Fiscal Policy and Interest Rates: The Role of Sovereign Default Risk

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Abstract

Recent events have highlighted the potential importance of nonlinear effects of fiscal variables (notably debt and deficits) on interest rates: While in times when government solvency is not a concern the standard crowding-out effects are of moderate magnitude, in times when default risk becomes an issue the interest rate effects can become very large. This paper provides new evidence on the magnitude of these effects. For the case when default risk is not a concern, it uses an arbitrage-free term structure model to estimate the dynamic effects of fiscal policy shocks on interest rates along the entire maturity spectrum. For the case when default risk becomes a concern (thereby violating a central assumption of the term structure model), I present evidence based on EMU government bond spread regressions on time-varying effects of national fiscal policies on spreads as well as the time-varying sensitivity of yield spreads to international risk aversion as a function of the state of fiscal policy.

JEL classification: E6, H6.

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1 Introduction

Much ink has been spilled on the topic of the relationship between fiscal policy, especially between government debt, deficits or government spending, and interest rates. A large body of empirical studies documents that an increase in government deficits or debt has either statistically insignificant effects on interest rates, or raises them by a statistically significant, but economically modest amount. As so often, most of this evidence is based on postwar U.S. data.

As recent events in the euro area, however, have made abundantly clear, there are situations in which interest rates react very sensitively to fiscal policy changes. The purpose of this paper is to study the recent empirical evidence. The main theme that emerges is that there are times and circumstances in which the effects of fiscal policy on interest rates can be very large indeed. These are presumably times when sovereign default risk becomes an issue. When no risk of sovereign default is perceived, the effects of government debt or deficits on interest rates seem to be significant, but modest. The paper develops models for either of these two situations and then asks what circumstances, including fiscal ones, trigger concerns about default risk in financial markets.

The paper brings together ideas and techniques from several different strands of literature. The first part of the paper focusses on evidence from the U.S., mostly prior to the onset of the financial crisis. It draws on the structural VAR literature to measure the effects (on interest rates, among other variables) of fiscal policies, as well as on the literature on affine term structure models to separate out the effects on expectations of future interest rates from risk premia. This part focusses on the U.S. because it uses techniques that are appropriate only in the absence of default risk and because it requires a rich set of zero-coupon government bond yields that are not available for a sufficiently long period and of sufficient quality for all euro area countries that are studied later.

A key issue regarding fiscal policies by member countries of the European Monetary Union (EMU) is the extent to which those policies affect the area-wide level of interest rates and to what extent they affect yield spreads among EMU government bonds (defined vis-à-vis German government bonds). Based on data prior to 2008, the literature arrived at somewhat conflicting results. Whereas Faini (2006) finds in panel regressions for EMU members for the sample 1979-2002 that there are large spillover effects from individual member countries' fiscal policies to the area-wide level of interest rates, Manganelli and

Wolswijk (2009), studying the sample 1999 to 2008, conclude that government credit risk as measured by bond ratings is still being priced.

In hindsight, it is difficult to avoid the impression that default risk was underpriced prior to 2008. For example, the average spread of 10-year Greek government bonds over German ones was 25 basis points; for Italian bonds it was about the same magnitude, for Portugese 15 basis points. Regardless of how one views the period of compressed EMU government bond spreads that lasted until early March of 2008, it is important to ask whether we can identify a threshold or estimate the nonlinear process by which the interest rate effects of fiscal policy become amplified. In section 3 I develop one such model, but clearly on this issue much work remains to be done.

2 Fiscal policy and interest rates in the U.S. prior and during the crisis

This section provides evidence on the interest rate effects of fiscal policy in a situation where it is plausible to assume that default risk was perceived by investors to be negligible. Specifically we focus on the United States over the period since the 1980s. There is of course already an enormous empirical literature on the interest rate effects of fiscal policy in the United States (Gale and Orszag 2002 provide an excellent survey).

Most of this literature is based on reduced-form regressions of interest rates on fiscal variables and possibly additional regressors. In this section I instead present estimates of an arbitrage-free term structure model in which (most of) the factors are observable macroeconomic variables, including fiscal variables. The law of motion of these factors is modelled as a VAR. Bonds at various maturities are priced under the restriction that their market prices of risk are linear functions of the factors. There are several advantages of this framework:

- There is strong evidence against the expectations hypothesis of the term structure. Therefore it is of interest to decompose yields into risk-neutral (expectations) components and risk premia, and to ask how fiscal policy affects interest rates, whether mostly through changes in expected future short-term interest rates or through changes in risk prices.
- We can study how yields at all maturities (not only at one maturity as in reduced-form

regressions) respond to fiscal policies while simultaneously imposing the assumption of no-arbitrage, which seems plausible in as deep and liquid a market as that for U.S. Treasury securities.

- By using a VAR as law of motion, the results can also be related to the literature of identified fiscal policy shocks starting with Blanchard and Perotti (2002). Thus we can compute impulse responses to tax and spending shocks of yields for the full range of maturities.

Dai and Philippon (2006) have estimated an affine term structure model along the lines just described. A central point of the papers by Gale and Orszag (2002) and Laubach (2009) is that the response of interest rates (especially of long maturities) to fiscal policy depends on expectations about the *future course* of fiscal policy. In the term structure model(s) these expectations are generated by the estimated VAR. However, the VAR-based expectations that the econometrician estimates in hindsight may be quite different from the expectations of investors at the time when the government bonds were priced.[1] I therefore use CBO projections for fiscal variables as additional observable variables in estimation, as in Canzoneri et al. (2002) and Laubach (2009), using the method developed in Kim and Orphanides (2005). Because the CBO projections refer to the federal government, like Favero and Giavazzi (2007) and unlike Blanchard and Perotti (2002) I focus on measures of fiscal policy for the federal sector only.[2]

Favero and Giavazzi (2007) and Perotti (2007) find evidence for a change in responses of macro variables to fiscal policy between samples ending before and starting after about 1980. This change has been interpreted as evidence for a change in the reaction of fiscal policy with regard to stabilization of the debt/GDP ratio. Moreover, the CBO projections are available only from 1976. The analysis therefore starts in 1980. I consider estimates

[1] In a different context, Piazzesi and Schneider (2009) also make the point that investor expectations may have been quite different from those imputed by the econometrician.

[2] Favero and Giavazzi (2007) point out that VARs in the tradition of Blanchard and Perotti (2002) may produce inaccurate parameter estimates and hence inaccurate impulse response functions because of not including government debt in the state space and thereby ignoring any systematic response of fiscal policy to stabilize the debt/GDP ratio. Because of its inherent nonlinearity, including their debt accumulation equation in the state transition equation is not feasible. In future work I hope to explore the effects of including an approximate linear government budget constraint in the model, using again CBO projections for the debt/GDP ratio as additional information in estimation.

based on samples ending before the onset of the crisis and those including the crisis period. Despite the extreme deterioration in the U.S. fiscal outlook, long-term yields have remained low presumably because of safe-haven demand for Treasury securities. To estimate the effect of the increase in government debt on interest rates, it is therefore important to quantify the extent to which (presumably temporary) safe-haven demand has held down yields.

