_t^d and produce a differentiated good $Y_t^x(i)$ using a linear technology. Wholesalers in the export sector sell their goods under monopolistic competition to export retailers, who use the differentiated goods $Y_t^x(i)$ to produce the composite final good Y_t^x.

We consider a variable demand elasticity in the three sectors, indexed by $k = (d, m, x)$, by assuming a flexible variety aggregator à la Kimball (1995):

$$\left[\int_0^1 G\left(\frac{Y_t^k(i)}{Y_t^k};\lambda_{p,t}^k\right)di\right] = 1$$

such that the domestic retailers demand function for differentiated goods is:

$$Y_t^k(i) = Y_t^k G'^{-1}\left[\frac{P_t^k(i)}{P_t^k}\varkappa_{p,t}^k\right] \tag{13}$$

where:

$$\varkappa_{p,t}^k \equiv \int_0^1 G'\left(\frac{Y_t^k(i)}{Y_t^k};\lambda_{p,t}^k\right)\frac{Y_t^k(i)}{Y_t^k}di$$

The optimization problem of wholesalers firms that are allowed to re-optimize their prices reads:

$$\max_{\widetilde{P}_t^k(i)} E_t \sum_{j=0}^{\infty}\left(\beta\xi_p^k\right)^j \vartheta_{t+j}\left[\widetilde{P}_t^k(i)X_{t,t+j}^k - MC_{t+j}^k\right]Y_{t+j}^k(i)$$

$$\text{s.t. (13) and } X_{t,t+j}^k = \left\{\begin{array}{cc} 1 & for\ j=0 \\ \Pi_{l=0}^j\left(\pi_{t+l-1}^k\right)^{\iota_p^k}\pi_*^{1-\iota_p^k} & for\ s=1,...,\infty \end{array}\right\}$$

where $MC_t^d = P_t^i$, $MC_t^m = e_t P_t^*$ and $MC_t^x = P_t^d/e_t$ are the nominal marginal costs of the domestic, import sector and export sector wholesalers, respectively. The term $\left(\beta\xi_p^k\right)^j\vartheta_{t+j}$ denotes the stochastic discount factor of the firm, where ξ_p^k is the Calvo probability of price adjustment. $\lambda_{p,t}^k = e^{\varepsilon_{p,t}^k}$ are *i.i.d.* stochastic processes defining the time-varying markups[8] and $X_{t,t+j}^k$ denote price indexation functions.

The first order condition for the optimality problem above is given by:

$$E_t\sum_{j=0}^{\infty}\left(\xi_p^k\beta\right)^j\vartheta_{tt+j}Y_{t+j}^k(i)\left[\widetilde{P}_t^k(i)X_{t,t+j}^k + \left(\widetilde{P}_t^k(i)X_{t,t+j}^k - MC_{t+s}^k(i)\right)\frac{1}{G'^{-1}\left(\nu_t^k\right)}\frac{G'\left(\theta_{t+j}^k\right)}{G''\left(\theta_{t+j}^k\right)}\right] = 0 \tag{14}$$

where $\theta_t^k = G'^{-1}\left(\nu_t^k\right)$, $\nu_t^k = \frac{P_t^k(i)}{P_t^k}\varkappa_{p,t}^k$, and the aggregate domestic price indexes read:

$$P_t^k = \left(1-\xi_p^k\right)P_t^k(i)G'^{-1}\left[\frac{P_t^k(i)}{P_t^k}\varkappa_{p,t}^k\right] + \xi_p^k P_{t-1}^k\left(\pi_{t-1}^k\right)^{\iota_p^k}\pi_*^{1-\iota_p^k}G'^{-1}\left[\frac{P_{t-1}^k\left(\pi_{t-1}^k\right)^{\iota_p^k}\pi_*^{1-\iota_p^k}}{P_t^k}\varkappa_{p,t}^k\right] \tag{15}$$

[8] We assume i.i.d. mark-up shocks in order to enhance the identifiability of the price equations. For a more in dept explanation of this point, see the estimation section below and Giuli and Tancioni (2012).

1.3 Financial sector and default risks

1.3.1 Financial intermediaries and private default risk

In each period t a continuum of monopolistically competitive banks receives deposits $D_t(i)$ from the households and supplies loans $L_t(i)$ to banks in the retail sector at the nominal interest rate $R_t^l(i)$. Retail banks purchase differentiated loans from the monopolistically competitive banks and aggregate them in the single composite loan $L_t = \left[\int_0^1 L_t(i)^{(\Lambda_t^l-1)/\Lambda_t^l}\right]^{\Lambda_t^l/(\Lambda_t^l-1)}$, purchased by the intermediate good producer firms at the interest rate R_t^l for anticipated wage payments $W_t n_t$.The term $\Lambda_{p,t+j}^l$ represents the stochastic loan demand elasticity in the credit sector, which is assumed to follow the AR(1) stochastic process $\Lambda_t^l = \Lambda^{l(1-\rho_\Lambda)}\Lambda_{t-1}^{l(\rho_\Lambda)}e^{\varepsilon_{\Lambda,t}}$.

Intertemporal cost minimization implies that the optimal loan demand is given by $L_t(i) = \left(R_t^l(i)/R_t^l\right)^{-\lambda_t^l}L_t$. At the end of each period, the monopolistically competitive bank pays back the interest-augmented initial deposits $R_t D_t(i)$ and ownership profits to households. The representative monopolistically competitive bank maximizes its profit function facing Rotemberg-type costs for adjusting the interest rate on loans:

$$\max_{D_t(i),IB_t,R_t^l} E_t\sum_{s=0}^{\infty}\beta^s\frac{\Lambda_{t+s}P_t}{\Lambda_t P_{t+s}}\left[\left(1-p_t^{d,p}\right)R_{t+s}^l(i)\,L_{t+s}(i) - R_{t+s}D_{t+s}(i) - R_{t+s}IB_{t+s}(i) - \frac{\kappa_b}{2}\left(\frac{R_{t+s}^l(i)}{R_{t+s-1}^l(i)}-1\right)^2 L_{t+s}(i)\right] \quad (16)$$

subject to the credit balance sheet constraint:

$$D_{t+s}(i) + IB_{t+s}(i) = L_{t+s}(i) + Q_{t+s}(i)$$

where $IB_t(i)$, $Q_t(i) = \phi^q D_t(i)$ and ϕ^q denote interbank borrowing, the bank amount and the bank ratio of reserves respectively, and κ_b in (16) denotes the Rotemberg adjustment cost parameter.

The observed strong co-movement between government bond and lending rates indicates that the market valuation of sovereign debt assets affects the private sector credit conditions[9]. In order to capture this relation, we assume a non zero default probability in the private sector, described by the following cumulative density function:

$$p_t^{d,p} = \frac{1-\exp\left[-\varphi^{s,p}\left(p_t^{d,g}\right)^{\phi^{s,p}}\right]}{1-\exp\left[-\left\{\varphi^{s,p}+\left(1-p_t^{d,g}\right)^{\phi^{s,p}}\right\}\right]} \quad (17)$$

where $\varphi^{s,p}$ and $\phi^{s,p}$ are the scale and the shape parameters of the private sector default c.d.f., respectively, such that:

$$p_t^{d,p} = \left\{\begin{array}{lll} 1 & if & p_t^{d,g}=1 \\ 0 & if & p_t^{d,g}=0 \end{array}\right\}$$

Equation (17) expresses to which degree the probability of default of sovereign debt $p_t^{d,g}$ spills-over the private sector. Given values for the scale and the shape parameters in (17), our preferred formulation ensures

[9] Harjes (2011) provides evidence about these spill-over effects.

a flexible and accurate representation of the actual relations between private sector credit and government bond spreads emerging in country-specific time series data.

Note that, compared to the formulation adopted in Corsetti *et al.* (2013), who assume a direct log-linear relation between government and credit rate spreads, we model the underlying relation between the sovereign debt and private sector default probabilities.

From the optimality condition of the monopolistically competitive bank, the following lending rate equation is obtained:

$$\left[R^l_t\right] : R^l_t(i) = \frac{1}{\left(1 - p^{d,p}_t \left(1 - z^p \digamma\nu_t\right)\right)} \left[\frac{\lambda^l_t}{\lambda^l_t - 1} R_t - \frac{\kappa_b}{\lambda^l_t-1}\left[\left(\frac{R^l_t(i)}{R^l_{t-1}(i)} - 1\right)\frac{R^l_t(i)}{R^l_{t-1}(i)} - \beta \frac{P_t\Lambda_{t+1}}{P_{t+1}\Lambda_t}\left(\frac{R^l_{t+1}(i)}{R^l_t(i)} - 1\right)\frac{R^l_{t+1}(i)}{R^l_t(i)}\frac{L_{t+1}}{L_t(}\right.\right. \tag{18}$$

where z^p is the share of the Gordon's firm value $\digamma\nu_t = \left[p^d_t(i)\, y_t(i) - r^k_t k_t(i) - w_t n_t(i)\right] / \left[r^k_t - (\mu - 1)\right]$, determining the Loss Given Default (LGD) $1 - z^p \digamma\nu_t$ of the bank in the case of counterparty default[10]. The above expression highlights that, in our setting, the lending rate is determined by the risk free rate, the mark-up and the cost of adjusting the interest rate as in the standard literature considering imperfect credit markets (Curdia and Woodford, 2010; Gerali *et al.,* 2010), as well as by the survival rate of the private sector firms and the LGD.

1.3.2 The sovereign default risk

Along the lines of the analysis in Corsetti *et al.* (2013), we do not model the event of default as the result of a strategic decision (Eaton and Gersovitz, 1981; Yue, 2010; Arellano, 2008; Mendoza and Yue, 2012), but relate the sovereign default probability to two fundamental triggers addressed in the literature (Edwards, 1986; Manasse and Roubini, 2009; De Grauwe and Ji 2013): *i)* the government debt to GDP ratio B_t/Y_t and *ii)* the NFA position to GDP ratio A_t/Y_t. Our preferred specification for the sovereign default probability is defined by the cumulative distribution function:

$$p^{d,g}_t = \left\{1 - \exp\left[-\varphi^{s,g}\left(\lambda_b \frac{B_t}{Y_t} + \lambda_a \frac{A^-_t}{Y_t}\right)\right]\right\}^{\phi^{s,g}} \tag{19}$$

such that, $\frac{\partial p^{d,g}_t}{\partial B_t} > 0$, $\frac{\partial p^{d,g}_t}{\partial A_t} < 0$ and:

$$p^{d,g}_t = \left\{\begin{array}{lll} 1 & if & \frac{B_t}{Y_t} = +\infty \cap \frac{A^-_t}{Y_t} = +\infty \\ 0 & if & \frac{B_t}{Y_t} = \frac{A_t}{Y_t} = 0 \end{array}\right\}$$

where A^-_t is the net foreign indebtedness.

From the optimality condition for deposits and domestic bond holdings, and since $R^g_t = R_t q^b_t$, the following

[10] Instead of considering the standard Gordon's firm value model, we consider the value of the entire production and supply chain, that is, the value of the intermediate and final sector firm.

no arbitrage condition must hold:

$$R_t = R_t q_t^b \left[\left(1 - p_{t+1}^{d,g} \right) + z^g p_{t+1}^{d,g} \right] \tag{20}$$

where $z^g = \sigma_z \frac{\phi^i}{\phi^{i*}} \frac{Y}{Y^*}$ is the recovery rate on government bond in the case of sovereign debt default. The parameters ϕ^i and ϕ^{i*} denote the domestic and foreign contribution to a hypothetical international insurance institution (e.g. the IMF) and σ_z is the efficiency parameter defining the relation between contribution and insurance coverage (e.g. the quota of SDRs to the IMF).

Given the positions above and considering the no arbitrage condition (20), the interest rate spread on government bonds reads:

$$q_t^b = \frac{1}{\left[1 - (1 - z^g) \, p_t^{d,g} \right]} \tag{21}$$

where the government bond premium q_t^b emerges as a result of a non zero probability $p_t^{d,g}$ of sovereign debt default.

Note that, aside from the consideration of the net foreign assets position, our preferred specification of the sovereign default risk depart from the one adopted in Corsetti *et al.* (2013) in two main respects: first, we do not consider a fiscal limit, i.e. an upper bound for the debt to GDP ratio, on the grounds that such a limit is neither theoretically nor empirically identifiable. Second, in line with the empirical literature, we consider the debt to GDP *ratio* in the place of the debt *level*, in order to take into account the crucial role of the GDP dynamics in the definition of the sovereign default risk addressed in the literature (Levy-Yeyati and Panizza 2011; Mendoza and Yue, 2012), relate the analysis more closely to the available empirical literature, addressing the debt to GDP ratio as a fundamental measure of the capacity of the government to service its debt, and consider the evolution of the NFA position to GDP ratio as an additional trigger of sovereign the default probability (Edwards, 1986; De Grauwe and Ji, 2013). Note also that the consideration of the debt to GDP ratio implies that the size and the sign of the default risk channel crucially depends on the size of the fiscal multipliers. When fiscal multipliers are large, fiscal contractions can lead to transitory but persistent increases in the debt to GDP ratio, activating a default risk channel operating in an opposite - pro-cyclical - direction than predicted.

Figure 1 depicts, for different levels of the debt to GDP ratio and of the sensitivity parameter λ_b, the behavior of the default probability function (panel a) and of the government bond spread (panel b), considering a parameterization which is consistent with the data of the five economies in the analysis. The shape parameter $\phi^{s,g}$ is fixed to a value of 20, whilst the scale parameter $\varphi^{s,g}$ is fixed such that, given an elasticity coefficient $\lambda_b = 0.5$, the observed intersections between the debt to GDP ratio and the government bond spread for each country belong to the default probability surface.

FIGURE 1 about here

1.4 The labor market

The matching process is described by a standard Cobb-Douglas matching technology:

$$m_t = \sigma_m \upsilon_t^{\sigma_n} u_t^{1-\sigma_n} \tag{22}$$

where σ_m is the matching efficiency parameter, υ_t is the number of vacancies and $u_t = 1 - n_{t-1}$ denotes the unemployment rate once the labor force stock has been normalized to one. The chosen timing in the unemployment relation shows that individuals entering the labor force stock activate their job search immediately, whilst workers that loss their job in t are not able to search for a new one in the same period of the separation event. Given the job filling rate $q_t = m_t/\upsilon_t$ and the job finding rate $s_t = m_t/u_t$, the labor market tightness can equivalently be defined as $\theta_t = \upsilon_t/u_t$ or $\theta_t = s_t/q_t$.

Under the assumption of exogenous separation, the employment law of motion is described by the following dynamic equation

$$n_t = (1-\rho)\, n_{t-1} + m_t \tag{23}$$

where ρ is the separation rate.

