The Price Relevance of Fiscal Developments

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ABSTRACT
We use Seemingly Unrelated Regressions 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 and 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 behavior 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. 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 Buitert (2002). In addition, most available empirical assessments, provided by, for instance, Canzoneri and Diba (1996), Canzoneri, Cumby, and Diba (2001a, 2001b), Cochrane (1998), Woodford (1995) and

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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 & Fuerst, 2000); and a ‘strong form’ of the FTPL, as in LSW, whereupon fiscal policy affects the price level independently of the money supply.

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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 Estimation (SURE) 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.

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; and (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 fulfillment 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 \( P \):

\[
\frac{B_t}{P_t} = \sum_{j=0}^{\infty} \frac{s_{t+j}}{(1 + r)^{j+T}},
\]

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

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

In a non-Ricardian regime, Equation (1) is fulfilled if after the government has chosen a sequence for primary balances, the price level adjusts endogenously. If Equation (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 SURE methods with an iteration procedure over the estimated disturbance covariance matrix and parameter estimates that converge to stable maximum likelihood results (see Zellner, 1962, 1963; Zellner & Huang, 1962 for further details). The following system with two equations is estimated:

$$ltbond_{it} = \alpha_1^1 + \alpha_1^t + + \beta_1^1 capb_{it} + \beta_1^1 stockret_{it} + \epsilon_1^1,$$  (3)

$$pit = \alpha_2^2 + \alpha_2^t + \beta_2^1 ltbond_{it} + \beta_2^1 gap_{it} + \epsilon_2^2.$$  (4)

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, $pit$, as a function of the country-specific cost of government financing, $ltbond_{it}$, and controls for the business cycle by including the output gap, $gap_{it}$. $\alpha_1^1$, $\alpha_1^t$ and $\alpha_2^1$ denote time and country-fixed effects, respectively, in Equations (3) and (4). Time effects are included to capture unobserved heterogeneity across countries and time-unvarying factors such as geographical variables; country-fixed effects are added to control for global common shocks. $\epsilon_1^1$ and $\epsilon_2^2$ are iid error term disturbances in Equations (3) and (4), respectively. Since serial correlation can arise in this environment due to the presence of individual effects (Baltagi, 1980), Equations (3) and (4) are estimated with HAC robust standard errors clustered at the country level to allow us to make valid inferences.

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). Indeed, in the face of several theoretical underpinnings, notably the FTPL (see notably Leeper, 1991; Woodford, 1994) we could expect, in principle, that the price level would not be simply a pure monetary phenomenon and thus be affected also by fiscal conditions. On the other hand, we also expect that those borrowing costs tend to be higher, the higher are the fiscal imbalances, an effect

$^2$ In the classical linear SURE model, one usually assumes that the errors are iid over time with mean zero and homoscedastic variance $\Sigma = E(\epsilon_{it} \epsilon_{it}') (X)$ (with $X$ being the vector of regressors). Furthermore, $\Sigma$ is assumed to be positive definite. As in standard univariate models, nonspherical disturbances can be accommodated by either modelling the residuals or computing robust covariance matrices.

$^3$ We thank an anonymous referee for raising this point.
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, and assess, also in this vein, any evidence supporting the FTPL.

### 3.2. Dynamic Approach: Computing Impulse Response Functions

To estimate the dynamic distributional impact of fiscal developments (long-term bond yield) on long-term bonds yield (prices) over the short and medium term, we follow Jorda’s (2005) method. This method consists of estimating impulse response functions (IRFs) directly from local projections. For each period \( k \), we estimate the following regression:

\[
Y_{i,t+k} - Y_{i,t} = \alpha_i^k + Time_i^k + \sum_{j=1}^{l} \gamma_{i}^k \Delta Y_{i,t-j} + \beta_k S_{i,t} + X'_{i,t-1} \sigma_{i,t}^k + \varepsilon_{i,t}^k, \tag{5}
\]

