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The Price Relevance of Fiscal Developments

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Abstract

We use SURE estimation methods to assess the link between prices, bond yields and the fiscal behavior. A first equation determines the country-specific cost of government financing via the long-term government bond yield, as a function of budget balance positions. A second equation links the price level to the cost of government financing. Our results for 15 EU countries in the period 1980Q1-2013Q4, show that: improvements in the fiscal stance lead to persistent falls in sovereign yields; higher sovereign yields are reflected in upward price movements; improvements in the fiscal stance in recession times lead to short-term decreases in yields; better fiscal stance in expansions induce downward movement in bond yields only after 8 quarters.

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1. INTRODUCTION

The relevance of fiscal developments for price behaviour and inflation can be traced back to some recent theoretical work linked to the so-called Fiscal Theory of the Price Level (FTPL), initially made popular by Leeper (1991), Sims (1994) and Woodford (1994, 1995). On the other hand, this discussion links further back to Sargent and Wallace (1975), and to the controversy of using rules to control the nominal interest rate, which may lead to price level indeterminacy.[1] In this case, Leeper-Sims-Woodford (hereafter LSW) argue that it will be then up to the government budget constraint to play a key role in the determination of the price level. In other words, fiscal policy may have a relevant role in determining the price level, and then inflation would not be "always and everywhere a monetary phenomenon".

Nevertheless, several authors argued against such theoretical possibility, notably McCallum (1999, 2001), McCallum and Nelson (2005), and Buiter (2002). In addition, most available empirical assessments, provided by, for instance, Canzoneri, and Diba (1996), Canzoneri, Cumby and Diba (2001a,b), Cochrane (1998) and Woodford (1995), and Afonso (2008), point to the lack of adherence to the idea that the price level may be determined via the intertemporal government budget constraint, given that governments turn out to be rather Ricardian. In other words, primary government budget balances react to government debt to ensure fiscal solvency, and money and prices are determined by money supply and demand, implying the existence of an active monetary policy, and a passive fiscal policy. Still, Rother (2004) reports that activist fiscal policy may have relevant effects on inflation volatility.

This paper adds to the literature by applying Seemingly Unrelated Regressions (SURE) estimation methods to a set of two core specifications linking prices, sovereign bond yields and fiscal developments. The first equation determines the country-specific cost of government financing via the long-term government bond yield, as a function of budget balance positions, and other relevant determinants. The second equation links the price level to its determinants, notably the cost of government financing and the business cycle.

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[1] In this context, we can mention a "weak form" of the FTPL, due to Sargent and Wallace (1981), where fiscal policy is exogenous, and impinges on the price level via the money supply (see, Carlstrom, and Fuerst, 2000); and a "strong form" of the FTPL, as in Leeper-Sims-Woodford, whereupon fiscal policy affects the price level independently of the money supply.

Our main results show that: i) improvements in the fiscal stance lead to persistent falls in sovereign yields; ii) higher sovereign yields are reflected in upward price movements; iii) improvements in the fiscal stance in recession times lead to short-term decreases in yields, followed by a correction after 10 quarters; iv) better fiscal stance in expansion times induce downward movement in bond yields only after 8 quarters.

The remainder of the paper is organized as follows. Section 2 briefly presents the theoretical framework. Section 3 presents the data and the econometric methodology. Section 4 discusses the empirical results and the last section concludes.

2. THEORETICAL FRAMEWORK

The idea of non-Ricardian regimes rests on the hypothesis that primary budget balances could be determined by the government without taking into account the level of government debt. In that vision of the world, money and prices would then need to adjust to the level of government debt to guarantee the fulfilment of the government intertemporal budget constraint, a passive monetary policy.

Therefore, in the context of a non-Ricardian regime, the fiscal authority may autonomously decide on the budget balance and government debt, influencing the determination of the price level, while the monetary authority would set endogenously the money supply and take the price level from the government budget constraint. In practice we would see an influence of fiscal developments on the price level, either indirectly via the effects of the sovereign bond long-term interest rate on inflation, or via the fiscal effects on the sovereign bond yield and the yield itself.

According to LSW in a non-Ricardian regime, the government budget constraint determines a unique price level (P):

$$\frac{B_t}{P_t} = \sum_{j=0}^{\infty} \frac{s_{t+j}}{(1+r)^{j+1}} \qquad (1)$$

where, B_t – nominal government liabilities (including debt and money base); s_t – real primary government budget surplus (with seigniorage revenue); r – real interest rate, constant by hypothesis, and with the usual transversality condition (no-Ponzi game condition)

$$\lim_{j \to \infty} \frac{B_{t+j}}{(1+r)^{j+1}} = 0 . \qquad (2)$$

In a non-Ricardian regime, (1) is fulfilled if after the government has chosen a sequence for primary balances, the price level adjusts endogenously. If (1) is fulfilled for any price level, then it will be fiscal policy to adjust implying a Ricardian regime. Therefore, this discussion can have relevant policy implications given notably the empirical possibility that fiscal developments do impinge on the price level and on inflation.

3. DATA AND ECONOMETRIC METHODOLOGY

3.1 Static Approach: estimating a panel data system of equations

We employ Seemingly Unrelated Regressions estimation methods with an iteration procedure over the estimated disturbance covariance matrix and parameter estimates that converge to stable Maximum Likelihood (ML) results (see Zellner, 1962, 1963; and Zellner and Huang, 1962 for further details). The following system with two equations is estimated:

$$
\begin{cases}
ltbond_{it} = \alpha_i^1 + \alpha_t^1 + + \beta_0^1 capb_{it} + \beta_1^1 stockret_{it} + \varepsilon_{it}^1 & (3) \\
p_{it} = \alpha_i^2 + \alpha_t^2 + \beta_0^2 ltbond_{it} + \beta_1^2 gap_{it} + \varepsilon_{it}^2 . & (4)
\end{cases}
$$

The first equation determines the country-specific cost of government financing $ltbond_{it}$, defined as the country's long-term bond yield, as a function of structural budget balance positions, that is, the cyclically adjusted primary balance, $capb_{it}$, and the stock market index, $stockret_{it}$. The second equation defines the price level, p_{it}, as a function of the country-specific cost of government financing, $ltbond_{it}$, and controls for the business cycle by including the output gap, gap_{it}.