1.4.1 Workers value functions

Let $W_t(w_t)$ be the worker value of being matched to a job evaluated at the wage w_t and U_t be the value of being unemployed at time t. The value of the employment/unemployment states are the following:

$$\begin{aligned} W_t(w_t) &= (1-\tau_t^n)\frac{w_t}{P_t} - \frac{\chi_t}{\Lambda_t} + \beta E_t\left[\frac{\Lambda_{t+1}}{\Lambda_t}\left[(1-\rho)\left[\theta_w W_{t+1}(w_t) + (1-\theta_w)W_{t+1}(w_{t+1}^*)\right] + \rho U_{t+1}\right]\right] && (24) \\ U_t &= (1-\tau_t^n)b_t^u + \beta E_t\left[\frac{\Lambda_{t+1}}{\Lambda_t}\left[s_{t+1}\left(\theta_w W_{t+1}(w_t) + (1-\theta_w)W_{t+1}(w_{t+1}^*)\right) + (1-s_{t+1})U_{t+1}\right]\right] && (25) \end{aligned}$$

where θ_w is the Calvo parameter defining the probability of being unable to re-optimize the wage in $t+1$, Λ_t is the Lagrange multiplier and w_t^* is the re-optimized wage. From equations (24) and (25) the net value of being employed, i.e. the worker' surplus $W_t(w_t) - U_t$, is obtained.

1.4.2 Firms value functions

Let $J_t(w_t)$ be the asset value of a job evaluated at the wage w_t:

$$J_t(w_t) = (1-\tau_t^p)(\zeta_t - R_t^t \frac{w_t}{P_t^d}) + (1-\rho)\widetilde{\beta}\mu E_t\left[\frac{\Lambda_{t+1}}{\Lambda_t}\left(\theta_w J_{t+1}(w_t) + (1-\theta_w)J_{t+1}(w_{t+1})\right)\right] \tag{26}$$

where P_t^d is the domestic price index, τ_t^p denotes the business profits tax rate and $\zeta_t = (1-\alpha)P_t^i Y_t/n_t$ the marginal productivity of labor.

Given the value of a vacancy:

$$J_t^v = -\kappa + q_t \left[\theta_w J_t(w_{t-1}) + (1-\theta_w) J_t(w_t^*)\right] \tag{27}$$

and imposing the free entry condition, $J_t^v = 0$, the vacancy posting condition is obtained

$$\frac{\kappa}{q_t} = \theta_w J_t(w_{t-1}) + (1-\theta_w) J_t(w_t^*) \tag{28}$$

1.4.3 Nash wage bargaining

Given the the worker' surplus $W_t(w_t) - U_t$, the firm's asset value of a job $J_t(w_t^*)$ and the union's bargaining power ς, the Nash-bargaining solution is given by $\varsigma(1-\tau_t^n) J_t(w_t^*) = (1-\varsigma)\left(1-\tau_t^p\right)\left[W_t(w_t^*) - U_t\right]$. Plugging the value functions in the latter equation, the optimal real wage reads:

$$\begin{aligned}
w_t^* \quad = \quad & \Theta_t \left[\varsigma\zeta_t + (1-\varsigma)\left(b^u + \frac{\chi_t}{\Lambda_t}\right)\right] + \frac{1}{(1-\tau_t^p)}\Theta_t\varsigma\,(1-\rho)\,\widetilde{\beta}\mu E_t\left[\frac{\Lambda_{t+1}}{\Lambda_t}\frac{\kappa}{q_{t+1}}\left(1 - \Xi_t\frac{\Upsilon_{t+1}^n}{\Upsilon_{t+1}^p}\right)\right] \\
& + \Theta_t \sum_{j=1}^{\infty}\frac{\Lambda_{t+j}}{\Lambda_t}\left((1-\rho)\widetilde{\beta}\mu\gamma_w\right)^j \left\{(1-\varsigma)\,E_t\left[\Upsilon_{t+1}^n\left((w_{t+1}^* - w_t^*) - \frac{s_{t+1}}{1-\rho}(w_{t+1}^* - w_t)\right)\right]\right. \\
& \left. + \,\varsigma E_t\left[\Upsilon_{t+1}^p\left(R_{t+1}^t\frac{w_{t+1}^*}{p_{t+1}^d} - R_t^t\frac{w_t^*}{p_t^d}\right) - \left(\Upsilon_{t+1}^p - \Xi_t\Upsilon_{t+1}^n\right)\left(R_{t+1}^t\frac{w_{t+1}^*}{p_{t+1}^d} - R_t^t\frac{w_t}{p_t^d}\right)\right]\right\}
\end{aligned} \tag{29}$$

where we have used the transformations $\Upsilon_t^i = (1-\tau_t^i)/(1-\tau_{t-1}^i)$, for $i = (n,p)$, $\Xi_t = (1-\rho-s_t)/(1-\rho)$, $\Theta_t \equiv 1/\left[1-\varsigma\left(1-1/p_t^d\right)\right]$, $p_t^d = P_t^d/P_t$, and w_t is the average real wage $w_t = [\theta_w w_{t-1} + (1-\theta_w)w_t^*]$. Note that, for $\tau_t^i = 0$ the real wage equation (29) resolves in a standard Nash wage equation (Gertler and Trigari, 2009).

1.5 Government policies

1.5.1 The monetary authority

The Central Bank sets the nominal interest rate $R_t \equiv 1 + r_t$ according to a contemporaneous rule considering inflation, output and output growth deviations from the respective steady state values. The policy instrument is adjusted gradually, giving rise to interest rate smoothing:

$$\frac{R_t}{\overline{R}} = \left(\frac{R_{t-1}}{\overline{R}}\right)^{\rho^R}\left[\left(\frac{\pi_t}{\overline{\pi}}\right)^{\psi_1}\right]^{1-\rho^R}\left(\frac{Y_t}{Y_{t-1}}\right)^{\psi_2} + \epsilon_t^r \tag{30}$$

where ρ^R defines the degree of interest rate smoothing, ψ_1 and ψ_2 are the feedback coefficients to CPI inflation π_t[11], and output growth, respectively. The stochastic term ϵ_t^r denotes the monetary policy shock, which is assumed to be white noise $\epsilon_t^r = e^{\varepsilon_t^r}$. Similar to money-growth rules, implementation of this policy rule does

[11] CPI inflation is obtained as a weighted average considering domestic and imported price variations, i.e.: $\pi_t = \left[(1-\nu)\left(p_t^d\pi_t^d\right)^{1-\eta} + \nu\left(p_t^m\pi_t^m\right)^{1-\eta}\right]^{\frac{1}{1-\eta}}$.

not require knowledge about the natural rate of interest or of the level of potential output, both of which are unobserved[12].

The fact that the countries being considered in this study all joined a common currency and a centralized monetary policy since 1999 (2001 for Greece) implies that, at the estimation stage, a regime break has to be taken into account. To implement such a structural break, we will consider a permanent observed exogenous shock acting as a multiplicative regime-shift dummy variable on all the three monetary policy coefficients.

1.5.2 The fiscal authority

By expressing government consumption, government transfers, hiring subsidies and unemployment benefits in terms of domestic goods, the government budget constraint in real terms reads:

$$
\begin{aligned}
& \frac{P_t^d}{P_t}\left[G_t + I_t^g + (1-\tau_t^n)\, b_t^u\,(1-n_t)\right] + \phi^i Y_t + TR_t + \left[\left(1-p_t^{d,g}\right) + z^g p_t^{d,g}\right]\frac{B_{t-1}}{\pi_t} + d_t^{c,g}\frac{B_{t-1}}{\pi_t} \\
= \; & \frac{B_t}{R_t^g} + \tau_t^c C_t + \tau_t^n w_t n_t + \tau_t^k\left[r_t^k u_t^k - a(u_t^k) - \delta\right] K_{t-1}^{p,r} + \tau_t^p\left(\zeta_t - w_t\right)
\end{aligned}
$$

where $d_t^{c,g} = p_t^{d,g}(1-z_t)$ is the unit cost of sovereign default, $G_t = G_{t-1}^{\rho_g} Y_t^{(1-\rho_g)\eta_{gy}} D_t^{\eta_{gd}} e^{\varepsilon_{g,t}}$ and $TR_t = TR_{t-1}^{\rho_{tr}} Y_t^{(1-\rho_{tr})\eta_{try}} D_t^{\eta_{trd}} e^{\varepsilon_{tr,t}}$ are the partial adjustment stochastic processes for government expenditures for consumption and transfers, respectively, with D_t denoting the government financial need, and $\varepsilon_{g,t}$, $\varepsilon_{tr,t}$ *i.i.d.* shocks.

The government financial need D_t is the following:

$$
\begin{aligned}
D_t \;\equiv\; & \frac{P_t^d}{P_t}\left[G_t + I_t^g + (1-\tau_t^n)\, b_t^u\,(1-n_t)\right] + \phi^i Y_t + TR_t + \frac{B_{t-1}}{\pi_t} + \left[\left(1-p_t^{d,g}\right) + z^g p_t^{d,g}\right]\frac{B_{t-1}}{\pi_t} \\
& + d_t^{c,g}\frac{B_{t-1}}{\pi_t} - \tau_t^c C_t - \tau_t^n w_t n_t - \tau_t^k\left[r_t^k u_t^k - a(u_t^k) - \delta\right] K_{t-1}^{p} - \tau_t^p\left(\zeta_t - w_t\right) \qquad (31)
\end{aligned}
$$

A fraction ψ_τ of D_t is financed with distortionary taxation on consumption, labor income, capital and on business profits, such that:

$$
\psi_\tau\left(D_t - D\right) = \left(\tau_t^c - \tau^c\right) C_t + \left(\tau_t^n - \tau^n\right) w_t n_t + \left(\tau_t^k - \tau^k\right) K_{t-1}^p\left[r_t^k u_t^k - a\left(u_t^k\right) - \delta\right] + \left(\tau_t^p - \tau^p\right)\left(\zeta_t - w_t\right) \quad (32)
$$

whilst the remaining fraction is financed by issuing government bonds:

$$
\frac{B_t - B}{R_t^g} = \left(1-\psi_\tau\right)\left(D_t - D\right) \qquad (33)
$$

We assume that the different tax rates are partially adjusted by choosing the vector of government tax

[12] The hypothesis that the central bank targets trend output instead of the output that would have prevailed in the absence of nominal rigidities has been adopted in the empirical literature (e.g. Del Negro *et al.*, 2006; Adolfson *et al.*, 2007) and is consistent with the main objective of our analysis, which is basically empirical.

instruments $\boldsymbol{\omega} = \left[\omega^c \omega^n \omega^k \omega^p\right]'$, where $\omega^c + \omega^n + \omega^k + \omega^p = 1$.

$$\omega^c \psi_\tau \left(D_t - D\right) = \left(\overline{\tau}_t^c - \tau^c\right) C_t \tag{34}$$

$$\omega^n \psi_\tau \left(D_t - D\right) = \left(\overline{\tau}_t^n - \tau^n\right) w_t n_t \tag{35}$$

$$\omega^k \psi_\tau \left(D_t - D\right) = \left(\overline{\tau}_t^k - \tau^k\right) \frac{k_{t-1}^p}{\mu} \left[r_t^k u_t^k - a\left(u_t^k\right) - \delta\right] \tag{36}$$

$$\omega^p \psi_\tau \left(D_t - D\right) = \left(\overline{\tau}_t^p - \tau^p\right) \left(\zeta_t - w_t\right) \tag{37}$$

where $\overline{\tau}_t^i$, $i = c, n, k, p$, denotes the systematic component on the revenue side, which relates to the stochastic tax rate considering a first order autoregressive stochastic wedge $\eta_t^{\tau^i}$ denoting the discretionary component, such that $\tau_t^i = \overline{\tau}_t^i \eta_t^{\tau^i}$, with $\eta_t^{\tau^i} = \eta_{t-1}^{\tau^i \rho_{\tau^i}} e^{\varepsilon_{\tau^n,t}}$.

An optimal rule is considered for government investment expenditures. The fiscal authority is assumed to choose the public capital stock K_t^g and public investment I_t^g by maximizing the distance between output Y_t and the financial need, i.e.:

$$\begin{aligned} &\max_{K_t^g, I_t^g} E_t \sum_{j=t}^{\infty} \beta^{t+j} \frac{\Lambda_{t+j}}{\Lambda_t} \left[Y_{t+j} - D_{t+j}\right] \\ \text{s.t. } Y_t \quad &= \quad (\xi_t^a)^{(1-\xi)} (K_{t-1}^g)^{\xi} (K_t)^{\alpha(1-\xi)} \left[\mu^t n_t\right]^{(1-\alpha)(1-\xi)} \\ K_t^g \quad &= \quad (1-\delta^g) K_{t-1}^g + q_t^{i^g} \left[1 - S^g\left(\frac{I_t^g}{I_{t-1}^g}\right)\right] I_t^g \end{aligned}$$

where δ^g is the public capital depreciation rate and $S^g(\frac{I_t^g}{I_{t-1}^g})$ denotes the government investment adjustment cost function, with curvature parameter ψ^{ig}. The first order conditions for government capital and investment are, respectively:

$$\beta E_t \left[(1-\delta^g) \Lambda_{t+1}^{k^g} q_t^{k^g} + \Lambda_{t+1} \xi (\xi_{t+1}^a)^{(1-\xi)} (K_t^g)^{\xi-1} (K_{t+1})^{\alpha(1-\xi)} \left(\mu^{t+1} n_{t+1}\right)^{(1-\alpha)(1-\xi)}\right] - \Lambda_t^{k^g} = 0$$

$$\beta E_t \left[q_{t+1}^{i^g} \Lambda_{t+1}^{k^g} S^{g\prime}\left(\frac{I_{t+1}^g}{I_t^g}\right)\left(\frac{I_{t+1}^g}{I_t^g}\right)^2\right] + \Lambda_t^{k^g} q_t^{i,g} \left[1 - S^g\left(\frac{I_t^g}{I_{t-1}^g}\right) - S^{g\prime}\left(\frac{I_t^g}{I_{t-1}^g}\right)\left(\frac{I_t^g}{I_{t-1}^g}\right)\right] - \frac{P_t^d}{P_t} \Lambda_t = 0$$

where $\Lambda_t^{k^g}$ is the shadow price of government capital and $q_t^{i^g} = q_{t-1}^{i^{g \rho_{ig}}} e^{\varepsilon_{i^g,t}}$ is a stochastic process for the government investment-specific shock.