2.1 An affine term structure model with fiscal factors

The model used to estimate the dynamic effects of fiscal policy shocks consists of a reduced-form description of the relationships between major macroeconomic and fiscal variables, and a specification of the stochastic discount factor that ensures that pricing of bonds at various maturities is arbitrage-free. Two key assumptions underlying this pricing framework are the bonds don't pay coupons, and that they are free of default risk.

Specifically, let

$$x_t \equiv [\tilde{r}_t \; \tilde{f}_t \; \tilde{\pi}_t \; q_t \; \bar{\pi}_t]'$$

denote the state vector comprising the detrended short rate $\tilde{r}_t$ (where the meaning of "detrended" will be explained shortly), a demeaned fiscal policy measure $\tilde{f}_t$, detrended inflation $\tilde{\pi}_t$, a measure q_t of real activity relative to potential, and trend inflation $\bar{\pi}_t$. The first four of these variables are assumed to follow a VAR(2), whereas trend inflation is assumed to follow an exogenous random walk:

$$X_t = \mu + \Phi X_{t-1} + U_t, \;\; t = 1, \ldots, T$$

where

$$X_t \;\; \equiv \;\; \begin{bmatrix} x_t \\ x_{t-1} \end{bmatrix}, \; U_t \equiv \begin{bmatrix} u_t \\ 0 \end{bmatrix}, \; \Phi = \begin{bmatrix} \phi_1 & 0 & \phi_2 & 0 \\ 0 & 1 & 0 & 0 \\ I & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \end{bmatrix} \tag{1}$$

Trend inflation $\bar{\pi}_t$ is a latent factor, which, however, will be tightly constrained in the estimation by a survey-based measure of long-horizon inflation expectations. The remaining four factors are observable macro variables.

In specifying the stochastic discount factor that prices bonds at different maturities, I am following the large literature of affine term structure models (e.g. Duffee, 2002). Let

$$\xi_{t+1} = \exp\left(-\frac{1}{2}\lambda_t'\omega^{-1}\lambda_t - \lambda_t'\sigma\epsilon_{t+1}\right)$$

denote the Radon-Nikodym derivative that converts the data-generating to the risk-neutral probability measure, such that

$$\begin{aligned} P_{nt} &= E_t[M_{t+1}P_{n-1,t+1}] = E_t[e^{-r_t}\xi_{t+1}P_{n-1,t+1}] \\ &= E_t^{rn}[e^{-r_t}P_{n-1,t+1}] \end{aligned}$$

where E_t^{rn} denotes expectation under the risk-neutral measure and λ_t are the prices associated with the macroeconomic risks. These risks are given by the (as yet to be identified) fundamental innovations ϵ_t, which are assumed to be i.i.d. standard normal. The reduced-form innovations u_t have covariance matrix ω.

A key assumption is the specification of the prices of risk λ_t as a general linear function of the states:

$$\lambda_t = \lambda_0 + \lambda_1 X_t \tag{2}$$

This specification plays an important role in enabling the model to explain the observed failures of the expectations hypothesis (Dai and Singleton, 2002) and to forecast yields (Duffee, 2002). I follow Dai and Philippon (2006) in assuming that the stochastic discount factor depends only on current innovations ϵ_t, but that the prices of risk can depend on both current and lagged states. Therefore λ_t is a 5×1 vector and λ_1 a 5×10 matrix. Certain other assumptions are being imposed to conserve on the otherwise very large number of parameters to be estimated; details are provided in the appendix. It should be noted that the flexibility provided by the general specification (2) comes at the cost that it is unclear how to generate this specification from preferences of a representative investor.

Besides the VAR specification (1) for the states, the stochastic discount factor and the prices of risk (2), the model is completed by a specification for the one-period nominal risk-free interest rate

$$y_t^1 = \delta_0 + \delta_1 X_t \tag{3}$$

where I assume that y_t^1 loads only on current states x_t.[3] As shown in the appendix, the model then implies that the yield on a nominal zero-coupon bond with n periods to maturity is a linear function

$$y_t^n = a_n + b_n' X_t \tag{4}$$

where the coefficients $a_n,\ b_n$ are determined recursively.

[3] To reconcile this with r_t being part of the state vector which follows a VAR(2), I follow Dai and Philippon and use the 3-month Treasury bill yield as y_t^1, but the federal funds rate as r_t.

2.2 The use of survey expectations and fiscal projections in estimation

The parameters of the term structure model developed above, i.e. the VAR parameters ϕ_1, ϕ_2 and the unique elements of the covariance matrix ω, the parameters δ_0 and δ_1 of the short-rate equation , and the parameters λ_0 and λ_1 of the risk price specification are estimated by maximum likelihood. Because the term structure model implies the exact linear relationships (4) between the states and the yields, with k states and only one latent factor it is necessary to add measurement error to at least $k-1$ yields to avoid stochastic singularity. The n-period yield is therefore assumed to equal

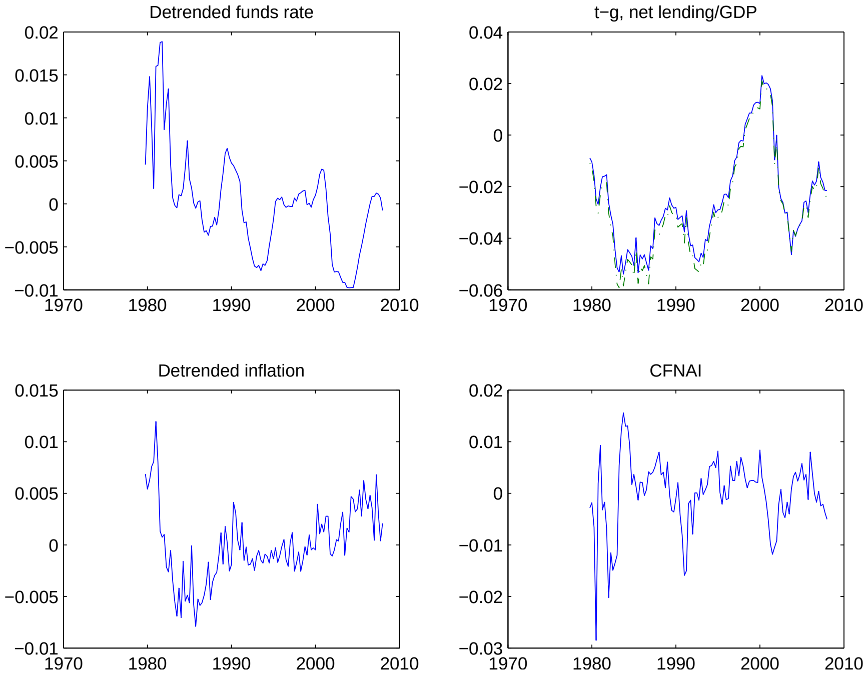
$$y_t^n = a_n + b_n' X_t + \epsilon_t^n$$

By contrast, the macroeconomic variables are assumed to be observed without error. Observed inflation is simply the sum of trend inflation and detrended inflation, $\pi_t = \tilde{\pi}_t + \bar{\pi}_t$, where the mean of trend inflation equals that of observed inflation, whereas the observed short rate r_t equals the sum of the mean real short-term interest rate, trend inflation and the detrended short-term real rate: $r_t = \bar{rr} + \tilde{r}_t + \bar{\pi}_t$. Finally, the observed fiscal measure is equal to its mean and the deviation from that mean: $f_t = \bar{f} + \tilde{f}_t$. The time series of the first four elements of x_t are shown in Figure xx.

