THE INTEREST RATE EFFECTS OF GOVERNMENT DEBT MATURITY

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Abstract

Using an empirical model, this paper finds that shortening the average maturity of US Treasury debt held outside the Federal Reserve by one year reduces the five-year forward 10-year yield by between 130 and 150 basis points. Based on a pre-crisis period, these estimates suggest that portfolio balance effects are unlikely to reflect only post-crisis market conditions. These findings also offer a partial explanation for the Greenspan conundrum: the fact that long-term interest rates in the mid-2000s rose less than expected after a rise in the Fed fund rate may have been due, to some extent, to the concomitant shortening of government debt maturity.

JEL classification: E43; E52; E63

Keywords: long-term interest rate; sovereign debt management; portfolio balance effects; quantitative easing.

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1. Introduction

Have large-scale central bank purchases of bonds reduced Treasury yields in the aftermath of the Great Financial Crisis? Were portfolio rebalancing effects quantitatively relevant even before the crisis? Answering these questions affirmatively requires some caution. For one, the channels through which these purchases work are not entirely clear. Buying long-term Treasuries and other assets may simply be viewed as strengthening the central bank's commitment to keeping the policy rate at the near-zero level for a longer period (see eg Krishnamurthy and Vissing-Jorgensen (2011); Bauer and Rudebusch (2014)). The longer the public believes the central bank will keep the Federal funds rate low, the lower long-term interest rates (signalling or expectation effect). Moreover, most quantitative estimates of the impact of central bank purchases are based on samples comprising the post-crisis period. Capital constraints on banks and other financial firms, worries about the creditworthiness of wholesale market counterparties and uncertainty about future regulations may all inhibit arbitrage by the private sector. Hence, any effects of asset purchases may largely reflect the difficult market conditions prevailing in the post-crisis period and may therefore vanish once market conditions normalise.

In order to shed some light on this issue, we investigate the empirical relevance of the maturity effects of US Federal debt *over a pre-crisis sample*. Our empirical strategy draws on Laubach (2009), who found significant effects of the prospective budget deficit and prospective debt-to-GDP ratio on forward long-term yields. We extend this analysis by including the average maturity of the Federal debt held outside the Federal Reserve. In addition, we check that our findings are consistent with those obtained from identical regressions but using an estimate of the long-run term premium (Hördahl and Tristani (2014)) as the dependent variable. Focusing on a period that precedes the start of the crisis improves the chance of identifying the supply effects of federal debt, which could otherwise be obfuscated by post-crisis special or transient factors that are currently depressing long-term yields.

The analysis has uncovered three main results. First, a one-month increase in the average maturity of debt outstanding held outside the Federal Reserve is associated with a rise of 12-13 basis points in the five-year forward 10-year rate and a rise of 10-13 basis points in the 10-year term premium. These estimates are consistent with the existence of significant portfolio balance effects. These effects rest on the existence of preferred-habitat investors that tend to demand specific maturities and the failure of arbitrage to eliminate price differences across maturities (Vayanos and Vila (2009) and Greenwood and Vayanos (2014)).[1]

Second, a shortening of the maturity of public debt since the early 2000s accounts for most of the reduction in the observed forward long rate and the term premium. Thus, it may partly explain what Greenspan called a conundrum – that is, the failure of long-term interest rates to rise in the face of a tightening of the Federal funds rate in 2005.

Third, the average maturity of Federal debt makes the inflows into US Treasuries from the foreign official sector superfluous in explaining long-term rates. That is, when the former variable is added as an explanatory variable, the estimated coefficient on inflows becomes statistically insignificant and close to zero.[2] We interpret this result as indicating that US debt managers have generally accommodated the demand for shorter maturities from the foreign official sector by reducing average issuance maturity.

In addition to Laubach (2009), our paper is related to Greenwood and Vayanos (2014), who find that the relative supply of long-dated securities is positively related to excess returns and a number of recent papers that attempt to quantify the effects of Federal Reserve interventions in the bond market (eg Gagnon et al (2011), Doh (2010), D'Amico and King (2012), D'Amico et al (2012) , Meaning and Zhu (2011, 2012)).[3] Our findings corroborate the results in these studies using a different empirical framework as well as different variables. The advantage of using the five-year forward 10-year interest rate as dependent variable is that this should be little influenced by expected changes in the policy rate.

1 The econometric effect of an increase in the debt-to-GDP ratio is found to be approximately two basis points, which is similar to what found by Laubach (2009).

2 Using the same data and specification as in Warnock and Warnock (2009), we find the adding average maturity to the set of regressors renders capital inflows into treasuries statistically insignificant (see Annex Table A5).

3 Swanson (2011) recently revisited "Operation Twist" in the 1960s, finding larger effects than earlier studies. Joyce et al (2011) find supporting evidence of portfolio balance effects following quantitative easing in the United Kingdom.

In addition, our findings confirm that sizeable portfolio rebalancing effects had been at work *even before* the financial crisis, thus suggesting that a shortening of public debt maturity could have been one of the factors that contributed to the Greenspan conundrum and the prolonged downward drift in long-term interest rates prior to the crisis.

The paper is organised as follows. Section 2 explains why the relative supply of short-dated and long-dated public debt to the market matters for long-term interest rates. Section 3 describes the empirical analysis and its results. Section 4 asks whether changes in debt maturity can help to explain the Greenspan conundrum. Section 5 assesses the potential impact of Federal Reserve purchases of Treasuries following the crisis. Section 6 concludes.

2. Theoretical considerations

Until the mid-1980s, the prevailing orthodoxy among economists followed Keynes, Tobin and Milton Friedman in viewing portfolio balance effects as key to understanding how monetary policy worked.[4] From the late 1980s to the onset of the crisis, however, many monetary economists had come to view portfolio balance effects as irrelevant or empirically too small to justify their inclusion in formal models of monetary policy. Indeed, in the standard New Keynesian model all that matters for demand determination of aggregate demand is the path of current and future short-term interest rates; changes in the relative supplies of financial instruments, including money and government debt, have no role in shaping the yield curve (Eggertsson and Woodford, 2003). Since the crisis, however, these assumptions have been called into question by the balance sheet policies pursued by central banks.

Finding better theoretical foundations for portfolio balance effects has therefore become a top research priority. One promising research avenue is the development of investors' preferred habitat theory within a modern model of the yield curve. Building on the preferred habitat hypothesis

[4] See eg Zampolli (2012) for an overview. Tobin (1961, 1969) provided formal models based on imperfect asset substitutability. See also Tobin (1963) on public debt management.

(Culbertson (1957), Modigliani and Sutch (1966, 1967)), Vayanos and Vila (2009) and Greenwood and Vayanos (2014) have developed a formal model that combines two ingredients. The first is the existence of heterogeneous preferences about the maturities that some agents want to hold. For example, pension funds may be keen to "lock in" a long-term interest rate for their assets. Or they may face regulatory restrictions that require them to match their asset duration to that of their liabilities. The second ingredient is the ability of arbitrageurs to undertake maturity transformation across the yield curve, which depends, among other things, on their risk aversion and their availability of capital.

The theory predicts that increasing the average maturity of public debt, for a given debt size, should raise long-term interest rates relative to the path of future short-term rates.[5] The initial effect of an expansion in the supply of long-term bonds, given unchanged demand, is that bond prices would fall. Arbitrageurs would then buy the cheaper long-term bonds (reversing part of the initial decline) in exchange of short-term bonds (whose price would rise). The initial gap in prices would not be eliminated completely because arbitrageurs demand a risk premium to cover the interest rate risk from holding a larger stock of long-term bonds. Their ability and willingness to bear this risk may also vary over time.

The effect of lengthening debt maturity is to raise all interest rates, including the short-term ones. This may seem surprising given that an increase in maturity should lead to a relative scarcity of short-term bonds (and possibly even to a reduction in their yield). However, as shown by Greenwood and Vayanos (2014), the larger exposure of arbitrageurs to interest rate risk raises the market price of risk, affecting bond prices at all maturities. Thus, for sufficiently low levels of arbitrageurs' risk aversion, changes in the market price of short rate risk dominate the effects of local demand and supply conditions.

[5] These effects require a failure of Ricardian equivalence; otherwise, private agents would simply buy any additional amount of bonds issued by the government with no need for interest rates to induce such behaviour. However, borrowing, liquidity and informational constraints, among others, are important factors that make non-Ricardian effects quantitatively important.

An important point to note is that, as a first approximation, what should matter for the determination of long-term interest rates is the net supply of government bond to the market, not on how much the central bank holds on its balance sheet.[6] The size and maturity composition of such net supply can be altered through central bank operations, such as the ones that the Federal Reserve conducted since the beginning of the Great Financial Crisis, or through the debt issuance policies and buy-back operations conducted by the debt manager, which in the United States is a separate entity from the central bank. Before the crisis, debt held outside the Federal Reserve and its average maturity had been driven mostly by fiscal policy and debt management decisions, closely tracking changes in the overall debt (Graph 1).[7]

3. Econometric framework

In our econometric analysis, we extend Laubach (2009) in three ways. First, we focus not only on the size of government debt, but also on its average maturity. Another important difference is that we employ government debt held outside the Federal Reserve (see above). Both the size and the average maturity of public debt held in the market proxy for investors' exposure to interest rate risk. Second, we conduct an extensive analysis of parameter stability and check the robustness of our results with additional control variables. Finally, as a cross check on the results regarding the forward rates, we also consider an estimate of the 10-year term premium (Hördhal and Tristani, 2014), which by construction excludes the expected future short-term rate.