with \( k = 1, \ldots, 12 \) (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; \( S_{i,t} \) denotes our shock variable, that is, either the CAPB or long-term bond yield, depending on the equation under scrutiny, in country \( i \) at time \( t \); \( X'_{i,t-1} \) is a vector of control variables to be added later (including the output gap and stock market developments); \( \alpha_i^k \) are country-fixed effects added to capture unobserved heterogeneity across countries and time-unvarying factors; \( Time_i^k \) is a time trend; \( \gamma_{i}^k \) and \( \sigma_{i,t}^k \) are coefficients to be estimated for the lagged dependent variable and set of controls, respectively; \( \varepsilon_{i,t}^k \) is a disturbance term satisfying usual assumptions and \( \beta_k \) measures the impact of \( S_{i,t} \) for each future period \( k \). The lag length \( (l) \) is set at 2 as selected by the Akaike information criteria, but our findings are strongly robust to different lag-structures.\(^4\) Equation (2) is estimated using Beck and Katz’s (1995) panel-corrected standard error (PCSE) estimator. IRFs are obtained by collecting the estimated \( \beta_k \) with confidence intervals computed using \( \beta_k \)'s standard errors.\(^5\)

We are aware of alternative ways of estimating dynamic impacts but, as we explain, those are inferior options. The first possible alternative would be to estimate a panel vector autoregression. However, this is generally considered a ‘back-box’ since all relevant regressors are considered endogenous. Moreover, one has to know the exact order in which they enter in the system. Since economic theory rarely provides such an ordering, the Choleski decomposition is often used as a solution of limited value for providing structural information to a vector autoregression (VAR). Moreover, a major limitation of the VAR approach is that it has to be estimated to low-order systems. Since all effects of omitted variables are in the residuals, this may lead to big distortions in the IRFs, making them of little use for structural interpretations (see e.g. Hendry, 1995). In addition, all measurement errors or misspecifications also induce unexplained information left in the error terms, making interpretations of the IRFs even more difficult (Ericsson, Hendry, & Mizon, 1997). One should bear in mind that due to its limited number of variables and the aggregate nature of the shocks, a VAR model should be viewed as an approximation to a larger structural

\(^4\) Results are not shown for reasons of parsimony but are available upon request.

\(^5\) The presence of a lagged dependent variable and country-fixed effects could bias the estimation of \( \gamma_{i}^k \) and \( \beta_k \) in small samples (Nickell, 1981). However, in our case, this is not a problem since the finite sample bias is around 0.007 (that is, \( 1/T \)), where \( T \) is 136.)
system. In contrast, the approach used here does not suffer from these identification and size-limitation problems and, in fact, has been suggested by Auerbach and Gorodnichenko (2013), *inter alia*, as a sufficiently flexible alternative.

A second alternative of assessing the dynamic impact of fiscal consolidation episodes would be to estimate an autoregressive distributed lag model of changes in inequality and consolidation episodes and to compute the IRFs from the estimated coefficients (Cerra & Saxena, 2008; Romer & Romer, 1989). Note that the IRFs obtained using this method, however, tend to be lag-sensitive, therefore undermining the overall stability of the IRFs. Moreover, the statistical significance of long-lasting effects can result from one-type-of-shock models, particularly when the dependent variable is very persistent, as the Gini (Cai & Den Haan, 2009). Contrarily, in the local projection method we do not experience such an issue since lagged dependent variables enter as control variables and are not used to derive the IRFs. Lastly, estimated IRFs’ confidence intervals are computed directly using the standard errors of the estimated coefficients without the need for Monte Carlo simulations.

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.\(^6\) 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 & Newbold, 1974). This situation arises when series are not stationary.\(^7\) 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, Pesaran, and Shin (2003) (IPS) test and the Maddala and Wu (1999) (MW) test, 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 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

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\(^6\) Sample selection was dictated by data availability.

\(^7\) 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 (Banerjee, 1999).
Table 1. Price dynamics and fiscal policy: system estimations.