1.6 Model closure

Given the presence of intertemporally optimizing households $j \in [0, 1-\phi^h]$ and of rule-of-thumb households $j \in (1-\phi^h, 1]$, aggregate consumption and government transfers are given by:

$$C_t = \left(1 - \phi^h\right) C_t^r + \phi^h C_t^{nr} \tag{38}$$

and

$$TR_t = \left(1 - \phi^h\right) TR_t^r + \phi^h TR_t^{nr} \tag{39}$$

where, given $d = TR_t^{nr}/TR_t^r$, the fraction of government transfers to Ricardian and non Ricardian households are, respectively: $TR_t^r(i) = \frac{TR_t}{1+\phi^h(d-1)}$ and $TR_t^{nr}(i) = \frac{dTR_t}{1+\phi^h(d-1)}$.

Since only Ricardian households hold bonds and accumulate capital, aggregate variables are related to the vector of Ricardian-specific variables as follows:

$$X_t = \left(1 - \phi^h\right) X_t^r$$

where $X_t = [I_t, K_t^p, K_t, B_t, B_t^*]'$.

Market clearing for the foreign bond market and the final goods market requires that at the equilibrium the following two equations for net foreign assets evolution and aggregate resources are satisfied:

$$\frac{e_t B_{t+1}^*}{\Phi_t R_t^* q_t^{b*}} = e_t P_t^x \left(C_t^x + I_t^x\right) - e_t P_t^* \left(C_t^m + I_t^m\right) + e_t B_t^* \tag{40}$$

and:

$$C_t^d + C_t^x + I_t^d + I_t^x + G_t + I_t^g + \frac{\kappa_b}{2}\left(\frac{R_{t+s}^l(i)}{R_{t+s-1}^l(i)} - 1\right)^2 L_{t+s}(i) \leq Y_t - a(u_t) K_{t-1}^p - \kappa_t \upsilon_t \tag{41}$$

where $C_t^x + I_t^x = \left[\frac{P_t^x}{P_t^*}\right]^{-\eta_*} Y_t^*$ are total exports, with η_* denoting the foreign demand elasticity parameter[13].

The stationary representation of the model is obtained by scaling the real variables with respect to the trending technology process. The scaled model is then log-linearized around the deterministic steady state, taking into account that the presence of a deterministic term in the productivity growth process affects the coefficients of the dynamic equations.

The resulting log-linearized model is composed of 55 structural equations and of 22 shock processes, of which eight are assumed to be first order autoregressive and the remaining 14 are assumed to be *i.i.d.*. The economic relations are described by 67 structural parameters (including the fiscal and monetary policy rules coefficients), whilst the stochastic component of the model is defined by 30 coefficients (22 for the standard deviations of shocks and eight for the autoregressive coefficients)[14].

1.7 The foreign economy

Foreign output (y_t^*), inflation (π_t^*), short and long-term interest rates $\left(r_{s,t}^* \text{ and } r_{b,t}^*, \text{ respectively}\right)$ are exogenous to the variables of the small domestic economy and their evolution is described by a fourth-order

[13] At the estimation stage we will also consider an additive stochastic process ϱ_t in the aggregate resources constraint, i.e. a first order autoregressive measurement error $\varrho_t = \varrho_{t-1}^{\rho_\varrho} e^{\varepsilon_{\varrho,t}}$. Such a shock is generally considered in the empirical literature in order to enhance the estimates when these include output and all its components appearing in the model.

[14] We denote as structural parameters those defining preferences, technology, elasticities, real and nominal rigidities in the good and labor markets, as well as the coefficients describing the monetary and fiscal policy reaction rules. The seven autoregressive coefficients are those describing the memory of the technology process around the deterministic trend, of the structural shock on government investments, on exports, the home bias, the uncovered interest parity, the long-term interest rate spread and the memory of a measurement error included in the aggregate constraint.

structural Bayesian B-VAR, where contemporaneous correlations are defined by the structure of the stochastic component matrix $\mathbf{B}$. Formally:

$$\mathbf{A}(L)\begin{bmatrix}\pi_t^*\\ \Delta y_t^*\\ r_{s,t}^*\\ r_{b,t}^*\end{bmatrix} = \mathbf{B}\begin{bmatrix}\varepsilon_t^{\pi^*}\\ \varepsilon_t^{y^*}\\ \varepsilon_t^{r_s^*}\\ \varepsilon_t^{r_b^*}\end{bmatrix}, \quad \mathbf{A}_0 = \mathbf{I}_4, \quad \boldsymbol{\varepsilon}_t \sim N(\mathbf{0}, \mathbf{I}_4) \tag{42}$$

$$\mathbf{B} = \begin{bmatrix} b_{11} & 0 & 0 & 0\\ 0 & b_{22} & 0 & 0\\ b_{31} & b_{32} & b_{33} & 0\\ b_{41} & b_{42} & b_{43} & b_{44}\end{bmatrix}, \quad \mathbf{BB}' = \mathbf{\Omega}$$

The assumptions on the contemporaneous correlations matrix $\mathbf{B}$ are consistent with the hypothesis that output and inflation do not respond contemporaneously to the other shocks in the system (Adolfson *et al.*, 2008)[15], and that the 10-years government bond rate is post-recursive with respect to the short-term interest rate.

The SVAR system adds four linear stochastic equations to the economic and stochastic relations of the domestic economy model, resulting in a total of 81 equations and 26 shocks.

2 Bayesian estimation

The rich parameterization of the model precludes the estimation of the entire parameter space, because of the poor empirical identifiability of medium and large scale DSGE models (Canova and Sala, 2009; Iskrev, 2010*a*,b; Koop *et al.*, 2011). Even if log-linearized around the deterministic steady state, these structures are in fact characterized by relevant nonlinearities in parameter convolutions, such that the likelihood generated by the model can be uninformative, i.e. multimodal or flat with respect to some parameter values. On these premises, only the subset of the parameter space that satisfies the theoretical and empirical identification conditions is estimated using the Bayesian method, whilst for the remaining subset we adopt dogmatic priors specified according to the available country-specific evidence and to conventional calibration values.

A Bayesian approach is adopted also for the estimation of the foreign variables SVAR, in this case considering a partially modified Minnesota priors specification approach. The choice of using the Bayesian method for the estimation of the SVAR is based on recent results showing its good properties both within sample and in terms of minimization of the predictive variance of the resulting model (Banbura *et al.*, 2010).

2.1 Data issues and measurement equations

To enhance the empirical identification of the widest fraction of the structural parameters space, we use a large set of domestic and foreign quarterly variables to estimate the country-specific models.

[15] Consistently with the results in Adolfson *et al.* (2011), the over-identifying restriction that output does not respond contemporaneously to the price shock is not rejected by the data at the standard 5% criterion.

Considering the domestic economies, 22 observables are considered: (log differences of) of real per capita GDP[16] (Δy_t^{obs}), consumption (Δc_t^{obs}), investment (Δi_t^{obs}), imports (Δm_t^{obs}), exports(Δx_t^{obs}), the real wage (Δw_t^{obs}), real government expenditures for consumption (Δg_t^{obs}), investment $\left(\Delta i_t^{g,obs}\right)$ and transfers (Δtr_t^{obs}); the tax rate on labor income $\left(\tau_t^{n,obs}\right)$, on business profits $\left(\tau_t^{p,obs}\right)$, on capital $\left(\tau_t^{k,obs}\right)$ and on consumption $\left(\tau_t^{c,obs}\right)$; the unemployment rate (u_t^{obs}), the (quarterly) rates of change of the price deflators for consumption $\left(\pi_t^{c,obs}\right)$, import $\left(\pi_t^{m,obs}\right)$, export $\left(\pi_t^{x,obs}\right)$ and for the domestic sector $\left(\pi_t^{y,obs}\right)$; the nominal effective exchange rate (e_t^{obs}), the (quarterly) short-term interest rate, the 10-years government bond rate and the lending rate to non financial corporations $\left(r_{s,t}^{obs}, r_{b,t}^{obs} \text{ and } r_{l,t}^{obs} \text{ respectively}\right)$. All real variables are referred to the base-year 2005.

Considering the variables for the foreign sector, the log difference of real output $\left(y_t^{*,obs}\right)$ is obtained from the real world output index (base-year 2005) and short and long-term interest rates $\left(r_{s,t}^{*,obs} \text{ and } r_{b,t}^{*,obs}, \text{ respectively}\right)$ are obtained as weighted averages of the corresponding figures for the US and the EMU area, with weights given by the relative importance of the two economic areas in domestic capital movements. The foreign price deflator $\left(\pi_t^{*,obs}\right)$ is obtained from the real effective exchange rate definition equation using observed data on domestic inflation, the nominal and the real effective exchange rates. A total of 26 variables is thus considered in the country-specific estimates[17].

All data are taken from official sources and cover the period 1980:1-2012:4[18]. Real variables of the private domestic sector, their deflators and the nominal short and long-term interest rates are taken from the OECD-Economic Outlook database. Nominal and real effective exchange rate indexes, defined at the base-year 2005, the world real output index (2005 = 100) and the lending rates to nonfinancial corporations are taken from the IMF-International Financial Statistics database. Data for government expenditures and revenues are, for the quarterly frequency $(1999 - 2012)$, from the IMF Government Financial Statistics database and, for the yearly frequency, from the OECD-Tax Statistics database and from the IMF Finance Statistics Yearbook [19].

Before linking the observed variables to the theoretical counterparts, some of the latter are transformed in order to get full consistency with the statistical definitions. In particular, the transformations take into account that, differently from the statistical aggregates, consumption and investment in the theoretical model are composites of domestic and imported goods and output also includes the hiring cost and that related to changes in the capital utilization rate.

Further transformations are needed in order to make the data consistent with the theoretical steady states and in particular with the model property of balanced growth (μ), a theoretical prediction which is not supported by the evidence in all the countries being considered, in particular for export and import shares.

[16] Per capita variables are obtained considering the labor force as the normalizing variable.

[17] To the best of our knowledge, the use of such a high number of observables in the estimates is unprecedented in the literature on empirical DSGE models.

[18] Because of the lack of quarterly time series prior to 1990 for Ireland and to 2000 for Greece, quadratic interpolation methods are applied to yearly observations to obtain the quarterly figures 1980:1-1989:4 and 1980:1-1999:4 for Ireland and Greece, respectively.

[19] Even in this case, since quarterly data are available only after 1999:1, adjustments to changing definitions and quadratic interpolation methods are applied to yearly observations in order to obtain the quarterly frequency for the preceding time span. A detailed description of the data manipulation is provided in a technical appendix of the paper, available upon request from the authors

More specifically, the positive/negative excess trends in real variables are removed by considering sample deviations from the steady state output growth rate μ in the measurement equations of all the real variables in the system, such that the theory-consistent stationary great ratios are restored.

Formally, considering the vector of real per capita variables $\mathbf{x}_t' = (c_t,\, i_t,\, m_t,\, x_t,\, w_t,\, g_t,\, i_t^g,\, tr_t,\, y_t^*)$, of tax rates $\boldsymbol{\tau}_t' = \left(\tau_t^n,\, \tau_t^p,\, \tau_t^k,\, \tau_t^c\right)$, of inflation rates $\boldsymbol{\pi}_t' = (\pi_t^c,\, \pi_t^m,\, \pi_t^x,\, \pi_t^y,\, \pi_t^*)$, of short-term, bond and lending interest rates $\mathbf{r}_{s,t}' = \left(r_{s,t},\, r_{s,t}^*\right)$, $\mathbf{r}_{b,t}' = \left(r_{b,t},\, r_{b,t}^*\right)$ and $r_{l,t}$, the 26 measurement equations linking the linearized model variables to the respective observables read as follows:

$$\begin{bmatrix} \Delta y_t^{obs} \\ \Delta \mathbf{x}_t^{obs} \\ \boldsymbol{\tau}_t^{obs} \\ u_t^{obs} \\ \boldsymbol{\pi}_t^{obs} \\ \mathbf{r}_{s,t}^{obs} \\ \mathbf{r}_{b,t}^{obs} \\ r_{l,t}^{obs} \\ e_t^{obs} \end{bmatrix} = \begin{bmatrix} \widetilde{y}_t - \widetilde{y}_{t-1} + \log \mu \\ \widetilde{\mathbf{x}}_t - \widetilde{\mathbf{x}}_{t-1} + \log \mu + \log \boldsymbol{\mu}_{xy} \\ \widetilde{\boldsymbol{\tau}}_t + \boldsymbol{\tau} \\ \widetilde{u}_t + u \\ \widetilde{\boldsymbol{\pi}}_t + \log \boldsymbol{\pi} \\ \widetilde{\mathbf{r}}_{s,t} - \log \overline{\boldsymbol{\beta}}^{(.,*)} + \log \boldsymbol{\pi}^{(c,*)} \\ \widetilde{\mathbf{r}}_{b,t} - \log \overline{\boldsymbol{\beta}}^{(.,*)} + \log \boldsymbol{\pi}^{(c,*)} + \mathbf{q}^{b(.,*)} \\ \widetilde{r}_{l,t} - \log \overline{\beta}^{(.,*)} + \log \pi^c + p^{d,p} \\ \widetilde{e}_t + \log e \end{bmatrix} \tag{43}$$

where the coefficients μ_{xy} denote the excess trend (or excess growth rate) of each observed generic real per capita variable in $\mathbf{x}_t^{obs}$ from the real per capita GDP growth rate, μ. $\boldsymbol{\tau}$, $-\log \overline{\boldsymbol{\beta}}$, $\boldsymbol{\pi}$, $\mathbf{q}^b$, $q^{d,p}$ and s denote the (steady state) tax rates, the domestic and foreign real interest rates, the inflation rates, the domestic and foreign government bond rate spreads, the lending rate default probability and the nominal effective exchange rate, respectively, and u denotes the steady state unemployment rate.

2.2 Calibrated parameters

Calibrated values are chosen taking into account both sample and extraneous evidence when informative for the theoretical parameters, and conventional values when such information is missing.

We impose 29 dogmatic priors on the 67-dimensional structural parameters space. Absent country-specific information, 17 structural parameters are fixed to common values across countries. These are the steady-state mark-up coefficients λ_p^d, λ_p^m and λ_p^x, fixed to the conventional value of 1.2, consistent with prior demand elasticities for domestic, import and export sector firms equal to 6; the Kimball endogenous demand elasticity parameters κ_ϵ^d, κ_ϵ^m and κ_ϵ^x, fixed to the conventional value of 10 (Eichenbaum and Fisher, 2007; Smets and Wouters, 2007); the parameter defining the fraction of government transfers to Ricardian and non Ricardian households d, fixed to 1, consistent with an hypothesis of equally distributed transfers; the three parameters defining the partial indexation mechanism for the domestic, import and export sectors , i.e. ι_p^d, ι_p^m and ι_p^x, respectively, all fixed to zero in order to allow for an interpretation of the (observed) frequency of price changes in terms of (theoretical) price re-optimization[20]; the exchange rate sensitivity to the net foreign assets to GDP

[20] Under the hypothesis of indexation, prices are changed period by period, ruling out any intepretation of the observed

ratio $\widetilde{\phi}_a$, fixed to the arbitrary small value of 1^{-3}([21]); the private and government capital depreciation rates, δ and δ^g, respectively, both fixed to the conventional value of 0.025; the steady-state mark-up coefficient for the credit sector λ^b, fixed to the value of 1.025, consistent with a demand elasticity parameter equal to 40 (*INS CITAZIONE*); the shape parameter for the government default probability function $\phi^{s,g}$ in (19), fixed to 20 in order to capture the recent observed nonlinear relation between fundamentals and government bond spreads; the scale parameter for the private sector default probability $\varphi^{s,p}$ in (17), fixed to 5 to initialize the estimation of the corresponding shape parameter $\phi^{s,p}$ in a neigborhood of a unit prior value, consistent with the relatively stable relation between the lending and the government bond rate spreads observed in the data; the world contribution to the IMF parameter ϕ^{i*}, fixed to 0.008 according to the observed total SDR (in USD) to world GDP ratio[22].