On the one hand, we want to check whether a direct effect on inflation of the borrowing costs of the government is present, via equation (4). On the other hand, we also expect that those borrowing costs tend to be higher the higher are the fiscal imbalances, an effect that is specified via equation (3). Therefore, in this SURE framework, it is possible to test for both direct and indirect effects of fiscal developments on the price level.

3.2 Dynamic Approach: computing Impulse Response Functions

In order to estimate the impact of fiscal developments (long term bond yield) on long term bonds yield (prices) over the short and medium run, we follow the method proposed by Jorda (2005), which consists of estimating impulse response functions (IRFs) directly from local projections. For each period k the following equation is estimated on quarterly data:

$$Y_{i,t+k} - Y_{i,t} = \alpha_i^k + Time_t^k + \sum_{j=1}^{l} \gamma_j^k \Delta Y_{i,t-j} + \beta_k X_{i,t} + \varepsilon_{i,t}^k \qquad (5)$$

with $k=1,...,12$ (in quarters) and where Y represents one of our dependent variables as indicated in Equations (3) and (4), long-term bond yields and the price level, respectively; $X_{i,t}$ denotes either the CAPB or long term bond yield, depending on the equation under scrutiny, in country i at time t; α_i^k are country fixed effects; $Time_t^k$ is a time trend; and β_k measures the impact of $X_{i,t}$ for each future period k. Since fixed effects are included in the regression the dynamic impact should be interpreted as compared to a baseline country-specific trend. In the main results, the lag length (l) is set at 2, even if the results are extremely robust to different numbers of lags included in the specification (see robustness checks and sensitivity presented in the next section). Equation (5) is estimated using the panel-corrected standard error (PCSE) estimator (Beck and Katz, 1995).

Impulse response functions are obtained by plotting the estimated β_k for $k=1,...,12$, with confidence bands computed using the standard deviations of the estimated coefficients β_k. While the presence of a lagged dependent variable and country fixed effects may in principle bias the estimation of γ_j^k and β_k in small samples (Nickell, 1981), the length of the time dimension mitigates this concern.[2] The robustness checks for endogeneity confirm the validity of the results.

An alternative way of estimating, for instance, the dynamic impact of fiscal developments is to estimate an ARDL equation and to compute the IRFs from the estimated coefficients (Romer and Romer, 1989; and Cerra and Saxena, 2008). However, the IRFs derived using this approach tend to be sensitive to the choice of the number of lags this making the IRFs potentially unstable. In addition, the significance of long-lasting effects with ARDL models can be simply driven by the use of one-type-of-shock models (Cai and Den Haan, 2009). This is particularly true when the dependent variable is highly persistent, as in our analysis. In contrast, the approach used here does not suffer from these problems because the coefficients associated with the lags of the change in the dependent variable enter only as control variables and are not used to derive the IRFs, and since the

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[2] The finite sample bias is in order of 1/T, where T in our sample is 136.

structure of the equation does not impose permanent effects. Finally, confidence bands associated with the estimated IRFs are easily computed using the standard deviations of the estimated coefficients and Montecarlo simulations are not required.

4. EMPIRICAL ANALYSIS

For the empirical analysis we have considered 15 European Union countries (Austria, Belgium, Estonia, Finland, France, Germany, Greece, Ireland, Italy, Netherlands, Portugal, Slovenia, Slovak Republic, Spain and the UK) throughout the period 1980Q1-2013Q4.[3] We get our data from the Eurostat via Datastream.

4.1. Panel Unit Roots

Prior to presenting and discussing our main empirical results, one concern when working with time-series data is the possibility of spurious correlation between the variables of interest (Granger and Newbold, 1974). This situation arises when series are not stationary.[4] Given the notoriously low power of individual country-by-country tests for unit roots and cointegration, it is preferable to pool the time series of interest and conduct panel analysis. We employ three different types of panel unit root tests: two first generation tests, namely the Im et al. (2003) test (IPS); the Maddala and Wu (1999) test (MW) and one second generation test – the Pesaran (2007) CIPS test. The latter is associated with the fact that previous tests do not account for cross-sectional dependence of the contemporaneous error terms and failure to consider it may cause substantial size distortions in panel unit root tests (Pesaran, 2007). Tables A1 and A2 in the Appendix show the results and reveal that the unit root null hypothesis can be generally rejected (with the exception of public debt, which – when mentioned – will be used in first differences).

4.2. Baseline Results

In Table 1 we report the baseline results for the price and yield specifications. We observe that an improvement in the government's fiscal balance (corrected by the cycle) leads to a fall in long-term bond yields, therefore signaling a credible fiscal strategy and path and less concerns

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[3] Sample selection was dictated by data availability.
[4] The advantage of panel data integration is twofold: firstly, the tests are more powerful than the conventional ones: secondly, cross-section information reduces the probability of a spurious regression (Barnerjee, 1999).

about long-term sustainability. Moreover, higher bond yields are triggered by inflationary pressures and larger output gaps. Highly positive output gaps are traditionally associated with over-heating and significant price rises. Our results are robust to single equation estimation (via fixed effects) and system of equations estimation (SURE).[5] In addition, the short and medium-term impacts of the budget balance on long term bonds are shown in Figure 1 for the baseline regression without controls and for one where the stock market index and the output gap are added as regressors. Each figure shows the estimated impulse response function and the associated one standard error bands (dotted lines), where the horizontal axis measures quarters.

