

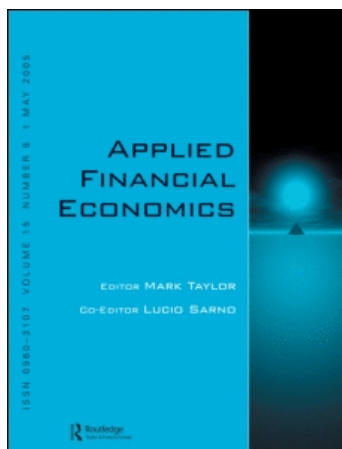
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Explaining the US bond yield conundrum

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We analyse if and to what extent fundamental macroeconomic factors, temporary influences or more structural factors have contributed to the low levels of US bond yields over the last few years. For that purpose, we start with a general model of interest rate determination. The empirical part consists of a cointegration analysis with an error-correction mechanism. We are able to establish a stable long-run relationship and find that the behaviour of bond yields, even during the last years, can be well explained by macroeconomic and structural factors. Alongside the more traditional determinants like core inflation, monetary policy and the business cycle, we also include foreign holdings of US Treasuries. The latter should capture the frequently mentioned structural effects on long-term interest rates. Finally, our bond yield equation outperforms a random walk model in different forecasting exercises.

I. Introduction

Long-term interest rates in Europe and in the US fell to all-time lows in the last few years. And despite temporary ups and downs, they still have been trading at historically low levels in 2006 and 2007, especially in the US. In his February 2005 testimony before the Committee on Banking, Housing and Urban Affairs of the US Senate, former Fed Chairman Alan Greenspan (2005) asserted: 'For the moment, the broadly unanticipated behaviour of world bond markets remains a conundrum. Bond price movements may be a short-term aberration, but it will be some time before we are able to better judge the forces underlying recent experience.' In the monthly report of April 2005, the European Central Bank (ECB) also stated that macroeconomic fundamental factors alone cannot explain the development

of long-term interest rates and pointed to structural factors that are behind recent bond market developments. 'A number of changes in the regulatory environment for pension funds and life insurance corporations appear to be under way in the euro area and the US, which aim to reduce the problems of mismatches between the duration of their assets and liabilities. It is generally perceived that these regulatory changes will favour the purchase of bonds over other asset classes by pension funds and life insurance corporations' (ECB, 2005, p. 23). As a result of these changes and anticipatory effects of the proposed legislation, there may have been an increase in the structural demand for bonds of longer maturities from institutional investors, which contributed to a bullish market.

While some of these more structural factors point to a possible permanent change in long-term real

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interest rates, there are hints that other, more temporary market influences related to speculative behaviour, may have played a role as well. The alleged widespread use of carry trades – borrowing at low short-term interest rates and investing in higher yielding, longer-term maturities – appears to exploit market trends, and thus may have amplified the downturn in long-term interest rates. Speculative flows of this sort, however, are likely to be reversed at some point and hence should not have a permanent effect on the level of long-term interest rates. Furthermore, as Bernanke *et al.* (2004) pointed out, the massive purchases of government bonds by Asian central banks probably have had a significant impact on long-term bond yields in the US. According to the ECB (2006), quantitative estimates about the yield impact of foreign reserve accumulation ranges from 30 to 200 basis points.

Over and above that, recent empirical papers find tentative evidence that structural factors are at work. Clostermann and Seitz (2005) construct a traditional US-bond yield model driven by monetary policy, the business cycle and inflation expectations. Although the out-of-sample performance of their model was good, they conclude that, ‘there are hints of some instability in the last years.’ Kozicki and Sellon (2005, p. 29) suggest ‘that the key factor behind the conundrum is a large reduction in the term premium’ (see also Kim and Wright, 2005; Backus and Wright, 2007), whereas Taboga (2007) counters this view by arguing that a decline in the real natural rate of interest and a structural reduction in inflation expectations are more important. Bandholz (2006) shows that the unexplained part of his US bond model especially increases to values not seen in the past when data for 2006 are included. Rudebusch *et al.* (2006) confirm this result on the basis of empirical no-arbitrage macro-finance models of the term structure when they state ‘that the recent behaviour of long-term yields has been unusual – that is, it cannot be explained within the framework of the models.’

On the other hand, Fels and Pradhan (2006) reject the existence of a conundrum. They state that ‘a drop of bond yields below their fair value such as the one seen last year did not represent a break with past pattern and, as such, did not require a new paradigm to explain it. In fact, our statistical tests suggest that the relationship between bond yields and our three

fundamental factors (i.e. the real federal funds rate, inflation expectations and inflation volatility, BCS) did not change significantly in recent years. And, as in previous episodes of overvaluation in the bond market, actual bond yields eventually corrected since the fall of 2005, rising towards their fundamental fair value.’ Thornton’s (2007) analysis also suggests that there was no break in the relationship between short- and long-term interest rates. He argues that the change in this relationship is due to the Fed having changed its policy to targeting the funds rate in the late 1980s. That is, the change in the relationship between the funds rate and long-term rates is an example of Goodhart’s Law.