The remaining 12 dogmatic priors for structural parameters are fixed considering country-specific evidence. These are the trend growth parameter μ, fixed considering the sample growth rate of per capita GDP, the discount factor β, calibrated considering the country-specific trend growth and the average real interest rate, the home bias parameter $(1-\nu)$, fixed according to the country-specific sample evidence on import shares, the separation rate ρ, fixed to the country estimates provided by Hobijn and Sahin (2009), the parameter defining the frequency of wage re-optimization θ_w, fixed to the country estimates provided in Druant *et al.* (2012), and the parameter defining the unemployment benefit b^u, fixed according to the country-specific replacement rates provided in the OECD-LFS data base (Christoffel *et al.*, 2009). The private capital share α, the matching efficiency parameter σ_m and the labor disutility scale parameter χ are calibrated such that the labor share, the unemployment rate and the job finding rate steady-state values evaluated at the prior parameterization match the sample counterparts for each country[23]. Considering the country-specific dogmatic priors for the financial sector parameters, the contribution to the IMF parameter ϕ^i is set according to the country SDR quota (in Euro) to GDP ratios, whilst the international insurance efficiency parameters σ_z is fixed such that the debt repayment rate parameter z^g in (*20*) matches the country-specific sample SDR quota. The country-specific scale parameter of the government default probability function $\varphi^{s,g}$ is fixed in the following manner: given the country-specific z^g parameter and the sample government bond rate differential q^b (evaluated with respect to the short-term interest rate), the country-specific government default probability $p^{d,g}$ is obtained from equation (*21*). The latter univocally determines the country-specific scale parameter $\varphi^{s,g}$ from (*19*), given the common shape parameter $\phi^{s,g}$, the sample debt to GDP ratio and a prior value for the government default probability λ_b.

Finally, the coefficients in the system of measurement equations (43), i.e. those in the vector of deviations from GDP trend $\boldsymbol{\mu}_{xy}$, in the vectors of tax rates $\boldsymbol{\tau}$, of inflation rates $\boldsymbol{\pi}$, of domestic and foreign real interest rates and bond rate spreads, $-\log\overline{\boldsymbol{\beta}}$ and $\mathbf{q}_b$, respectively, and the long-run nominal effective exchange rate e,

frequencies of price changes in terms of frequencies of price re-optimizations.

[21] Such a small value ensures the satisfaction of the stability conditions (Lundvik, 1992; Schmitt-Grohé and Uribe, 2001) while minimizing the exchange rate persistence induced by its "technical" relation with the NFA evolution.

[22] We assume full equivalence between the amount of resources devoted to the IMF and SDR quotas.

[23] Sample data for the job finding rate are obtained by elaborating the information in the OECD Labor Force Survey data-base series "Unemployment by duration".

are fixed to the respective sample means.

The seven exclusion restrictions for the identification of the foreign variables' SVAR, i.e. the zero restriction for b_{12}, b_{13}, b_{14}, b_{21}, b_{23}, b_{24} and b_{34} add further seven dogmatic priors. Table 1 summarizes the common and country-specific dogmatic priors adopted in model estimation for the structural parameters.

TABLE 1 ABOUT HERE

2.3 Priors for estimated parameters

The subset of (38) structural model parameters who is not affected by evident identification problems, the 34 coefficients defining the stochastic component (30 for the domestic economy model and 4 for the foreign SVAR) and the 73 coefficients of the SVAR system (nine for the elements of the $\mathbf{B}$ matrix and 64 for the vector autoregressive component) are estimated with the Bayesian method[24].

Outside the Calvo price parameters, the prior distributions are common across countries and are specified following the standard practice: *i)* the shape of the probability density functions is the gamma and the inverted gamma for parameters theoretically defined over the $\mathbb{R}^+$ range, the beta for parameters defined in a $[0-1]$ range and the normal for priors on parameters theoretically defined over the $\mathbb{R}$ range; *ii)* prior means and standard deviations are defined on the basis of sample information (when available), or considering the results of previous analyses[25]. In order to enhance the estimation of parameters subject to weak empirical identifiability, informative priors are adopted such that a certain degree of curvature in the log-kernel is obtained.

The prior means for the Calvo parameters of the domestic, import and export sectors, (ξ_p^d, ξ_p^d ξ_p^d, respectively) are specified according to the country-specific micro-evidence provided in Druant *et al.* (2012)[26], i.e. 0.71 for Greece, 0.75 for Ireland, 0.69 for Portugal and 0.70 for Italy and Spain. Since the available information does not distinguish across sectors, we adopt a relatively high value for the prior standard deviation, equal to 0.1. A weak gamma-distributed prior with mean 1.5 and standard deviation 0.4 is adopted for the import and export Armington elasticities η and η^* (Adolfson *et al.*, 2008; Christiano *et al.*, 2011).

Considering the modified UIP equation, the autoregressive coefficient $\widetilde{\phi}_s$ is assumed to be beta-distributed with prior mean 0.5 and prior s.d. 0.15, whilst for the country risk adjustment coefficient $\widetilde{\phi}_r$ we basically follow Christiano *et al.* (2011), assuming a (more) diffuse gamma distribution with prior mean 1.25 and prior s.d. 0.5.

The private and public investment adjustment cost parameters ψ^i and ψ^{ig} are assumed to be normally

[24] Operationally, posterior modes are obtained by maximizing the log-posterior kernel (resulting from the prior distribution and the conditional distribution approximated by the Kalman filter) with respect to the model parameters, and posterior distributions are obtained from the Metropolis-Hastings Monte Carlo Markov chain (MCMC) numerical integration algorithm. Two chains of 500k iterations are considered.

[25] The standard practice of considering results from previous studies is not free of limitations, since the validity domain of prior evidence is not independent of the model being considered.

[26] The Kimball curvature, Calvo and mark-up (or demand elasticity) parameters are not separately identifiable, as testified by the results of preliminary identification checks at the prior values (Iskrev, 2010*a*,*b*). We adopt the standard practice of fixing the Kimball and mark-up parameters to ensure the empirical identification of the estimated Calvo parameters.

distributed around a prior mean of 5 with a prior s.d. of 2.5, and the utilization rate curvature parameter ψ^k is assumed to be beta-distributed with prior mean 0.5 and prior s.d. 0.15 (Christiano *et al.* 2011).

Concerning the preference parameters, the consumption curvature parameter σ_c is assumed to be normally-distributed with a prior mean of 2 and a prior s.d. of 0.1, whilst the external habits parameter is assumed to be beta-distributed and centered around 0.8 with a prior s.d. of 0.1. The prior for the fraction of liquidity constrained households is rather diffuse, with mean 0.25 and s.d. 0.10[27].

Considering the labor market-specific parameters, a relatively weak beta-distributed prior with mean 0.5 and s.d. 0.15 is assumed for the matching function share parameter σ_n and the union's relative bargaining power parameter ς. The prior for the hiring cost parameter κ is assumed to be gamma-distributed with mean 0.05 and s.d. 0.01, a prior mean value consistent with a hiring cost to GDP ratio $\frac{\kappa v}{Y}$ close to 1%.

Considering the financial sector parameters defining the government and private sector default probabilities and interest rate spreads, a gamma-distributed prior with mean 0.5 and s.d. 0.25 is adopted for the sensitivity coefficients λ_b and λ_a (to the debt and net foreign assets to GDP ratios, respectively), and the shape parameter for the private sector default probability function $\phi^{s,p}$ is assumed to be normally-distributed with a prior mean of 1 and a prior s.d. of 0.5. These mean values are set jointly with the dogmatic priors on the other financial sector parameters and ensure exact correspondence between the steady state government bond and lending rate spreads and their sample counterparts. The parameter defining the fraction of borrowed wage bill ϑ_b is assumed to be beta-distributed with prior mean 0.5 and s.d. 0.25, whilst the lending rate adjustment cost parameter κ_b is assumed to be gamma-distributed with prior mean 3 and s.d. 1.5 (Gerali *et al.,* 2010). The very diffuse prior distributions adopted for these parameters reflect our imprecise prior opinions, and imply that their posterior estimates will be dominated by the conditional distribution.

Concerning the monetary policy parameters, the interest rate smoothness coefficient ρ^R is assumed to be beta-distributed with prior mean 0.75 and prior s.d. 0.2, the inflation response parameter ψ_1 is assumed to be normally distributed with prior mean 2 and s.d. 0.2, whilst the output growth sensitivity parameter ψ_2 is assumed to be beta-distributed with prior mean (s.d.) of 0.25 (0.1). The three shift parameters accounting for the monetary policy structural break in the smoothness coefficient and in the feedback coefficients are assumed to be normally distributed with zero prior mean and s.d. equal to 0.2.

Considering the fiscal policy parameters, a beta-distributed prior with mean 0.75 and s.d. 0.15 is adopted for the autoregressive components ρ_{τ^c}, ρ_{τ^n}, ρ_{τ^k} and ρ_{τ^k} in the tax rates partial adjustment equations, and ρ_g, ρ_{tr} in the government consumption and transfers equations, respectively. For the coefficients denoting the sensitivity of these expenditure components to output, η_{gy} and η_{try}, an informative and normally distributed prior with mean 1 and s.d. 0.1 is adopted, consistent with the hypothesis of long-run balanced growth of public expenditures. A weakly informative beta-distributed prior with mean 0.05 and s.d. 0.02 is chosen for the parameters η_{gd} and η_{trd}, defining the sensitivity of public consumption and transfers to the government financial need. The latter prior is equivalent to that chosen for the sensitivity of the tax rates to the financial

[27] The preference parameters, even if separately identifiable in our setting, are not fully variation-free. The choice of a relatively tight prior for the consumption curvature parameter enhances the identifiability of the other parameters.

need ψ_τ, basically following the calibration value adopted in Drautzburg and Uhlig (2011). Finally, a weakly informative beta-distributed prior with mean 0.25 and s.d. 0.10 is adopted for the tax instruments ω^c, ω^n and ω^k, whilst ω^p is restricted to be equal to $1-\left(\omega^c+\omega^n+\omega^k\right)$.

Considering the stochastic component of the models, the prior opinions for the autoregressive coefficients of the seven persistent shock processes (i.e., ρ_{ξ^a}, ρ_{i^g}, $\rho_{\tilde{\phi}}$, ρ_{q_b}, ρ_{ϱ}, ρ_ν and ρ_x) are commonly described by a weakly informative beta-distributed prior with mean 0.75 and s.d. 0.15[28]. For the standard errors of the 26 innovations, we assume a prior mean of 0.01 with two degrees of freedom for all shocks, except those multiplying convolutions of parameters whose values are outside the $\left[10^{-1},\ 10\right]$ range, that are scaled accordingly.

The prior opinions on the estimated structural parameters are summarized in the first column of the result Table 2 (panels a-f).

The elicidation of priors for the foreign variables' SVAR is based on the partially modified Minnesota priors approach (Doan *et al.*, 1984; Litterman, 1986; Sims and Zha, 1998) suggested by Banbura *et al.* (2010). Accordingly, priors are specified under the hypothesis of independent AR(1) processes (random walks for variables close to non-stationarity), with prior variabilities decreasing in the power of the lag order of the SVAR i (net of an overall shrinkage parameter λ, calibrated according to the number of variables in the system) and scaled considering the variables' error variance ratios σ_m^2/σ_n^2, the latter approximated by the estimated residuals of univariate autoregressive representations. Formally, the prior moments for the 73 coefficients of the fourth-order SVAR (42) are specified as follows:

$$E\left[(\mathbf{A}_i,\mathbf{B})_{mn}\right]=\begin{array}{l}\vartheta \text{ for } i=1,\ m=n\\ 0 \text{ otherwise}\end{array}\quad,\quad V\left[(\mathbf{A}_i,\mathbf{B})_{mn}\right]=\begin{array}{l}\frac{\lambda^2}{i^2} \text{ for } m=n\\ \frac{\lambda^2}{i^2}\frac{\sigma_m^2}{\sigma_n^2} \text{ otherwise}\end{array}\qquad(44)$$

where the values for the first-order autoregressive coefficients ϑ are obtained from the estimates of independent AR(1) processes.

2.4 Posterior mean estimates

Table 2 summarizes the priors and the posterior mean estimates. Panels *a*-*b*-*c*-*d* consider the model economy, the financial sector, the monetary policy and the fiscal policy parameters, respectively. Panels *e* and *f* report the estimates of the 34 parameters defining the persistence and the size of the 26 exogenous stochastic components, respectively[29].

According to the estimated posterior mode standard deviations and the implied pseudo t-values, the structural parameter estimates, aside from ϑ_b, all appear significant for each of the countries being considered.

[28] The autoregressive coefficients ρ_ν and ρ_x denote the persisitency of the stochastic component in the import and export equations, respectively. Analitically, the first component defines a stochastic home bias parameter, and the second a stochastic elasticity of substitution between foreign and domestic goods. The two stochastic components enter the log-linear representation of the model additively, such that they do not influence the empirical identifiability of the preference parameters.

[29] Mode checks and multivariate M-H convergence plots signal that the estimation process performs correctly for all countries. The mode estimates intersect the log posterior kernel at its maximum for all parameters. The multivariate diagnostics signal that the estimates are stable both within (over replications) and particularly between chains. Posterior densities confirm these encouraging indications, signaling a close to normal shape and a reasonable distance from prior densities (or a more concentrated distribution), signalling that the estimated parameters are empirically identified. These results are available upon request from the authors.

The exogenous innovations are all significant according to their standard errors and a relevant degree of autocorrelation is obtained for the subset of autoregressive processes.