The key function of the VAR is to generate expectations of the future states and, through (3), expectations of the future one-period yield. Yet expectations generated by a VAR estimated *ex post* on a given data sample can be poor guides to expectations held by investors at a given point in time. For example, if we were to estimate a VAR using actual inflation instead of decomposing it into trend inflation $\bar{\pi}_t$ and detrended inflation, the VAR would generate long-horizon inflation expectations, and thereby long-horizon expectations of short-term nominal interest rates, that are not nearly volatile enough over our sample (Kozicki and Tinsley, 2001). Similarly, long-horizon expectations of fiscal variables would not show nearly enough volatility compared to long-horizon projections such as the ones prepared regularly by the Congressional Budget Office (CBO). To address this shortcoming, I therefore include survey measures of long-horizon inflation expectations and CBO projections as information variables in the estimation, and impose that the VAR-implied expectations at the respective horizons are equal to these survey measure and projections plus some i.i.d. measurement errors. This assumption implies linear relationships of the form

$$\bar{\pi}_t^{svy} \quad = \quad \zeta_\pi' X_t + \epsilon_t^{svy} \tag{5}$$

Figure 1: Data in the state space



$$f^{CBO}_{t+p,t} = \zeta'_{f,p} X_t + \epsilon^{CBO}_{t,p} \tag{6}$$

where $\bar{\pi}^{svy}_t$ denotes survey-based long-horizon inflation expectations and $f^{CBO}_{t+p,t}$ the p-year-ahead CBO projection of the fiscal variable f_t (details of the data are discussed in the appendix). The coefficients ζ_π and $\zeta_{f,p}$ are functions of the VAR parameters.

With these assumptions, the state space model consists of the transition equation given by the VAR (1) and a measurement equation

$$Y_t = A + BX_t + \epsilon_t$$

in which the vector of observables is given by

$$Y_t = [y^1_t \dots y^N_t \; r_t \; f_t \; \pi_t \; q_t \; \bar{\pi}^{svy}_t \; f^{CBO}_{t+k,t}]'$$

where N denotes the longest maturity included in estimation.

2.3 Identifying fiscal policy shocks

The major challenge in assessing the effects of fiscal policy on interest rates is the endogeneity of fiscal policy measures like spending, revenues or the deficit to other economic variables and shocks. In Laubach (2009) I tried to overcome this problem in the context of reduced-form regressions by focussing on the effects of long-horizon (5-year-ahead) projections of deficits, debt, spending and revenues on proxies for expectations at the same horizon of long-term interest rates. The implicit assumption in this strategy is that changes in fiscal policy measures, especially the deficit/GDP ratio, projected at long horizons reflect exclusively exogenous changes to fiscal policy.

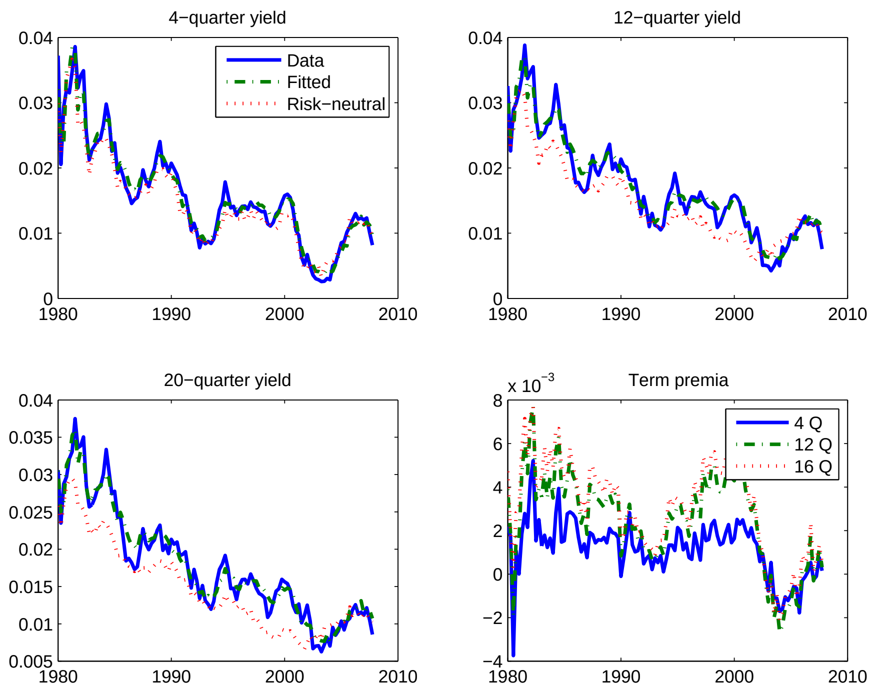
Given that the transition equation of the state vector is a VAR, I am instead following the methodology developed in Blanchard and Perotti (2002) and extended in Perotti (2004) for identifying exogenous shocks to government spending and taxes. In short, the key assumption underlying their strategy is that within the quarter, the fiscal authorities are not able to respond in a discretionary manner to economic news. Hence, the only contemporaneous responses are those implied by the "automatic fiscal stabilizers," i.e. the elasticities of spending and taxes with respect to the macro variables (output, prices, interest rates) included in the VAR. The relations between the VAR's reduced-form residuals u for log real taxes τ and log real spending g and their structural shocks ε can then be written as

$$\begin{aligned} u_t^\tau &= \eta^{\tau,r} u_t^r + \eta^{\tau,\pi} u_t^\pi + \eta^{\tau,q} u_t^q + \varepsilon_t^\tau \\ u_t^g &= \eta^{g,r} u_t^r + \eta^{g,\pi} u_t^\pi + \eta^{g,q} u_t^q + \varepsilon_t^g \end{aligned}$$

The elasticities η can be calibrated from institutional information on tax codes and benefits rules, allowing to identify the structural shocks.