<table>
<thead>
<tr>
<th>Equations</th>
<th>FE (Equation (1))</th>
<th>FE (Equation (2))</th>
<th>SURE (Equation (1))</th>
<th>SURE (Equation (2))</th>
</tr>
</thead>
<tbody>
<tr>
<td>Estimation</td>
<td>(1)</td>
<td>(2)</td>
<td>(3)</td>
<td>(4)</td>
</tr>
<tr>
<td>Long-term yield</td>
<td>0.155***</td>
<td>0.249***</td>
<td>0.249***</td>
<td>0.164***</td>
</tr>
<tr>
<td>CPI</td>
<td>(0.0101)</td>
<td>(0.013)</td>
<td>(0.013)</td>
<td>(0.005)</td>
</tr>
<tr>
<td>Long-term bond</td>
<td>−0.053***</td>
<td>−0.035***</td>
<td>−0.021***</td>
<td>−0.062**</td>
</tr>
<tr>
<td>CAPB</td>
<td>(0.0188)</td>
<td>(0.011)</td>
<td>(0.0081)</td>
<td>(0.0299)</td>
</tr>
<tr>
<td>Stock Market</td>
<td>−3.355***</td>
<td>0.035**</td>
<td>−2.023**</td>
<td>0.028***</td>
</tr>
<tr>
<td>(2013)</td>
<td>(0.2013)</td>
<td>(0.0083)</td>
<td>(1.060)</td>
<td>(0.009)</td>
</tr>
<tr>
<td>Output Gap</td>
<td>1608</td>
<td>1526</td>
<td>1526</td>
<td>638</td>
</tr>
<tr>
<td>Observations</td>
<td>0.627</td>
<td>0.562</td>
<td>0.661</td>
<td>0.219</td>
</tr>
<tr>
<td>R^2</td>
<td>0.627</td>
<td>0.562</td>
<td>0.661</td>
<td>0.219</td>
</tr>
</tbody>
</table>

Notes: Estimation by panel FE with robust standard errors and 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. FE regressions include time effects omitted for reasons of parsimony. CPI, consumer price index.

*Significance at 10% levels.
**Significance at 5% levels.
***Significance at 1% levels.
less concerns 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 (FE)) and system of equations estimation (SURE). 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 IRF and the associated one standard error bands (dotted lines), where the horizontal axis measures quarters.

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 6 bp 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 (2015). On the other hand, higher sovereign yields are also reflected in upward price movements. The positive price effect of higher nominal long-term yields can be seen as an indication of both an increase in the long-term section of the yield curve, alongside future higher economic growth expectations.

The previous set of results are then subjected to several robustness checks. To begin with, Equation (2) is re-estimated by excluding the time trend and including time FE to control for specific time shocks (such as those affecting world interest rates). The results for this specification remain statistically insignificant and broadly unchanged (Figure 2(a)).

Secondly, there is a potential bias from estimating Equation (5) using country-fixed effects as the error term could have a non-zero expected value as a result of the interaction of FE and country-specific fiscal developments (Teulings & Zubanov, 2010). This would then result in a bias in the coefficient estimates, which is a function of \( k \). With this in mind, Equation (5) was re-estimated by excluding country-fixed effects from the analysis. Results reported in Figure 2(b) suggest that this bias is negligible.

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

9 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 Harmonized Index of Consumer Prices instead of the CPI results are qualitatively similar).
Thirdly, 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).

Fourthly, 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(c) and confirm the robustness of the previous findings.

**Figure 2.** Sensitivity and robustness of IRFs.

Note: Dotted lines equal one standard error confidence bands. See main text for more details.
In addition, in order to address further potential endogeneity concerns, we re-estimate Equation (5) by means of a GMM estimator (Arellano & 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.10 Results in Figure 3 are qualitatively in line with our previous findings.

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

\[
Y_{i,t+k} - Y_{i,t} = \alpha_i k + Time_t^k + \sum_{j=1}^{l} \gamma_j^k \Delta Y_{i,t-j} + \beta_{rec}^{k} \cdot Y(z) \cdot S_{i,t} + \beta_{exp}^{k} \cdot (1 - Y(z)) \cdot S_{i,t} \\
+ X'_{i,t-1} \sigma_{i,t}^k + \epsilon_{i,t},
\]

with

\[
Y(z_{it}) = \frac{\exp(-\gamma z_{it})}{1 + \exp(-\gamma z_{it})}, \gamma > 0
\]

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. The remainder of the variables and parameters are as in Equation (5). This method is equivalent to Granger and Terasvirta’s (1993) smooth transition autoregressive model, whose advantage relative to estimating VARs for each regime is that it uses a larger number of observations to estimate the IRFs, thus increasing stability and precision.