To find out whether fundamental macroeconomic, temporary or more structural factors have been at work or whether there is even no conundrum at all, we proceed as in Clostermann and Seitz (2005). First, we discuss theoretically, which fundamentals should determine bond yields. These are essentially the three macroeconomic factors identified by Dewachter and Lyrio (2006) and Diebold *et al.* (2006): monetary policy, inflation expectations and the business cycle.¹ These variables are augmented by foreign holdings of US Treasuries, a structural factor that is essential in the context of explaining the yield developments in 2005/06.² In the third section, we estimate an interest rate model for 10-year (10Y) US Treasury notes and analyse whether there are hints of unexplained interest rate developments and of overvaluations of the bond market in recent years. In doing so, we also derive a ‘fair value’ for the yield of 10Y Treasuries, which we compare to actual developments to get an idea of the magnitude of the evolving disequilibrium. This helps to answer the question whether the bond market overvaluation from 2004 to 2006 has been unusually strong in a historical context. Furthermore, we perform some out-of-sample forecasting exercises of our preferred model and compare it to a random walk model.

The existing empirical literature approaches the problem of bond yield determination in four different ways. The first strand of literature looks for fundamental factors as explanatory variables (see, e.g. Mehra, 1995; Brooke *et al.*, 2000; Caporale and Williams, 2002; Durré and Giot, 2005). The second approach uses high-frequency (in most cases daily) data to analyse the reaction of yields to news or announcements (see, e.g. Demiralp and Jordà, 2004;

¹ The first three variables are also the main determinants of nominal yields on 1-year Treasury bills in Mehra (1995) in a sample from 1955 to 1994.

² This is in line with Warnock and Warnock (2005) and consistent with the savings-glut-hypothesis (see, e.g. Tatom, 2007). In contrast to the Warnock–Warnock article, we do not need this variable before 2005, which means that the traditional macrofactors do a good job in capturing the evolution of US bond yields until mid-2005.

Monticini and Vaciago, 2005). The third kind of models discusses the international transmission of shocks with respect to bond markets (see, e.g. Ehrmann *et al.*, 2005). And, finally, the fourth approach combines bond yield modelling strategies from a finance and macroeconomic perspective to get a comprehensive understanding of the whole term structure of interest rates (e.g. Dewachter and Lyrio, 2006; Diebold *et al.*, 2006). Our view is a synthesis of especially one and three, but also partly borrows from four.

II. What Determines Interest Rates? Some Theory

Generally, interest rates should be determined by the supply of and the demand for loanable funds and their determinants including the production opportunities in the economy, the rate of time preference, risk aversion and the relative returns of alternative investments. Ideally, this would necessitate a dynamic and stochastic general equilibrium model of the economy with supply and demand conditions derived from first principles.³ So far, however, dynamic stochastic general equilibrium models with an elaborated financial sector are still in their infancy.

Therefore, and in line with other studies, our analysis starts with a general model for the term structure of interest rates:

$$r_t^l = r_t^s + rp(l, c_t) \quad (1)$$

where r_t^l is the real long-term rate, r_t^s is the real short-term rate, l and s denote the terms of the bonds, c_t is a set of variables that influences investors' risk attitudes and rp is the function defining that influence which gives us the term or risk premium in r_t^l (Caporale and Williams, 2002, p. 121).⁴

To make Equation 1 suitable for empirical work, we need information on the specifics of rp and the partial derivatives of r^l with respect to $r^s(\beta)$ and $rp(\gamma)$. Following Breedon *et al.* (1999), Caporale and Williams (2002) and others, c_t is a catchall variable for risks arising from macroeconomic policy developments. Specifically, we define

$$r_t^l = \beta r_t^s + \gamma rp(l, bc_t, etc_t) \quad (2)$$

where bc is a variable capturing the state of the business cycle. In 'etc' different further factors influencing the macroeconomic environment could

be subsumed. In this direction, Caporale and Williams (2002) as well as Paesani *et al.* (2006) analyse the fiscal position. Jordá and Salyer (2003) ask whether the liquidity situation helps to explain bond yields (see also ECB, 2005, p. 23). Durré and Giot (2005) investigate whether stock market variables are responsible for bond market developments. We decide to include an indicator variable, which has already been considered in the past, but in a different way than we do (Frey and Moët, 2005; Warnock and Warnock, 2005; Wu, 2005). It captures structural demand by foreigners for US Treasuries (d). A more detailed description and discussion is provided in section 'The data'.

Equations 1 and 2 are specified in real terms. Two problems arise in this context (Caporale and Williams, 2002, p. 122). First, real rates are not directly observable but have to be proxied for empirical work. Second, the strength of the effect of expected inflation on nominal long-term rates (i^l) is ambiguous. It might be a one-to-one relationship if the Fisher effect holds. This is the case in all models in which the real interest rate does not depend on monetary variables and monetary neutrality holds. It is violated, however, in models where an increase in expected inflation lowers the real interest rate (e.g. Tobin, 1965). Even a greater than one-to-one relationship is possible as in Tanzi (1976). Therefore, we modify Equation 2 and leave the exact response of i^l to expected inflation open.

$$i_t^l = \beta_1 i_t^s + \beta_2 \pi_t^e + \gamma rp(l, bc_t, d_t) \quad (3)$$

where $i^l(i^s)$ is the nominal long-term (short-term) interest rate and π^e is expected inflation.

Equation 3 has several testable economic implications and allows the testing of various hypotheses. For example, the pure expectations hypothesis would imply $\gamma = \beta_2 = 0$ and $\beta_1 = 1$. If expected inflation and interest rates are stationary and the Fisher effect holds either for the long-term or the short-term interest rate, $(\beta_1 + \beta_2) = 1$. If there is an exogenous rise in demand for US bonds (d), γ_d would be negative. Finally, the coefficient on bc may be positive or negative depending on whether the supply of or the demand for bonds changes more with altered business cycle conditions.

This framework allows us to test empirically whether macroeconomic factors and/or structural factors and/or temporary factors are important determinants of interest rates. However, proper inference can only be drawn within an appropriate

³ See for a prototype model in this spirit Christiano *et al.* (2005).