Since taxes and spending are not included separately in the VAR but only through the surplus $f_t = \tau_t - g_t$, I follow Dai and Philippon (2006) and calculate the structural "surplus shocks" by using $\eta^{f,q} = \eta^{\tau,q} - \eta^{g,q}$ etc. To provide some additional information on properties of the VAR, below I also report impulse response functions to a "monetary policy shock" that is identified in the usual recursive manner by assuming that the funds rate responds contemporaneously to all variables in the VAR, but that fiscal policy, inflation and real activity do not respond within the quarter to r_t. More details on the calibration of the elasticities η is provided in the appendix.

Figure 2: Actual and fitted yields and term premia



2.4 Results

2.4.1 Estimation

Despite the imposition of restrictions especially on the risk price parameters λ, the model is fairly highly parameterized, with a total of 78 parameters to be estimated.[4] As in other studies in this literature, I therefore rely on finding good starting values for the VAR parameters ϕ_1, ϕ_2 and ω and δ_1 and only then estimate the price-of-risk and other parameters.[5]

A natural way of assessing the model's properties is the in-sample fit of the yields.

[4]There are 32 parameters in the matrices ϕ_1 and ϕ_2 in (1), 10 unique elements in the covariance matrix ω, 25 parameters in λ_1, five elements in the vector δ_1 in (3) and six measurement error variances on the yields. The intercepts μ and δ_0 are calibrated from the sample means of the respective variables, as are the standard deviations of the innovations to trend inflation (20 bps) and the measurement errors on the survey expectations of inflation (20 bps) and the CBO projections (100 bps). The intercepts λ_0 of the risk prices (2) seem to be poorly identified, so I fix them at zero.

[5]Parameter estimates are available upon request. Inference on especially the price-of-risk parameters produces conflicting results. Although the t statistics on the individual elements of λ_1 are very small, a likelihood ratio test of the hypothesis that $\lambda_1 = 0$ has a p value that is zero at machine precision.

Figure 3: Trend inflation and long-horizon survey expectations

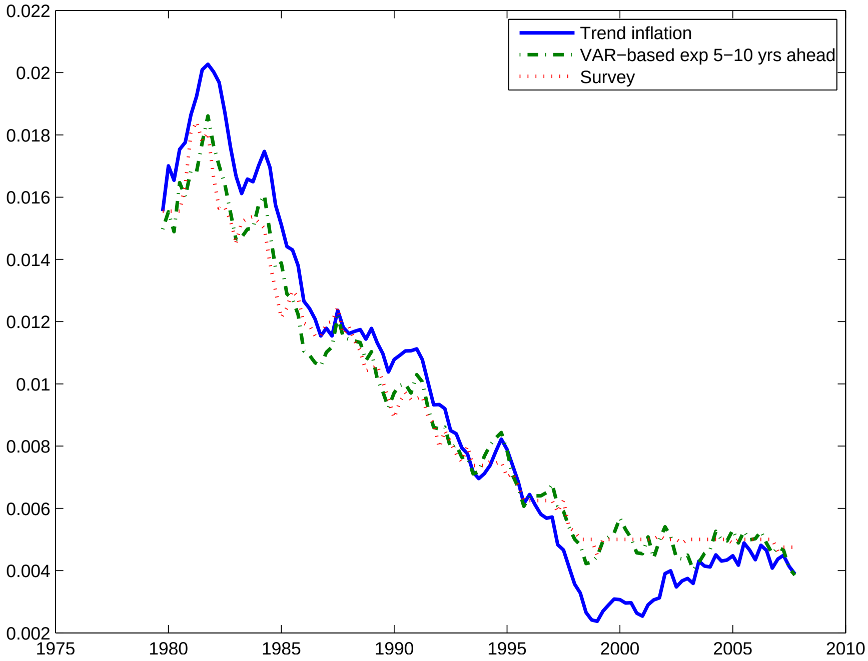
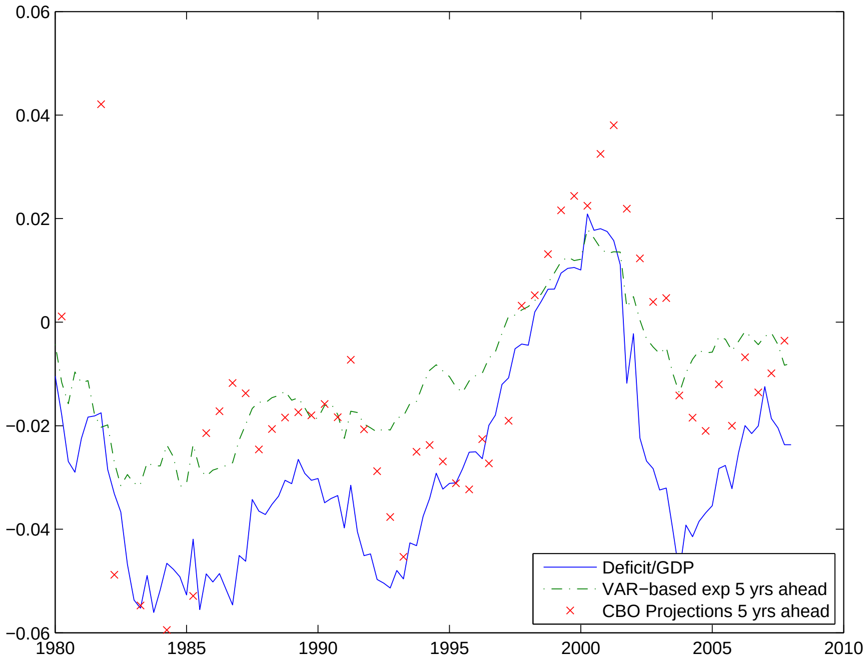


Figure 2 shows for three maturities the historical and fitted yields. These are visually very close, reflecting the fact that the estimated standard deviations of measurement error range from 14 to 16 basis points for any of the maturities included. The dotted line in each of the first three panels presents yields that are obtained from setting λ_1 to zero, in which case the expectations hypothesis holds. The difference between the fitted yields and those dotted lines is shown in the lower right panel of Figure 2. These are the historical series of the term premia on nominal bonds. In the immediate aftermath of the "great inflation," these stood at nearly 8% for 5-year Treasury bonds, but then declined rapidly over the course of the 1980s and early 1990s, rose again during the late 90s before falling to slightly negative levels during the "conundrum" period of 2004-05.