The posterior mean values for the model economy parameters are generally close to the respective modal values and indicate reasonable estimates based on our prior opinions and results in the literature. Evident exceptions are the unconventionally high posterior estimates obtained for the private and public capital adjustment cost parameters ψ^{i} and ψ^{ig}, on average more than the double of the prior mean values, implying milder investment and capital responses than those obtainable under standard calibration values. Furthermore, the curvature parameter for the capital utilization rate ψ^{k} is estimated to be very high and distant from the prior in all countries. These results imply slow adjustments on both the investment and the capital utilization sides, thus - other things being equal - high persistence in model dynamics.

TABLE 2a ABOUT HERE

A relevant degree of cross-country heterogeneity is obtained with respect to the parameter defining the fraction of liquidity constrained households ϕ^{h}, that are quite high for Portugal (0.49), basically in line with the EZ estimates in Coenen and Straub (2005) and Forni *et al.* (2009) for Italy (0.36), Ireland (0.24) and Spain (0.24) and quite low for Greece (0.13). These differences are expected to affect the size of the fiscal multipliers, since a higher degree of rule-of-thumb behavior is reflected in a more direct link between current income and private consumption, i.e. in the breakdown of Ricardian equivalence (Galì *et al.*, 2007).

The posterior mean estimates of the Calvo parameters in the domestic, import and export sectors, ξ_p^d, ξ_p^m and ξ_p^x, respectively, are generally higher than the prior opinions based on survey evidence and the conventional values used in the literature. This result basically reflects the flat slope of the NKPCs, which is more pronounced than that implied by the joint consideration of the Calvo frequency micro-estimates and of the conventional calibration values for the mark-up (or demand elasticity) parameters[30].

The estimated Armington elasticity η, and in particular η^{*}, are generally smaller than the prior and denote a differentiated pattern across countries. A similar consideration holds true for the risk premium parameter $\widetilde{\phi}_r$, which is estimated to be slightly above unit only for Spain and Italy, thus ruling out a direct emergence of the forward premium puzzle in the remaining countries.

The labor market parameters show a certain degree of variability across countries, particularly for the union's relative bargaining power parameter ς, estimated to be higher than the conventional value of 0.5 for all countries except Italy (0.34). The posterior mean estimates for the hiring cost parameter κ and the matching function share parameter σ_n are not distant from priors, except for the former parameter in the case of Ireland ($\kappa = 0.032$).

[30] For the countries considered in this study, the introduction of endogenous demand elasticities does not solve the micro-macro dichotomy in the estimate of the NKPC slope coefficients (Eichenbaum and Fisher, 2005).

TABLE 2b ABOUT HERE

Concerning the financial sector parameters, the coefficient capturing the elasticity of the government default probability to the debt to GDP ratio λ_b is estimated to be well above the prior in all countries, ranging from a minimum of 0.66 for Spain to a maximum of 1.64 for Portugal. The elasticity to the net foreign assets to GDP ratio λ_a is on average smaller and more in line with the prior, ranging from a minimum of 0.31 for Greece to a maximum of 0.62 for Spain. The estimated shape parameter of private sector default probability function is more homogeneous across countries and on average twice the prior size. A high degree of heterogeneity is estimated for the lending rate adjustment cost parameter κ^b, ranging from a minimum of 1.13 for Portugal to a maximum of 21.5 for Spain. An evaluation of the elasticity of the government and private sector default probabilities (thus of the government bond and lending rate spreads) to the debt and net foreign assets to GDP ratios cannot be directly obtained from these parameters. Table three reports the expected variation in the government bond and lending rate spreads consistent with a 20 percentage points temporary increase in the debt to GDP ratio and in the net foreign assets to GDP ratio in the different countries.

TABLE 3 ABOUT HERE

TABLE 2c ABOUT HERE

Considering the estimated monetary policy coefficients adjusted for the break implied by the shift to the single currency, a low degree of policy activism emerges. The size of the policy rate response to inflation ψ_1 is low in all countries, ranging from a minimum of 1.05 for Spain to a maximum of 1.28 for Portugal, whilst the output growth response coefficient ψ_2 ranges from a minimum of 0.05 for Greece to a maximum of 0.12 for Italy. Joint with the estimated high degrees of inertial behavior (the coefficient ρ^R is always well above 0.8), these results indicate a particularly mild monetary policy response to variations in inflation and output, potentially dampening its counter-cyclical effects under standard fiscal expansions.

It is interesting to note that the posterior estimates of the three shift parameters accounting for the monetary policy structural break are negative in all countries being considered, signalling that the shift to a common currency and a centralized authority targeting average EZ inflation and output has implied a reduced degree of monetary policy activism with respect to the single economies' macroeconomic developments[31].

TABLE 2d ABOUT HERE

Finally, the posterior estimates for the fiscal policy coefficients confirm the high degree of inertia on both

[31] Detailed results on the monetary policy break estimates are reported in a technical appendix available upon request from the authors.

the expenditure and the revenue sides, with estimated autoregressive coefficients well above the conventional calibration value of 0.9 (Perotti, 2005). It is interesting to note that the posterior estimates for the parameter denoting the sensitivity of the tax rates to the government financial need ψ_{τ}, even if low and distant from the prior, are basically consistent with the Galì and Perotti (2003) estimates for OECD countries. interestingly, the estimated sensitivities of government consumption and transfers to the financial need (η_{gd} and η_{trd}, respectively) are on average higher and more heterogeneous across countries, with a minimum size close to 0.01 for Ireland and a maximum size close to 0.08 for Greece. The parameter defining the link between long-run expenditure and output levels (η_{gy} and η_{try}) are always not significantly different from unity, such that the hypothesis of balanced growth in the fiscal variables, for the sample being considered, cannot be rejected.

TABLE 2e ABOUT HERE

TABLE 2f ABOUT HERE

3 Policy simulations

In this section we provide a comparative analysis of the country-specific expected effects from the implementation of five financially equivalent contractionary fiscal policies: *i)* a persistent, albeit not permanent, reduction in government consumption; *ii)* an equally persistent reduction in government transfers; *iii)* a reduction in government investment; *iv)* a generalized increase of indirect tax rates (on labor incomes, business profits and capital gains); *v)* an increase in the consumption tax rate. These policies are evaluated by simulating the model stochastically (thus assuming that they are unanticipated) and considering the parameterization obtained at the country-specific posterior mean estimates.

The different simulations are made comparable by calibrating the size of each policy shock to be equivalent to a 1% of GDP on impact and by homogenizing their persistence considering a common memory coefficient of 0.75, consistent with a one year average duration of the policy shock.

By construction, each policy measure implies government budget and debt variations, thus changes in the tax rates and in the structure of public expenditure. However,in order to enhance the understanding of the simulation results, we only consider the estimated systematic components in the revenue equations, i.e., the specific elasticity of the tax rates to the financial need, whilst the expenditure side is assumed to be fully exogenous by setting the elasticities of the expenditure components to the financial need and to GDP to zero.

The policy simulations are performed assuming both a standard environment, i.e. one in which the monetary policy reacts to inflation and output growth deviations from target according to the estimated values of the Taylor rule feedback coefficients, and a recessionary environment in which the economies are operating in a liquidity trap. To implement such a scenario, we calibrate a negative preference shock implying an eight-quarters period non positive equilibrium interest rate for each country, and impose the zero-lower-bound (ZLB) condition.

3.1 Government purchase and direct tax shocks: into the mechanics of the risk channel

Before discussing the results of the specific austerity measures, it is worth providing some details on the dynamics activated by two alternative policy interventions on expenditure and revenues, i.e. a 1% GDP negative government consumption shock and a 1% GDP positive shock to direct taxes (on labor income, business profits and capital gains), depicted in Figures 2 and 3, respectively. The latter multiple shock is obtained considering that the the 1% GDP fiscal contraction is obtained by incresing the specific tax rates according to the estimated policy instruments weights ω^i, $i = n, p, k$.

To clarify the functioning of the transmission mechanics under the hypothesis of a default risk channel, the 20 quarters ahead impulse responses of GDP, the debt level, of the debt to GDP ratio, of the NFA evolution and of the government bond and lending rate spreads are reported. These are normalized such that the GDP response has an interpretation in terms of the dynamic monetary fiscal multiplier (i.e. the expected monetary variation in GDP from a one euro budget variation), the debt to GDP ratio response depicts the deviation from its steady state in terms of GDP percentage points, and the responses of the spreads refer to annualized basic points.

FIGURE 2 ABOUT HERE

Considering the government consumption contraction, a first outcome that merits to be highlighted is the modest variability of the output response across countries, reflecting the low sensitivity of the dynamic multiplier of this measure to the heterogeneity in the estimated parameterization. This is due to the fact that government purchases affect output mainly directly, inducing only second-round effects on price and wage dynamics. The peak response is negative and reached on impact, and denotes a monetary multiplier ranging from values slightly above 1 for Greece, Italy and Spain, to 1.35 for Portugal and above 1.8 for Ireland. These results are fully consistent with the available average European estimates (Coenen and Straub, 2005; Forni *et al.*, 2009). In the standard times scenario, there are no evident signals of the operation of a sovereign debt channel, since the size of the country-specific multipliers are basically aligned with those obtainable from equally parameterized country models in which the default risk channel effects are eliminated.

As expected, the fiscal contraction leads to a reduction in the bond level in all countries, signalling that the positive response of government expenditure, due to the rise in unemployment benefits payments, and the negative response of revenues, due to the tax rate cuts implicit to their endogenous specification, are not sufficient to reverse the positive effects of the fiscal contraction on the level of debt.

However, since the fiscal contraction leads to a more than proportional decrease in output, the debt to

GDP ratio temporarily increases in all the PIIGS countries, with a dynamic pattern which is substantially dominated by the negative output response. The highest increase of the debt ratio, close to 1.7% of GDP, is obtained on impact for Ireland, consistently with the negative output response; the smallest, close to 0.45% of GDP, is obtained on impact for Spain. Conditional to our model and to the estimated parameterization, fiscal austerity plans implemented with government purchase cuts are thus expected to be self-defeating in the short-term.

In line with the expectations, the NFA response is positive in all countries, with evident cross-country heterogeneity. The effects are stronger in Ireland and Portugal, consistent with the deeper output contraction and, in the case of Ireland, with the higher estimated elasticity coefficient of imports, leading to even deeper reductions in imported goods.

The moderate but positive response of both the interest rate spreads in all countries signals that the improved NFA position relative to GDP is not enough in couterbalancing the pressure on sovereign default risk due to the increase in the debt to GDP ratio. In other terms, the size of the elasticity of default risk and thus of the bond spread to the variation in the debt to GDP ratio is high enough to dominate the counteracting effects implicit in the improvement in the NFA positions.

These results signal that, conditional to a negative government consumption shock, the default risk channel operates in the opposite direction than predicted in the analysis of Corsetti *et al.* (2013). Moreover, the size of the interest rate spreads response is very limited, signalling that, according to our model estimates, the default risk channel is basically irrelevant. Aside from the role played by the estimated small size of the elasticity of default risk to the macroeconomic fundamentals, the main responsible for these results is the consideration of the debt to GDP ratio in the place of the debt level, whose response to a fiscal contraction is positive for sufficiently large fiscal multipliers.

The effects of a contractionary direct taxes shock are only qualitatively similar to those obtained considering a financially equivalent government consumption reduction. The fiscal contraction has negative and persistent effects on real output for all countries, even if the implicit peak multipliers are substantially smaller than those obtained with the government purchase shock, a result which is basically in line with the abundant SVAR-based empirical literature on fiscal multipliers since the seminal analysis of Blanchard and Perotti (2002). Moreover, the output dynamic multiplier is heterogeneous across countries, mainly because of the different fractions of liquidity constrained households estimated in the different countries. The fraction of rule-of-thumb households is in fact estimated to be particularly low for Greece, reflecting the low correlation between private consumption and current net incomes in the sample. Considering the recent evolution of the Greek economy, it is highly probable that the fraction of liquidity contrained households increased strongly. We have verified that, by including a dummy variable controlling for the recessionary periods, the estimated degree of liquidity constraints increases by nearly 18 percentage points for Greece.

Following the tax rates shock, the debt level decreases temporarily in all countries but Ireland, partly because of the higher unemployment response and the resulting increase in unemployment benefit payments. As a result of the debt and the GDP dynamics, a moderate but persistent surge in their ratio emerges.

Even in the case of a revenue-based fiscal contraction, our results indicate that the hypothesis of expansionary fiscal contractions is not empirically relevant, such that the implementation of austerity plans can be self-defeating in the short to medium term, and that the hypothesis of a sovereign risk channel, if effective, operates in the opposite direction when evaluated conditional to fiscal shocks.

It is interesting to note that, under the tax-based fiscal contraction, the positive response of the net foreign asset position obtained in all countries is always significantly larger than that obtained under the expenditure-based contraction, despite the smaller drop in economic activity. This implies that the response of imports is much stronger, a result signalling that the tax reduction induces a significant variation in the relative price of the domestic production, i.e. a real exchange rate devaluation. The internal devaluation is triggered by the increased tax pressure, implying an immediate contraction of the after tax incomes and of the consumption expenditures of liquidity-constrained households. Even if the resulting decrease in labor supply tends to counterbalance the deflationary pressure, the latter tends to prevail.

Concerning the effects of the fiscal retrenchment on the sovereign default probability on bond and lending rate spreads, the impulse responses clearly show that the contractionary tax policy, similarly to the contractionary expenditure policy, stimulates a moderate increase in the government bond and lending rate spreads.

Two key indications from the analysis of the conditional dynamics emerge: *first*, the relation between sovereign debt, net foreign position and interest rate spreads is rather weak in all the peripheral EZ economies considered in the study, such that the recent surge in government bond and lending rate premia in these countries should be mainly attributed to idiosyncratic factors only loosely related with macroeconomic fundamentals (De Grauwe and Ji, 2013); *second*, the hypotesis that - when monetary policy is unconstarined - a sovereign risk channel can mitigate the contractionary effects of fiscal consolidations or even - in a liquidity trap constrained regime - lead to an economic expansion, is not empiriclly supported when considering a short to medium term perspective, since fiscal contractions are temporarely but persistently self-defeating, irrespective of the policy instrument being considered.

The explanation for the different results of our analysis as compared to those in Corsetti *et al.* (2013) relies heavily on the measure of indebtness considered in the definition of the default probability. The use of the debt level basically constrains the direction of change of the default probability to the one of the policy. The use of the debt to GDP ratio, which is generally accepted as a more appropriate measure of fiscal health, does not impose such a restriction and highlights the role of the size of the fiscal multipliers, determining the direction of the variation in the debt to GDP ratio and thus of the default probability.

3.2 Fiscal contractions in unconstrained and constrained monetary policy regimes

The relative efficacy of alternative fiscal measures in different countries depends both on the different degrees of nominal and wage rigidity and on the interaction between fiscal and monetary policy regimes. In particular, an aggressive monetary policy increases the expected effects of fiscal measures targeted to induce a price deflation through the reduction of the labor cost, and dampens those of policies stimulating the general economic activity, because of their inflationary implications.

