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Abstract:
We analyze 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 behavior of bond yields, even during the last two years, can well be explained. Alongside the more traditional macroeconomic 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.

Keywords: bond yields, interest rates, cointegration, inflation, forecasting

JEL: C32, E43, E47

Deutscher Abstract:
In dem vorliegenden Papier untersuchen wir, ob und in welchem Ausmaß fundamentale makroökonomische Faktoren, temporäre Einflüsse und/oder strukturelle Faktoren zum niedrigen Niveau der Renditen in den USA in den letzten Jahren beigetragen haben. Dafür gehen wir von einem allgemeinen Zinsbestimmungsmodell aus. Die empirische Umsetzung verwendet eine Kointegrationsanalyse mit einem Fehlerkorrekturmechanismus. Es gelingt uns, eine stabile Langfristbeziehung für die Renditen aufzustellen, mit der wir die Entwicklung der Renditen, auch in den letzten Jahren, befriedigend nachvollziehen können. Neben den mehr traditionellen Faktoren wie Kerninflation, Geldpolitik und Konjunktur, berücksichtigen wir auch die ausländische Nachfrage nach US-Staatsanleihen. In letzterer dürften sich strukturelle Einflüsse auf die Renditen niederschlagen. Mit der präferierten Renditegleichung kann ein Random Walk in verschiedenen Prognoseszenarien geschlagen werden.

Explaining the US Bond Yield Conundrum*

1. Introduction

Long-term interest rates in Europe and in the US fell to all-time lows in the last few years. And despite a rebound, they still have been traded at historically low levels in 2006, especially in the US. This is even more remarkable as the economic environment during the same time has been unfavorable for Treasuries: The US economy has so far been growing above trend, the Fed has raised its target rate several times and core inflation has been increasing since 2004.

In his February 2005 testimony before the Committee on Banking, Housing, and Urban Affairs of the U.S. Senate, Alan Greenspan asserted: "For the moment, the broadly unanticipated behavior 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 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 United States, 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 favor the purchase of bonds over other asset classes by pension funds and life insurance corporations." (ECB, 2005, 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 interest rates, there are hints that other, more temporary market influences related to speculative behavior may have played a role, too. The alleged widespread use of so-called 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

* We thank seminar and workshop participants of the ROME network, the Institute of the German Economy and the Bayerische HypoVereinsbank as well as D. Dakshaw and B. Schnatz for helpful comments. The views expressed here are those of the authors and do not necessarily represent those of the Bayerische HypoVereinsbank AG.

reversed at some point and hence should not have a permanent effect on the level of long-term interest rates. In addition, 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.

Moreover, recent empirical papers find tentative evidence that structural factors are at work. Clostermann and Seitz (2005) developed 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 instabilities in the last years". Kozicki and Sellon (2005, 29) suggest "that the key factor behind the conundrum is a large reduction in the term premium." Bandholz (2006) shows that the unexplained part of his US bond model increases to values not seen in the past when data for 2006 are included. Moreover, Rudebusch et al. (2006) confirm on the basis of empirical no-arbitrage macro-finance models of the term structure "that the recent behavior 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) find that a conundrum no longer exists. 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 volatiltiy, 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."

To find out whether fundamental macroeconomic, temporary or more structural factors have been at work we proceed as in Clostermann and Seitz (2005). First, we discuss which fundamentals should theoretically 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 which are augmented by structural factors.[1] Next, we estimate an interest rate model for ten-year (10Y) US Treasury notes and analyze 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"

[1] The first three variables are also the main determinants of nominal yields on one-year Treasury bills in Mehra (1995) in a sample from 1955 to 1994. This is exactly the macroeconomic interpretation of Dewachter and Lyrio (2006) of the three latent factors often found in standard finance models of the yield curve.

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 in 2005/06 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; Caporale and Williams, 2002; Brooke et al., 2000; Durré and Giot, 2005). The second approach uses high-frequency (in most cases daily) data to analyze the reaction of yields to news or announcements (see, e.g., Monticini and Vaciago, 2005; Demiralp and Jordà, 2004). 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 modeling 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 1 and 3, but also partly borrows from 4.