The small size of the measurement error standard deviations is remarkable in light of the common finding that pure macro-factor term structure models tend to produce standard deviations on the order of 50 basis points (e.g. Mönch, 2008). Although strictly speaking the model does include one latent factor, it is important to point out that this factor is tighly linked to an observable series, the long-horizon survey expectations of inflation. As shown in Figure 3, the latent factor (the solid blue line) follows the survey expectations

Figure 4: Surplus/GDP, VAR and CBO 5-year-ahead projections: Model with CBO projections

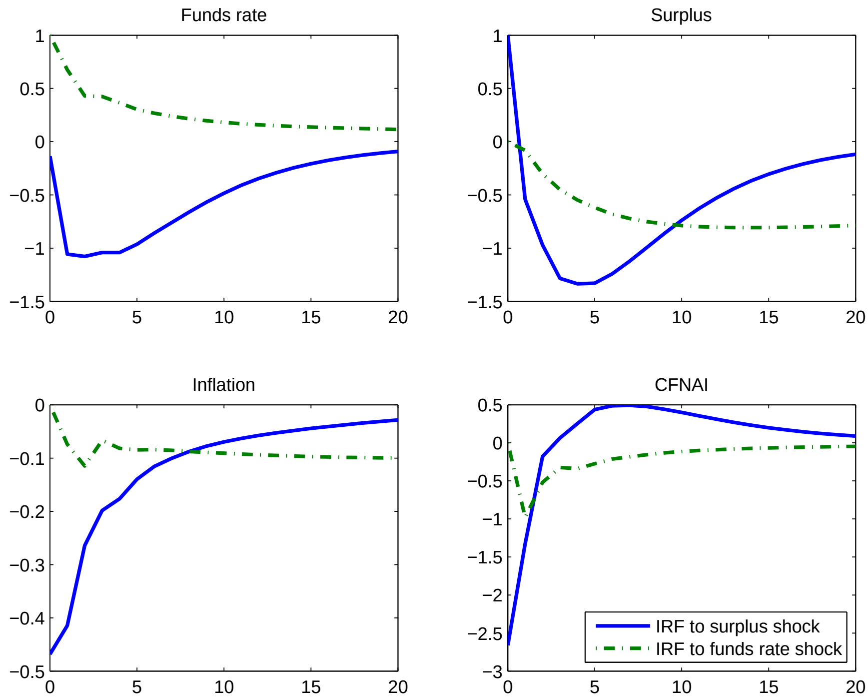


(the dotted red line) relatively closely, because the match between the model-implied (the dashed green line) and the survey-based long-horizon inflation expectation is forced to be close by calibrating the measurement error on the survey expectations to 20 basis points.[6]

The role of the CBO projections in estimating the model is illustrated in Figure 4. The solid line shows the actual surplus/GDP ratio wheres the `x` symbols indicate 5-year-ahead projections of the CBO in the quarter when the CBO released the projection (usually in January and July or August). For the sample 1980:1 to 2007:4, a total of 51 projections is available. In the estimation I use the 3-year and 5-year-ahead projections of the surplus/GDP ratio. Because I am using the so-called "baseline projections" which, as the CBO emphasizes, are no meant to be best forecasts but are by statute based on the continuation of current policies, I deliberately set the standard error on the measurement error to a large 100 basis points because investors may have disagreed with the baseline projections. Nonetheless, including the CBO projections helps to impart substantially more variability

[6]The good properties of long-horizon inflation expectations to substitute for a latent "level factor" have also been found by Joyce et al. 2010.

Figure 5: IRFs of state variables



to the model-implied surplus/GDP projections (the dashed line in Figure 4 for the 5-year horizon).

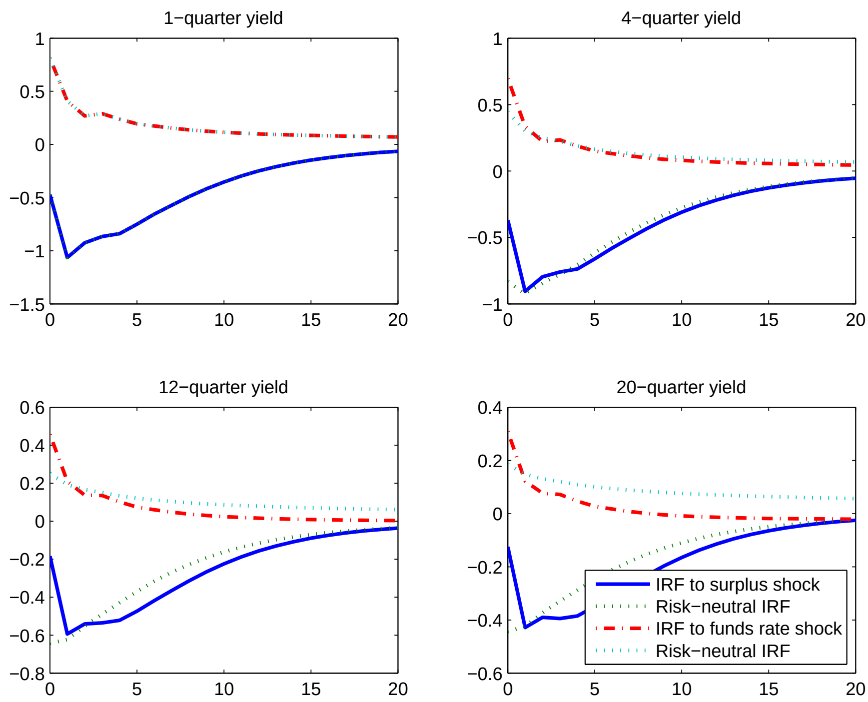
2.4.2 The interest rate effects of fiscal shocks

Figures 5 and 6 present the main results of this section, i.e the impulse responses of the states and of yields at various maturities to a surplus shock and, for comparison with the large literature on measuring monetary policy, also to a funds rate shock.[7]

As shown in the upper right panel of Figure 5, an exogenous fiscal tightening of 1% of GDP is followed by a persistent deficit for the following twelve quarters. According to the VAR, the exogenous fiscal contraction leads to immediate, sharp declines in real activity and inflation, and these declines in turn drive the budget balance into negative territory through the automatic stabilizers. In response to the declines in real activity and inflation, the short rate, shown in the upper left panel, declines by about one percentage point for

[7]Computing confidence bands around the IRFs using conventional bootstrap methods would be computationally extremely burdensome.

Figure 6: IRFs of yields with and without term premia

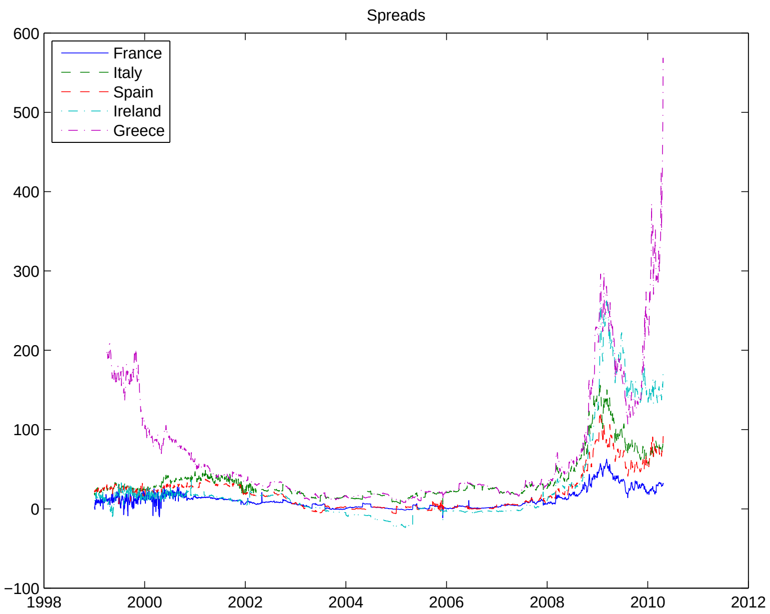


several quarters before gradually returning to its original level.