Table 1 – Price Dynamics and Fiscal Policy: System Estimations

| Estimation	FE		FE		SURE		SURE	
Equations	(Eq.1) Long term yield	(Eq.2) CPI	(Eq.1) Long term yield	(Eq.2) CPI	(Eq.1) Long term yield	(Eq.2) CPI	(Eq.1) Long term yield	(Eq.2) CPI
	(1)	(2)	(3)	(4)	(5)		(6)	
Long term Bond		0.155*** (0.0101)		0.249*** (0.013)		0.164*** (0.005)		0.301** (0.122)
CAPB	-0.053*** (0.0188)		-0.035*** (0.011)		-0.021*** (0.0081)		-0.062** (0.029)	
Stock Market			-3.355*** (0.2013)				-2.023** (1.060)	
Output Gap				0.035*** (0.0083)				0.028*** (0.009)
Observations	*1608*	*1526*	*1,082*	*1111*	*1526*	*1526*	*638*	*638*
R-squared	*0.627*	*0.562*	*0.724*	*0.661*	*0.264*	*0.361*	*0.219*	*0.124*

Note: Estimation by panel fixed effects (FE) with robust standard errors and seemingly unrelated regression (SURE). The former includes two equations estimated separately; the latter includes one system of two equations estimated jointly. Standard errors are in parenthesis. Constant term was omitted for reasons of parsimony. Fixed effects regressions include time effects omitted for reasons of parsimony. *, **, *** denote significance at 10, 5 and 1% levels.

In general, an improvement in the fiscal stance leads to a persistent fall in the yield of sovereign bonds. Long-term sovereign bond yields fall by about 2-3 bp in the short term (after 3 quarters) and by nearly 6bp in the medium term (after 12 quarters). This is consistent with results notably from Heppke-Falk and Hüfner (2004), Manganelli and Wolswijk (2009), and by Afonso and Guimarães (2014). On the other hand, higher sovereign yields are also reflected in upward price movements.[6]

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[5] The use of alternative estimators, such as 2SLS or 3SLS (not shown but available upon request), yielded qualitatively similar results.

[6] We have also checked (not shown) that the impact of an increase in the debt-to-GDP ratio leads to a short and medium-term increase in bond yields and that the positive impact of the latter on prices is invariant to the chosen proxy (using the HIPC instead of the CPI results are qualitatively similar).

Figure 1. Baseline Impulse Response Functions
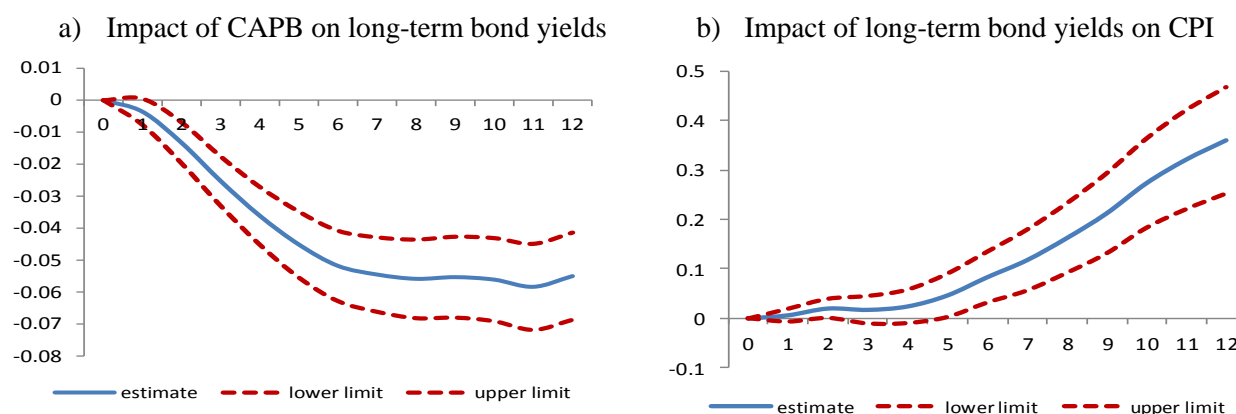
a) Impact of CAPB on long-term bond yields   b) Impact of long-term bond yields on CPI

Note: Dotted lines equal one standard error confidence bands. See main text for more details.

In order to check the robustness of the results, Equation (5) is re-estimated by including time fixed effects to control for specific time shocks, as those affecting world interest rates. The results for this specification remain statistically significant and broadly unchanged (Figure 2 panel (a)).

Moreover, as shown by Tuelings and Zubanov (2010), a possible bias from estimating Equation (5) using country-fixed effects is that the error term of the equation may have a non-zero expected value, due to the interaction of fixed effects and country-specific fiscal developments. This would lead to a bias of the estimates that is a function of k. to address this issue and check the robustness of our findings, Equation (5) was re-estimated by excluding country fixed effects from the analysis. The results reported in Figure 2 panel (b) suggest that this bias is negligible (the difference in the point estimate is small and not statistically significant).

Estimates of the impact of fiscal developments on long term bond yields could be biased because of endogeneity, as unobserved factors influencing the dynamics of public finances may also affect the probability of the occurrence of a consolidation episode. In particular, a significant deterioration in economic activity, which would affect unemployment, may determine an increase in the public debt ratio via the budgetary effect of the automatic stabilizers, and therefore increase the probability of consolidation. To address this issue, Equation (5) was augmented to control for the output gap and stock market developments. The results of this exercise are reported in Figure 2 panel (c) and confirm the robustness of the previous findings.