2. 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 (depending on technological developments), 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.[2] So far, however, DSGE 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:

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

where r^l_t is the real long-term rate, r^s_t 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

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

defining that influence which gives us the term or risk premium in r^l_t (Caporale and Williams, 2002, 121).[3]

To make (1) suitable for empirical work, we need information on the specifics of *rp*. Following Breedon, Henry and Williams (1999), Caporale and Williams (2002) and others, c_t is a catch-all variable for risks arising from macroeconomic policy developments. Specifically, we define

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

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) analyze the fiscal position. Jordá and Salyer (2003) ask whether the liquidity situation helps to explain bond yields (see also ECB, 2005, 23). Durré and Giot (2005) investigate whether stock market variables are responsible for bond market developments. We decided to include an indicator variable which has already been considered in the past, but in a different way than we do (see Warnock and Warnock, 2005; Wu, 2005; Frey and Moëc, 2005). It refers to a changing structural demand by foreigners for US Treasuries. A more concrete description and discussion are provided in section 3.1.

(1) and (2) are specified in real terms. Two problems arise in this context (Caporale and Williams, 2002, 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 change (2) and leave the exact response of i^l to expected inflation open.

(3) $i_t^l = ß_1 i_t^s + ß_2 \pi_t^e + \gamma rp(l, bc_t, etc_t)$

where i^l (i^s) is the nominal long-term (short-term) interest rate and π^e is expected inflation. This suggests estimation of the following equation:

[3] (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).

(4) $$i_t^l = \alpha_o + \beta_1 i_t^s + \beta_2 \pi_t^e + \gamma_1 bc_t + \sum_{i=2}^{n} \gamma_i etc_{t,i} + \varepsilon_t$$

where α_0 is a constant and ε_t is a white-noise error term.

Equations (3) and (4) have several testable economic implications and allow the testing of various hypotheses. For example, the pure expectations hypothesis implies $\alpha_0 = \gamma_i = \beta_2 = 0$ ∇i and $\beta_1 = 1$. If the Fisher effect holds either for the long-term or the short-term interest rate, $(\beta_1+\beta_2)=1$. For $\alpha_0 \neq 0$, we would have a constant term-premium model. If an exogenous higher demand for US bonds (*d*) occurs, we would get $\gamma_2 < 0$. 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.

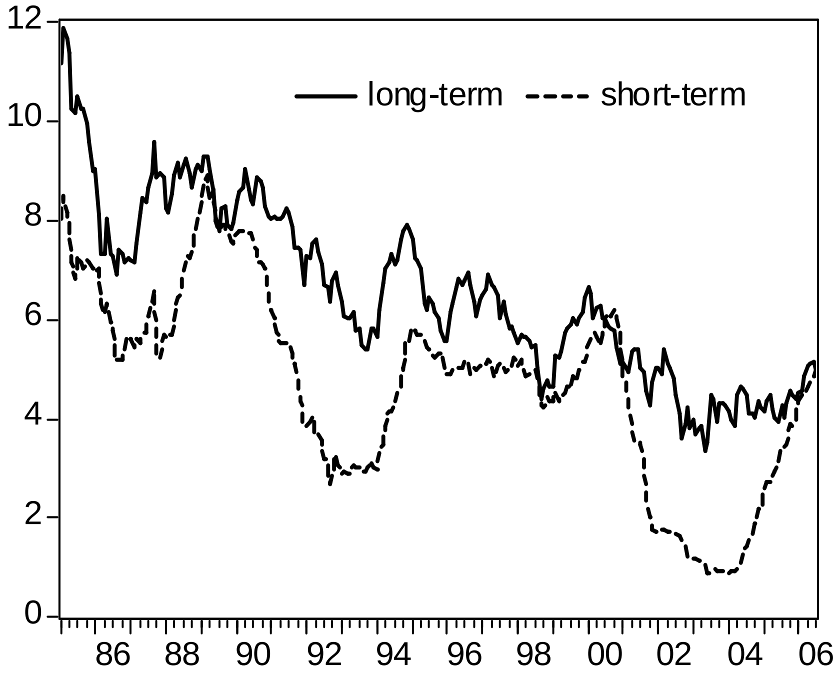
This framework allows us to test empirically whether macroeconomic factors and/or structural factors and/or temporary factors are important determinants of interest rates. However, proper inference can only be drawn within an appropriate econometric framework. This will be discussed in the next section.

3. Estimation

3.1 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, 405). The two interest rates are shown in figure 1.

Figure 1: Long-term and short-term interest rates



We measure inflation expectations with core inflation, i.e. the annual change of headline CPI excluding food and energy prices to capture the underlying price trend (see figure 2).[4] As a measure for the state of the business cycle, we use the Institute for Supply Management's manufacturing index (*ism*, see figure 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.