The fact that the labor market targeted fiscal policies being evaluated are expected to be implemented in economies operating well below their potential, as is the case of the countries considered in this study, suggests to extend the analysis to the situation of a binding ZLB. In these circumstances, a deflationary fiscal policy cannot be accommodated by the automatic response of the monetary authority, since the nominal interest rate cannot be reduced further (Eggertsson et al., 2013)[32]. On the contrary, an expansionary and inflationary fiscal policy, until it does not succeed in taking the economy out of the liquidity trap, will not face the same counteracting effects originating in the stabilizing response of the monetary policy during standard times (Christiano et al. 2011; Eggertsson, 2011; Eggertsson and Krugman, 2012). Tables 5 and 6 replicate, for a below potential-liquidity trap economic environment, the information on the fiscal multipliers and on the employment effects of the alternative policies provided by Tables 3 and 4 for the economies operating at their potential output levels. Since strongly negative output multipliers are often found, one row reporting the peak negative multiplier is added in Table 5.

The consideration of a liquidity trap environment affects the efficacy of the labor market targeted fiscal policies in different directions in the short and in the long term. Considering the hiring cost subsidization policy, the short-term output multipliers are significantly negative in all countries but Greece, (between -0.04 for Spain and -2.6 for Portugal), whilst the long-term peak output multipliers are increased and delayed further (between 0.5 for Ireland and 3.2 for Greece). Qualitatively similar results are obtained considering the subsidization of the wage of the new hires of labor, for which the short term multipliers are again negative (between -0.03 for Spain and -2.5 for Portugal), whilst in the long run their peak values are confirmed to be increased (between 0.4 for Ireland and 5.2 for Greece). The employment effects are instead always positive, even if the stronger peak employment reduction is in general delayed further as compared to the standard time simulations.

The transmission mechanics explaining these results is the same described for the simulations assuming a not binding ZLB environment. Even in this case, the subsidization policy generates a deflation through the real wage contraction. The main difference here is that, for the eight periods in which the ZLB binds, the monetary authority cannot accommodate the policy with a nominal interest rate reduction, such that the resulting increase in the real interest rate is of the same size of the price deflation. The transitory but sizeable negative output response amplifies the real wage contraction and the deflation during the liquidity trap period.

As the economy recovers, the monetary authority decreases the policy rate by a larger amount than in a not binding ZLB environment, because of the stronger deflation, and firms are willing to hire more workers, because of the stronger real wage contraction. This justifies the expansion following the transitory but persistent depression activated by the labor market policies.

Notwithstanding the amplified and delayed long run output responses, and with the exception of Greece, the labor market targeted policies are confirmed to be inferior to a fiscal policy expansion based on government consumption. As expected, the output and employment effects of fiscal expansions based on government

[32] Eggertsson *et al.* (2013), by simulating a monetary model calibrated to average EZ data, show that a permanent reduction in product and labor market markups (a structural policy in authors' terms), can have contractionary short term effects when the economy is in a liquidity trap.

expenditures are significantly increased, with the peak government consumption output multipliers in the range $1.7 - 3.3$, and the unemployment reduction within -0.8% and -1.3%. When the ZLB binds, the counteracting response of the monetary authority does not take place until the economy is out of the liquidity trap. In this circumstance, the real interest rate tends to decrease with the increased inflation, adding a positive private expenditure response to the government stimulus.

TABLE 5 ABOUT HERE

TABLE 6 ABOUT HERE

It is interesting to note that, under a binding ZLB, fiscal expansions based on tax rate cuts are counterproductive in all countries in the short term, and basically ineffective in the long run. This result is only apparently surprising. On the one hand, a labor tax cut increases the after tax current income, leading to both increased labor supply and to increased consumption demand in the fraction of liquidity constrained households. On the other, the increased labor supply induces a real wage and thus marginal cost contraction, activating a deflationary pressure. Since only a minor fraction of households are liquidity constrained, the deflation stimulated by the reduced tax pressure prevails such that, given the fixed policy rate, an increase in the real interest rate emerges, leading to reduced private expenditures[33].

4 Conclusions

We develop, estimate and simulate a model characterized by government bond and lending rate spreads originating in the sovereign default risk triggered by internal and foreign debt positions. The consideration of an endogenous default risk channel introduces interesting elements for the conduct of fiscal policy in highly indebted economies, especially when the economy is stuck at the ZLB. In principle, for increasing levels of debt and for small sized fiscal multipliers, a fiscal retrenchment can even be espansionary, given the induced reduction in the domestic and foreign debt positions, triggering a reduction in sovereign and private default risk and thus of the interest rate spreads.

The analysis, developed at the country-level for a selection of peripheral EZ economies (the PIIGS), is based on the simulation of the country-specific response of output and employment to financially equivalent contractionary fiscal policies affecting government expenditure and revenues. *Results show that, contrary to some conclusions in the recent literature and the policy recommendations within the European EP and YG programmes, the labor market targeted fiscal measures, in a short term perspective, are not superior to more*

[33]The mechanics behind this result has been explained in detail by Eggertsson (2010) in a simplified model setting assuming full Ricardian equivalence. In his comment to the Eggertsson's (2010) paper, Christiano (2010) provides some useful insights and identifies two major ingredients for the deflationary pressure to emerge following a tax cut: *i)* the persistence of the deflationary pressure, i.e. the presence of relevant price rigidities; *ii)* the sensitivity of expenditures to the real interest rate, i.e. the empirical relevance of the Euler consumption equation. Our results, emerging in an extended structural model setting estimated on country data, provide evidence in support to Eggertsson's result giving an empirical assesment of both key factors.

standard fiscal instruments in the management of the business cycle. The analysis also indicates that, even in a longer term perspective and aside Greece, the output multiplier of government consumption is higher than that from hiring costs and newly hired workers' subsidization. Considering the employment effects, these policies prove to be clearly superior to more standard fiscal expansions only in the long term and at the Greece and Ireland model parameter estimates.

The consideration of a liquidity trap environment reinforces these conclusions, as both output and employment multipliers of government expenditures are significantly increased. On the contrary - and with the exception of Greece - the output multiplier of the labor market targeted measures are strongly negative in the short term, and their peak effects are reached with an increased delay as compared with the standard environment simulations.

These results basically highlight the importance of the fiscal-monetary policy coordination in the business cycle management, an option which might be out of reach during a deep recession.

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TABLE 1 - DOGMATIC PRIORS: STRUCTURAL PARAMETERS

Parameter	Greece	Ireland	Italy	Portugal	Spain
β	0.994	0.997	0.996	0.998	0.995
α	0.265	0.220	0.333	0.210	0.220
δ	0.025	0.025	0.025	0.025	0.025
δ^g	0.025	0.025	0.025	0.025	0.025
ν	0.206	0.656	0.210	0.262	0.202
ρ	0.028	0.042	0.021	0.039	0.061
σ_m	0.910	0.200	0.950	0.500	1.500
χ	0.100	1.000	4.000	0.200	0.600
b^u	0.650	0.650	0.630	0.720	0.610
θ_w	0.750	0.800	0.850	0.770	0.750
λ_p^i	1.200	1.200	1.200	1.200	1.200
κ_ϵ^i	10.00	10.00	10.00	10.00	10.00
ι_p^i	0.000	0.000	0.000	0.000	0.000
μ	0.999	1.007	1.002	1.003	1.002
$\widetilde{\phi}_a$	0.001	0.001	0.001	0.001	0.001
d	1.000	1.000	1.000	1.000	1.000
λ_b	1.025	1.025	1.025	1.025	1.025
$\phi^{s,g}$	20.00	20.00	20.00	20.00	20.00
$\varphi^{s,g}$	0.490	0.565	0.509	0.614	0.759
$\varphi^{s,p}$	5.000	5.000	5.000	5.000	5.000
ϕ^i	0.027	0.022	0.013	0.023	0.011
ϕ^{i*}	0.008	0.008	0.008	0.008	0.008
σ_z	0.228	0.621	0.560	0.340	0.461

Note: $i = d, m, x$.

TABLE 2*a* - PRIOR DISTRIBUTIONS AND POSTERIOR MEAN ESTIMATES: MODEL ECONOMY

	Prior distribution		Posterior mean				
	Density	Mean (s.d.)	Greece [c.i.]	Ireland [c.i.]	Italy [c.i.]	Portugal [c.i.]	Spain [c.i.]
ξ_p^d	$\mathcal{G}$	$0.69-0.75^*$	0.933	0.892	0.879	0.848	0.924
		(0.10)	$[0.927-0.939]$	$[0.875-0.911]$	$[0.870-0.888]$	$[0.832-0.865]$	$[0.916-0.933]$
ξ_p^m	$\mathcal{G}$	$0.69-0.75^*$	0.856	0.878	0.848	0.925	0.789
		(0.10)	$[0.819-0.899]$	$[0.843-0.913]$	$[0.816-0.883]$	$[0.903-0.948]$	$[0.753-0.824]$
ξ_p^x	$\mathcal{G}$	$0.69-0.75^*$	0.821	0.843	0.820	0.875	0.901
		(0.10)	$[0.784-0.860]$	$[0.805-0.886]$	$[0.774-0.878]$	$[0.842-0.915]$	$[0.895-0.906]$
σ_c	$\mathcal{N}$	2.00	2.074	2.012	1.909	2.024	2.078
		(0.10)	$[1.919-2.227]$	$[1.843-2.178]$	$[1.774-2.048]$	$[1.858-2.193]$	$[1.923-2.240]$
h	$\mathcal{B}$	0.70	0.828	0.905	0.818	0.910	0.845
		(0.10)	$[0.791-0.866]$	$[0.872-0.940]$	$[0.784-0.853]$	$[0.876-0.941]$	$[0.814-0.877]$
ϕ^h	$\mathcal{B}$	0.25	0.127	0.239	0.358	0.490	0.238
		(0.10)	$[0.081-0.174]$	$[0.119-0.360]$	$[0.276-0.443]$	$[0.377-0.598]$	$[0.165-0.311]$
η	$\mathcal{G}$	1.50	1.050	1.514	0.445	0.601	1.092
		(0.40)	$[0.865-1.218]$	$[0.855-2.167]$	$[0.313-0.580]$	$[0.438-0.764]$	$[0.914-1.273]$
η^*	$\mathcal{G}$	1.50	0.526	0.826	0.852	0.527	0.607
		(0.40)	$[0.400-0.658]$	$[0.651-0.998]$	$[0.748-0.955]$	$[0.409-0.647]$	$[0.469-0.736]$
$\widetilde{\phi}_s$	$\mathcal{B}$	0.50	0.827	0.834	0.942	0.883	0.960
		(0.15)	$[0.725-0.954]$	$[0.694-0.957]$	$[0.904-0.981]$	$[0.836-0.934]$	$[0.934-0.986]$
$\widetilde{\phi}_r$	$\mathcal{G}$	1.25	0.948	0.878	1.010	0.886	1.247
		(0.50)	$[0.885-1.002]$	$[0.778-0.970]$	$[0.964-1.057]$	$[0.841-0.923]$	$[1.097-1.396]$
ψ^i	$\mathcal{N}$	5.00	13.04	12.40	11.37	10.80	10.89
		(2.50)	$[10.14-15.90]$	$[10.01-14.82]$	$[9.389-13.33]$	$[8.669-12.95]$	$[8.65-13.20]$
ψ^{ig}	$\mathcal{N}$	5.00	12.30	15.43	15.11	6.765	13.96
		(2.50)	$[9.380-15.08]$	$[12.66-18.26]$	$[12.57-17.55]$	$[4.275-9.413]$	$[11.33-16.53]$
ψ^k	$\mathcal{B}$	0.50	0.987	0.645	0.970	0.972	0.949
		(0.15)	$[0.979-0.996]$	$[0.608-0.683]$	$[0.958-0.982]$	$[0.953-0.992]$	$[0.921-0.976]$
σ_n	$\mathcal{B}$	0.50	0.553	0.571	0.525	0.584	0.501
		(0.10)	$[0.447-0.671]$	$[0.427-0.707]$	$[0.373-0.690]$	$[0.450-0.712]$	$[0.332-0.662]$
ς	$\mathcal{B}$	0.50	0.658	0.780	0.337	0.628	0.571
		(0.10)	$[0.582-0.736]$	$[0.720-0.841]$	$[0.245-0.426]$	$[0.559-0.706]$	$[0.484-0.660]$
κ	$\mathcal{G}$	0.05	0.041	0.034	0.047	0.043	0.050
		(0.01)	$[0.028-0.055]$	$[0.024-0.044]$	$[0.032-0.062]$	$[0.032-0.056]$	$[0.034-0.065]$

*: denotes the range of values for the country-specific values in Druant et al. (2009).