⁴ Equation 1 already shows that, if economic surprises are minimal and there are no reasons to revise expected future short-term rates, then there should be no trend in long-term rates (see also Poole, 2005).

econometric framework. This will be discussed in the next section.

III. Estimation

The data

In what follows, we estimate an equation for yields of 10Y US Treasuries from the mid-1980s until mid-2006. Thus, we concentrate mainly on the Greenspan era. On the right-hand side, we distinguish between long-run influences and determinants of short-run dynamics. This split is done by economic reasoning and unit root tests. The short-term interest rate is the 3-month money market rate. Both interest rates are end-of-month data. End-of-month data have the advantage of incorporating all information of the respective month and, in contrast to monthly averages, do not introduce smoothness into the data, which in turn leads to autocorrelation in the residuals (Gujarati, 1995, p. 405). Time series of the two interest rates are shown in Fig. 1.

We measure inflation expectations by using core inflation, i.e. the annual change of headline Consumer Price Index (CPI) excluding food and energy prices, to capture the underlying price trend (Fig. 2).⁵ As a measure for the state of the business cycle, we use the Institute for Supply Management's manufacturing index (*ism*, see Fig. 3). It has the advantage (and this is especially important for forecasting exercises) of not being revised and of being available with only a short publication lag.

Our '*d*'-variable captures structural factors.⁶ As mentioned above, higher foreign demand for US Treasuries, due to (i) demand from Asian central banks, (ii) the recycling of petrodollars, (iii) the strong interest of institutional investors and (iv) liquidity-driven demand due to world-wide expansionary monetary policies could be responsible for the low level of US bond yields during the last years. To quantify the influence of these factors, we include official and private foreign holdings of US Treasuries ('Treasury Securities') in percent of overall

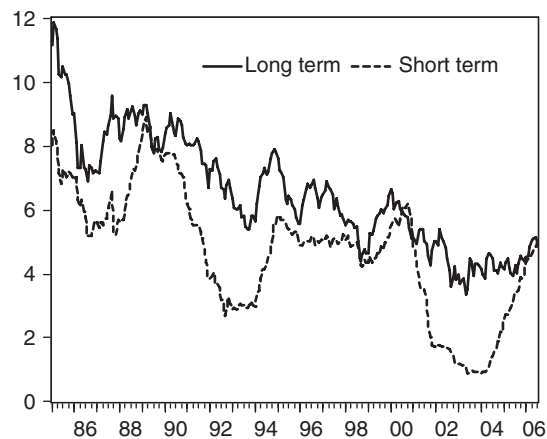


Fig. 1. Long-term and short-term interest rates

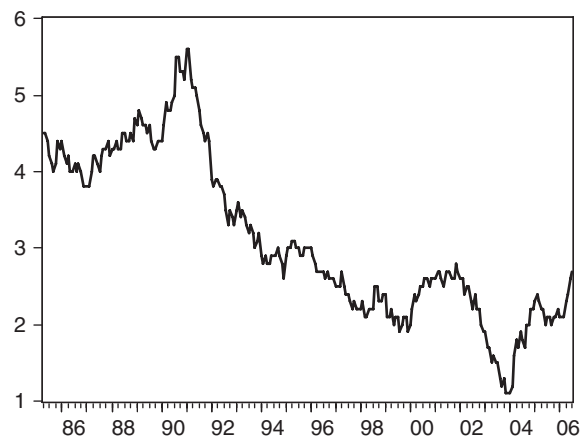


Fig. 2. Core inflation

federal debt ('total liabilities').⁷ Figure 4 shows that, since the beginning of the Japanese FX market intervention in 2002, the external debt of the US Treasury has increased considerably. Overall, the volume of Treasuries held by foreigners nearly doubled between 2002 and 2006 from USD 1100 billion to USD 2000 billion. This is equivalent to about 35% of Federal Government's total liabilities.

Our sample of monthly data runs from 1986:1 to 2006:6. The business cycle variable *ism* is in

⁵ We get slightly worse statistical results with the headline CPI measure. An alternative to our preferred measure of inflation expectations would be the difference between conventional and inflation-indexed bonds (TIPS). However, as the first TIPS have only been issued by the US Treasury in the late 90s, their use would significantly shorten our sample.

⁶ We tried several other '*etc*'-variables (e.g. the public debt and deficit situation, liquidity measures, stock market variables) which do not help to explain bond yields. Mehra (1995) also finds that fiscal policy measures do not affect bond yields once one controls for the effects of inflation expectations, monetary policy and real growth. In contrast, Paesani *et al.* (2006, p. 4), who disregard output developments, conclude for Germany, Italy and the US that 'a more sustained debt accumulation leads at least temporarily to higher long-term interest rates.'

⁷ Wu (2005) shows that it is not convincing to concentrate only on purchases of US Treasury securities by foreign central banks.