The solid blue lines in the panels of Figure 6 show the responses of yields at four different maturities to the surplus shock, whereas the dashed red lines provide some information on the transmission of a monetary shock through the yield curve. The dotted lines in each case, labelled "risk-neutral IRFs," provide the impulse responses under the assumption that the prices of risk are zero, with the difference between the solid and dotted lines showing the contribution of risk premia to the overall response. As the maturity increases, the response of the yields is muted. At the 5-year maturity, the tightening leads to a 40 basis point decline in the yield that persists for about four quarter and graduatlly dissipated over the following 12 quarters.

Combining Figures 5 and 6, it is apparent that the mechanism by which exogenous fiscal tightening leads to a reduction in yields is by way of inducing a sharp contraction in real activity and inflation. The effects of exogenous fiscal policy measures on real activity are of course a subject of much recent controversy. The regression coefficient of detrended log real GDP to the real activity measure used here (the CFNAI) is about 1.25. Thus, the response

Figure 7: EMU 10-year government bond spreads vis-à-vis Germany



in the lower right panel of Figure 5 implies a "fiscal multiplier" that is very large but at the same time short-lived. In line with the arguments in Favero and Giavazzi, what seems to be missing is any influence from the implied debt accumulation to interest rates. The impulse response of the surplus/GDP ratio implies that the long-run effect of an exogenous fiscal tightening is an *increase* in debt/GDP, yet this effect is not captured by the VAR analysis. As mentioned before, inclusion of a debt-accumulation identity is not straightforward as the linearity of the law of motion of the state vector is necessary for the model's ability to derive closed-form bond-pricing formulas.

3 Fiscal policy and interest rate spreads in Europe before and since 2008

Developments in yield spreads between EMU government bonds since early 2008 have been rather dramatic. Figure 7 shows some of the data used in this section. These are yields of the currently outstanding government bond closest to 10-year maturity minus their German counterpart.

Although it would in principle be desirable to have the same framework explaining yields whether sovereign default risk is negligible or not – a key factor behind the dramatic rise in the spreads – there are several reasons why the arbitrage-free term structure model developed in the previous section would be difficult to apply to euro area yield spreads. First, while extensions of the framework to the case of bonds with default risk exist (e.g. Duffie and Singleton, 1999), they would be challenging to estimate on the relatively short post-EMU sample that is apparently subject to a regime change in 2008. Second, such models are data-demanding, requiring zero-coupon yields on government bonds for sufficiently many maturities for all the countries considered (for an application to three EMU member countries see e.g. Monfort and Renne, 2010). In this section I therefore pursue a more limited exercise in the spirit of the existing literature on EMU government bond spreads discussed in the introduction.

Focusing on *spreads* of government bond yields between other EMU member countries and Germany, instead of trying to model the *levels* of the individual interest rate series (as in Faini 2006) has the advantage that we do not need to take a stand on the determinants of euro area interest rates, but can focus directly on country-specific influences.

3.1 The relation between EMU government spreads and fiscal policy

Does the level of a country's deficit-to-GDP ratio or debt-to-GDP ratio, or both, affect the interest rate spread that it has to pay on a debt instrument of a given maturity over a comparable German yield? In the *time series*, the answer seems to be cleary "no". Consider the example of Italy. In the middle of 2002, Italy's debt/GDP ratio stood at 118%, its deficit/GDP ratio at 2.9%, and the spread of its 10-year bond over the 10-year Bund at 22 basis points. At the end of 2009, Italy's debt/GDP ratio was 115%, its deficit/GDP ratio 2.9% and the 10-year spread stood at 60 bps. Qualitatively similar observations can be made about many euro area countries: Although the fiscal position of several euro area countries improved over the period until late 2008, the spreads that they had to pay since late 2008 are dramatically higher.

What about the *cross-sectional* relationship between countries' fiscal position at a given moment in time and the spreads that they have to pay? Given a panel of ten euro member countries (GR, PT, ES, BE, NL, AT, FI, IE, IT, FR) I address this question through a sequence of 16 regressions at semi-annual intervals between May 2002 and November 2009. Each of these regressions uses as regressors a constant, the country's current debt/GDP ratio and the (1- to 2-year-ahead projection of) the deficit/GDP ratio as projected by the OECD:[8]

$$s_t^i = \beta_{0,t} + \beta_{1,t} f_{t+k,t}^i + \beta_{2,t} b_t^i + \epsilon_t^i,\ i = 1, \ldots, 9,\ t = \text{May } 02, \ldots, \text{Nov } 09$$

where s_t^i is country i's yield spread at date t, b_t^i the level of debt/GDP at date t and $f_{t+k,t}^i$ the OECD's projection of the surplus/GDP ratio.[9] [10]

Although the number of observations to be fitted at each date is small, the fit of these regressions is nonetheless surprisingly good. Of the 16 regressions, only three produce an adjusted R^2 of less than 0.5, whereas 11 produce an adjusted R^2 between 0.7 and 0.9. Figure 8 illustrates the fit of four of these regressions by plotting the fitted value on the horizontal against the actual spreads on the vertical axis. If all points were lying on the 90° line, the R^2 would be 1.

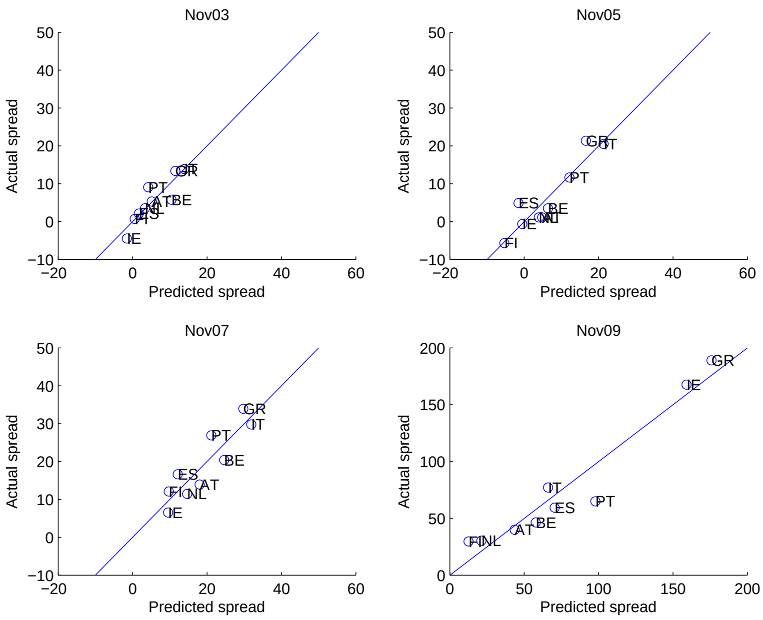
How can the (presumably) poor explanatory power of the fiscal variables for spreads in the time series be reconciled with the very good fit of these variables in the cross-section? The main explanation, as shown in Figure 9, is significant time variation in the magnitude of the parameters. The coefficient estimates on the surplus/GDP ratio and the debt/GDP ratio are shown as xs in the upper and lower panel, respectively, with the vertical bars indicating 90% confidence intervals. For the first 13 regressions (May 2002 to May 2008), the regression coefficient on the surplus/GDP ratio falls between 0 and -3, indicating a 3 bp increase in the spread per percentage point decline in the surplus/GDP ratio; half of the t statistics are 1.75 or higher. The coefficient estimates for the debt/GDP ratio are clustered between 0.1 and 0.3, with 11 out of 13 significantly different from 0 at the 5% level.