[6] The main qualification to this is the differing impact on expectations. Central banks, with their purchases, can also influence expectations of their future decisions about the policy rate or additional purchases. They may therefore have more leverage on interest rates than debt managers. But actions should first be judged on the basis of how they affect the portfolios of those agents who have to be induced, in the open market, to buy bonds.

[7] Since the beginning of this crisis, sovereign debt management has often worked in apparent conflict with the objectives of monetary policy (Blommestein and Turner, 2012).

3.1. Econometric specification and data description

Our specification is as follows:

$$f = \alpha\pi^e + \beta g^e + \delta d^e + \gamma m + \phi X + \varepsilon \tag{3.1}$$

where f is the five year forward 10-year interest rate or the 10-year term premium, π^e is a measure of long-term inflation expectations, g^e is a measure of future expected real output growth (or trend growth), d^e is a measure of expected future fiscal policy (either the debt-to-GDP ratio or the fiscal deficit), m is a measure of the average maturity of debt held outside the Federal Reserve, X is a list of other control variables and ε is an error.

The measure of inflation is the survey of long-horizon inflation expectations by market participants and professional forecasters (published by the Federal Reserve Bank of Philadelphia). The measures of expected future output growth, public debt and deficits are all 5-year-ahead projections of the respective variables published twice a year by the CBO.[8] The average maturity of federal debt held outside the central bank is expressed in months.[9]

The list of control variables includes a measure of risk appetite as well as a measure of short-term interest rate uncertainty. Specifically, we control for the stock market dividend yield on the assumption that investors who become more risk averse shift their portfolios away from equity and towards government bonds, thereby leading to a rise in the dividend yield and a corresponding fall in the bond rate. As a measure of short-term interest rate uncertainty, we take the 12-month rolling standard deviation of the 3-month Treasury bill rate.[10] Furthermore, to control for the effects of the business

8 Since 1985 the CBO has regularly released its projections twice a year. Hence, our sample is made up of half-yearly observations. Before 1985, its projections were published most of the time with yearly frequency and occasionally twice a year. This means that there are gaps in the early part of our sample.

9 Table FD5 in the US Treasury Bulletin (www.fms.treas.gov/bulletin/b2012_4fd.doc).

10 We also considered other measures of risk aversion such as the residuals from the consumption model of Lettau and Ludvigson (2001), as used in Laubach (2009), and different measures of risk such as the volatility of long-term interest rates calculated, like in Warnock and Warnock (2009), as the rolling 36-month standard deviation of changes in long rates; the realised standard deviation of the SP500 return and the VIX index

cycle, we consider the three-month Treasury bill rate, the real-time output gap,[11] the unemployment rate and the Aruoba-Diebold-Scotti index (Aruoba et al (2009)). We also check that our results are not driven by either domestic or foreign central bank operations. That is, we also control for the total holdings of Treasuries by the Federal Reserve and for purchases by the foreign official sector.[12]

3.3. Estimation method and properties of the data

We estimate equation (3.1) by OLS. Given that the errors exhibit serial correlation, we compute and report Newey-West standard errors. We assume error serial correlation of lag three. Preliminary investigation has revealed that the univariate AR model that best fits the errors has generally three lags, which is sufficiently large to allow for quite general processes governing the residuals. Regardless, our estimation results are generally robust to changes in the lag. Before proceeding to the discussion of the results, three potential problems need to be addressed: stationarity; endogeneity; and instability.

3.3.1. Stationarity

Measures of long-term interest rates have been declining since the early 1980s along with inflation and inflation expectations. Other variables among the regressors also seem to exhibit non-stationary behaviour. We therefore test for unit roots and cointegration. Annex 1 provides the details. For the interest rate variables and inflation expectations the tests generally point to the presence of a unit root. For the fiscal variables the evidence is more mixed. Cointegration tests generally support the assumption

(available only as of 1986). In a preliminary investigation these variables generally did not turn out to be significant or did not lead to well-specified models.

11 This is constructed using real-time estimates from the Federal Reserve Board staff before 1997 and CBO estimates after 1997.

12 Official inflows into US Treasuries are taken from TIC data and are available only as of 1978. We take a one-year rolling average and scaled it by potential GDP, rather than actual GDP, to minimise the chance of business cycle fluctuations spuriously affecting the estimates. We also scale Federal Reserve holdings of Treasuries by potential GDP.

that interest rates are cointegrated with inflation expectations even when fiscal variables are added to the basic regression.[13]

It is important to note that in small samples such as ours there is no sure way of distinguishing between a series exhibiting a unit root and a stationary one with a high degree of persistence. That is, the tests have very low power and cannot be relied upon entirely. In our view, plots of the series and judgment, along with the tests, support the stationary assumption on which the regressions are based.

3.3.2. Endogeneity

The second potential problem is the endogeneity of the regressors. In particular, the maturity structure, which is the main focus of our analysis, may not be exogenous to macroeconomic developments. Several authors (eg Kuttner (2006), Gagnon et al (2011)) have suggested that OLS estimates of regressions of yields or spreads should be *biased downward*, thereby offering a conservative estimate of the true portfolio rebalancing effects of public debt. The reason is that sovereign debt managers tend to reduce issuance maturity when long-term yields or term spreads increase, thereby inducing a negative correlation between dependent and explanatory variables.[14]

A number of other reasons suggest that the endogeneity bias is either small or negative. For one, some of the changes in the maturity structure of the public debt over our sample are clearly driven by changes in legislation and hence could be considered exogenous with respect to interest rates. For example, at the start of our sample in 1976 the Congress lifted the 4¼% ceiling on the coupon that could be offered on 10-year Treasuries – in effect allowing the issuance of 10-year bonds, which led to a gradual increase in average maturity in the following years. Later, following several years of large fiscal surpluses in the second half of 1990s, the Treasury undertook a number of buyback operations between

[13] Laubach (2009) also shows that the five-year ahead 10-year interest rate is cointegrated with expected long-term inflation and that adding the fiscal variables do not alter this conclusion.

[14] Hoogduin, Öztürk and Wierts (2011) report evidence that euro area countries increase the share of short-term debt when the short-term interest rate falls. Blommestein and Turner (2012) show that US debt issuance tends to shorten when the Federal funds rate falls.

2000 and 2002. As short-term notes were expiring without being replaced by new issuance, the average maturity rose. The buyback operations were decided to bring average maturity down. In 2001 the Treasury also discontinued the issue of the 30-year bond (see eg Garbade and Rutherford (2007)). Second, the US debt managers have explicitly ruled out timing the market and opportunistic behaviour. Their choices about issuance maturity has often been driven by the objective of reducing fiscal risks associated with rising levels of debt. Finally, another reason why the endogeneity bias may be small is that we focus on the forward long-term rate and the term premium, which should not be much influenced by the short-term interest rate. To the extent that debt managers issue longer-term debt in response to a narrowing of the yield spread, we would expect the endogeneity bias to be smaller in regressions in which the short-term rate, and hence the yield spread, does not play a relevant role.

3.3.3. Instability of estimates

The third potential problem is the instability of the estimated coefficients. Our sample runs from 1976 and hence covers different monetary and fiscal policy regimes. Over this long period estimates may vary too much, thus misleading about the true size of the effects. To minimise this risk, we carried out some preliminary structural break analysis (see Annex 2 for the details). Namely, we assume that all coefficients except that of inflation expectations (including the intercept) in our baseline specification could be subject to a structural break at an unknown date.[15] By varying the possible break date, we compute a sequence of F statistics on the break coefficients, which can then be used to construct a test for the existence of a structural break within the sample. The test statistics that we consider are: the largest value of the F statistic (the Quandt or SupF statistic); and the ExpF and the AvgF statistics suggested by Andrews and Ploberger (1994).[16] Given the possible presence of heteroscedastic errors, we rely on

[15] Our baseline specification includes long-term inflation expectation, expected growth, the dividend yield, interest rate volatility, expected debt and average maturity. Given the assumption that the forward rate is cointegrated with long-term inflation expectations, we maintained the assumption that this latter variable has no structural break. However, we also checked the existence of a break in inflation expectations and found no evidence for it. The dataset used for the parameter stability analysis runs from 1976 to 2011.

[16] Approximate p values for these statistics are provided by Hansen (1997). We also vary the set of coefficients assumed to be subject to a structural break obtaining consistent results.

minimising the residual variance to identify the date at which the break occurs. We first find strong evidence of a break in the first half of 1986. Repeating the test procedure over the split sample, we find another clear break in the second half of 2008, corresponding to the most acute phase of the financial crisis. We therefore proceed to present results for the pre-crisis sample 1976H1–2008H1 taking into account the break in 1986H1.

4. Results

In what follows we present two sets of results which differ in the dependent variable used in the regression. The first set concerns the five-year-forward 10-year interest rate while the second set regards the 10-year term premium.

4.1. Five-year forward 10-year interest rates

The structural break analysis leads to the estimation of a general specification (shown in Annex Table A4) in which all coefficients but inflation expectations are allowed to change at the selected break date of 1986H2. The examination of this equation suggests that the coefficients of expected debt and average maturity remain constant before and after the break date, whereas those of the other variables do change. In particular, expected future growth, the dividend yield and the short-term interest rate volatility turn out to be significant only before the break date (t<86H2) and statistically insignificant thereafter (t>=86H2). The validity of these restrictions is also confirmed by an F test (see the bottom of Annex Table A4). The shift in the intercept is strongly significant, capturing the unusually high interest rate levels in the early part of the sample, which cannot be completely explained by changes in the explanatory variables (and for which it is hard to find other satisfactory explanatory variables).