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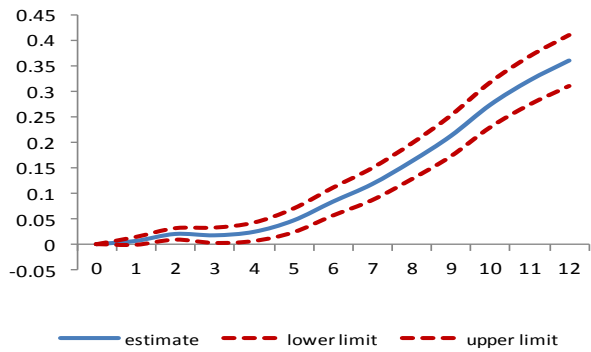**Figure 2. Sensitivity and Robustness of Impulse Response Functions**
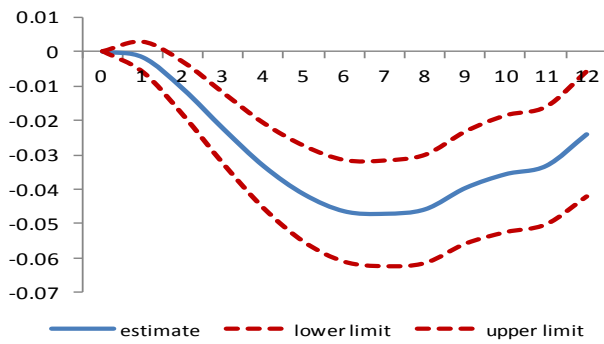
Impact of CAPB on long-term bond yields   Impact of long-term bond yields on CPI
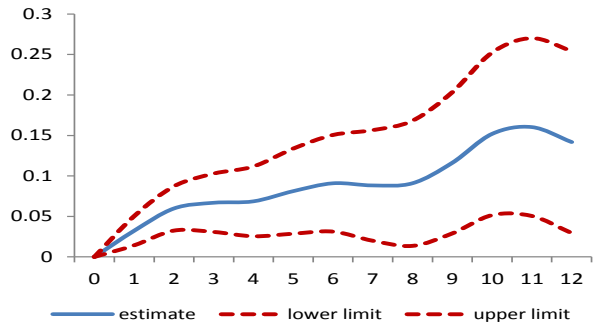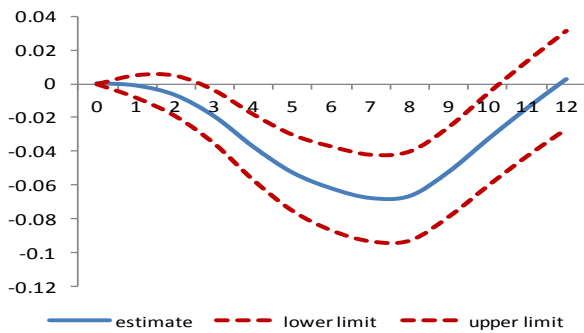
a) Including country and time effects
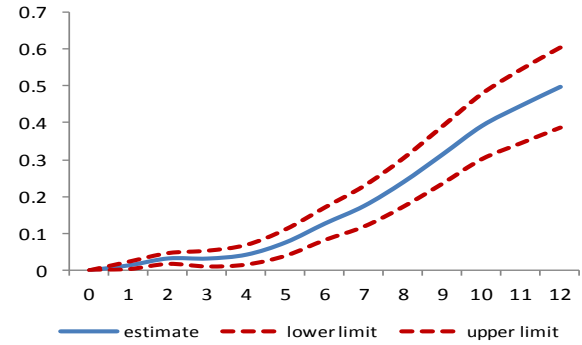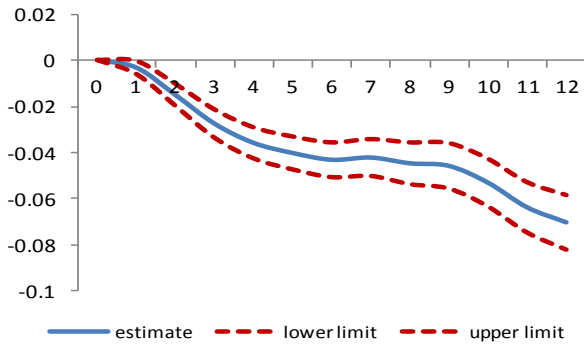
b) No country effects

c) Controlling for stock market and output gap

Note: Dotted lines equal one standard error confidence bands. See main text for more details.

As an additional sensitivity check, Equation (5) was re-estimated for different lags (l) of changes in the Gini coefficient. The results confirm that previous findings are not sensitive to the choice of the number of lags (results are not shown for reasons of parsimony but are available upon request). In addition, in order to deal with endogeneity concerns we re-estimate Equation (5) by means of a GMM estimator (Arellano and Bover, 1995). This estimator is particularly relevant when series are very persistent and the lagged levels may be weak instruments in the first

differences. In this case, lagged values of the first differences can be used as valid instruments in the equation in levels and efficiency is increased by running Equation (5) by means of a system GMM estimator.[7] Results in Figure 3 are qualitatively in line with our previous findings.