TABLE 2b - PRIOR DISTRIBUTIONS AND POSTERIOR MEAN ESTIMATES: FINANCIAL SECTOR

	Prior distribution		Posterior mean				
	Density	Mean	Greece	Ireland	Italy	Portugal	Spain
		(s.d.)	[c.i.]	[c.i.]	[c.i.]	[c.i.]	[c.i.]
ϑ_b	$\mathcal{B}$	0.50	0.201	0.218	0.037	0.270	0.027
		(0.20)	[0.002 – 0.452]	[0.001 – 0.475]	[0.000 – 0.081]	[0.007 – 0.533]	[0.000 – 0.050]
λ_b	$\mathcal{G}$	0.50	1.093	1.627	1.295	1.643	0.664
		(0.25)	[0.668 – 1.502]	[1.153 – 2.079]	[0.823 – 1.765]	[1.219 – 2.061]	[0.363 – 0.961]
λ_a	$\mathcal{G}$	0.50	0.311	0.428	0.429	0.462	0.618
		(0.25)	[0.050 – 0.535]	[0.098 – 0.748]	[0.080 – 0.766]	[0.119 – 0.782]	[0.262 – 0.963]
$\phi^{s,p}$	$\mathcal{N}$	1.00	1.957	1.933	1.959	2.106	1.733
		(0.50)	[1.678 – 2.235]	[1.604 – 2.263]	[1.611 – 2.292]	[1.778 – 2.420]	[1.402 – 2.063]
κ_b	$\mathcal{G}$	3.00	3.389	1.516	5.782	1.131	21.55
		(1.50)	[0.500 – 6.382]	[0.240 – 2.853]	[1.665 – 9.695]	[0.421 – 1.798]	[18.87 – 23.77]

TABLE 2c - PRIOR DISTRIBUTIONS AND POSTERIOR MEAN ESTIMATES: MONETARY AUTHORITY

	Prior distribution		Posterior mean				
	Density	Mean	Greece	Ireland	Italy	Portugal	Spain
		(s.d.)	[c.i.]	[c.i.]	[c.i.]	[c.i.]	[c.i.]
ρ^R	$\mathcal{B}$	0.75	0.844	0.900	0.884	0.879	0.877
		(0.20)	[0.825 – 0.865]	[0.879 – 0.922]	[0.869 – 0.898]	[0.861 – 0.896]	[0.863 – 0.891]
ψ_1	$\mathcal{N}$	2.00	1.174	1.220	1.115	1.279	1.045
		(0.20)	[1.129 – 1.222]	[1.099 – 1.334]	[1.053 – 1.173]	[1.203 – 1.351]	[1.019 – 1.072]
ψ_2	$\mathcal{B}$	0.10	0.051	0.052	0.123	0.055	0.134
		(0.05)	[0.024 – 0.081]	[0.027 – 0.075]	[0.094 – 0.149]	[0.033 – 0.076]	[0.105 – 0.163]

TABLE 2*d* - PRIOR DISTRIBUTIONS AND POSTERIOR MEAN ESTIMATES: FISCAL AUTHORITY

	Prior distribution		Posterior mean				
	Density	Mean	Greece	Ireland	Italy	Portugal	Spain
		(s.d.)	[c.i.]	[c.i.]	[c.i.]	[c.i.]	[c.i.]
ω^c	$\mathcal{B}$	0.25	0.392	0.550	0.223	0.423	0.484
		(0.10)	[0.274 – 0.516]	[0.427 – 0.675]	[0.144 – 0.307]	[0.318 – 0.528]	[0.380 – 0.585]
ω^n	$\mathcal{B}$	0.25	0.604	0.244	0.703	0.580	0.477
		(0.10)	[0.478 – 0.720]	[0.130 – 0.363]	[0.600 – 0.812]	[0.459 – 0.701]	[0.370 – 0.581]
ω^k	$\mathcal{B}$	0.25	0.007	0.205	0.031	0.001	0.008
		(0.10)	[0.003 – 0.011]	[0.096 – 0.309]	[0.010 – 0.051]	[0.000 – 0.001]	[0.003 – 0.012]
ψ_τ	$\mathcal{B}$	0.05	0.013	0.012	0.013	0.019	0.013
		(0.02)	[0.009 – 0.016]	[0.008 – 0.016]	[0.010 – 0.016]	[0.012 – 0.026]	[0.010 – 0.015]
ρ_{τ^c}	$\mathcal{B}$	0.75	0.969	0.977	0.955	0.953	0.979
		(0.15)	[0.944 – 0.995]	[0.961 – 0.995]	[0.919 – 0.991]	[0.918 – 0.987]	[0.963 – 0.996]
ρ_{τ^n}	$\mathcal{B}$	0.75	0.986	0.989	0.988	0.988	0.964
		(0.15)	[0.974 – 0.997]	[0.981 – 0.998]	[0.979 – 0.998]	[0.978 – 0.999]	[0.940 – 0.990]
ρ_{τ^k}	$\mathcal{B}$	0.75	0.975	0.981	0.978	0.987	0.981
		(0.15)	[0.956 – 0.994]	[0.969 – 0.994]	[0.959 – 0.997]	[0.977 – 0.999]	[0.968 – 0.995]
ρ_{τ^p}	$\mathcal{B}$	0.75	0.977	0.972	0.961	0.992	0.970
		(0.15)	[0.960 – 0.995]	[0.951 – 0.993]	[0.931 – 0.993]	[0.985 – 0.999]	[0.947 – 0.993]
ρ_g	$\mathcal{B}$	0.75	0.924	0.958	0.960	0.976	0.981
		(0.15)	[0.881 – 0.966]	[0.932 – 0.982]	[0.931 – 0.991]	[0.959 – 0.995]	[0.966 – 0.998]
ρ_{tr}	$\mathcal{B}$	0.75	0.974	0.973	0.971	0.916	0.984
		(0.15)	[0.953 – 0.995]	[0.960 – 0.986]	[0.953 – 0.990]	[0.870 – 0.960]	[0.971 – 0.999]
η_{gy}	$\mathcal{N}$	1.00	0.927	0.974	1.035	1.038	1.044
		(0.10)	[0.770 – 1.081]	[0.808 – 1.137]	[0.873 – 1.203]	[0.868 – 1.199]	[0.871 – 1.219]
η_{try}	$\mathcal{N}$	1.00	0.999	1.016	1.027	1.022	1.034
		(0.10)	[0.826 – 1.152]	[0.849 – 1.185]	[0.862 – 1.192]	[0.864 – 1.190]	[0.870 – 1.194]
η_{gd}	$\mathcal{B}$	0.05	0.032	0.020	0.020	0.021	0.011
		(0.02)	[0.013 – 0.049]	[0.009 – 0.030]	[0.012 – 0.028]	[0.010 – 0.031]	[0.007 – 0.016]
η_{trd}	$\mathcal{B}$	0.05	0.078	0.063	0.019	0.023	0.017
		(0.02)	[0.041 – 0.114]	[0.048 – 0.077]	[0.012 – 0.026]	[0.011 – 0.034]	[0.010 – 0.023]

TABLE 2e - PRIOR DISTRIBUTIONS AND POSTERIOR MEAN ESTIMATES: AR(1) COEFICIENTS OF SHOCKS

	Prior distribution		Posterior mean				
	Density	Mean	Greece	Ireland	Italy	Portugal	Spain
		(s.d.)	[c.i.]	[c.i.]	[c.i.]	[c.i.]	[c.i.]
ρ_{ξ^a}	$\mathcal{B}$	0.75	0.994	0.942	0.912	0.936	0.924
		(0.15)	[0.989 – 0.999]	[0.926 – 0.960]	[0.895 – 0.930]	[0.914 – 0.957]	[0.910 – 0.938]
ρ_{i^g}	$\mathcal{B}$	0.75	0.904	0.778	0.154	0.193	0.804
		(0.15)	[0.855 – 0.951]	[0.758 – 0.797]	[0.059 – 0.245]	[0.084 – 0.294]	[0.734 – 0.876]
$\rho_{\bar{\phi}}$	$\mathcal{B}$	0.75	0.736	0.837	0.890	0.982	0.915
		(0.15)	[0.332 – 0.926]	[0.776 – 0.901]	[0.865 – 0.916]	[0.973 – 0.993]	[0.889 – 0.940]
ρ_{q_b}	$\mathcal{B}$	0.75	0.947	0.953	0.906	0.925	0.923
		(0.15)	[0.915 – 0.984]	[0.923 – 0.982]	[0.878 – 0.935]	[0.894 – 0.958]	[0.895 – 0.952]
ρ_{ϱ}	$\mathcal{B}$	0.75	0.933	0.995	0.911	0.955	0.934
		(0.15)	[0.891 – 0.999]	[0.990 – 0.999]	[0.886 – 0.936]	[0.925 – 0.989]	[0.915 – 0.953]
ρ_{ν}	$\mathcal{B}$	0.75	0.966	0.973	0.910	0.943	0.945
		(0.15)	[0.948 – 0.987]	[0.951 – 0.994]	[0.885 – 0.936]	[0.901 – 0.985]	[0.926 – 0.965]
ρ_{x}	$\mathcal{B}$	0.75	0.973	0.982	0.894	0.940	0.904
		(0.15)	[0.965 – 0.983]	[0.968 – 0.996]	[0.866 – 0.923]	[0.888 – 0.989]	[0.880 – 0.927]
ρ_{λ_b}	$\mathcal{B}$	0.50	0.631	0.863	0.593	0.905	0.840
		(0.20)	[0.520 – 0.742]	[0.799 – 0.932]	[0.473 – 0.718]	[0.860 – 0.953]	[0.780 – 0.917]

TABLE 2*f* - PRIOR DISTRIBUTIONS AND POSTERIOR MEAN ESTIMATES: S.D. OF SHOCK PROCESSES

	Prior distribution		Posterior mean				
	Density	Mean	Greece	Ireland	Italy	Portugal	Spain
		(s.d.)	[c.i.]	[c.i.]	[c.i.]	[c.i.]	[c.i.]
$\varepsilon_{\tau^n,t}$	$\mathcal{G}^{-1}$	0.01	0.008	0.003	0.005	0.006	0.003
		(2.00)	[0.007 – 0.008]	[0.003 – 0.004]	[0.004 – 0.005]	[0.006 – 0.007]	[0.003 – 0.004]
$\varepsilon_{\tau^p,t}$	$\mathcal{G}^{-1}$	0.01	0.003	0.003	0.003	0.002	0.004
		(2.00)	[0.002 – 0.003]	[0.003 – 0.004]	[0.003 – 0.003]	[0.002 – 0.003]	[0.004 – 0.004]
$\varepsilon_{\tau^k,t}$	$\mathcal{G}^{-1}$	0.01	0.004	0.019	0.007	0.001	0.004
		(2.00)	[0.004 – 0.005]	[0.017 – 0.021]	[0.006 – 0.008]	[0.001 – 0.001]	[0.003 – 0.004]
$\varepsilon_{\tau^c,t}$	$\mathcal{G}^{-1}$	0.01	0.004	0.005	0.002	0.004	0.004
		(2.00)	[0.003 – 0.004]	[0.004 – 0.005]	[0.002 – 0.002]	[0.004 – 0.005]	[0.003 – 0.004]
$\varepsilon_{g,t}$	$\mathcal{G}^{-1}$	0.01	0.025	0.028	0.019	0.024	0.011
		(2.00)	[0.022 – 0.027]	[0.025 – 0.031]	[0.017 – 0.021]	[0.021 – 0.026]	[0.010 – 0.012]
$\varepsilon_{tr,t}$	$\mathcal{G}^{-1}$	0.01	0.079	0.027	0.014	0.021	0.013
		(2.00)	[0.071 – 0.087]	[0.024 – 0.030]	[0.012 – 0.015]	[0.019 – 0.023]	[0.011 – 0.014]
$\varepsilon_{i^g,t}$	$\mathcal{G}^{-1}$	0.1	0.112	0.198	0.994	0.787	0.124
		(2.00)	[0.088 – 0.133]	[0.164 – 0.231]	[0.822 – 1.157]	[0.477 – 1.094]	[0.092 – 0.155]
$\varepsilon_{\xi^a,t}$	$\mathcal{G}^{-1}$	0.01	0.012	0.019	0.010	0.013	0.009
		(2.00)	[0.010 – 0.013]	[0.016 – 0.021]	[0.009 – 0.011]	[0.011 – 0.015]	[0.008 – 0.010]
$\varepsilon_{r,t}$	$\mathcal{G}^{-1}$	0.01	0.003	0.004	0.002	0.002	0.003
		(2.00)	[0.003 – 0.003]	[0.004 – 0.005]	[0.002 – 0.002]	[0.002 – 0.002]	[0.003 – 0.003]
$\varepsilon^d_{p,t}$	$\mathcal{G}^{-1}$	0.50	7.185	2.747	1.309	0.735	2.836
		(2.00)	[5.873 – 8.433]	[1.815 – 3.668]	[1.086 – 1.530]	[0.555 – 0.918]	[2.128 – 3.520]
$\varepsilon^m_{p,t}$	$\mathcal{G}^{-1}$	0.5	3.605	1.744	2.789	12.42	1.565
		(2.00)	[1.749 – 5.686]	[0.739 – 2.655]	[1.535 – 3.923]	[5.287 – 19.68]	[1.019 – 2.082]
$\varepsilon^x_{p,t}$	$\mathcal{G}^{-1}$	0.5	2.998	2.002	1.814	3.048	1.096
		(2.00)	[1.717 – 4.252]	[0.960 – 3.160]	[0.818 – 3.004]	[1.412 – 5.025]	[0.740 – 1.441]
$\varepsilon_{q_b,t}$	$\mathcal{G}^{-1}$	0.01	0.004	0.004	0.002	0.002	0.002
		(2.00)	[0.003 – 0.004]	[0.003 – 0.004]	[0.002 – 0.002]	[0.002 – 0.002]	[0.002 – 0.003]
$\varepsilon_{q_i,t}$	$\mathcal{G}^{-1}$	0.5	0.284	0.855	0.208	0.320	0.269
		(2.00)	[0.218 – 0.343]	[0.682 – 1.021]	[0.170 – 0.245]	[0.252 – 0.385]	[0.214 – 0.326]
$\varepsilon_{\widetilde{\phi},t}$	$\mathcal{G}^{-1}$	0.01	0.007	0.003	0.003	0.001	0.002
		(2.00)	[0.002 – 0.016]	[0.002 – 0.004]	[0.002 – 0.003]	[0.000 – 0.001]	[0.002 – 0.003]

TABLE 2e - (CONTINUED)

	Prior distribution		Posterior mean				
	Density	Mean	Greece	Ireland	Italy	Portugal	Spain
		(s.d.)	[c.i.]	[c.i.]	[c.i.]	[c.i.]	[c.i.]
$\varepsilon_{\xi^c,t}$	$\mathcal{G}^{-1}$	0.01	0.185	0.337	0.122	0.423	0.112
		(2.00)	[0.136 − 0.231]	[0.219 − 0.451]	[0.093 − 0.151]	[0.258 − 0.591]	[0.084 − 0.140]
$\varepsilon_{\xi^n,t}$	$\mathcal{G}^{-1}$	0.01	0.061	0.043	0.013	0.033	0.025
		(2.00)	[0.051 − 0.071]	[0.033 − 0.053]	[0.011 − 0.015]	[0.028 − 0.038]	[0.021 − 0.029]
$\varepsilon_{x,t}$	$\mathcal{G}^{-1}$	0.01	0.037	0.026	0.030	0.024	0.035
		(2.00)	[0.033 − 0.041]	[0.023 − 0.028]	[0.027 − 0.033]	[0.021 − 0.027]	[0.031 − 0.039]
$\varepsilon_{cpi,t}$	$\mathcal{G}^{-1}$	0.01	0.010	0.011	0.007	0.005	0.012
		(2.00)	[0.009 − 0.011]	[0.010 − 0.012]	[0.006 − 0.008]	[0.004 − 0.005]	[0.011 − 0.013]
$\varepsilon_{\nu,t}$	$\mathcal{G}^{-1}$	0.01	0.028	0.032	0.028	0.021	0.033
		(2.00)	[0.025 − 0.031]	[0.028 − 0.035]	[0.025 − 0.031]	[0.019 − 0.024]	[0.029 − 0.037]
$\varepsilon_{\varrho,t}$	$\mathcal{G}^{-1}$	0.01	0.010	0.013	0.007	0.007	0.009
		(2.00)	[0.009 − 0.011]	[0.011 − 0.014]	[0.007 − 0.008]	[0.006 − 0.008]	[0.008 − 0.010]
$\varepsilon^*_{dp,t}$	$\mathcal{G}^{-1}$	0.005	0.006	0.006	0.006	0.006	0.006
		(2.00)	[0.006 − 0.007]	[0.006 − 0.007]	[0.006 − 0.007]	[0.006 − 0.007]	[0.006 − 0.007]
$\varepsilon^*_{y,t}$	$\mathcal{G}^{-1}$	0.005	0.006	0.006	0.006	0.006	0.006
		(2.00)	[0.005 − 0.006]	[0.005 − 0.006]	[0.005 − 0.006]	[0.005 − 0.006]	[0.005 − 0.006]
$\varepsilon^*_{r,t}$	$\mathcal{G}^{-1}$	0.005	0.002	0.002	0.002	0.002	0.002
		(2.00)	[0.002 − 0.002]	[0.002 − 0.002]	[0.002 − 0.002]	[0.002 − 0.002]	[0.002 − 0.002]
$\varepsilon^*_{rl,t}$	$\mathcal{G}^{-1}$	0.005	0.001	0.001	0.001	0.001	0.001
		(2.00)	[0.001 − 0.001]	[0.001 − 0.001]	[0.001 − 0.001]	[0.001 − 0.001]	[0.001 − 0.001]
$\varepsilon_{\lambda_b,t}$	$\mathcal{G}^{-1}$	0.01	0.133	0.184	0.068	0.103	0.079
		(2.00)	[0.117 − 0.149]	[0.164 − 0.204]	[0.059 − 0.077]	[0.092 − 0.114]	[0.071 − 0.088]