[8]The projections are taken from issues 71 to 86 of the OECD's *Economic Outlook*, which were released each year in May and in November. I use the interest rate spreads on the last trading day of the month of the release, and the projected surplus/GDP ratio for the final quarter of the OECD's projection horizon, which is the fourth quarter of the following calendar year in the *EO*s released in May and the fourth quarter two years ahead in the November releases. At this point I am still missing the debt/GDP data for France, hence the results are based on only nine countries.

[9]The use of the OECD's projection for the *total* instead of the primary surplus raises the risk of reverse causality, especially for countries with very high debt/GDP levels. In future work I hope to obtain the OECD's projections for the primary surplus/GDP.

[10]As presumably the ultimate object of interest is some measure of fiscal sustainability, I have also considered regressions in which the *product* of the projected surplus/GDP and debt/GDP is included. For 11 of the 16 regressions these fit the data similarly well, but for ease of interpretation of the results I am focussing here on the linear regression.

Figure 8: Actual vs. fitted spreads from regressions on fiscal variables



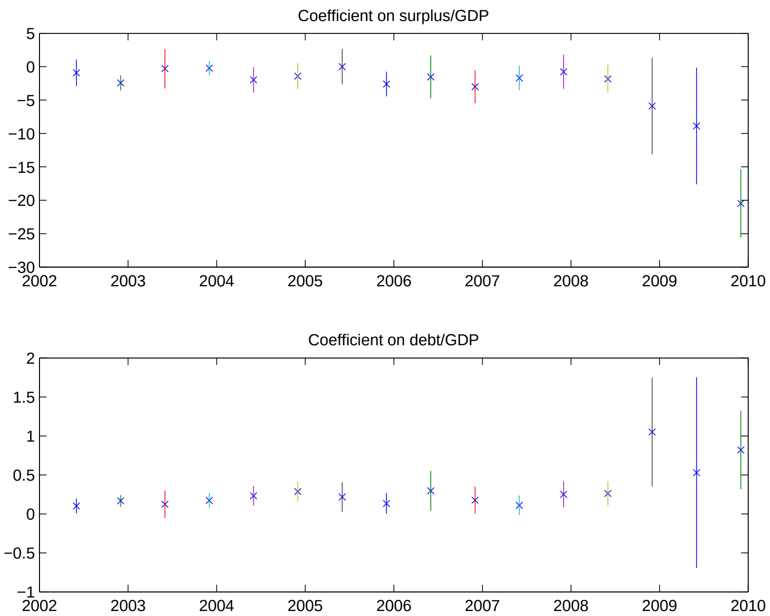
For the final three regressions (November 2008 and May and November 2009), the coefficients have risen dramatically in absolute magnitude. As of November 2009, a percentage point decline in the surplus/GDP ratio raises the spread by 20 bps, and each percentage point increase in the debt/GDP ratio raises it by 0.8 basis points.

3.2 A common-components framework

The regression results presented in the previous subsection suggest that, even before the start of the period of high and volatile EMU spreads, fiscal variables *did* explain EMU government bond spreads, only that the spreads demanded to compensate for different deficit and debt levels were tiny. An important implication of these results is that one and the same combination of deficit/GDP and debt/GDP would nowadays demand a higher spread than prior to the onset of the financial crisis. The recent sharp increase in spreads thus reflects not only a deterioration in countries' fiscal position, but also a higher price for bearing a given amount of risk. One may therefore want to disentangle the changes in spreads due to variation in some common factor from those due to changes in the country-specific fiscal positions.

Previous studies (e.g. Manganelli and Wolswijk) have emphasized the importance of

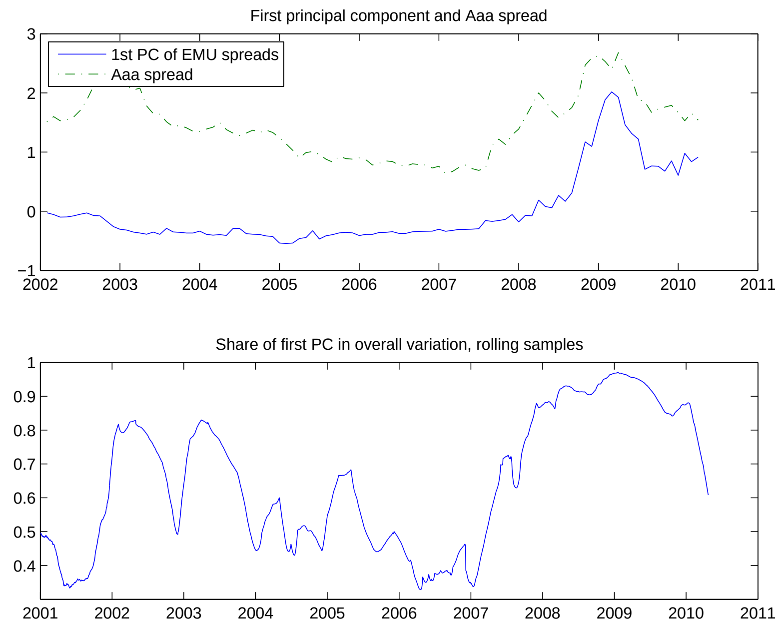
Figure 9: Evolution of regression coefficients



a common factor in interest rate spreads: The first principal component in the monthly data on interest rate spreads used by Manganelli and Wolswijk explains 86% of the total variation. Using daily data of spreads between 10-year government bonds for ten countries I find that, for the sample from the beginning of 1999 until April 23, 2010 the first principal component accounts for 89% of the total variation in the data, for the sample from the beginning of 2001 (when Greek yield spreads had roughly converged to their range until 2008) until February 2008 it explains about 84%. These numbers are little changed when excluding Greece and Portugal from the country sample.