Figure 2: Core inflation

[4] 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.

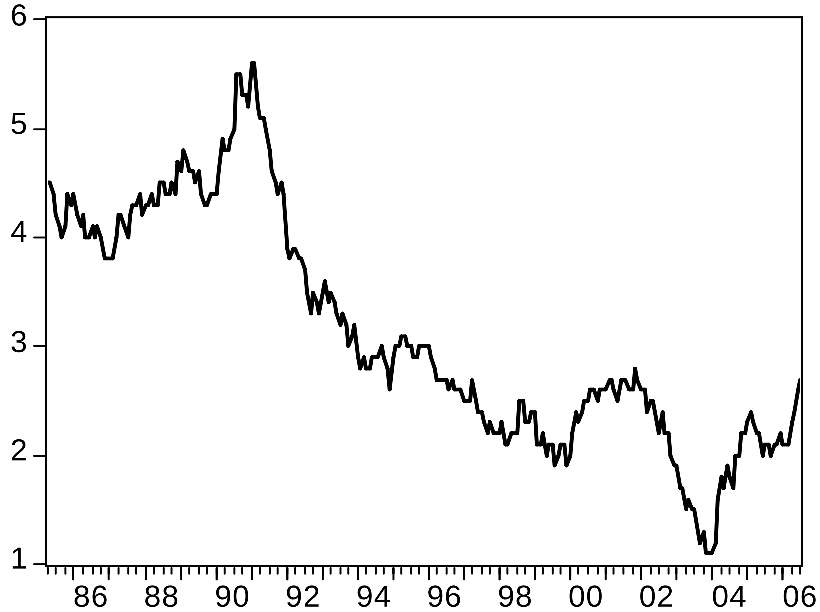
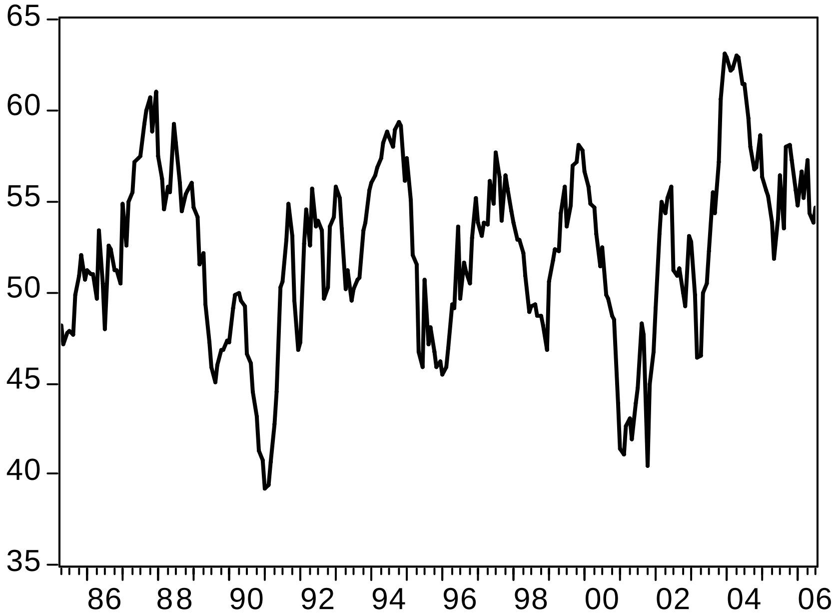