The tested-down version of the previous equation is shown in the first column of Table 1. With one exception, all coefficients in column 1 have the expected sign and appear of reasonable size:

- A one percentage rise in long-horizon inflation expectations adds about one percentage point to the 10-year yield.

- A one percentage point rise in the debt-to-GDP ratio five years ahead is associated with about 2 basis points increase in the forward rate, a finding that is very close to what found by Laubach (2009) in his regressions as well as in his calibration of a small neoclassical growth model.
- Greater volatility in the short-term rate drives up the long-term rate.
- Greater risk aversion (as proxied by the dividend yield) drives down long-term interest rates.
- Higher trend economic growth (as proxied by five-year-ahead output growth rate) reduces the long-term interest rate.

The latter result may appear at odds with economic theory. One would have expected the long-term interest rate to be positively related to economic growth. Yet, the coefficient on the five-year-ahead output growth rate has a negative sign and a large magnitude. This variable is unrelated to the other regressors: when it is dropped the estimated coefficients of the other variables are very little changed and the loss of fit of the model (measured by the adjusted R-squared) is minimal (Column 2). One possible explanation is that CBO projections for output five years ahead are not a good proxy for trend growth. For example, a negative sign may be an indication that forecasters were projecting future trend growth on the basis of current macroeconomic conditions including the long-term interest rate, downgrading future trend growth when interest rate were unusually high, other things equal. Further investigation reveals that trend growth appears to be negatively correlated with the forward rate in the early part of the sample. When we re-run the regression from 1980 (not shown here), thus excluding this early part of the sample, we find that the coefficient on trend growth becomes positive albeit statistically insignificant. Hence we conclude that the negative and significant coefficient on trend growth is likely to be an artefact of the data. We therefore regard the regression in Column 2 as our baseline case.

Public debt and its maturity are significantly and positively associated with the forward long-term interest rate. A one percentage point rise in the debt-to-GDP ratio five year ahead is associated with about 2 basis points increase in the forward rate. Lengthening the maturity of public debt (held outside the Federal Reserve) by one month is associated with a rise of between 12 and 13 basis points (the 95% confidence interval is from 11 to 15 basis points). This is equivalent, other things equal, to a rise in the forward rate of almost 150 basis points for each year of increase in the average maturity.

The effects of maturity are robust to dropping the dividend yield (Column 3) as well as interest rate volatility (Column 4), although in the latter case the effect of debt becomes slightly smaller and statistically insignificant. This shows that controlling for the volatility of short-term interest rate in the early part of the sample is necessary to identify the effects of debt. Column 5 shows that the results remain robust even if a break is not allowed in the control variables (although of course the fit of the model diminishes). In this case, the dividend yield is statistically insignificant and expected future growth continues to have a negative but statistically insignificant coefficient.[17]

Finally, the results concerning the debt and its maturity are robust to re-estimating the regression over the post-break period 1986H2-2008H1 (Column 6). Note that when we add the control variables, these turn out to be statistically insignificant (Column 7), thereby confirming previous results about the existence of a break in the coefficients associated with these set of variables.

Table 2 shows that the findings are robust to additional control variables. For ease of comparison Column 1 in Table 2 reports our baseline regression (corresponding to Column 2 in Table 1). Both the short-term interest rate (Column 2 in Table 2) and the real-time output gap (Column 3 in Table 2) have no effect on the forward rate and cause very little change in the other coefficients.

When Federal Reserve holdings of Treasuries are added, it is also found to be statistically insignificant – that is, once account is taken of changes in the maturity of Treasuries outside the central bank (Column 4). Interestingly, however, when maturity is dropped from the regression, Federal Reserve holdings of Treasuries (which approximate the size of the their balance sheet before the crisis) has a strong negative, and statistically significant, effect on the forward rate (Column 5). However, there is no

17 Laubach (2009) finds that trend growth has a negative but statistically insignificant coefficient. So, when we do not allow for breaks in the sample our results about trend growth are consistent with his.

plausible explanation in our view for a causal link running from their balance sheet to the forward long-term rate for the sample period of our regressions. Indeed, from the early 1980s to the beginning of the financial crisis, the Federal Reserve had been expanding the size of its balance sheet, in response to a higher demand for currency (which was rising more rapidly than nominal GDP) at the same time as long-term interest rates and expected inflation were falling. Controlling for the maturity of public debt held by the private sector – which is affected by the central bank operations – eliminates this spurious and counterintuitive effect.

When we add foreign official purchases of Treasuries (Column 6), we find that this variable is statistically insignificant.[18] When average maturity is dropped, the effect of official inflows becomes negative and statistically significant and its magnitude is similar to the one found by Warnock and Warnock (2009). These authors estimate that an increase of about 2 percentage points of (lagged) GDP in 2004 contributed to reducing the 10-year interest rate by some 80 basis points. Our estimate leads to a similar conclusion: expressed in terms of potential GDP, inflows in 2004 peaked at 1.77%, which multiplied by our estimate of –0.522 leads to a negative contribution of about 90 basis points. The crucial point, however, is that we find this result when we do not control for public debt maturity.[19]

[18] Note that the series of official sector inflows into Treasuries is available as of 1978. Given that we are using a 12-month moving average, the first available observation is 1979H1. Taking into account the gaps in the data due to the irregular releases of the CBO, the estimation sample has now three fewer observations.

[19] The results shown in Warnock and Warnock (2009) are based on a different sample period (1984–2005) and on monthly frequency. In addition, the dependent variable is the spot 10-year rate instead of the forward rate; foreign inflows are scaled by lagged GDP rather than potential GDP; and the set of other regressors also differs. We downloaded the dataset used in Warnock and Warnock (2009) and tried to replicate their results using the average maturity of privately-held public debt (at monthly frequency) as an additional control variable. Also in this case, we found that the official inflows are no longer significant while average maturity is (see Annex Table A5). Specifically, we are able to replicate exactly the results for the nominal 10-year yield in columns (1), (2) and (5) of Table 1 in Warnock and Warnock (2009), but not for column (6). The reason is probably due to insufficient number of observations in the dataset. Nevertheless, the estimates are very similar. (We did not try to replicate columns (3) and (4) for the real 10-year yield.) When we add the average maturity, both the latter variable and capital inflows are statistically significant in the full sample and the coefficients on capital inflows increase substantially. However, in the shorter samples starting in 1987 and 1994, capital inflows become statistically insignificant, while average maturity is significant. We presume that the different result obtained in the full sample is due to the existence of a structural break not accounted for in the original analysis of Warnock and Warnock (2009) (see also Annex 2).

Our findings are consistent with the idea that debt managers may have varied the maturity structure of public debt in partial response to the preferences of foreign official investors for shorter maturities. Of course, average maturity could capture not only the accommodation of rising foreign official inflows but possibly also domestic factors, including the accommodation of an increasing demand for more liquid assets arising from an expanding and leveraging financial system.[20]

We conduct further robustness checks in Table 3. We add both the unemployment rate and the ADS index of business cycle conditions (Aruoba et al (2009)) to the baseline regression (Column 1). Neither variable alters the original findings (Columns 2 and 3). Interestingly, the ADB index turns out to have a negative effect (statistically significant at the 10% level) on the forward rate, suggesting that the term premium is countercyclical. We also add a time trend to check whether there are common trends driving the main results. The time trend is not statistically significant and the main results are upheld (Column 4).[21] Finally, we checked whether replacing the five-year ahead with a current measure of public debt makes a difference. The latter turns out to be statistically significant, but the coefficient on average maturity remains statistically significant and of similar magnitude (Column 5). A similar conclusion is reached when the five-year ahead debt measure is replaced by the current fiscal deficit (Column 6).

4.2. 10-year term premium

The term premium is the part of long-term interest rate that is not affected by the expected future short-term interest rate. As such, it offers a useful cross-check on the estimates obtained using the

20 Using the five-year forward five-year interest rate as well as the 10-year forward five-year rate as dependent variables gives estimates of the effects of debt and its average maturity which are very similar to the ones shown in Table 1 (details available on request).

21 We have also run the regressions using the real forward rate (ie the difference between the nominal forward rate and inflation expectations) as well as in first differences, finding similar results (details available on request).

forward rate. Table 4 shows regressions that employ the term premium estimated by Hördahl and Tristani (2014) as the dependent variable. Given data availability, our sample starts in 1990.[22]

Our baseline equation (Column 1) includes trend growth, the dividend yield and the volatility of Treasury bill rates as control variables in addition to five-year ahead debt and average maturity. Columns 2 and 3 report the same regressions dropping some of the insignificant control variables. Note that long-term inflation expectations are generally not significant. This is consistent with the relative stability of inflation enjoyed by the US economy since the 1990s. Furthermore, the effect of debt is statistically insignificant. For this reason, we also run regressions with the CBO five-year ahead projection of budget deficit in place of the five-year ahead debt (Columns 4–6). This is statistically significant and shows that a one percentage point increase in the expected deficit leads to 9 basis point rise in the term premium.

Another noticeable finding is that, perhaps surprisingly, the dividend yield is positively correlated with the term premium. We took this variable as a proxy for investors' risk aversion. In fact, over the sample period that starts in 1990, the positive correlation of this variable seems consistent with investors shifting their portfolios towards equities on the wake of a positive dividend growth expectation. Given competing higher returns from equities, bond investors would have required a larger premium to stick to government bonds.

Across different specifications the effect of average maturity is between 10 and 13 basis points, which is in the ball park of what was found in regressions of the five-year forward 10-year rate. The fact that we find similar estimates suggests that the estimated effects on the forward rate mostly reflect changes in the term premium.