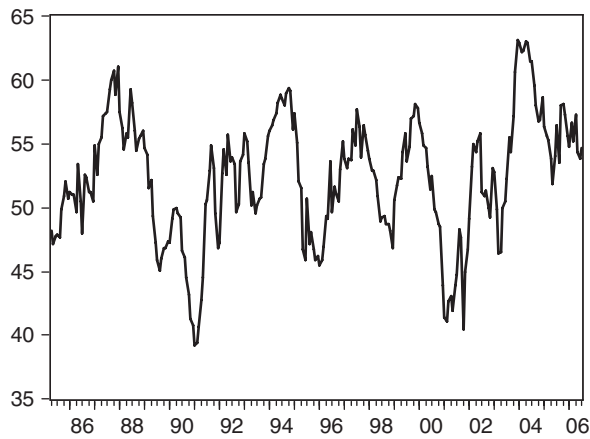
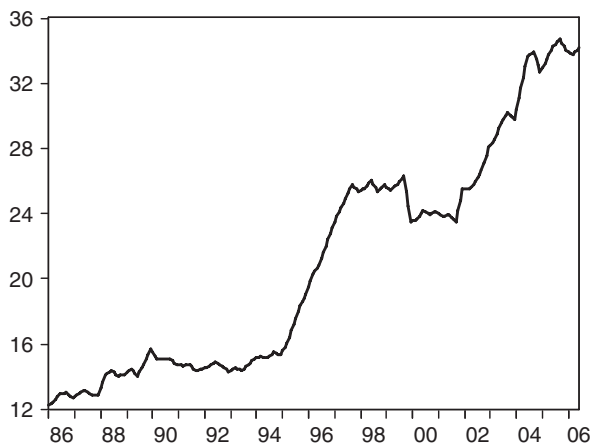
Fig. 3. *ism* index

Fig. 4. Foreign holdings of US Treasuries in percentage of federal debt outstanding

logarithms and the difference operator Δ refers to first (monthly) differences.⁸

Econometric analysis

Standard unit root tests suggest that most of our variables are $I(1)$ in levels and stationary in first differences.⁹ The only exception is the '*ism*' index, which (in line with theoretical considerations) is identified as a stationary variable. Owing to the nonstationarity of the time series, the nominal long-term yield is estimated within a Vector Error

Correction Model (VECM) based on the procedure developed by Johansen (1995, 2000). This approach seems to be particularly suited to verify the long-term equilibrium (cointegration) relationships on which the theoretical considerations are based.¹⁰ The empirical analysis starts with an unrestricted VECM, which takes the following form:

$$\Delta y_t = \Pi y_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta y_{t-i} + \Psi x_t + \eta + \varepsilon_t \quad (4)$$

where y_t represents the vector of the nonstationary variables i^l , i^s , Π^e and d . ε denotes the vector of the independently and identically distributed residuals, Ψ is the coefficient matrix of exogenous variables (x_t) and η the vector of constants. The number of cointegration relationships corresponds to the rank of the matrix π . Granger's representation theorem asserts that if the coefficient matrix π has reduced rank $r < n$, then there exist $(n \times r)$ matrices α (the loading coefficients or adjustment parameters) and β (the cointegrating vectors) each with rank r (number of cointegration relations) such that $\pi = \alpha\beta'$ and $\beta'y_t$ is $I(0)$. The cointegration vectors represent the long-term equilibrium relationships of the system. The loading coefficients denote the importance of these cointegration relationships in the individual equations and the speed of adjustment following deviations from long-term equilibrium.

The lag order (k) of the system is determined by estimating an unrestricted vector autoregression (VAR) model in levels and using the information criteria suggested by Schwarz (SC) and Hannan–Quinn (HQ). All criteria recommend a lag length of 2 (Table 1). The number of cointegration vectors is verified by determining the cointegration rank with the trace-test and the max-eigenvalue-test. Both tests suggest one cointegration relationship, i.e. one equilibrium relationship between the nonstationary variables i^l , i^s , Π^e and d . (Table 2).¹¹

Therefore it seems reasonable to restrict the VECM to one cointegration relationship and, as the above mentioned unit root tests suggest, to include the indicator for the expected stance of the business cycle '*ism*' as a stationary (nonmodelled exogenous) variable (with a lag length of 0 to 1) into the system.

⁸ All data are available upon request and can alternatively be downloaded at: http://freenet-homepage.de/clostermann/data_us_bonds.xls

⁹ Test results in detail are available from the authors upon request.

¹⁰ A similar approach is used by Lange (2005) in the case of Canada. Rao (2007) compares our econometric method with others to distinguish between short- and long-run relationships. He finds that often there are only minor differences in the estimates.

¹¹ According to the theory put forward in Section II, three cointegration relations are possible: one between the short and the long rate, one between the long rate and inflation expectations and one which has to be interpreted as a bond rate equation.

Table 1. Lag length tests

Lag	SC	HQ
0	13.62868	13.59482
1	-0.93903	-1.10831
2	-1.35017	-1.65487
3	-1.11267	-1.55279
4	-0.92327	-1.49882
5	-0.68674	-1.39771
6	-0.41889	-1.26528
7	-0.14877	-1.13059
8	-0.01870	-1.13594

Hence, a VECM with the following structure is estimated:

$$\begin{pmatrix} i_t^l \\ i_t^s \\ \pi_t^e \\ d_t \end{pmatrix} = \Gamma_1 \begin{pmatrix} \Delta i_{t-1}^l \\ \Delta i_{t-1}^s \\ \Delta \pi_{t-1}^e \\ \Delta d_{t-1} \end{pmatrix} + \begin{pmatrix} \alpha^{il} \\ \alpha^{is} \\ \alpha^\pi \\ \alpha^d \end{pmatrix} (1 \quad \beta^{is} \quad \beta^\pi \quad \beta^d) \begin{pmatrix} i_{t-1}^l \\ i_{t-1}^s \\ \pi_{t-1}^e \\ d_{t-1} \end{pmatrix} + \psi \begin{pmatrix} ism_t \\ ism_{t-1} \end{pmatrix} + \eta + \varepsilon_t \tag{5}$$

The long-run relationship of this system – after the cointegration coefficients have been normalized to the long-term interest rate i^l – is

$$i^l = \beta^{is} \cdot i^s - \beta^\pi \cdot \pi^e - \beta^d \cdot d$$

where the β 's reflect the long-term coefficients.