Figure 3: ISM index

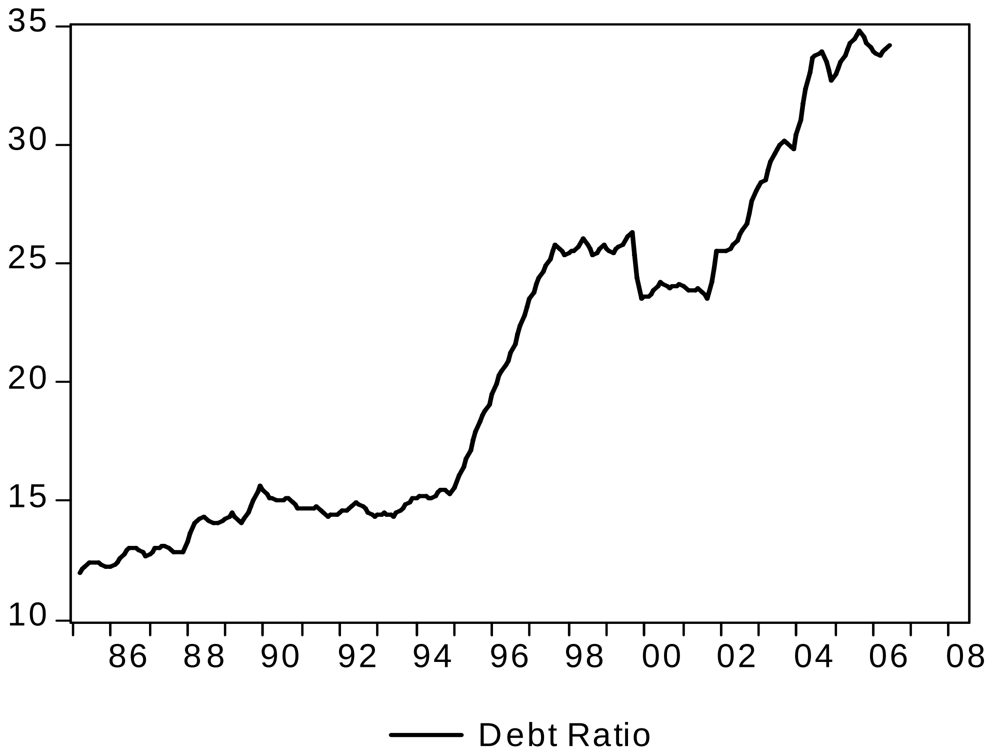


Our "*d*"-variable captures structural factors.[5] As mentioned above, higher foreign demand for US Treasuries, due to (i) demand from Asian central banks, (ii) the recycling of petrodollars, (iii) the strong interest of institutional investors and (iv) liquidity-driven demand due to world-wide expansionary monetary policies, could be responsible for the low level of US

[5] We tried several other 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, 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."

bond yields during the last two years. To quantify the influence of these factors, we include official and private foreign holdings of US Treasuries ("Treasury Securities") in percent of the overall federal debt ("total liabilities").[6] The following figure 4 shows that, since the beginning of the Japanese FX market intervention in 2002, the external debt of the US in the form of Treasuries has increased considerably. Overall, the volume of Treasuries held by foreigners nearly doubled between 2002 and 2006 from USD 1,100 bn to USD 2,000 bn. This is equivalent to about 35% of Federal Government`s total liabilities.

Figure 4: Foreign holdings of US Treasuries in % of federal debt outstanding



Our sample of monthly data runs from 1986:1 to 2006:6. All variables, except interest rates, the inflation rate and the foreign debt ratio, are in logarithms. The difference operator Δ refers to first (monthly) differences.[7]

3.2 Econometric analysis

Standard unit root tests suggest that most of our variables are *I(1)* in levels and stationary in first differences.[8] The only exception is the "*ism*" index which (in line with theoretical

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

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

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

considerations) is identified as a stationary variable. Owing to the non-stationarity 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. The empirical analysis starts with an unrestricted *VECM* which takes the following form:

$$(5) \quad \Delta y_t = \Pi y_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta y_{t-i} + \Psi x_t + \eta + \varepsilon_t ,$$

where y_t represents the vector of the non-stationary 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 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 (nxr) matrices α (the loading coefficients or adjustment parameters) and $ß$ (the cointegrating vectors) each with rank r (number of cointegration relations) such that $\Pi = \alpha ß'$ and $ß' y_t$ 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 *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 (see table 1).

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

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 non-stationary variables i^l, i^s, π^e and d. (see table 2).

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
Trace test indicates 1 cointegrating eqn(s) at the 0.05 level				
* denotes rejection of the hypothesis at the 0.05 level				
**MacKinnon-Haug-Michelis (1999) p-values				

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
Max-eigenvalue test indicates 1 cointegrating eqn(s) at the 0.05 level				
* denotes rejection of the hypothesis at the 0.05 level				
**MacKinnon-Haug-Michelis (1999) p-values				