[22] According to this model-based calculation the term premium has been around -100 basis points since early 2012. Part of this unusual negative reading, however, reflects an exceptional flight-to-quality or a flight-to-liquidity after the crisis. Chadha (2012) shows that such effects have been important historically.

4.3. How does the model fit data during the crisis?

It is interesting to assess how the empirical model fits actual data during the crisis period. Graph 2 shows the five-year forward 10-year rate along with the predicted values from the respective baseline regressions. In the first stage of the crisis the forward rate was significantly above what the model predicts. The strength of liquidation forces may well explain why rates did not decline. The strong demand for liquidity in that period was met by several large Federal Reserve programmes.

In the subsequent stage of the crisis (from the end of 2008 onward), long-term rates fell well below their predicted values. Graph 2 shows that, given the significant rise in both the expected future level of Federal government debt and the average maturity of debt held outside the Federal Reserve between 2008 and the present, the 5-year forward 10-year yield should have risen from 4% in 2008 to over 5% (dashed blue). It actually fell to 3% (red line), below the lower 95% confidence interval (yellow shading).

This large discrepancy reflects the existence of important factors not captured by the empirical model. First of all, continued asset purchases by the Federal Reserve, reinforced by several prominent speeches, could have created the conviction in markets that the central bank will try to keep long-term rates from rising above a certain ceiling seen as incompatible with a rapid return to full employment.

Another reason is that several new, non-monetary factors seem to have increased the demand for government bonds. New prudential regulations, mark-to-market accounting rules, actuarial conventions, etc have induced banks, insurance companies, pension funds and other financial intermediaries to hold a higher proportion of their assets in government bonds (Turner, 2011). Furthermore, the post-crisis decline in unsecured interbank lending and higher swap margin requirements have led to increased demand for collateral in financial transactions in wholesale markets.

The impact of these non-monetary factors may ultimately wane, but the timing of this is highly uncertain. At some point financial institutions will have reconstituted their capital and liquidity buffers. Furthermore, financial institutions that hold a large share of their portfolio in US Treasuries are unlikely to be able to meet the return expectations of their clients (eg to pay satisfactory pensions) and will face increasing pressure to invest in higher-yield assets. Should pre-crisis empirical regularities reassert themselves, the rise in long-term rates could be substantial.

Our estimates can help to throw some light on: (1) the decline in long-term interest rates in the 2000s and especially the decoupling of long-term rates from short-term rates; and (2) the effects of quantitative easing during the financial crisis. We turn to these issues in turn.

5. The low interest rates of mid-2000s

In February 2005, Alan Greenspan lamented that long-term interest rates had continued to fall even though the Federal funds rate had been raised by 150 basis points to 2.5 percent. In his view, there was no obvious explanation, and he famously called this a "conundrum". In subsequent months, the Federal Reserve continued to raise the Federal funds rate, which reached 5.25 percent in July 2006 and remained at that level until July 2007. But the 10-year long-term rate did not increase as much as it had in previous tightening episodes, being offset by a sizeable decline in the term premium.[23]

Our estimates suggest that an important reason for the decline in the term premium that occurred during the early part of the 2000s might have been the shortening of the maturity of public debt in the market. Average maturity reached a peak of over 70 months in the final months of 2001; it then steadily declined to reach a trough of 56 months in March 2005, soon after Greenspan's remarks; and remained very close to an average of almost 58 months until July 2007 (Graph 1). Based on our estimates, a decline in average maturity of over 12 months is equivalent to a reduction of over 150 basis points in the five-year forward 10-year rate.

[23] In January 2002 the 10-year rate was about 5 percent and fell to about 4 percent before the Federal Reserve started tightening and then gradually rose to the 4.8 per cent until mid-2007. Over the same period, the five-year forward 10-year rate and the estimated 10-year premium had declined by around 90 basis points by the time the Federal Reserve started to raise the Federal funds rate and continued to decline thereafter. By the summer of 2007 they were respectively 1.2 and 1.5 percentage points below the level reached in January 2002.

Such a reduction matches most of the fall in both variables over the same period. This is shown in Graph 3. The red line plots the cumulative change in the forward rate since January 2002 (2002H1), when average maturity was 70 months; the dashed blue line shows the estimated contributions of average debt maturity and expected future debt.[24] The negative contribution of lower average maturity was, to some small extent, offset by the positive contribution of large expected debt.[25]

It may seem that the reduction in maturity cannot explain the Greenspan conundrum, as most of the shortening of maturity occurred before the Federal Reserve started to tighten the Federal fund rate. To the extent that maturity is an important factor determining the term premium, however, the fact that it remained relatively constant until the middle of 2007 makes it one of the factors that could explain why long-term interest rates failed to rise in the face of rising short-term rates.

One can ask why debt managers did not raise average maturity in response to the narrowing in yield spreads (or to term premia below the historical average) to minimise expected debt servicing costs. There might be at least two complementary reasons. The first is that debt had declined to relatively low levels in the early 2000s after several years of large fiscal surpluses. With low levels of debt, the refinancing risk associated with having to roll over a large amount of short-term debt at high interest rates was small. Blommestein and Turner (2012) found that, as the Federal budget deficit rises, US debt issuance tends to lengthen so that scheduled repayments can be spread over a longer period. The second reason is a strong demand for shorter (and more liquid assets) by foreign investors – especially foreign central banks – and by domestic financial institutions which needed relatively safe short-term

[24] The model uses half-yearly observations of average maturity sampled at the month in which the CBO releases its projections. Based on these data the semester in which average maturity peaked is 2002H1, which corresponds to January 2002. Looking at the entire monthly series of average maturity, the peak is reached in September 2001 at 73 months.

[25] By contrast, the measure of long-term inflation expectations (not shown) was very stable and hence contributed almost nothing to the changes in the dependent variables.

assets for their refinancing operations. With the expectation of low debt in the future, debt managers may have simply opted for accommodating the increasing demand for shorter maturities.[26]

6. The effects of QE

Since the start of the crisis several studies have attempted to assess the quantitative importance of central bank purchases of government bonds and related assets. Event studies have generally confirmed that central bank actions have had significant effects on various long-term interest rates on announcement. However, interest rates have increased in some post-announcement periods and it is unclear whether they would have been even higher had the central bank not intervened.

Research has therefore focused on disentangling the effects of central bank actions from other possible determinants of long-term interest rates. Studies using various sample periods, data and empirical strategies differ on the magnitude of these effects. For example, Krishnamurthy and Vissing-Jorgensen (2011) suggest that most purchases are likely to have worked by lowering the expectation of future short-term policy rates. Similarly, Hamilton and Wu (2012) find that the portfolio balance effects of public debt supply are relatively small. Larger effects have been found by Gagnon et al (2011), D'Amico and King (2012) and D'Amico et al (2012). This latter study in particular finds that the purchases of Treasuries by the Federal Reserve contributed to reducing the long-term rate by about 80 basis points in the first two large-scale purchasing programmes. In line with these more recent studies, our estimates also suggest that portfolio rebalance effects of public debt might have been sizeable before the start of the crisis. For that reason, they cannot be attributed to the signalling effects that central bank balance sheet policies may have had during the crisis.

[26] We re-ran our regressions ending the sample in 2001, when average maturity peaked, as well as in 1995 before outstanding debt began to decline. In both cases the estimated coefficients on average maturity are very similar to that estimated over the full sample (details available on request).

Both the magnitude and maturity of the federal debt changed significantly since QE1 was announced in November 2008. Between then and the end of 2012, marketable debt (including Federal Reserve holdings) rose 28.5 percentage points of GDP. The Federal Reserve has absorbed about 7 percentage points of this increase.[27] By buying very long-dated bonds, the Federal Reserve also lowered the average maturity of debt held outside the central bank by about 7 months.[28]

Table 5 shows how much higher the five-year ahead 10-year rate and the 10-year term premium would have been if public debt held outside the central bank had been 7 percentage points higher and average maturity 7 months longer. Specifically, the absorption of 7 percentage points of debt translates into a 12–15 basis points lower forward rate and 0–8 basis points lower term premium. A 7-month lower average maturity translates into an approximately 81–100 basis points lower forward rate and 67–89 basis points lower term premium. Combining the two effects, Fed purchases since November 2008 may have therefore contributed to lowering the five-year forward 10-year rate by approximately 90–115 basis points and the 10-year term premium by approximately 70–95 basis points. These estimates are not too far from those of D'Amico et al (2012) which, using a different approach and data, reach the conclusion that the first two large-scale purchasing programmes have reduced long-term yields by about 80 basis points (hence without including the effects of the most recent Maturity Extension Program).[29]

[27] December 2012 is the latest available observation.

[28] The average maturity of marketable debt is taken from Quarterly Refunding Report available on the US Treasury website. With the exception of the most recent figures, averages are rounded to nearest integer. Similarly, average maturity of debt held "by the public" (that is, outside the Federal Reserve) in the Table FD5 is also rounded to the nearest integer. Hence, taking the difference between reported figures may overestimate the change in mean maturity attributable to Federal Reserve actions. The average maturity of outstanding marketable debt was 65 months at the end of 2012. It has remained approximately unchanged from this level from the end of September 2012. The average maturity of debt held outside the Federal Reserve was 55 months at the end of September 2012.