Figure 3. Endogeneity-corrected Impulse Response Functions (GMM estimation)
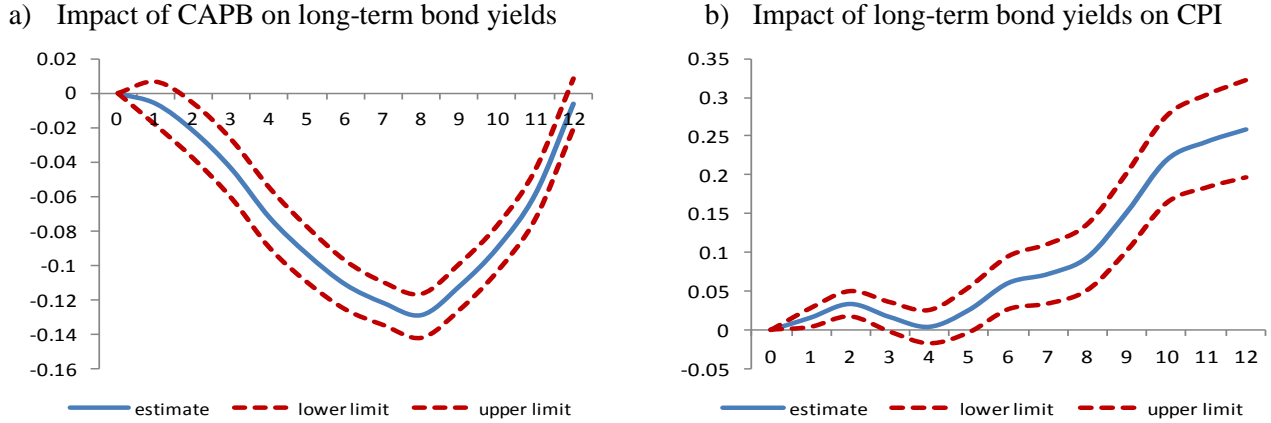
a) Impact of CAPB on long-term bond yields   b) Impact of long-term bond yields on CPI

Note: Dotted lines equal one standard error confidence bands. See main text for more details.

To explore whether long term bond yields vary depending on the phase of the business cycle, the following alternative regression will be estimated:

$$Y_{i,t+k} - Y_{i,t} = \alpha_i^k + Time_t^k + \sum_{j=1}^{l} \gamma_j^k \, \Delta Y_{i,t-j} + \beta_k^{rec} \cdot Y(z) \cdot X_{i,t} + \beta_k^{exp} \cdot (1 - Y(z)) \cdot X_{i,t} + \varepsilon_{i,t}^k \,(6)$$

with $Y(z_{it}) = \frac{\exp(-\gamma z_{it})}{1+\exp(-\gamma z_{it})}, \quad \gamma > 0$

where z is an indicator of the state of the economy (using the output gap computed by means of the HP filter with a smoothing parameter of 1600 applied to real GDP) normalized to have zero mean and unit variance.[8] The remainder of the variables and coefficients are defined as in Equation (5).

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[7] The list of instruments includes the first and second lags of all the right-hand-side variables. The null of Hansen J-test for over-identifying restrictions is not rejected, meaning that the model specification is correct and all over-identified instruments are exogenous. The tests for serial correlation also point to the absence of second-order serial correlation in the residuals.

[8] This approach is equivalent to the smooth transition autoregressive (STAR) model developed by Granger and Teravistra (1993). The main advantage of this approach relative to estimating structural VARs for each regime is that it considers a larger number of observations to compute the impulse response functions, thus making the responses more stable and precise.

Figure 4. State-contingent Impulse Response Functions: Recessions vs. Expansions
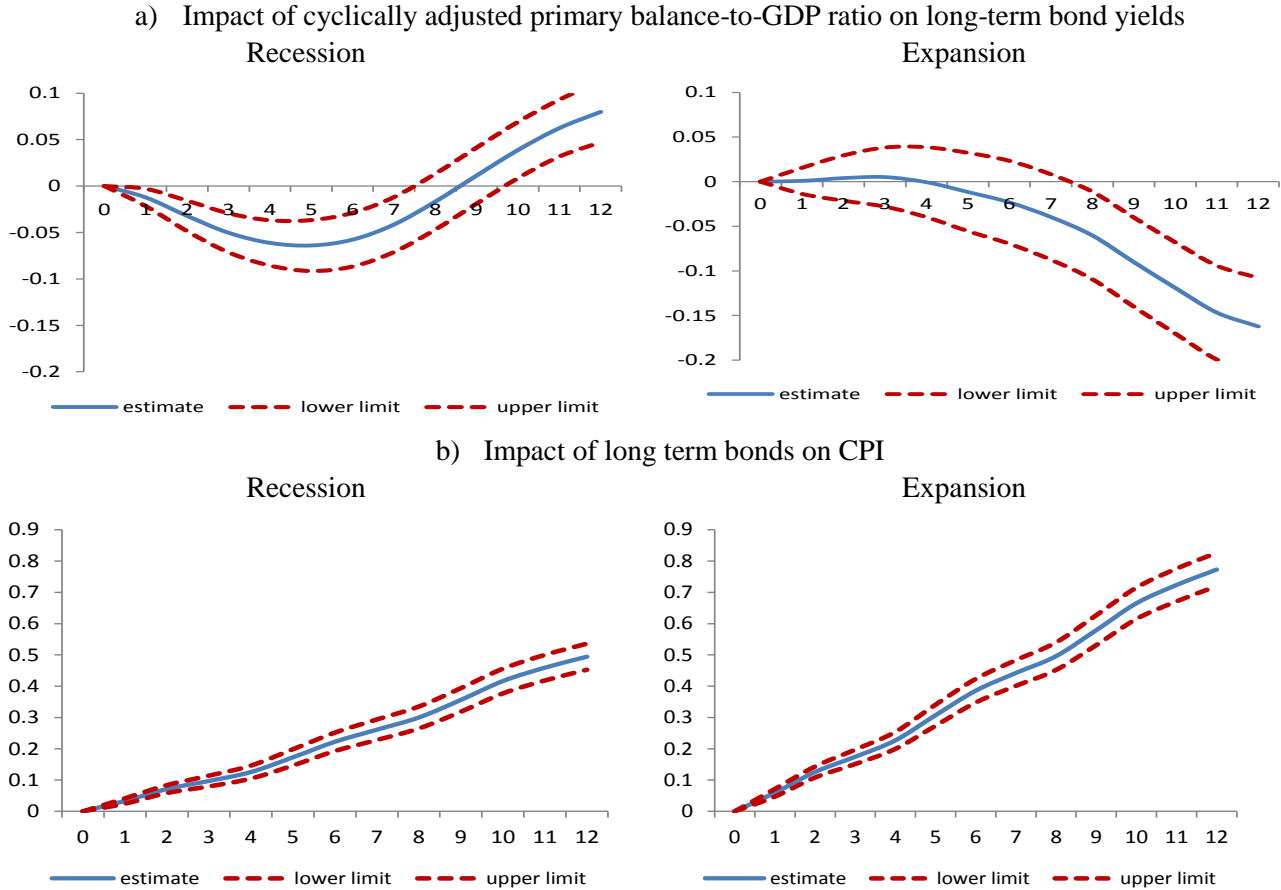
a) Impact of cyclically adjusted primary balance-to-GDP ratio on long-term bond yields