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 (non-modeled exogenous) variable (with a lag length of 0 to 1) into the system. Hence, a *VECM* with the following structure is estimated:

$$(6)\qquad \begin{pmatrix} i_t^l \\ i_t^s \\ \pi_t^e \\ d_t \end{pmatrix} = \Gamma_1 \begin{pmatrix} \Delta i_{t-1}^l \\ \Delta i_{t-1}^s \\ \Delta \pi_{t-1}^e \\ \Delta d_{t-1} \end{pmatrix} + \begin{pmatrix} \alpha^{il} \\ \alpha^{is} \\ \alpha^{\pi} \\ \alpha^{d} \end{pmatrix} \begin{pmatrix} 1 & \beta^{is} & \beta^{\pi} & \beta^{d} \end{pmatrix} \begin{pmatrix} i_{t-1}^l \\ i_{t-1}^s \\ \pi_{t-1}^e \\ d_{t-1} \end{pmatrix} + \psi \begin{pmatrix} ism_t \\ ism_{t-1} \end{pmatrix} + \eta + \varepsilon_t \,.$$

The long run relationship of this system – after the cointegration coefficients have been normalized to the long-term interest rate i^l – *is* obtained from ($i^l - \beta^{is} \cdot i^s - \beta^{\pi} \cdot \pi^e - \beta^d \cdot d$), where the βs reflect the long-term coefficients.

To interpret the long-term relationship 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 are corrected solely through 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).[9] 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). Summing up, the following regression results for the *VECM* ensue (see table 3). For reasons to be mentioned below, the results and their implications are not discussed here but in the context of equation (8).

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

Cointegrating Eq:	CointEq			
i^l_{-1}	1.00000			
i^s_{-1}	-0.34096			
	[-5.35907]			
π^e_{-1}	-0.54466			
	[-3.27147]			
d_{-1}	0.08133			
	[3.64565]			
Constant	-4.87872			
Error Correction:	Δi^l	Δi^s	$\Delta \pi^e$	Δd
α	-0.19123	0.00000	0.00000	0.00000
	[-6.06881]	[NA]	[NA]	[NA]
Δi^l_{-1}	0.15323	0.08281	-0.01582	0.01699
	[2.36096]	[1.62024]	[-0.49044]	[0.45743]
Δi^s_{-1}	-0.18083	0.03621	0.00656	-0.11276
	[-2.02968]	[0.51606]	[0.14809]	[-2.21123]
$\Delta \pi^e_{-1}$	0.09165	-0.05428	-0.02915	-0.13977
	[0.71014]	[-0.53407]	[-0.45440]	[-1.89215]
Δd_{-1}	0.08131	0.02928	0.00029	0.71839
	[1.06432]	[0.48670]	[0.00769]	[16.4290]
Constant	-1.22078	-1.09858	-0.15435	-0.17138
	[-4.79346]	[-5.47767]	[-1.21949]	[-1.17570]
log(ism)	0.04843	0.02177	-0.00272	-0.01146
	[5.13196]	[2.92931]	[-0.58086]	[-2.12069]
$\log(ism_{-1})$	-0.02567	-0.00100	0.00552	0.01516
	[-2.76646]	[-0.13728]	[1.19766]	[2.85433]
R-squared	0.19572	0.20017	0.02051	0.53541
S.E. equation	0.27806	0.21897	0.13819	0.15915
F-statistic	8.55167	8.79494	0.73586	40.50041

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 non-linear approach of Stock (1987)

[9] When 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.

where the error correction model and the cointegration relation are estimated simultaneously.[10] Thus, we estimate the following equation:

$$\text{(7)} \qquad \Delta i_t^l = \alpha \cdot (i_{t-1}^l - ß \cdot z_{t-1}) + \sum_{j=1}^{m} \gamma_j \cdot \Delta i_{t-j}^l + \sum_{j=0}^{m} \varphi_j \cdot \Delta z_{t-j} + \sum_{j=0}^{m} \psi_j \cdot x_{t-j} + \varepsilon_t$$

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 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.[11] To obtain the standard errors and the t-statistics of the long-run coefficients *ß*, we estimate the Bewley transformation of the model (West, 1988).

The bracket term of (7) 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-modeling 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)

$$\text{(8)} \qquad \begin{aligned} \Delta i_t^l = & -\underset{(5.8)}{0.25} \cdot (i_{t-1}^l - \underset{(6.9)}{0.33} i_{t-1}^s - \underset{(4.6)}{0.56} \pi_{t-1}^e + \underset{(6.9)}{0.07} d_{t-1} + \underset{(2.3)}{8.61}) \\ & + \underset{(2.4)}{0.15} i_{t-1}^l + \underset{(4.1)}{1.88} ism_t - \underset{(2.4)}{1.02} ism_{t-1} + \underset{(7.1)}{0.53} \Delta i_t^s - \underset{(2.3)}{0.20} \Delta i_{t-1}^s \end{aligned}$$

$R^2 = 0.32$; 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. This is indicative of some stability irrespective of the applied econometric methodology. In the long run, a rise in core inflation has almost a 1-to-1-effect on the long-term interest rate (assuming that a rise in expected inflation increases the short run interest rate one to one). This might confirm the existence of the Fisher effect and 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 paper obviously was

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10 As 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). Compared to Johansen's maximum likelihood procedure (Johansen, 1995; 2000) we restrict the number of cointegration relationships to one. But this seems justified according to the pre-tests within the Johansen framework.