[29] Unfortunately, the figures reported in the official sources are rounded to the first integer making it impossible to compute the exact difference in average maturities. We reckon that taking into account the possible largest rounding errors the difference between average maturities could be as low as 5¼ months. Assuming that the Federal Reserve contributed to alter the mean maturity of privately-held debt by only 5¼ months would give lower estimates of the overall impact: a reduction of 70–90 basis points in the five-year forward 10-year rate and a reduction of 50–75 basis points in the 10-year term premium.

7. Conclusion

In this paper we have presented evidence that changes in the maturity structure of public debt have had a statistically and economically significant effect on long-term interest rates. Such evidence is based on a pre-crisis period. Hence, it suggests that the magnitude of portfolio balance effects do not simply reflect some of the special factors that prevailed in the aftermath of the crisis. While portfolio balance effects may be stronger during or soon after a crisis, they seem to be relevant even in normal times. The evidence is also based on the five-year forward 10-year rate. Hence, it is unlikely to reflect only changes in monetary policy or signalling effects. Furthermore, the findings may also partly help explain the Greenspan conundrum and the decline in long-term interest rates prior to the financial crisis.

The recognition that portfolio balance effects are economically relevant even outside crisis period raises the important issue of coordination between central banks and sovereign debt managers. Central bank purchases of government bonds affect the volume and the maturity of bonds the market is induced to hold, and so might influence the shape of the yield curve. The decisions of government debt managers have very similar effects. The two agents of government do not, however, have the same objective for the yield curve – indeed it is primarily the central bank, not the debt manager, that focuses on the macroeconomic implications of changes in the yield curve (eg Blommestein and Turner (2012); Greenwood, Hanson, Rudolph and Summers (2014)).

Another important aspect that deserves further research is the extent to which any empirical estimates of the interest rate effects of maturity is a good guide to the future. Indeed, the link between average debt maturity and long-term rates or term premia will depend on the overall policy framework. In particular, it is likely to depend on how fiscal policy is expected to be run in the future. Given that public debt has grown to very high levels, fiscal policy may well be conducted differently than during the sample period. Different expected consolidation paths may mean that the same change in debt maturity could have a different impact on long-term rates. In addition, inflation expectations could be destabilised by radical change in debt management policy. All these issues merit further reflection.

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Annex 1: Tests of unit roots and cointegration

Our unit root tests (see Annex Table A1) support the assumption that the five-year forward 10-year rate, the constant maturity 10-year rate and inflation expectation exhibit a unit root. The three-month Treasury bill also appears to have a unit root although the evidence is less strong. As to the fiscal variables, the statistical evidence for the projected future debt-to-GDP ratio and the projected future deficit is also mixed with most tests indicating stationarity. The Augmented Dickey-Fuller tests cannot reject the null of a unit root but p values are generally lower than the interest rate variables. At the same time the Phillips-Perron test strongly rejects a unit root. The Kwiatkowski-Phillips-Schmidt-Shin (KPSS) test cannot reject the null of stationarity at the 10% significance level for projected debt, while it cannot reject the null for projected deficit at the 5% level.

Up to here, our results are in line with those of Laubach (2009), who also stress the very low power of unit root tests in small samples. One variable that is not in the original dataset is the average maturity of public debt. For this variable, one test indicates the presence of a unit root, while the others suggest stationarity.

Clearly, the sample is too short to rely solely on tests of unit roots, which have very low power. Plots of the series as well as a-priori knowledge suggest that the assumptions on which our baseline regression (2009) is based are plausible. There is also no strong reason a priori to believe that the average maturity of public debt is not a stationary variable.

We also have tested for the existence of cointegration among the basic variables that form our basic regressions. Annex Table A2 in the Annex shows no clear evidence of cointegration between the five-year forward 10-year interest rate and the measure of inflation expectation. While this may be due again to the small power of the tests in small samples, Johansen's test of cointegration cannot reject the hypothesis of cointegration (Annex Table A3). As shown by Annex Table A2, adding fiscal variables to the basic cointegration relationship does not overturn but strengthen the conclusion that the two series are cointegrated.

Annex 2: Structural break analysis

This appendix describes the procedure used to identify structural breaks in our sample. In searching for structural breaks we assume that all coefficients except that of inflation expectations in our baseline specification could be subject to a structural break at an unknown date. We rule out the coefficient on long-term inflation expectations, which is found to be cointegrated with the nominal forward rate (Annex 1).

In the presence of a structural break our baseline regression can be written in matrix form as follows:

(A2.1) $$f_t = \theta_0' x_t + \theta_1' x_t 1(t \leq t_b) + \theta_2' x_t 1(t > t_b) + \varepsilon_t$$

where the vector of regression coefficients $\theta_1$and θ_2 may include (as in our case) the intercept. The break date t_b is unknown and needs to be estimated. The vector θ_0 refers to the variables that are assumed to be fixed across regimes (in our case inflation expectations).

It is convenient to rewrite (A2.1) as:

(A2.2) $$f_t = \theta_1' x_t + \delta' x_t 1(t > t_b) + \varepsilon_t$$

We can test for the existence of a structural break in the coefficients using a Chow test for the null hypothesis that $\delta = \theta_2 - \theta_1 = 0$. If the break date is known, the distribution of this test follows a standard distribution. However, if the break date is unknown, the test distribution is non-standard because it depends on a parameter which is not identified under the null hypothesis. Andrews and Ploberger (1994) provides critical values for this test and Hansen (1997) has developed approximate p-values.

The other consequence of not knowing the break date is that we have to assume that each date in the sample (at least within a given interval) could be a possible break date. Hence, we compute a sequence of F statistics associated with different candidate break dates within the interval $t_0 + \pi_0 \leq t_b \leq t_1 - \pi_0$, where the dates t_0 and t_1 indicate the start and end dates of the sample, respectively, and

π_0 is a trimming parameter. The latter is chosen to ensure that there is a minimum number of observations in each regime and that the regression is well-behaved.

Our sample is 1976H1–2011H2 for a total of 63 observations.[30] We trim seven observations (or around 11% of available observations) both at the beginning and the end of the sample. We find that these are the minimum number of observations that could be left in both regimes without giving rise to collinearity.

The top panel in Graph A1 plots the F statistic value as a function of the candidate break date. The F statistic (computed using Newey-West standard errors) has a clear peak of 48.8886 at 1985H2. Using Hansen's approximate p-values the statistic is significant at 1% level.[31] In the presence of homoscedastic errors the peak in the F test would correspond to the trough in the residual variance. Yet the residual variance has a minimum at 1986H1. Hence we identify a break at 1986H1.

To check whether there are other breaks in the regression, we re-compute the sequence of F tests over the period 1986H1–2011H2 ensuring that there are sufficient number of observations in both regimes to avoid collinearity. We find that the F statistic tends to peak at the end of the sample and the residual variance has a global minimum in 2008H2.

Based on the structural break analysis we estimate our baseline regression over the sample 1976H1–2088H1 allowing for a break in all coefficients (bar that of inflation expectations) in 1986H1. The estimates are reported in Annex Table A4. Kejriwal and Perron (2008) show that in a cointegrated model with structural changes, which allows both stationary and integrated regressors, the limiting distribution of the estimates of the regression coefficients is the same as that obtained when the break dates are known. We therefore use a standard distribution to test the restriction shown at the bottom of Annex Table A4.

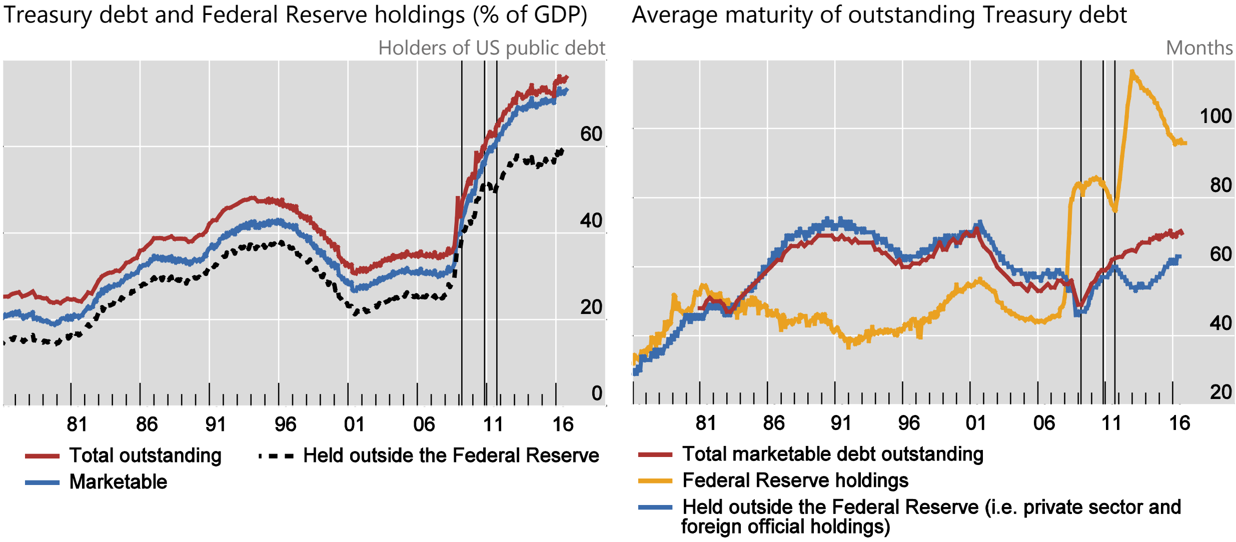
30 If all observations were semi-annual the sample should have a total of 72 observations. However, note that the sample before 1985 has a mix of annual and semi-annual observations.