To interpret the cointegration relation as an equation for the long-term interest rate, however, all variables except the long-term interest rate i^l have to be weakly exogenous, i.e. deviations from the long-term equilibrium should be corrected solely via movements of i^l . As mentioned above, the extent to which the individual variables adjust to the long-term equilibrium is captured by the α -values. In a formal test, the null of weak exogeneity of i^s , d and π^e ($\alpha^{is} = \alpha^d = \alpha^\pi = 0$) cannot be rejected at standard levels of significance ($\chi^2(3) = 1.85$, p -value = 0.60).¹² In contrast, the null of weak exogeneity of i^l has to be rejected at all levels of significance ($\chi^2(1) = 21.41$, p -value = 0.00). Table 3 summarizes the regression results for the VECM.

¹²When exogeneity is tested for each variable separately the conclusions do not change: i^s : $\chi^2(1) = 0.39$, π : $\chi^2(1) = 0.39$, d : $\chi^2(1) = 0.72$.

¹³As Banerjee *et al.* (1986) have shown, this single equation model is superior to the two-step procedure of Engle and Granger (1987) as it avoids the small sample bias. Furthermore, this approach still yields valid results in the case of structural breaks (Campos *et al.*, 1996).

Table 2. Test for the number of cointegration relationships in the VECM

Unrestricted cointegration rank test (Trace)				
Hypothesized no. of CE(s)	Eigenvalue	Trace statistic	Critical value	Prob.**
None*	0.1358	55.1177	47.8561	0.0090
At most 1	0.0522	18.0517	29.7971	0.5622
At most 2	0.0171	4.4279	15.4947	0.8661
At most 3	0.0002	0.0443	3.8415	0.8332
Unrestricted cointegration rank test (maximum eigenvalue)				
Hypothesized no. of CE(s)	Eigenvalue	Max-eigen statistic	Critical value	Prob.**
None*	0.1358	37.0659	27.5843	0.0023
At most 1	0.0522	13.6238	21.1316	0.3966
At most 2	0.0171	4.3836	14.2646	0.8168
At most 3	0.0002	0.0443	3.8415	0.8332

Notes: *Denotes rejection of the hypothesis at the 0.05 level. Max-eigenvalue test indicates 1 cointegrating eqn(s) at the 0.05 level. Trace test indicates 1 cointegrating eqn(s) at the 0.05 level.

**MacKinnon-Haug-Michelis (1999) p -values.

Owing to the weak exogeneity of the fundamentals, switching to a Single Equation Error Correction Model (SEECM; Engle *et al.*, 1983; Johansen, 1992) may improve the efficiency of the estimates. We test the existence of a stable long-run relationship within this approach according to an error-correction model, i.e. the significance of the error-correction term. To be more specific, we proceed with the single equation nonlinear approach of Stock (1987) where the error-correction model and the cointegration relation are estimated simultaneously.¹³ Thus, we estimate the following equation:

$$\begin{aligned}
 \Delta i_t^l &= \alpha \cdot (i_{t-1}^l - \beta \cdot z_{t-1} - \eta) \\
 &+ \sum_{j=1}^m \gamma_j \cdot \Delta i_{t-j}^l + \sum_{j=0}^m \varphi_j \cdot \Delta z_{t-j} \\
 &+ \sum_{j=0}^m \psi_j \cdot x_{t-j} + \varepsilon_t \tag{6}
 \end{aligned}$$

where z is the vector of $I(1)$ -variables i^s , d and π^e which enter the cointegration space, x is a vector of (stationary) regressors only entering short-run dynamics (in our case ism), α is the error correction

Table 3. Coefficients and test statistics of the VECM (*t*-values in brackets)

Error correction	Δi_t^l	Δi_t^s	$\Delta \pi_t^e$	Δd
α	-0.19123 [-6.06881]	0.00000 [NA]	0.00000 [NA]	0.00000 [NA]
Δi_{t-1}^l	0.15323 [2.36096]	0.08281 [1.62024]	-0.01582 [-0.49044]	0.01699 [0.45743]
Δi_{t-1}^s	-0.18083 [-2.02968]	0.03621 [0.51606]	0.00656 [0.14809]	-0.11276 [-2.21123]
$\Delta \pi_{t-1}^e$	0.09165 [0.71014]	-0.05428 [-0.53407]	-0.02915 [-0.45440]	-0.13977 [-1.89215]
Δd_{t-1}	0.08131 [1.06432]	0.02928 [0.48670]	0.00029 [0.00769]	0.71839 [16.4290]
Constant	-1.22078 [-4.79346]	-1.09858 [-5.47767]	-0.15435 [-1.21949]	-0.17138 [-1.17570]
log(<i>ism</i>)	0.04843 [5.13196]	0.02177 [2.92931]	-0.00272 [-0.58086]	-0.01146 [-2.12069]
log(<i>ism</i> ₋₁)	-0.02567 [-2.76646]	-0.00100 [-0.13728]	0.00552 [1.19766]	0.01516 [2.85433]
R^2	0.19572	0.20017	0.02051	0.53541
SE equation	0.27806	0.21897	0.13819	0.15915
F-statistic	8.55167	8.79494	0.73586	40.50041

Note: Cointegrating Eq: i_{t-1}^l , 1.00000; i_{t-1}^s , -0.34096, [-5.35907]; π_{t-1}^e , -0.54466, [-3.27147]; d_{t-1} , 0.08133, [0.364565]; Constant, -4.87872.

term and ε is a white-noise residual. The significance of α is assessed according to the critical values of Banerjee *et al.* (1998). Significance is taken as evidence of cointegration.¹⁴ To obtain the SEs and the *t*-statistics of the long-run coefficients β , we estimate the Bewley transformation of the model (West, 1988).