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.[12] Accordingly, the term structure is going 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, indicating that the effect via the supply of bonds is dominating (in line with Diebold et al., 2006). The contemporaneous reaction of i^l to *ism* is positive and highly significant. In the short-run, a contemporaneous 1% increase of the *ism* results in a 1.9 percentage point increase in i^l. This value, being greater than 1, implies that the nominal interest rate is on average more volatile than expectations about the future development of the business cycle. The significantly positive relationship between i^l and its first lag may be an indication that the interest rate is in the short run also driven by non-fundamental factors. This could be due to the market behavior of chartists and 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 1 percentage point lowers the bond yield by 7 basis points. In the last four years, 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 Warnock and Warnock (2005), Frey and Moët (2005) as well as Bernanke et al. (2004). 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 three 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-

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11 The 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.

12 According to Poole (2005) the average historical relationship between the short and the long rate is about 0.30.

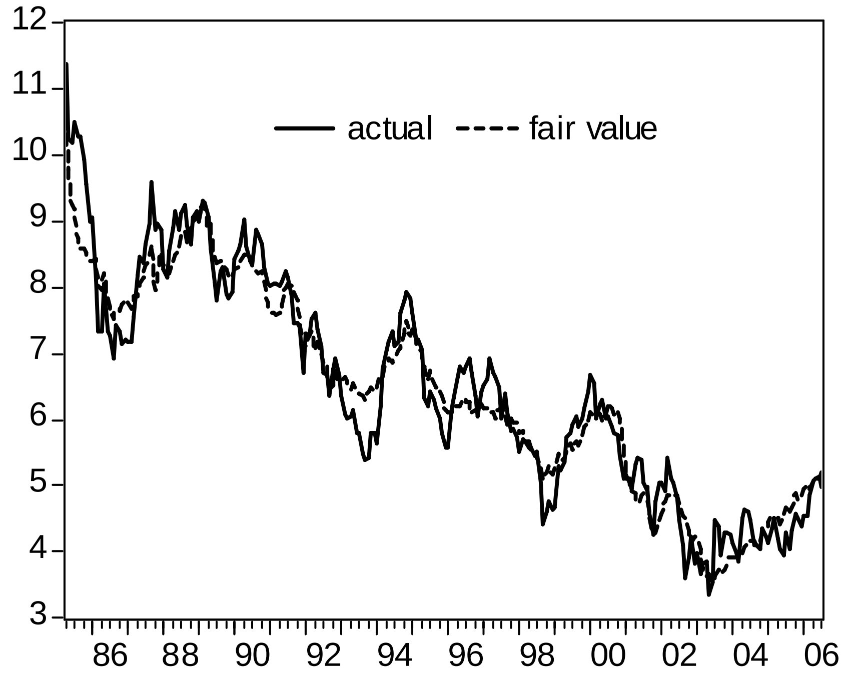
life of shocks of about two months. In other words, the gap between the long-term nominal interest rate and its equilibrium value is halved within two months after the occurrence of an exogenous shock. Within one year, the gap is accordingly reduced by over 97%.[13]

Breusch-Godfrey Lagrange multiplier tests (LM) do not indicate autocorrelation in the residuals (1st and 4th order). Nor can the Lagrange multiplier (ARCH) test for autoregressive conditional heteroskedasticity (1st and 4th order) identify any violations of the white-noise assumptions. In addition, the Jarque-Bera (JB) test confirms the normality of the residuals. And finally, CUSUM tests do not indicate 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 behavior 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 bond yields. Figure 5 shows that bond markets were overvalued in the course of 2005. But obviously the "disequilibrium" was not unusually high in a historical perspective. Our four variables seem to capture the evolution of bond yields quite well. Therefore, it is not necessary to revert to additional structural or technical factors. In contrast, the macro factors (i^s, π^e, *ism*) *alone* are not capable to explain the developments satisfactorily as would have been the case until mid 2005 (see Clostermann and Seitz, 2005).