31 The aveF statistic is 4.1563, which has an approximate p-value of 0.75 (with a trim parameter of 0.15). The expF statistic is 20.5525, which is significant at 1% level.

Graphs and tables

Evolution of US Federal debt

Graph 1

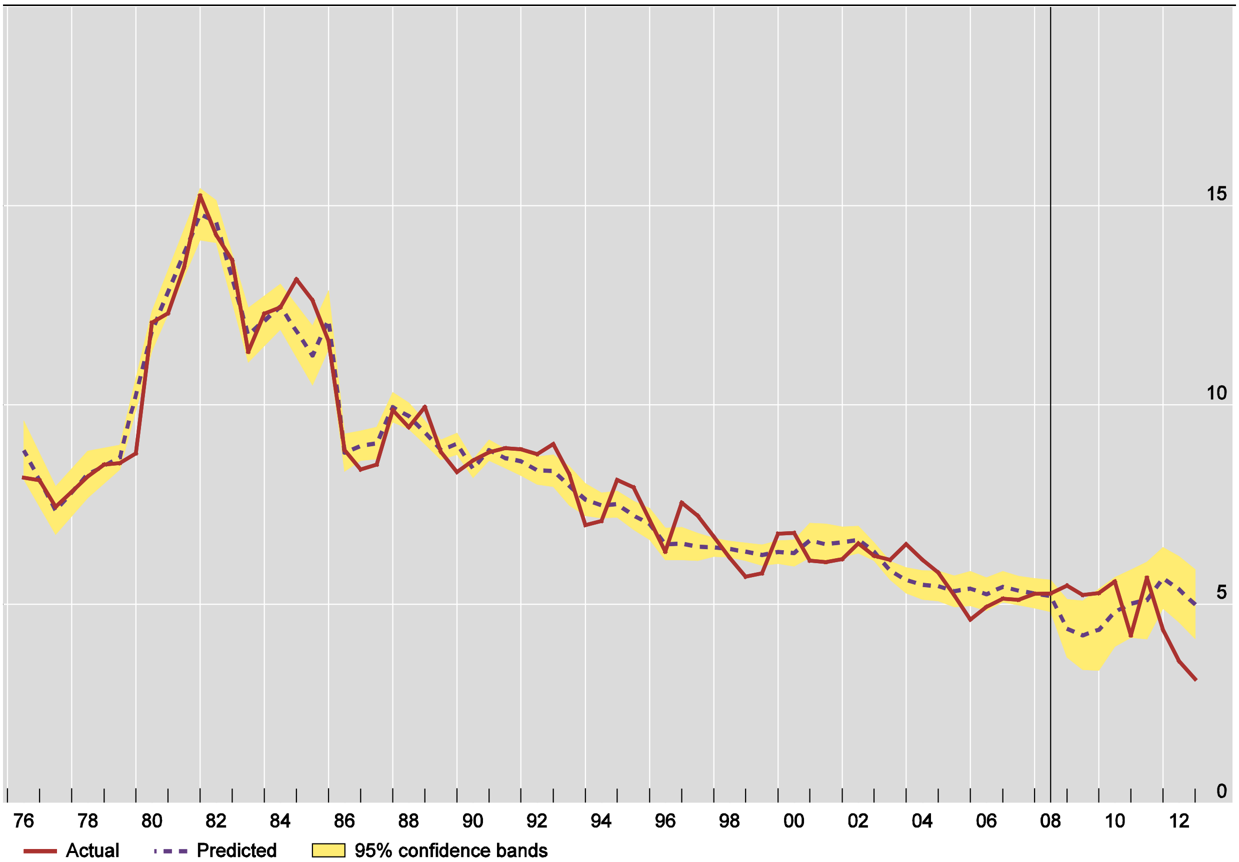


The vertical lines correspond to March 2009 (LSAP1), November 2010 (LSAP2) and September 2011 (MEP).

Sources: Datastream; US Treasury; national data; BIS calculations

Five-year forward 10-year rate: actual and predicted values[1]

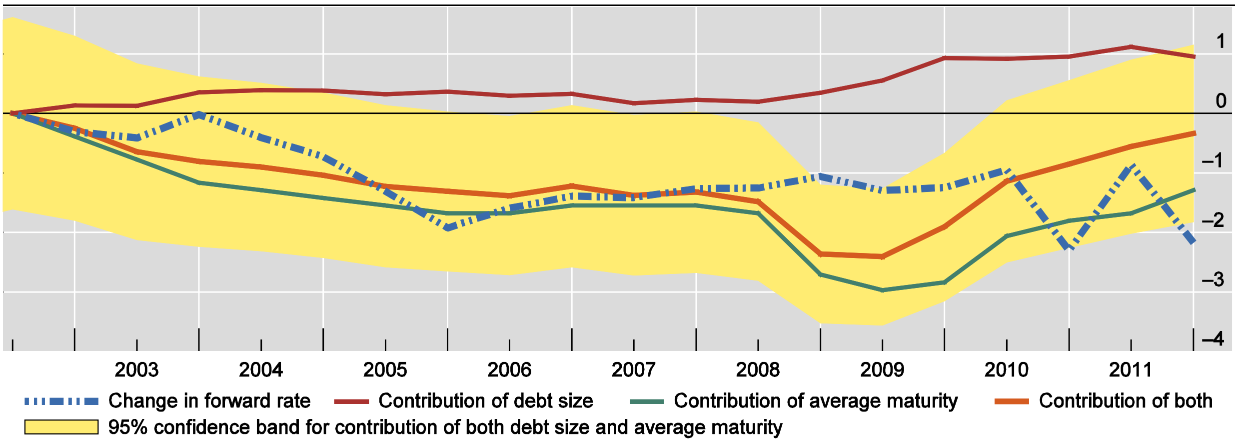
In per cent | Graph 2



[1] Predicted values are from a regression of the 5-year forward 10-year rate on average maturity of federal debt held outside the Federal Reserve and other regressors. Value to the right of the vertical line are out-of-sample predictions.

Contributions of public debt and its maturity to 5-year forward 10-year rate[1]

In percentage points | Graph 3



[1] .Predicted values from a regression of the 5-year forward 10-year rate. Cumulative change in the forward rate since January 2002 (2002H1), when average maturity was 70 months.

Five-year forward 10-year rate

Table 1

	1976H1-2008H1					1986H1-2008H1	
Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Inflation expectation	1.048***	0.999***	1.029***	1.138***	1.006***	1.018***	0.942***
	(0.070)	(0.068)	(0.082)	(0.156)	(0.132)	(0.074)	(0.087)
5-year ahead debt	0.017***	0.021***	0.017**	0.015	0.018**	0.021***	0.017**
	(0.006)	(0.005)	(0.007)	(0.010)	(0.008)	(0.005)	(0.008)
Average maturity	0.121***	0.129***	0.120***	0.132***	0.111***	0.118***	0.116***
	(0.013)	(0.012)	(0.012)	(0.015)	(0.010)	(0.016)	(0.017)
Tbill volatility (t<86H2)	2.997***	2.973***	2.296***				
	(0.250)	(0.257)	(0.442)				
Dividend yield (t<86H2)	-0.934***	-0.802***					
	(0.247)	(0.290)					
Trend growth (t<86H2)	-0.862***						
	(0.289)						
Trend growth					-0.231		-0.140
					(0.280)		(0.250)
Dividend yield					-0.019		0.110
					(0.114)		(0.091)
Tbill volatility					2.232***		0.601
					(0.450)		(0.856)
Observations	56	56	56	56	56	45	45
Adj R2	0.958	0.955	0.948	0.916	0.945	0.910	0.906

Notes: Newey-West standard errors in parentheses; *** p<0.01, ** p<0.05, * p<0.1. (t<86H2) indicates that a variable is multiplied by a dummy that takes the value of one before 1986H2 and zero thereafter. The regression includes a break dummy (t>=86H2).

Five-year forward 10-year rate: additional controls Table 2

Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Inflation expectation	0.999***	0.942***	1.007***	1.117***	0.778***	0.972***	1.139***
	(0.068)	(0.094)	(0.074)	(0.128)	(0.279)	(0.068)	(0.190)
Five-year ahead debt	0.021***	0.023***	0.028***	0.024***	0.019	0.019***	0.036***
	(0.005)	(0.005)	(0.006)	(0.005)	(0.012)	(0.005)	(0.012)
Average maturity	0.129***	0.126***	0.124***	0.143***		0.138***	
	(0.012)	(0.012)	(0.014)	(0.019)		(0.017)	
Dividend yield (t<86H2)	-0.802***	-0.834***	-0.961***	-0.828***	-0.312	-1.192***	-1.465***
	(0.290)	(0.275)	(0.239)	(0.276)	(0.621)	(0.187)	(0.344)
Tbill volatility (t<86H2)	2.973***	2.869***	3.125***	2.914***	3.174***	3.113***	3.367***
	(0.257)	(0.329)	(0.231)	(0.263)	(0.622)	(0.284)	(0.840)
Three-month bill rate		0.036					
		(0.055)					
Real-time output gap			0.049				
			(0.035)				
Fed holdings of Treasuries				0.304	-0.968**		
				(0.289)	(0.391)		
Official inflows into Treasuries						0.327	-0.522*
						(0.198)	(0.269)
Observations	56	56	53	56	56	53	53
Adj R2	0.955	0.955	0.952	0.956	0.892	0.960	0.901

Notes: Newey-West standard errors in parentheses; *** p<0.01, ** p<0.05, * p<0.1. (t<86H2) indicates that a variable is multiplied by a dummy variable that takes the value of one before 1986H2 and zero thereafter. The regression includes a break dummy (t>=86H2).