The bracket term of Equation 6 with the variables in levels describes the cointegration relationship that has been normalized to the long-term interest rate. The lag length (*m*) is restricted to a maximum of four. A general-to-specific-modelling is pursued with the so-called backward procedure, i.e. insignificant coefficients (error probability >5%) have been successively deleted. The final regression reads as (absolute *t*-values in brackets below coefficients)

$$\begin{aligned} \Delta i_t^l = & -0.25 \cdot \left(i_{t-1}^l - 0.33 i_{t-1}^s - 0.56 \pi_{t-1}^e \right. \\ & \left. + 0.07 d_{t-1} + 8.61 \right) \\ & + 0.15 \Delta i_{t-1}^l + 1.88 ism_t - 1.02 ism_{t-1} \\ & + 0.53 \Delta i_t^s - 0.20 \Delta i_{t-1}^s \end{aligned} \quad (7)$$

$R^2 = 0.32$; SE = 0.25; LM(1) = 0.04; LM(4) = 1.09; ARCH(1) = 0.10; ARCH(4) = 1.10; JB = 1.00; CUSUM: stable; CUSUM square: stable.

The coefficients of the long-run relationship show the theoretically expected signs and are statistically significant at standard levels. They largely resemble those of the Johansen procedure (Table 3). This is indicative of some stability irrespective of the applied econometric methodology. The estimates indicate

that in the long run a one percentage point increase in inflation expectations lift long-term interest rates by 56 basis points. If we assume that the Fisher hypothesis holds, i.e. higher inflation expectations increase short-term interest rates one-to-one, the total effect of rising inflation expectations on long-term interest rates is almost one-to-one. This is in line with Keeley and Hutchison (1986), who emphasize that this result could be due to monetary regime stability. The Greenspan era on which we concentrate in this article obviously was characterized by such stability. The short-term interest rate also exerts a highly significant positive impact. This result points to the important role of monetary policy and arbitrage in determining long-term rates. The coefficient on i^s indicates that a permanent rise in the short-term interest rate of, say, 100 basis points will result in an increase of the long-term interest rate of 33 basis points.¹⁵ Accordingly, the term structure tends to flatten with higher and to steepen with lower short-term rates (see also Diebold *et al.*, 2006). The less than proportional response of i^l to i^s in the US has also been detected by Ducoudré (2005). The overall impact of the business cycle, measured by *ism*, on i^l is positive and highly significant, indicating that the effect *via* the supply of bonds is dominating (in line with Diebold *et al.*, 2006). In the short run, a contemporaneous 1% increase of the *ism* lifts i^l by about two basis points. The significantly positive relationship between i^l and its first lag may be an indication that in the short run the interest rate is driven by nonfundamental factors as well. This could be due to the market behaviour of chartists and

¹⁴The conclusions of Pesavento (2004) indicate that such kind of tests, if suitably specified, perform better than other cointegration tests in terms of power in large and small samples and are also not worse or better in terms of size distortions.

¹⁵According to Poole (2005) the average historical relationship between the short and the long rate is about 0.30. Belke and Polleit (2007) show how the Fed set the short-term interest rate.

technical analysts (Nagayasu, 1999) whose interest rate forecasts are usually based on past interest rate movements.

The coefficient of the structural factor d is significantly positive. A value of 0.07 means that an increase of the debt ratio by one percentage point lowers the bond yield by seven basis points. From 2003 to 2006, the amount of Treasuries held by foreigners increased by about 10 percentage points. This alone would have had a downward impact of 70 basis points on bond yields. This result is in line with Bernanke *et al.* (2004), Frey and Moët (2005) as well as Warnock and Warnock (2005). Rudebusch *et al.* (2006) show that foreign ‘official’ purchases of US Treasuries alone seem to have played little or no role. Longstaff (2004), in contrast, argues that if US investors, who presumably may benefit more from the highly liquid Treasury market than many foreign holders of Treasury debt, suddenly begin to purchase Treasuries from these foreign holders, the yields on Treasuries should increase to reflect the increased popularity of holding Treasuries. However, he finds that this effect is only significant for maturities up to 3 years.

The coefficient of the error-correction term is negative and highly significant. Thus, one condition for long-run stability is satisfied. The parameter estimate of -0.25 suggests a half-life of shocks of about 2 months. In other words, the gap between the long-term nominal interest rate and its equilibrium value is halved within 2 months after the occurrence of an exogenous shock. Within 1 year, the gap is accordingly reduced by over 97%.¹⁶

Breusch–Godfrey Lagrange Multiplier tests (LM) do not indicate autocorrelation in the residuals (first and fourth order) and the LM (ARCH) test for autoregressive conditional heteroscedasticity (first and fourth order) cannot identify any violations of the white-noise assumptions. In addition, the Jarque–Bera (JB) test confirms the normality of the residuals. And finally, according to different CUSUM tests there are no signs of parameter or variance instability. This once again underscores the stability of the estimated relation.