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

Figure 5: The fair value of 10-year Treasury bonds (in %)



3.3 Forecast evaluation

In order to assess the quality of our single equation error correction model (*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 very popular benchmark in forecast evaluation. In line with 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 evolution 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^l_{t+h}) and its forecast value ($i^l_{t+h,t}$)

(9) $e_{t+h,t} = i^l_{t+h} - i^l_{t+h,t}$

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*".

The quality of the forecasts of the competing models is assessed using two criteria. The first is the root mean squared error (*RMSE*):

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

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.

Additionally, we calculate a so-called hit ratio (*HR*). It assesses the correct sign match and makes use of an indicator variable *J* which has the following properties

$$\begin{aligned} &if \;\; sign\left(i_{t+h}^l - i_t^l\right) = sign\left(i_{t+h,t}^l - i_t^l\right) \Leftrightarrow J = 1 \\ &if \;\; sign\left(i_{t+h}^l - i_t^l\right) \neq sign\left(i_{t+h,t}^l - i_t^l\right) \Leftrightarrow J = 0 \end{aligned}$$

Therefore, *HR* is defined as

$$(11) \quad HR_h = \left(\frac{1}{T}\sum_{t=1}^{T} J_t\right)\cdot 100 \,.$$

The higher the *HR*, the more often the forecast signals the correct direction of interest rate changes.[14] 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 in 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 *SEECMs* are lower and the signs of interest rate changes are more often correctly forecasted by the *SEECMs*. In the fully dynamic case, the predictions of the *SEECM* are also better than the *RWM*, but in many 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 *SEECMs*, 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.

Table 4: Forecast quality of different models

Forecast horizon	SEECM, Fully Dynamic				SEECM, Perf. Foresight			
Months ahead	RMSE	Probability	Hit Ratio	Probability	RMSE	Probability	Hit Ratio	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

4. Summary and conclusions

Our results reveal that the development of long-term bond yields in the US can be very well explained by three standard macroeconomic factors which are widely considered to be the minimum set of fundamentals needed to capture basic macroeconomic dynamics - monetary policy, the business cycle and inflation expectations - augmented by the share of Treasuries held by foreigners. The latter variable 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. Further macro variables are not needed to capture the evolution of bond yields from 2004 to 2006.

Our forecasting exercises show that we are able to outperform a random walk model. 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 side, the perfect foresight case neglects informational deficiencies. The random walk model which we use as a benchmark might be criticized as being too "naive" in that it can be improved by 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

[14] The direction-of-change statistic is one which is often used by practitioners.

in further evaluation metrics, e.g. a consistency criterion, to check the robustness of our results. This is left to future research.

References

Bandholz, H. (2006), US Treasuries: Cyclical Outweighing Structural Factors, HVB Friday Notes, 30 June 2006, 3-5.

Banerjee, A., J. Dolado, D.F. Hendry & G.W. Smith (1986), Exploring Equilibrium Relationships in Econometrics Through Static Models: Some Monte Carlo Evidence, Oxford Bulletin of Economics and Statistics 48, 253-277.

Banerjee, A., J. Dolado & R. Mestre (1998), Error Correction Mechanism Tests for Cointegration in a Single Equation Framework, Journal of Time Series Analysis 19, 267-283.

Bernanke, B.S., V.R. Reinhart & B.P. Sack (2004), Monetary Policy Alternatives at the Zero Bound: an empirical assessment, Finance and Economics Discussion Series of the Federal Reserve Board, 2004-48, September 2004.

Breedon, F., S.G.B. Henry & G.A. Williams (1999), Long-term Real Interest Rates: evidence on the global capital market. Oxford Review of Economic Policy 15, 128–142.

Brooke, M., A. Clare & I. Lekkos (2000), A Comparison of Long Bond Yields in the United Kingdom, the United States and Germany, Bank of England Quarterly Bulletin 40, 150-158.

Campos, J., N. Ericsson & D.F. Hendry (1996), Cointegration Tests in the Presence of Structural Breaks, Journal of Econometrics 70, 187-220.

Caporale, G.M. & G. Williams (2002), Long-term Nominal Interest Rates and Domestic Fundamentals, Review of Financial Economics 11, 119-130.

Cheung, Y.-W., M.D. Chinn & A.G. Pascual (2005), Empirical Exchange Rate Models of the Nineties: Are any fit to survive?, Journal of International Money and Finance 24, 1150-1175.