5-year forward 10-year rate: further robustness checks

Table 3

Variables	(1)	(2)	(3)	(4)	(5)	(6)
Inflation expectations	0.999***	1.017***	1.009***	1.093***	1.033***	0.945***
	(0.068)	(0.078)	(0.067)	(0.135)	(0.075)	(0.078)
Five-year ahead debt	0.021***	0.025***	0.022***	0.022***		
	(0.005)	(0.006)	(0.005)	(0.005)		
Average maturity	0.129***	0.128***	0.135***	0.131***	0.126***	0.140***
	(0.012)	(0.012)	(0.012)	(0.012)	(0.013)	(0.013)
Dividend yield (t<86H2)	-0.802***	-0.829***	-0.851***	-0.866***	-0.576**	-0.630**
	(0.290)	(0.261)	(0.271)	(0.268)	(0.277)	(0.281)
Tbill volatility (t<86H2)	2.973***	3.065***	2.966***	2.863***	2.912***	2.928***
	(0.257)	(0.248)	(0.234)	(0.269)	(0.333)	(0.263)
Unemployment rate		-0.070				
		(0.098)				
ADS index			-0.113*			
			(0.061)			
Time trend				0.009		
				(0.010)		
Current debt					0.032***	
					(0.012)	
Current deficit						0.110***
						(0.033)
Observations	56	56	54	56	56	56
Adj R2	0.955	0.955	0.954	0.955	0.950	0.952

Notes: Newey-West standard errors in parentheses; *** p<0.01, ** p<0.05, * p<0.1. (t<86H2) indicates that a variable is multiplied by a dummy variable that takes the value of one before 1986H2 and zero thereafter. The regression includes a break dummy (t>=86H2). The ADS index is an index of business cycle conditions.

10-year term premium 1990H1 - 2008H1 Table 4

Variables	(1)	(2)	(3)	(4)	(5)	(6)
Inflation expectation	0.148	0.139	0.524**	0.024	0.063	0.426**
	(0.346)	(0.303)	(0.235)	(0.315)	(0.225)	(0.200)
Five-year ahead debt	0.012	0.011	0.010			
	(0.008)	(0.007)	(0.007)			
Average maturity	0.115***	0.117***	0.096***	0.126***	0.127***	0.106***
	(0.026)	(0.023)	(0.020)	(0.024)	(0.020)	(0.019)
Trend growth	0.028			-0.064		
	(0.264)			(0.224)		
Dividend yield	0.212	0.220**		0.232*	0.213**	
	(0.133)	(0.102)		(0.121)	(0.093)	
Tbill volatility	0.418			0.407		
	(0.761)			(0.716)		
Five-year ahead deficit				0.093**	0.092***	0.090**
				(0.034)	(0.033)	(0.041)
Observations	37	37	37	37	37	37
Adj R2	0.834	0.843	0.835	0.844	0.853	0.844

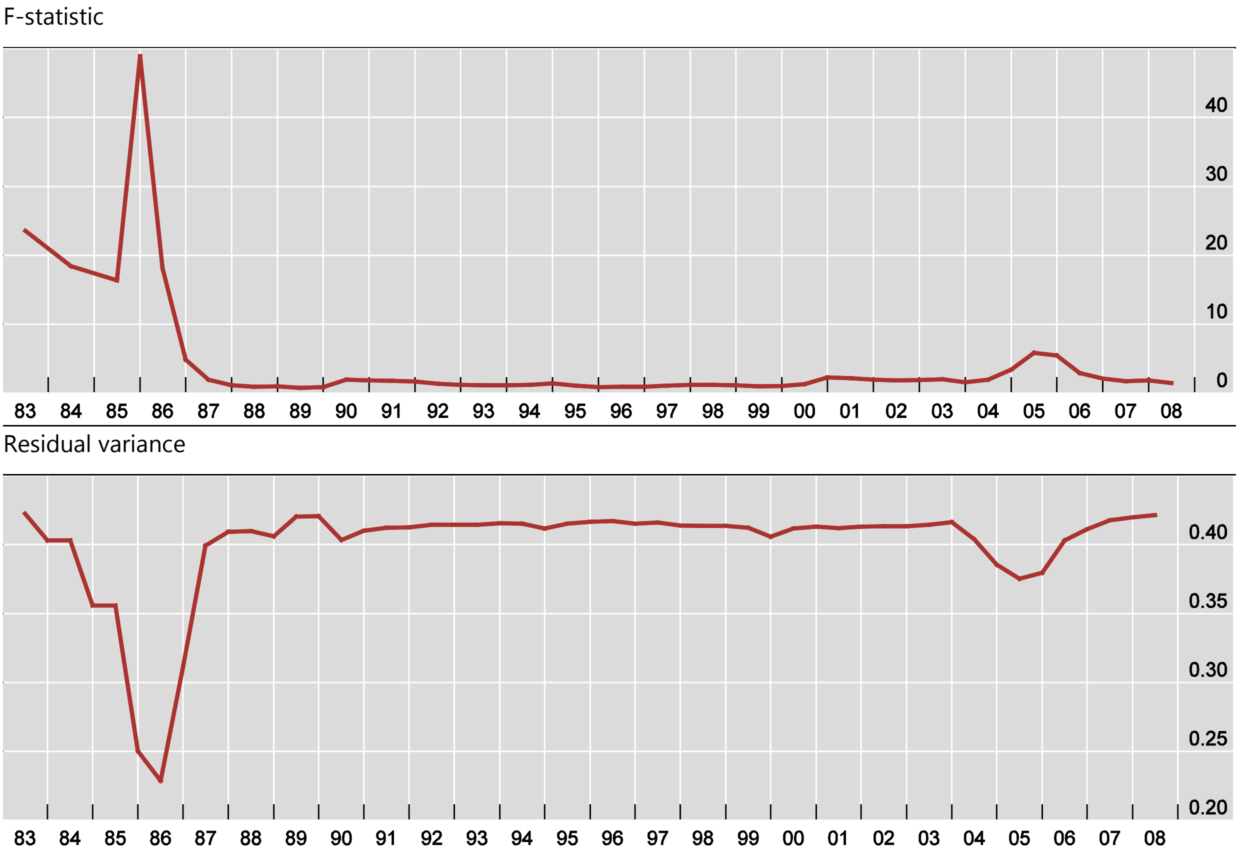
Notes: Newey-West standard errors in parentheses; *** p<0.01, ** p<0.05, * p<0.1.

Potential effects of central bank purchases of Treasuries since November 2008									Table 5
		5y forward 10y rate				10y term premium			
	Change	Marginal effect (range)		Total effect (range)		Marginal effect (range)		Total effect (range)	
Debt held outside the central bank (% of GDP)	7	*1.7*	*2.1*	12	15	*0*	*1.2*	0	8
Average maturity (months)	7	*11.6*	*14.3*	81	100	*9.6*	*12.7*	67	89
Total effect (bps)				93	115			67	97

Notes: Change in the first column refers to changes in debt held outside the Federal Reserve which could be attributed to central bank interventions since November 2008. The range is selected by taking the min and max estimated coefficients in Table 1-2 (forward rate) and Table 3 (term premium).

Structural break tests

Graph A1



Source: Authors' calculations.

Unit root tests

Annex Table A1

Variables	Test	Lag	Statistic	Approx. p	1% CV	5% CV	10% CV	T
5-year forward 10-year rate	ADF	1	-1.33	0.61	-3.56	-2.92	-2.60	62
	PP	3	-1.21	0.67	-3.56	-2.92	-2.60	63
	KPSS	3	1.12	--	0.74	0.46	0.35	63
Const. maturity 10-year yield	ADF	0	-1.38	0.59	-3.56	-2.92	-2.60	63
	PP	3	-1.35	0.61	-3.56	-2.92	-2.60	63
	KPSS	3	1.18	--	0.74	0.46	0.35	63
Inflation expectation	ADF	4	-0.98	0.76	-3.57	-2.92	-2.60	59
	PP	3	-0.61	0.87	-3.56	-2.92	-2.60	63
	KPSS	3	1.48	--	0.74	0.46	0.35	63
3-month treasury bill rate	ADF	2	-2.00	0.29	-3.57	-2.92	-2.60	61
	PP	3	-1.96	0.31	-3.562	-2.92	-2.595	63
	KPSS	3	0.95	--	0.74	0.46	0.35	63
Projected debt	ADF	2	-1.95	0.31	-3.63	-2.95	-2.61	43
	PP	3	-3.06	0.03	-3.61	-2.94	-2.61	46
	KPSS	3	0.45	--	0.74	0.46	0.35	46
Projected deficit	ADF	2	-1.75	0.40	-3.63	-2.95	-2.61	43
	PP	3	-4.00	0.00	-3.61	-2.94	-2.61	46
	KPSS	3	0.48	--	0.74	0.46	0.35	46
Average maturity	ADF	4	-3.13	0.02	-3.64	-2.96	-2.61	41
	PP	3	-1.87	0.35	-3.61	-2.94	-2.61	46
	KPSS	3	0.43	--	0.74	0.46	0.35	46

Notes: ADF is the Augmented Dickey Fuller test for the null of a unit root against the alternative of stationarity (no trend); PP is the Perron Philips test for the null of a unit root (no trend); and KPSS is the Kwiatkowski-Phillips-Schmidt-Shin test for the null of level stationarity. Approx. p-value is the MacKinnon's approximate p value of the computed statistic. Observations are half yearly and start in 1976. However, because there are gaps for some series in the early part of the sample, the test sample will start at a later date. T shows the number of observations available.