In the Introduction, we mentioned that some commentators argue that structural or uncommon factors are needed to explain the recent behaviour of bond yields. To examine whether the foreign debt ratio d captures these structural or uncommon factors adequately, we use the cointegration relation of our model to calculate a ‘fair value’ of

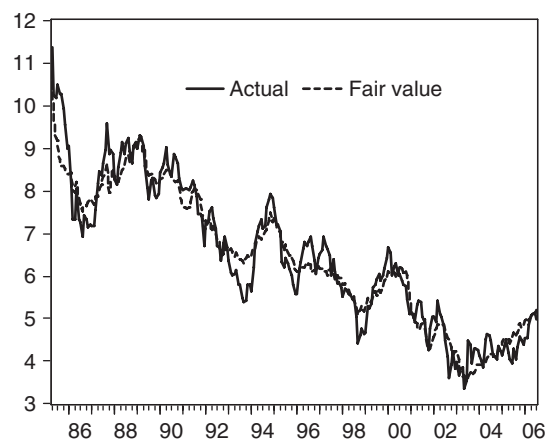


Fig. 5. The fair value of 10-year Treasury bonds (in %)

10Y Treasury yields. Figure 5 shows that bond markets were indeed overvalued in the course of 2005, but apparently this ‘disequilibrium’ was not unusually high in historical perspective. Hence, while the traditional macrofactors (i^s , Π^e , ism) ‘alone’ are no longer capable to explain the developments satisfactorily as would have been the case until mid-2005 (Clostermann and Seitz, 2005), our four variables seem to capture the evolution of bond yields throughout the entire estimation period quite well.

Visual inspection of Figs 1 and 4 may suggest that our structural variable (d) simply follows a deterministic trend, which captures the downward trend in the bond yield in the sample period. Therefore, the essential question is whether our foreign debt holdings variable contains more information than a deterministic trend. To answer this question, we first ran a battery of unit root and stationarity tests. All of them show that the debt variable captures more than a deterministic trend.¹⁷ The augmented Dickey–Fuller test, its generalized least squares variant developed by Elliot–Rothenberg as well as the Phillips–Perron test all indicate that the null of a unit root cannot be rejected at standard significance levels if a trend and a constant are included. On the other hand, the Kwiatkowski–Phillips–Schmidt–Shin test rejects the null of trend-stationarity (significance level 1%). This means that, in addition to a deterministic trend, the foreign debt holding variable also contains a stochastic trend and thus has more information than a deterministic trend alone.

¹⁶ The half-life is calculated as $\log(0.5)/\log(1 + \alpha)$.

¹⁷ The detailed test results are available upon request.

However, it is well known that the power of these tests to discriminate between trend stationarity and difference stationarity is rather weak, i.e. it is difficult to distinguish a deterministic trend from a stochastic trend. To take this into account, we add a deterministic trend to our baseline SEEC model. If we compare the estimates of the model with (Equation 7a) and without a trend (Equation 7), two main findings emerge: first, the coefficients of all variables remain mostly unchanged and are still significant. Second, the coefficient of the trend is not statistically significant. This corroborates that the information content of the debt variable is superior to the information content of a deterministic trend.

$$\begin{aligned} \Delta i_t^l = & -0.25 \cdot \left(i_{t-1}^l - 0.33 i_{t-1}^s - 0.56 \pi_{t-1}^e \right. \\ & \left. + 0.07 d_{t-1} + 0.006 \text{trend} - 2.13 \right) \\ & + 0.15 \Delta i_{t-1}^l + 1.86 \text{ism}_t - 1.02 \text{ism}_{t-1} \\ & + 0.53 \Delta i_t^s - 0.20 \Delta i_{t-1}^s \end{aligned} \quad (7a)$$

We also did some forecast encompassing tests (Diebold, 2004, p. 301) to see whether a forecast based on our SEEC model, which includes the foreign holding debt variable (F^{debt}), incorporates all relevant information of a competing forecast based on a SEEC model with a deterministic trend (F^{trend}) instead of the debt variable. Because the bond yield is an $I(1)$ -variable, we consider a forecast equation that models the ‘change’ in long-term interest rates during the next 24 months.

$$i_{t+24}^l - i_t^l = \alpha_{\text{debt}} \cdot (F_{t+24}^{\text{debt}} - i_t^l) + \alpha_{\text{trend}} \cdot (F_{t+24}^{\text{trend}} - i_t^l) + \alpha_o \quad (8)$$

Our test procedures yield the following results: First, the hypothesis $\alpha_{\text{debt}}=1$ and $\alpha_{\text{trend}}=0$ cannot be rejected (significance level 10%). This implies that the model which includes the debt variable encompasses the model that uses the trend. Second, we have to reject the hypothesis $\alpha_{\text{debt}}=0$ and $\alpha_{\text{trend}}=1$ (significance level 2%). Thus, the model with a trend does not encompass the model with the foreign debt holding variable.

In a nutshell, the foreign debt variable seems to be superior to a deterministic trend as it contains more information than the latter.¹⁸ By using this debt ratio, we are able to explain in economic terms the stochastic trend in bond yields.

¹⁸ Moreover, using only a deterministic trend as explanatory factor would have no reasonable economic interpretation and would also implausibly imply that bond yields can get negative in the future.

Forecast evaluation

In order to assess the quality of our SEECM in forecasting exercises, we compare it with a Random Walk Model (RWM). Following the influential article of Meese and Rogoff (1983), the RWM has become a popular benchmark in forecast evaluation. In line with the results of the unit root tests, the RWM is specified without a constant or trend.

We run two different kinds of out-of-sample forecasts of up to 12 months into the future. The first are fully dynamic forecasts which assume that the forecaster has no idea about the future path of the right-hand side variables and bases his predictions of these variables on simple univariate time-series models. Thus, the forecasts include only information that had actually been available at the time it was carried out. In contrast to this narrow information set, the second approach assumes that the forecaster knows the true values of the exogenous variables. Realistically, the actual forecasting environment should be somewhere between these two extreme cases.