b) Impact of long term bonds on CPI

Note: Dotted lines equal one standard error confidence bands. See main text for more details.

Results presented in Figure 4 panel (a) seem to suggest that improvements in the fiscal stance that took place in times of economic recessions led to a short-term decrease in long-term bond yields, followed by a correction after 9 quarters. On the other hand, in expansions, the overall impact in both the short and medium-term is not statistically different from zero. In panel (b) there seems to exist little difference in the impact of long term bonds on prices between recessions and expansions in the short run, but not in the medium run. During booms the positive impact of long-term bond yields on the price level is higher, relative to times of economic slack.

4.3. Robustness: Structural and Policy Variables

In order to control for additional relevant country features, we now assess whether the effect of fiscal behaviour on long-term bond yields and the effect of these on the price level depend on countries' structural and policy variables: the level of economic development (real GDP per capita), country size (population), indebtedness (debt-to-GDP ratio), and trade openness (exports plus

imports over GDP). To test whether the factors mentioned above affect the response of long-term bond yields to impulses on the CAPB and the response of CPI to impulses on long term bonds yields, Equation (5) is re-estimated using structural/policy variables' 2nd quartile as the threshold value to split the whole sample into two sub-samples that will be compared against the baseline.

Figure 5. The role of Structural Factors (PCSE): economic development
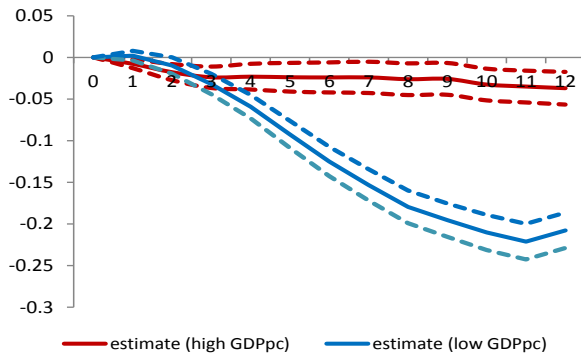
    a) Impact of CAPB on long-term bond yields   b) Impact of long-term bond yields on CPI
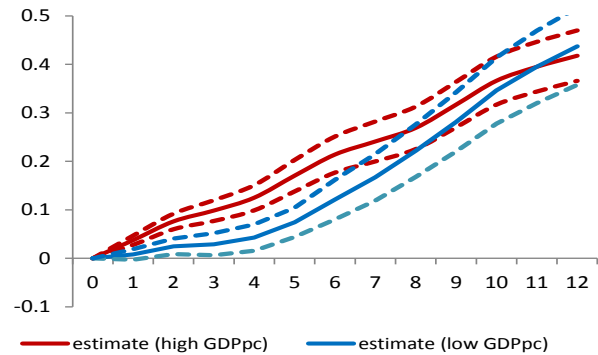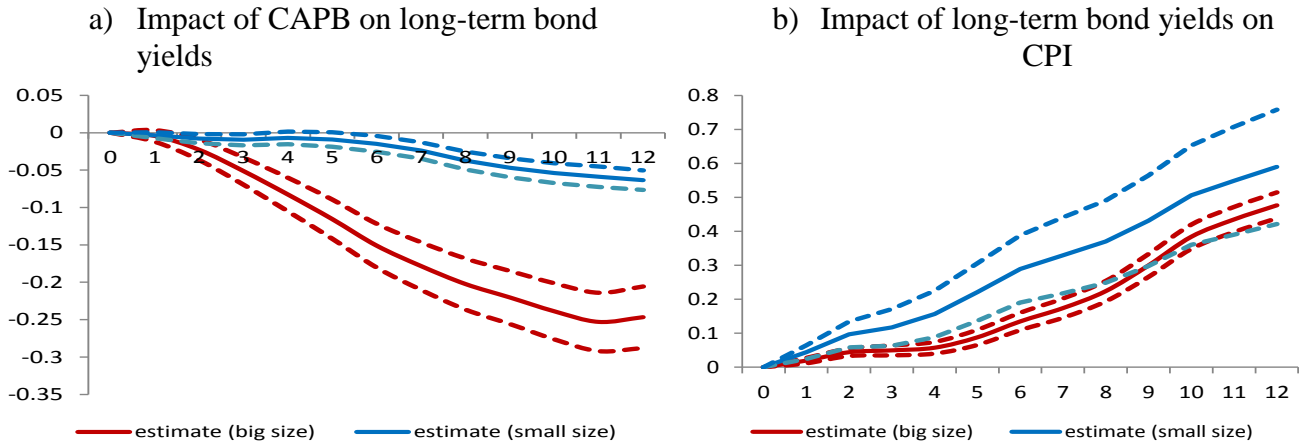

Note: Lines represent the impulse responses of long term bond yields to a CAPB shock (panel a)) or the impulse response of CPI to a long-term bond yield shock (panel b)). Blue (red) line represents the impulse response of those countries below (above) the corresponding threshold. The dotted lines denote the corresponding confidence bands. The threshold point for each structural (or policy) factor considered corresponds to the 2nd quartile (above/below). See main text for more details. Horizontal axis indicates years after the shock.