Cointegration tests

Annex Table A2

Variables	Test	Lag	Stat	1% CV	5% CV	10% CV	T	N
	ADF	1	-3.23	-4.08	-3.44	-3.11	62	2
5-yr forward 10-yr rate,	PP	3	-3.07	-4.08	-3.43	-3.11	63	2
inflation exp	KPSS	3	0.20	0.74	0.46	0.35	63	-
5-yr forward 10-yr rate,	ADF	1	-3.88	-4.16	-3.48	-3.14	44	2
inflation exp,	PP	3	-4.46	-4.15	-3.47	-3.14	46	2
5-yr ahead debt	KPSS	3	0.11	0.74	0.46	0.35	46	-
5-yr forward 10-yr rate,	ADF	1	-3.58	-4.16	-3.48	-3.14	44	2
inflation exp,	PP	3	-4.70	-4.15	-3.47	-3.14	46	2
5-yr ahead deficit	KPSS	3	0.14	0.74	0.46	0.35	46	-
5-yr forward 10-yr rate,	ADF	1	-4.50	-4.16	-3.48	-3.14	44	2
inflation exp, 5-yr ahead	PP	3	-4.69	-4.15	-3.47	-3.14	46	2
debt, avg maturity	KPSS	3	0.17	0.74	0.46	0.35	46	-
5-yr forward 10-yr rate,	ADF	1	-4.24	-4.16	-3.48	-3.14	44	2
inflation exp, 5-yr ahead	PP	3	-4.85	-4.15	-3.47	-3.14	46	2
deficit, average maturity	KPSS	3	0.18	0.74	0.46	0.35	46	
5-yr forward 10-yr rate,	ADF	1	-4.19	-4.16	-3.48	-3.14	44	2
inflation exp, trend	PP	3	-4.68	-4.15	-3.47	-3.14	46	2
growth, dividend yield	KPSS	3	0.08	0.74	0.46	0.35	46	-

Notes: This table reports residual-based tests of cointegration. For the Augmented Dickey Fuller (ADF) and the Phillip Perron (PP) tests, the critical values tabulated from the response surfaces tabulated by MacKinnon (2010) for the no trend case. T indicates the number of observations used to compute the test statistic. N indicates the number of I(1) variables in the estimated cointegrating relationship. The null hypothesis in the ADF and PP tests is the presence of a unit root in the regression. The Kwiatkowski-Phillips-Schmidt-Shin (KPSS) test tests for the null of level stationarity in the residuals. Observations are halfyearly and start in 1976. However, because there are gaps for some variables in the early part of the sample, the estimation sample starts at a later date.

Johansen's tests of cointegration – 5-year forward 10-year rate and inflation expectation

Annex Table A3

No of CE(s)	Eigenvalue	Trace stat	5% CV
None	0.292	21.97	15.49
At most 1	0.020	1.21	3.84

No of CE(s)	Eigenvalue	Max-Eigenvalue stat	5% CV
None	0.292	20.75	14.26
At most 1	0.020	1.21	3.84

Notes: Estimated VAR for differences has 3 lags. Number of observations is T=60.

5-year 10-year interest rate, structural break model		Annex Table A4
Variables	(1)	(2)
Inflation expectation	0.962***	1.048***
	(0.073)	(0.070)
Five-year ahead debt (t<86H2)	-0.008	
	(0.016)	
Five-year ahead debt (debt5) (t>=86H2)	0.017**	
	(0.008)	
Average maturity (t<86H2)	0.150***	
	(0.031)	
Average maturity (t>=86H2)	0.115***	
	(0.017)	
Trend growth (t<86H2)	-1.201***	-0.862***
	(0.224)	(0.289)
Trend growth (t>=86H2)	-0.144	
	(0.260)	
Dividend yield (t<86H2)	-0.849***	-0.934***
	(0.146)	(0.247)
Dividend yield (t>=86H2)	0.101	
	(0.090)	
Tbill volatility (t<86H2)	2.873***	2.997***
	(0.204)	(0.250)
Tbill volatility (t>=86H2)	0.529	
	(0.929)	
Intercept break (t>=86H2)	-10.723***	-10.524***
	(2.074)	(1.408)
Five-year ahead debt		0.017***
		(0.006)
Average maturity		0.121***
		(0.013)
Intercept	6.754***	5.992***
	(1.406)	(1.624)
Observations	56	56
Adj R2	0.956	0.958

Test:
1) debt5_l - debt5_r = 0; 2) avgmatfd5_l - avgmatfd5_r = 0; 3) gr5_r = 0
4) dwr_r = 0; 5) volm3_12m_r = 0
F(5, 43) = 1.19
Prob > F = 0.3296

Notes: Standard errors in parentheses; *** p<0.01, ** p<0.05, * p<0.1.

Warnock and Warnock (2009) model of the nominal 10-year Treasury yield

Annex Table A5

	January 1984 - May 2005						January 1984 - May 2005					
	(1)		(2)		(3)		(4)		(5)		(6)	
Variables	Coeff	t-stat	Coeff	t-stat	Coeff	t-stat	Coeff	t-stat	Coeff	t-stat	Coeff	t-stat
10-year ahead inflation expectations minus 3m eurodollar rate	0.628 ***	19.32	0.659 ***	19.47	0.725 ***	23.88	0.629 ***	18.97	0.636 ***	18.55	0.697 ***	22.71
1-year ahead minus 10-year ahead inflation expectation	0.431 **	2.02	0.613 ***	2.75	0.770 ***	3.59	0.486 **	2.24	0.538 **	2.36	0.675 ***	3.01
interest rate risk premium	4.351 ***	6.15	2.839 ***	3.62	6.076 ***	8.21	4.672 ***	6.74	4.292 ***	5.92	7.445 ***	11.07
1-year head real gdp growth	0.145 **	1.98	0.023	0.30	0.183 **	2.17	0.137 **	1.97	0.096	1.32	0.250 ***	2.95
structural budget balance (% of lagged GDP)	-0.137 ***	-4.52	-0.139 ***	-4.69			-0.134 ***	-4.51	-0.134 ***	-4.48		
12-month aggregate foreign flows into the US treasuries	-0.188 ***	-5.70	-0.313 ***	-5.96	-0.321 ***	-6.33						
12-month aggregate official foreign flows into the US treasuries							-0.399 ***	-6.00	-0.454 ***	-5.47	-0.451 ***	-5.30
average maturity of federal debt held by the public (months)			-0.046 ***	-3.27	-0.039 ***	-2.83			-0.014	-1.21	-0.006	-0.54
federal debt held by the public (% of GDP)					0.057 ***	6.87					0.054 ***	6.18
R-squared	0.76		0.77		0.79		0.76		0.76		0.78	
No of observations	257		257		257		257		257		257	

	August 1987 - May 2005						February 1994 - May 2005					
	(7)		(8)		(9)		(10)		(11)		(12)	
Variables	Coeff	t-stat	Coeff	t-stat	Coeff	t-stat	Coeff	t-stat	Coeff	t-stat	Coeff	t-stat
10-year ahead inflation expectations minus 3m eurodollar rate	0.614 ***	20.25	0.596 ***	22.07	0.707 ***	24.83	0.608 ***	16.24	0.571 ***	16.07	0.728 ***	18.59
1-year ahead minus 10-year ahead inflation expectation	-0.091	-0.50	-0.341 **	-2.03	-0.001	-0.01	0.074	0.28	-0.134	-0.53	-0.256	-1.08
interest rate risk premium	2.406 ***	3.57	2.419 ***	3.79	6.738 ***	12.41	13.196 ***	3.80	17.055 ***	5.32	10.956 ***	3.36
1-year head real gdp growth	-0.018	-0.30	0.049	0.82	0.216 ***	3.04	-0.078	-0.96	0.078	1.00	0.462 ***	5.02
structural budget balance (% of lagged GDP)	-0.216 ***	-8.21	-0.206 ***	-8.91			-0.215 ***	-4.30	-0.355 ***	-8.19		
12-month aggregate foreign flows into the US treasuries	-0.194 ***	-6.54	0.076	1.26	0.063	1.11	-0.248 ***	-3.83	0.044	0.64	0.070	1.11
12-month aggregate official foreign flows into the US treasuries												
average maturity of federal debt held by the public (months)			0.088 ***	4.88	0.093 ***	5.59			0.158 ***	6.94	0.125 ***	5.92
federal debt held by the public (% of GDP)					0.062 ***	8.38					0.094 ***	8.54
R-squared	0.87		0.89		0.90		0.86		0.89		0.91	
No of observations	214		214		214		136		136		136	

This table presents OLS estimates of the empirical model of the nominal 10-year Treasury yield in Table 1 in Warnock and Warnock (2009). The dependent variable is the nominal 10-year rate minus the the 3-month eurodollar rate. This allows to estimate directly all parameters but that on the 3-month rate without imposing a unit restriction on the sum of the coefficients on the 10-year inflation expectation and the 3-month rate. t-statistics are computed using standard errors that are robust to heteroskedasticity and serial correlation. ***, ** and * indicate significance at 1, 5 and 10 percent levels, respectively. Constants included but not reported. The sample is of monthly frequency. Columns (1), (4), (7) and (10) replicate columns (1), (2), (5) and (6) of Table 1 in Warnock and Warnock (2009). Columns (2), (5), (8) and (11) include average maturity as an additional control. Columns (3), (6), (9) and (12) additionally include debt held by the public in place of the structural fiscal balance.