The first principal component, shown for monthly spread data in the upper panel of Figure 10, has frequently been found to be correlated with proxies for "risk appetite" or "risk aversion," such as the spread between yields on corporate bonds of U.S. Aaa-rated issuers and yields on U.S. Treasury securities (Codogno et al. 2003) or (for a much shorter sample) the iTraxx index of non-financial CDS premia (Ejsing and Lemke 2009). The dominance of the first principal component in explaining variation in spreads seems to suggest that changes in investors' risk aversion, which comove all spreads, are the most important source of spread variability, while countries' riskiness remains broadly constant through time. But closer inspection reveals that the share of total variation explained by

Figure 10: Share of first principal component in total variation, 1-year rolling window



the first principal component varies substantially over time. Using again daily data on spreads, the lower panel of Figure 10 shows the share of the first principal component in explaining the total variation of the data estimated from 1-year rolling windows, beginning with the year 2000 (shown at the beginning of 2001) and ending with the year ending April 23, 2010.

As pointed out by Ejsing and Lemke (2009), from the point of view of the principal component analysis the rise in spreads is due not only to the sharp rise in the common component (or a proxy like the Aaa spread), but also to a substantial rise in countries' *loadings* on the measure of risk aversion following the announcement of bank rescue packages in October 2008. This increased sensitivity of spreads to changes in risk aversion is consistent with the finding above, that the coefficients on the deficit/GDP and debt/GDP ratios increased significantly since the onset of the crisis. It suggests a nonlinear effect of changes in risk aversion on spreads, or even the possibility of feedback loops: When risk aversion goes up, countries' debt burden and (all else equal) deficits increase, pushing the spread up even further.

To illustrate recent changes in the sensitivity of spreads to changes in risk aversion, Table 1 reports results from regressing each country's spread on a constant and the spread

Table 1: Comovement of government bond spreads with Aaa spread

Sample	02.01.07 - 29.02.08			03.03.08 - 12.11.09			13.11.09 - 23.04.10		
	β	t stat	$\bar{R}^2$	β	t stat	$\bar{R}^2$	β	t stat	$\bar{R}^2$
GR	0.164	15.97	0.789	1.334	9.572	0.505	-4.024	-3.997	0.368
PT	0.143	19.37	0.836	0.670	9.590	0.538	-1.079	-2.673	0.176
ES	0.090	13.55	0.742	0.474	9.916	0.517	-0.282	-2.172	0.110
BE	0.129	14.08	0.783	0.508	12.70	0.685	-0.050	-0.637	-0.001
NL	0.070	9.502	0.655	0.361	11.47	0.608	0.103	1.525	0.058
AT	0.065	11.38	0.704	0.532	8.306	0.449	0.133	1.429	0.053
FI	0.113	11.75	0.719	0.314	8.864	0.519	0.183	2.180	0.146
IE	0.184	23.01	0.870	0.876	4.489	0.192	0.329	2.211	0.161
IT	0.107	12.20	0.680	0.691	13.73	0.647	-0.080	-0.828	0.007
FR	0.039	5.726	0.383	0.256	11.00	0.573	-0.023	-0.332	-0.005

between Moody's index of yields on Aaa seasoned corporate bonds and 10-year US Treasury bond yields. The regressions are run for three samples:

- January 3, 2007 to February 29, 2008: A period during which EMU bond spreads still fluctuated close to their historical range over the previous years.
- March 3, 2008 to November 12, 2009: The period spanning the first significant rise in spreads around the collapse of Bear Stearns, the announcement of bank rescue packages in October 2008, up until the onset of a string of downgradings of the fiscal outlook for Greece.
- November 13, 2009 until the most recent observation: The period during which the prospect of sovereign default was at the center of attention.

For the first sample, for all but France the adjusted R^2 is 65% or higher, suggesting (in line with Codogno et al.) that the common component of "international risk aversion" was the main driver. This result is only slightly weaker for the second sample, but (as in Ejsing and Lemke) the slope coefficient estimates β, which in both samples are highly significant, rise dramatically – in many cases by a factor of about 10. In the third sample the relationship between the common risk factor and EMU bond spreads breaks down, suggesting the emerging predominance of country-specific factors.

To summarize, the results presented in this section suggest a framework for modelling EMU government bond spreads that features (i) time-varying risk aversion (an imperfectly observed common component), (ii) loadings on this common component that are functions of country characteristics (e.g. debt/GDP, deficit/GDP) as well as (iii) other country-specific effects.[11] While the nonlinear effect, that a deterioration in the fiscal position amplifies the sensitivity of the spread to a given (potentially exogenous) change in risk aversion, seems to be empirically important, it will be difficult to embed such a nonlinearity into a more formal dynamic model of arbitrage-free bond pricing as discussed in the previous section.

Before concluding, it should be noted that identifying the risk inherent in a sovereign bond with traditional measures such as the debt/GDP and deficit/GDP ratio has become more problematic due to the massive expansion of (explicit or implicit) contingent liabilities through government guarantees for their respective banking sectors. These guarantees do not affect traditional measures of deficits and debt, yet seem to have measurable effects on interest rates (Attinasi et al. 2009, Ejsing and Lemke 2009).

[11] In the spirit of Ejsing and Lemke I experimented with a time-varying coefficients model of the form

$$s_t^i = \mu^i + \beta_t^i s_t^{Aaa} + \epsilon_t^i,\ \beta_t^i = \beta_{t-1}^i + \eta_t^i$$

However, results proved to be very sensitive to whether the model was estimated in level form (as above) or in difference form, and generally lead to time variation in the βs completely explaining the variation in spreads. I leave this for future research.

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A The stochastic discount factor and bond pricing

Given the normality of u_{t+1}, the SDF

$$M_{t+1} = \exp\left(-r_t - \frac{1}{2}\lambda_t'\lambda_t - \lambda_t' u_{t+1}\right)$$

is log-normal too. To conserve on parameters, the market prices of risk

$$\lambda_t = \lambda_0 + \lambda_1 X_t$$

are restricted such that under risk-neutral dynamics

$$\begin{aligned} X_t &= \mu^Q + \Phi^Q X_{t-1} + U_t^Q \\ \mu^Q &\equiv \begin{bmatrix} \mu - \lambda_0 \\ 0 \end{bmatrix}, \ \Phi^Q \equiv \begin{bmatrix} (\phi_1 \ \phi_2) - \lambda_1 \\ I \qquad 0 \end{bmatrix} \end{aligned}$$

the states (and hence yields) follow VAR(1). This requires right 5×5 block of λ_1 to equal ϕ_2.

B The data

- Quarterly data, sample 70:1 to 07:4.
- Fama-Bliss zero-coupon yields 1, 4, 8, 12, 16, 20 quarters.
- Funds rate r_t, GDP deflator inflation π_t, CFNAI q_t.
- To use CBO projections, fiscal measures are federal sector. Use federal govt net lending from the NIPAs.
- Spliced series of long-horizon inflation expectations from FRB/US (treated as expectations of average inflation 5-10 years ahead).
- CBO projections 3 and 5 FY ahead (budget concepts close enough to NIPA govt net lending).