Starting with Figure 5 one observes that the lower the level of development, the higher the negative response of long-term bond yields to an improvement in the overall fiscal position. This can be linked to the fact that per capita GDP is usually a relevant determinant of sovereign ratings and low income countries might be seen by capital markets as more fiscally vulnerable to changes in the fiscal stance. [9] Moreover, the positive impact of long-term bond yields on the price level is higher in countries with smaller real GDP per capita, at least in the short run. Also, bigger countries experience a more sizeable negative response of long-term bond yields to a shock in the CAPB, relative to smaller countries (Figure 6), which can imply that for smaller economies, long-term yields are rather more exogenously determined. The positive spillover of high bond yields into higher prices is also higher in countries with less population, at least in the short run (the confidence bands cross one another around 7 quarters).
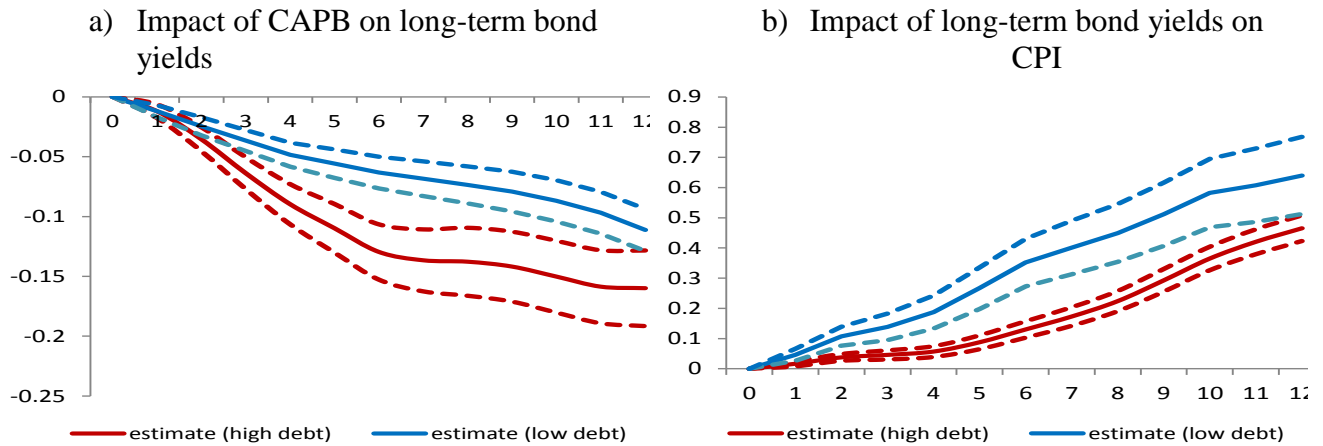
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[9] See, for instance, Afonso et al., (2012).

Figure 6. The role of Structural Factors (PCSE): size

a) Impact of CAPB on long-term bond yields

b) Impact of long-term bond yields on CPI

— estimate (big size)　　— estimate (small size)　　　　— estimate (big size)　　— estimate (small size)

Note: vide footnote figure 5. Mutatis mutandis.

Figure 7. The role of Policy Factors (PCSE): debt level

a) Impact of CAPB on long-term bond yields

b) Impact of long-term bond yields on CPI

— estimate (high debt)　　— estimate (low debt)　　　　— estimate (high debt)　　— estimate (low debt)

Note: vide footnote figure 5. Mutatis mutandis.

Turning to policy factors, countries with higher debt-to-GDP ratios tend to experience a sharper downward response of bond yields to an improvement in the fiscal position, compared to countries with lower debt (Figure 7). Hence, for more indebted economies, capital markets may perceive a higher gain in terms of future correction of fiscal imbalances, allowing the long-term yields to decrease as a premium to a so-called Ricardian behavior from the fiscal authority.

On the contrary, in countries with higher debt levels, an increase in bond yields does not translate into much higher prices, relative to countries with lower public indebtedness positions. Finally, trade openness also seems to play a role. The more open the country, the smaller (larger) response of bond yields (prices) to a shock in CAPB (bond yields) in the medium(short) run-Figure 8.