Results presented in Figure 4(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 nine quarters. On the other hand, in expansions, the overall impact in both the short and medium term is not statistically different from zero.

In addition, 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 behavior 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’ second quartile as the threshold value to split the whole sample into two sub-samples that will be compared against the baseline.

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

**Figure 4.** State-contingent IRFs: recessions versus expansions. (a) Impact of CAPB-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.
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.

Notes: 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 second quartile (above/below). See main text for more details. The horizontal axis indicates years after the shock.

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.

Note: Vide notes in Figure 5. Mutatis mutandis.

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 (see, for instance, Afonso, Furceri, & Gomes, 2012). 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 seven quarters).

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).

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 nonstationarity 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, modeled 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 eight 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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References


**Appendix**

**Table A1.** Price dynamics and fiscal policy: system estimations.

(a) Panel unit root test (IPS) (Im et al. 2003)

<table>
<thead>
<tr>
<th>Full</th>
<th>Stock market</th>
<th>CAPB</th>
<th>Output Gap</th>
<th>Long-term bond</th>
<th>CPI</th>
<th>Debt</th>
</tr>
</thead>
<tbody>
<tr>
<td>In levels</td>
<td>Lags</td>
<td>[t-bar]</td>
<td>Lags</td>
<td>[t-bar]</td>
<td>Lags</td>
<td>[t-bar]</td>
</tr>
<tr>
<td>1.00</td>
<td>-4.825**</td>
<td>1.73</td>
<td>-6.567**</td>
<td>2.44</td>
<td>-4.169**</td>
<td>2.17</td>
</tr>
</tbody>
</table>

(b) Panel unit root test (MW) (Maddala and Wu 1999)

<table>
<thead>
<tr>
<th>Full</th>
<th>Stock market</th>
<th>CAPB</th>
<th>Output Gap</th>
<th>Long-term bond</th>
<th>CPI</th>
<th>Debt</th>
</tr>
</thead>
<tbody>
<tr>
<td>In levels</td>
<td>Lags</td>
<td>$p_i$</td>
<td>(p)</td>
<td>$p_i$</td>
<td>(p)</td>
<td></td>
</tr>
<tr>
<td>0</td>
<td>2.878</td>
<td>0.942</td>
<td>81.700**</td>
<td>0.000</td>
<td>1.624</td>
<td>0.990</td>
</tr>
<tr>
<td>1</td>
<td>3.375</td>
<td>0.909</td>
<td>28.742**</td>
<td>0.000</td>
<td>21.164**</td>
<td>0.007</td>
</tr>
<tr>
<td>2</td>
<td>5.857</td>
<td>0.663</td>
<td>22.590**</td>
<td>0.004</td>
<td>5.411</td>
<td>0.713</td>
</tr>
</tbody>
</table>

Notes: (a) We report the average of the country-specific ‘ideal’ lag-augmentation (via akaike information criteria). We report the $t$-bar statistic, constructed as $t = \frac{1}{N} \sum t_i$ ($t_i$ are country augmented Dickey Fuller [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 $p_i = -2 \sum \log(p_i)$ ($p_i$ 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. Panel unit root test (CIPS) (Pesaran 2007)

<table>
<thead>
<tr>
<th>Full</th>
<th>Stock market</th>
<th>CAPB</th>
<th>Output Gap</th>
<th>Long-term bond</th>
<th>CPI</th>
<th>Debt</th>
</tr>
</thead>
<tbody>
<tr>
<td>In levels</td>
<td>Lags</td>
<td>$p_i$</td>
<td>(p)</td>
<td>$p_i$</td>
<td>(p)</td>
<td></td>
</tr>
<tr>
<td>0</td>
<td>-0.275</td>
<td>0.391</td>
<td>-4.990**</td>
<td>0.000</td>
<td>2.313</td>
<td>0.990</td>
</tr>
<tr>
<td>1</td>
<td>-0.197</td>
<td>0.422</td>
<td>-3.347**</td>
<td>0.000</td>
<td>-0.536</td>
<td>0.296</td>
</tr>
<tr>
<td>2</td>
<td>0.919</td>
<td>0.821</td>
<td>-2.546**</td>
<td>0.005</td>
<td>0.473</td>
<td>0.682</td>
</tr>
</tbody>
</table>

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