TABLE 3 - EXPECTED INCREASE IN BOND AND LENDING SPREADS - in basis points

20% Increase in	Spread on	Greece	Ireland	Italy	Portugal	Spain
$\frac{B_t}{Yt}$	r_t^g	213.0	19.6	36.9	39.8	85.4
$\frac{A_t^-}{Yt}$	r_t^g	5.6	1.1	2.3	2.2	23.6
$\frac{B_t}{Yt}$	r_t^l	14.0	1.4	0.6	1.3	1.4
$\frac{A_t^-}{Yt}$	r_t^l	0.4	0.1	0.0	0.1	0.4

Note: interest rate spread are expressed in basis points. The lending rate spread does not consider the mark-up

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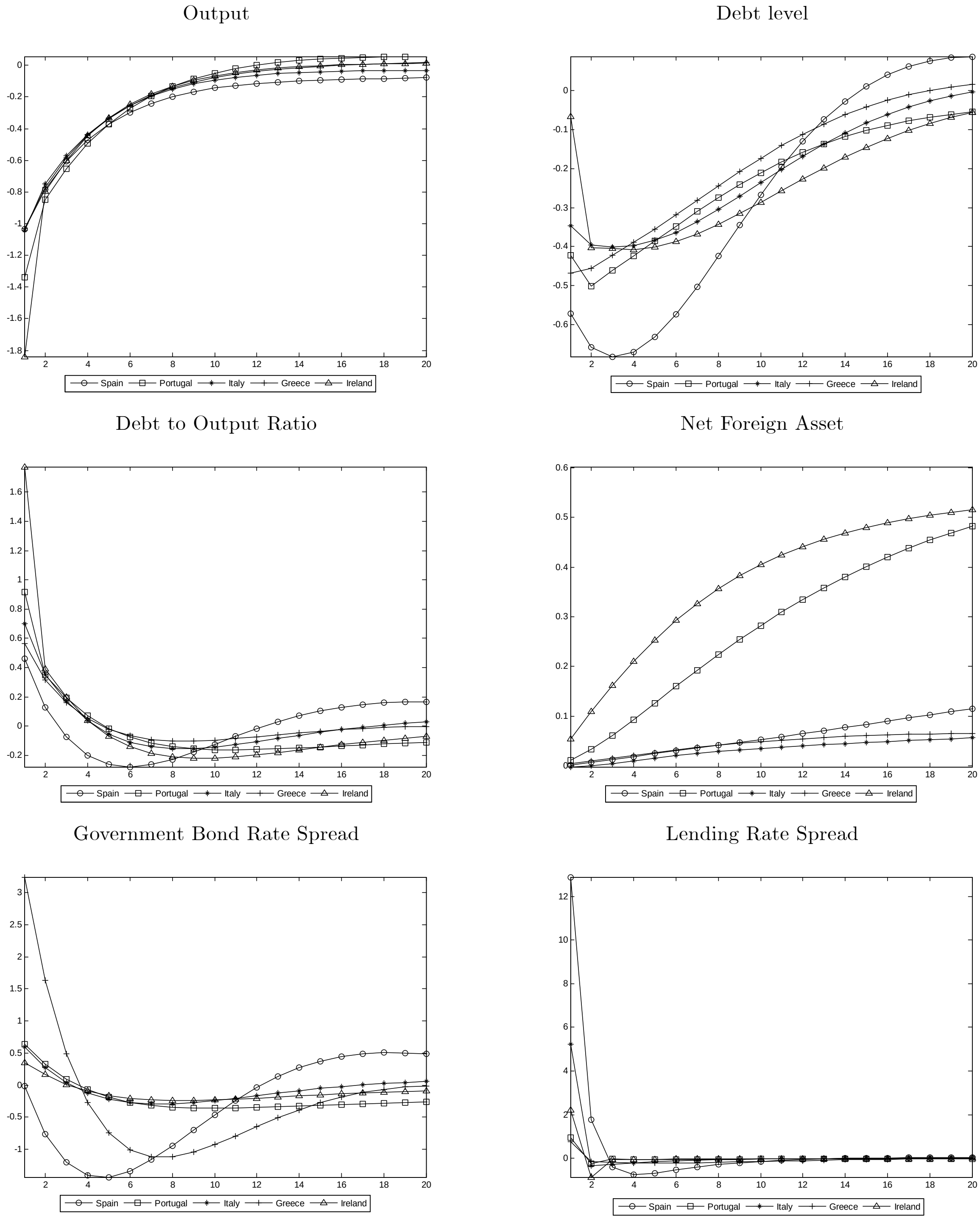
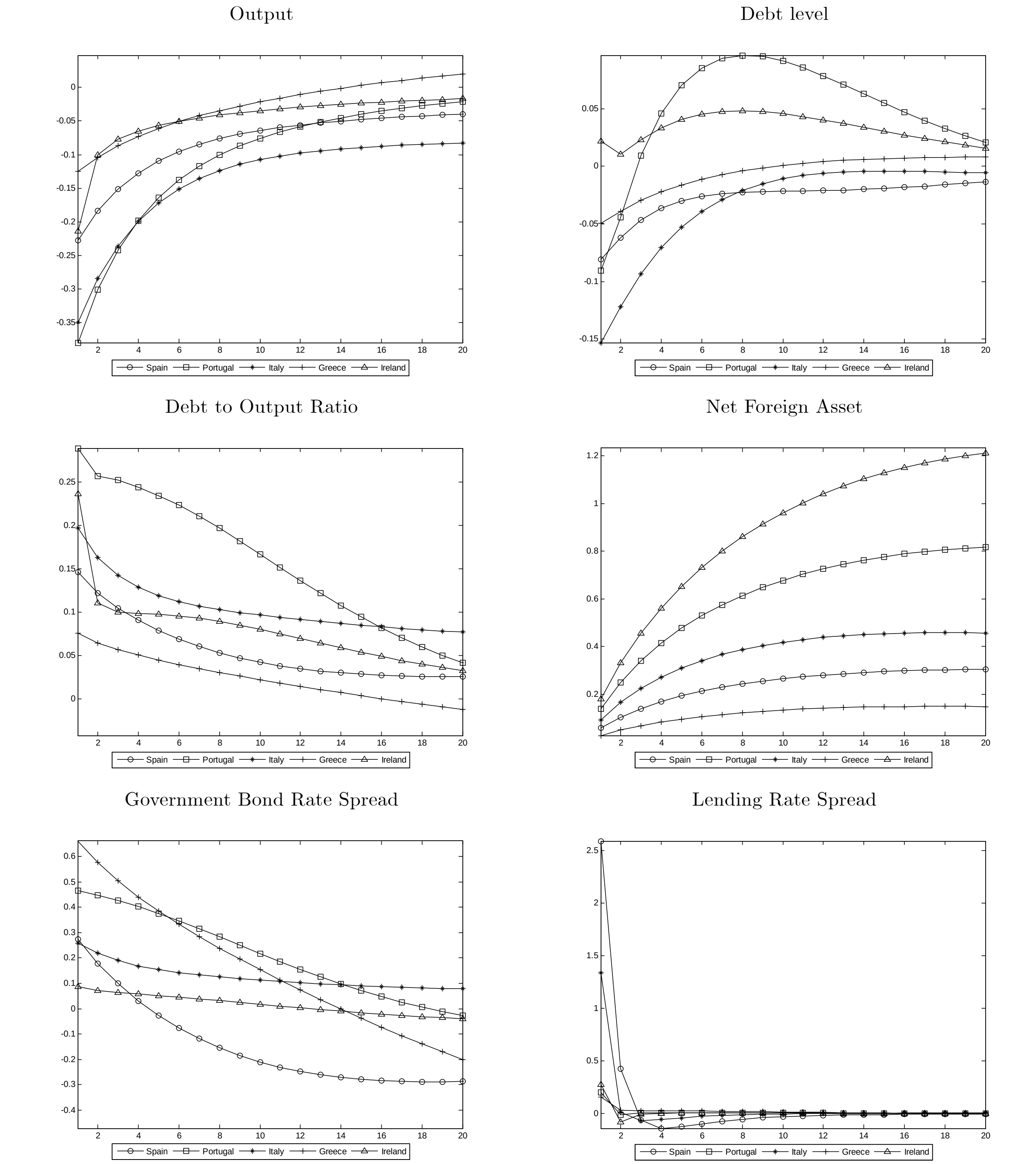
FIGURE 4 - RESPONSE TO A 1% GDP GOVERNMENT CONSUMPTION CONTRACTION



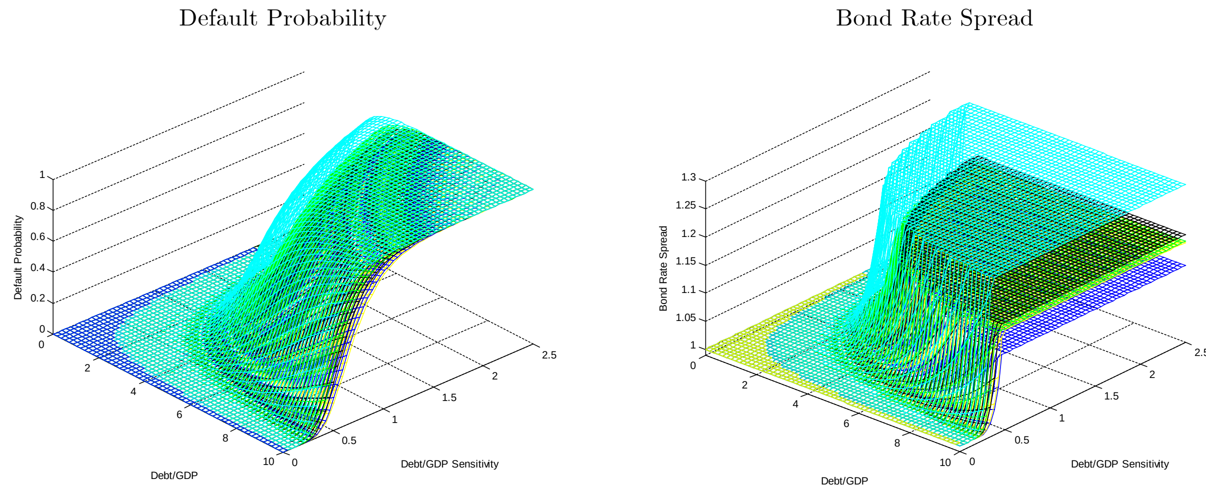
Notes: Impulse response of output (Y_t), bond (B_t), bond to output ratio (B_t/Y_t), net foreign asset (A_t),net foreign asset to output ratio (A_t/Y_t), government interest rate spread ($R_t^g - R_t$) and lending interest rate spread ($R_t^l - R_t$) to a one percent GDP government expenditure contraction in the periphery of the eurozone obtained at the posterior mean estimate. Government and lending interest rate spreads are expressed in basis points.

FIGURE 5 - RESPONSE TO A 1% GDP DIRECT TAX INCREASE



Notes: Impulse response of output (Y_t), bond (B_t), bond to output ratio (B_t/Y_t), net foreign asset (A_t),net foreign asset to output ratio (A_t/Y_t), government interest rate spread ($R_t^g - R_t$) and lending interest rate spread ($R_t^l - R_t$) to a one percent GDP direct taxes increase, such as enterprise, capital and labor income tax increases in the periphery of the eurozone obtained at the posterior mean estimate. Government and lending interest rate spreads are expressed in basis points.

FIGURE 4 - DEBT/GDP RATIO, SENSITIVITY PARAMETER, DEFAULT RISK AND BOND RATE SPREAD



Notes: In the figure the value of the sovreign default probability and interest rate spreads on government bonds are reported.The latter is based on both sovreign default probability and fiscal strain . The black line represent Ireland, the blue Greece, the cylan Spain, green Portugal and yellow Italy. For all the periphery countries fiscal strain binds before default occurs.

TABLE 4 - PEAK FISCAL MULTIPLIERS (quarter) - STANDARD TIMES AND ZLB

Instrument	Multiplier	Greece	Ireland	Italy	Portugal	Spain
Gov. consumption	Stardard Times	1.03 (1)	1.84 (1)	1.05 (1)	1.34 (1)	1.03 (1)
	ZLB	1.75 (2)	2.65 (2)	1.79 (2)	2.39 (2)	1.91 (2)
Gov. transfers	Stardard Times	0.08 (1)	0.12 (1)	0.27 (1)	0.41 (1)	0.17 (1)
	ZLB	0.12 (2)	0.17 (2)	0.44 (2)	0.52 (2)	0.26 (2)
Gov. investment	Stardard Times	0.45 (5)	0.36 (5)	0.35 (5)	0.77 (5)	0.38 (5)
	ZLB	0.33 (6)	0.30 (6)	0.38 (6)	0.52 (6)	0.43 (6)
Direct taxes	Stardard Times	0.12 (1)	0.21 (1)	0.35 (1)	0.38 (1)	0.23 (1)
	ZLB	0.06 (1)	0.18 (1)	0.33 (1)	0.18 (1)	0.19 (1)
Consumption.tax	Stardard Times	0.14 (2)	0.19 (1)	0.27 (1)	0.29 (1)	0.19 (1)
	ZLB	0.12 (2)	0.15 (1)	0.25 (1)	0.14 (1)	0.16 (1)

Notes: The ZLB binds for 8 quarters. The value of the monetary fiscal multiplier is reported.

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