The h -step-ahead forecast error ($e_{t+h,t}$) is calculated as the difference between the actual value of i^l at time $t+h$ (i_{t+h}^l) and its forecast value ($i_{t+h,t}^l$)

$$e_{t+h,t} = i_{t+h}^l - i_{t+h,t}^l \quad (9)$$

The forecasts are carried out recursively. The ‘first’ estimation period is 1986:1–1995:7 and the first forecast period runs from 1995:8 to 1996:7. The forecast ‘window’ is then successively extended month by month. Consequently, the next estimation period is 1986:1–1995:8 and the forecast period is from 1995:9 to 1996:8. And the last forecast period is from 2005:7 to 2006:6. In sum, we get 120 true out-of-sample forecast errors for each ‘ h ’.

To assess the quality of the forecasts of the competing models, we employ two criteria. The first is the Root Mean Squared Error (RMSE):

$$\text{RMSE}_h = \sqrt{\frac{1}{T} \sum_{t=1}^T e_{t+h,t}^2} \quad (10)$$

A smaller RMSE implies better forecast performance. A formal test based on the loss differential (Diebold and Mariano, 1995) provides information on the significance of the relative forecasts.

The second criterion is a so-called Hit Ratio (HR). It assesses the correct sign match and makes use of

Table 4. Forecast quality of different models

Forecast horizon Months ahead	SEECM, Fully dynamic				SEECM, Perf. foresight			
	RMSE	Probability	HR	Probability	RMSE	Probability	HR	Probability
1	26.76	0.50	54.17	0.38	24.46	0.01	62.50	0.00
2	38.40	0.43	55.00	0.32	32.56	0.00	72.50	0.00
3	45.16	0.68	58.33	0.13	34.53	0.00	75.83	0.00
4	51.46	0.52	61.67	0.03	36.43	0.00	75.00	0.00
5	55.48	0.35	56.67	0.14	37.38	0.00	79.17	0.00
6	57.12	0.22	60.00	0.08	37.90	0.00	83.33	0.00
7	58.60	0.19	65.00	0.01	37.99	0.00	80.00	0.00
8	60.76	0.25	60.00	0.10	38.13	0.00	82.50	0.00
9	62.71	0.20	67.50	0.00	38.18	0.00	80.00	0.00
10	65.72	0.15	69.17	0.00	37.98	0.00	81.67	0.00
11	68.38	0.09	66.67	0.00	37.53	0.00	82.50	0.00
12	71.49	0.05	70.00	0.00	37.34	0.00	81.67	0.00

an indicator variable J , which has the following properties:

$$\text{if } \text{sign}(i_{t+h}^d - i_t^d) = \text{sign}(i_{t+h,t}^d - i_t^d) \Leftrightarrow J = 1$$

$$\text{if } \text{sign}(i_{t+h}^d - i_t^d) \neq \text{sign}(i_{t+h,t}^d - i_t^d) \Leftrightarrow J = 0$$

Therefore, HR is defined as

$$\text{HR}_h = \left(\frac{1}{T} \sum_{t=1}^T J_t \right) \cdot 100 \quad (11)$$

The higher the HR, the more often the forecast signals the correct direction of interest rate changes.¹⁹ For example, a HR of 70% implies that in 70% of all cases the model predicts the correct sign of future interest rate changes. The significance relative to the RWM is again tested according to the test statistics developed by Diebold and Mariano (1995). Both forecast evaluation criteria, RMSE and HR, are discussed in Cheung *et al.* (2005).

Table 4 shows the two forecasting metrics as well as the p -values of the null that the SEECM and the RWM have equal forecasting accuracy. As is evident from this table, our model always outperforms the RWM significantly in the perfect foresight case, i.e. the average forecast errors of the SEECM are lower and the direction of interest rate changes are more often correctly forecasted by the SEECM. In the fully dynamic case, the predictions of the SEECM are also better than those of the RWM, but in some cases the differences are not significant. This is especially true for the RMSE where we are only able to beat the RWM significantly for the two longest forecast horizons ($h = 11, 12$). Overall, the results underpin the superiority of the SEECM,

especially for longer forecast horizons. Moreover, it is obvious that the SEECM does a better job the better the forecaster's predictive abilities with regard to the exogenous variables are.

IV. Summary and Conclusions

Our analysis reveals that the development of long-term bond yields in the US can be very well explained by standard macroeconomic variables and a structural factor. The macroeconomic factors which are widely considered to be the minimum set of fundamentals needed to capture basic macroeconomic dynamics are monetary policy, the business cycle and inflation expectations. In addition, the share of Treasuries held by foreign investors captures the structural factors often mentioned in the literature. These four variables are able to explain the movement of bond yields in a stable manner – even during the low interest-rate period of 2004 to 2006. Den Butter and Jansen (2004) find that the US yield is an important determinant of the German long-term bond yield. This means that the model presented here may be a useful input in a German bond yield model.

Our forecasting exercises show that we are able to outperform a RWM. In these tests, the fully dynamic approach assumes that the forecaster has no information at all about the exogenous variables. An assumption that is obviously conservative in real world applications. On the other hand, the perfect foresight case neglects informational deficiencies. The RWM, which we use as a benchmark, might be criticized as being too 'naive' in that it can be improved by

¹⁹ The direction-of-change statistic is commonly used by practitioners.

including more AR- and MA-terms. Nevertheless, it is standard in the literature (see, e.g. Cheung *et al.*, 2005). In this respect, one may be interested in further evaluation metrics, e.g. a consistency criterion, to check the robustness of our results. This is left to future research.

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