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Figure 8. The role of Policy Factors (PCSE): trade openness

a) Impact of CAPB on long-term bond yields

b) Impact of long-term bond yields on CPI
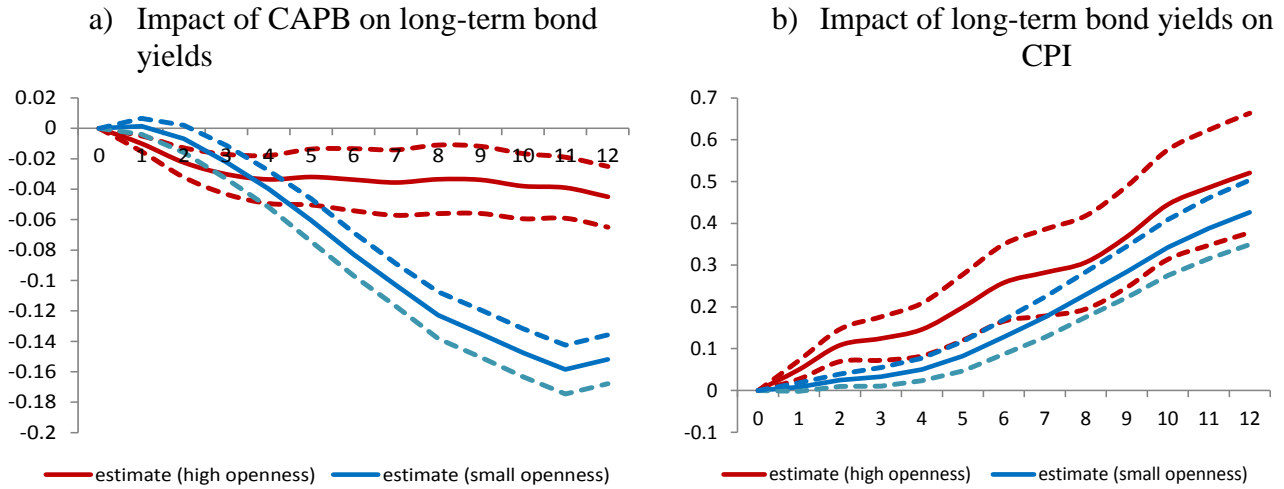
Note: vide footnote figure 5. Mutatis mutandis.

5. CONCLUSION

We have assessed the link between prices, sovereign bond yields and fiscal behavior for a set of 15 EU countries in the period 1980Q1-2013Q4. Our analysis strategy checked whether there is a direct effect on inflation of the borrowing costs of the government, via a first specification, and we then also study the effect of fiscal imbalances on the borrowing costs themselves, via a second equation, therefore, using estimation in a SURE framework.

In order to account for the possibility of non-stationarity in the panel, we have resorted to second generation unit root tests to account for cross-sectional dependence of the contemporaneous error terms. In fact, with the exception of public debt, which was used in first differences, the presence of unit roots was rejected.

Our main results show that: improvements in the fiscal stance lead to persistent falls in sovereign bond yields; higher sovereign yields are reflected in increasing price levels; improvements in the fiscal stance, modelled with the cyclically adjusted primary balance, in recession times lead to short-term decreases in sovereign bond yields; and improvements in the fiscal stance in economic expansions induce downward movements in sovereign bond yields only after 8 quarters.

In terms of robustness, we have also concluded, notably, that the lower the level of development, the higher the negative response of long-term bond yields to an improvement in the fiscal position. Moreover, the positive impact of long-term bond yields on the price level is higher

in countries with smaller real GDP per capita, at least in the short run. Also, bigger countries experience a more sizeable negative response of long-term bond yields to a shock in the cyclically adjusted primary balance, relative to smaller countries.

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Appendix

Table A1: First Generation Panel Unit Root Tests

Im, Pesaran and Shin (2003) Panel Unit Root Test (IPS) (a)

Full	Stock market		CAPB		Output Gap		Long term bond		CPI		Debt	
in levels												
	Lags	[t-bar]	Lags	[t-bar]	Lags	[t-bar]	lags	[t-bar]	Lags	[t-bar]	Lags	[t-bar]
	1.00	-4.825**	1.73	-6.567**	2.44	-4.169**	2.17	0.242	1.33	-2.052**	1.75	5.111

Maddala and Wu (1999) Panel Unit Root Test (MW) (b)

Full	Stock market		CAPB		Output Gap		Long term bond		CPI		Debt	
lags	p_λ	(p)					p_λ	(p)				
in levels												
0	2.878	0.942	81.700**	0.000	1.624	0.990	40.433**	0.010	128.866**	0.000	5.734	1.000
1	3.375	0.909	28.742**	0.000	21.164**	0.007	47.865**	0.001	52.680**	0.001	4.106	1.000
2	5.857	0.663	22.590**	0.004	5.411	0.713	38.197	0.17	40.985**	0.017	2.258	1.000

Notes: (a) We report the average of the country-specific "ideal" lag-augmentation (via AIC). We report the t-bar statistic, constructed as $t-bar = (1/N)\sum_i t_i$ (t_i are country ADF t-statistics). Under the null of all country series containing a nonstationary process this statistic has a non-standard distribution: the critical values are -1.73 for 5%, -1.69 for 10% significance level – distribution is approximately *t*. We indicate the cases where the null is rejected with **. (b) We report the MW statistic constructed as $p_\lambda = -2\sum_i \log(p_i)$ (p_i are country ADF statistic p-values) for different lag-augmentations. Under the null of all country series containing a nonstationary process this statistic is distributed $\chi^2(2N)$. We further report the p-values for each of the MW tests.

Table A2: Second Generation Panel Unit Root Tests

Pesaran (2007) Panel Unit Root Test (CIPS)

Full	Stock market		CAPB		Output Gap		Long term bond		CPI		Debt	
lags	p_λ	(p)					p_λ	(p)				
in levels												
0	-0.275	0.391	-4.990**	0.000	2.313	0.990	-0.152	0.440	-1.295**	0.098	0.394	0.653
1	-0.197	0.422	-3.347**	0.000	-0.536	0.296	-0.063	0.475	0.095	0.538	1.278	0.899
2	0.919	0.821	-2.546**	0.005	0.473	0.682	1.268	0.898	-0.662	0.254	2.523	0.994

Notes: Null hypothesis of non-stationarity. We further report the p-values for each of the CIPS tests.