Christiano, L., M. Eichenbaum & C. Evans (2005), Nominal Rigidities and the Dynamic Effects of a Shock to Monetary Policy, Journal of Political Economy 113, 1-45.

Clostermann, J. & F. Seitz (2005), Are Bond Markets Really Overpriced: The case of the US, University of Applied Sciences Ingolstadt, Working Paper Nr. 11, December.

Demiralp, S. & O. Jordà (2004), The Response of Term Rates to Fed Announcements, Journal of Money, Credit, and Banking 36, 387-405.

Dewachter, H. & M. Lyrio (2006), Macro Factors and the Term Structure of Interest Rates, Journal of Money, Credit, and Banking 38, 119-140.

Diebold, F.X. & R. Mariano (1995), Comparing Predictive Accuracy, Journal of Business and Economic Statistics 13, 253-265.

Diebold, F.X., G.D. Rudebusch & S.B. Aruoba (2006), The Macroeconomy and the Yield Curve: A Dynamic Latent Factor Approach, Journal of Econometrics 131, 309-338.

Ducoudré, B. (2005), Fiscal Policy and Interest Rates, OFCE Working Paper No. 2005-08, June.

Durré, A. & P. Giot (2005), An International Analysis of Earnings, Stock Prices and Bond Yields, ECB Working Paper No. 515, August.

Ehrmann, M., M. Fratzscher & R. Rigobon (2005), Stocks, Bonds, Money Markets and Exchange Rates: Measuring International Financial Transmission, NBER Working Paper 11166, March.

Engle, R. & C. Granger (1987), Cointegration and Error Correction: Representation, Estimation, and Testing, Econometrica 35, 251-276.

European Central Bank (2005), Monthly Bulletin April 2005.

European Central Bank (2006), The Accumulation of Foreign Reserves, Occasional Paper No. 43, February.

Fels, J. & M. Pradhan (2006), Fairy Tales of the US Bond Market, Morgan Stanley Research Global, July 26, 2006.

Frey, L. & G. Moët (2005), US Long-term Yields and FOREX Interventions by Foreign Central Banks, Banque de France Bulletin Digest no. 137, May 2005, 19-32.

Greenspan, A. (2005), Testimony Before the Committee on Banking, Housing, and Urban Affairs of the U.S. Senate on February 16, 2005.

Gujarati, D.N. (1995), Basic Econometrics, 3rd ed., McGraw-Hill, Singapore.

Johansen, S. (1995), Likelihood-based Inference in Cointegrated Vector Auto-regressive Models, Oxford University Press, Oxford, New York.

Johansen, S. (2000), Modelling of Cointegration in the Vector Autoregressive Model, Economic Modelling 17, 359-373.

Jordá, O. & K.D. Salyer (2003), The Response of Term Rates to Monetary Policy Uncertainty, Review of Economic Dynamics 6, 941-962.

Keeley, M.C. & M.M. Hutchison (1986), Rational Expectations and the Fisher Effect: implications of monetary regime shifts, Federal Reserve Bank of San Francisco, Working Papers in Applied Economic Theory 86-11.

Kozicki, S. & G. Sellon (2005), Longer-Term Perspectives on the Yield Curve and Monetary Policy, Federal Reserve Bank of Kansas City Economic Review 99, No 4, 5-33.

Longstaff, F.A. (2004), The Flight-to-Liquidity Premium in U.S. Treasury Bond Prices, Journal of Business 77, 511-526.

MacKinnon, J.G., A.A. Haug & L. Michelis (1999), Numerical Distribution Functions of Likelihood Ratio Tests for Cointegration, Journal of Applied Econometrics 14, 563-577.

Meese, R. & K. Rogoff (1983), Empirical Exchange Rate Models of the Seventies: do they fit out of sample?, Journal of International Economics 14, 3-24.

Mehra, Y.P. (1995), Some Key Empirical Determinants of Short-Term Nominal Interest Rates, Federal Reserve Bank of Richmond Economic Quarterly 81/3, 33-51.

Monticini, A. & G. Vaciago (2005), Are Europe's Interest Rates led by Fed Announcements, Working Paper, University of Exeter.

Nagayasu Y. (1999), Japanese Effective Exchange Rates and Determinants: A Long-Run Perspective, in: R. MacDonald & J. Stein: Equilibrium Exchange Rates, Norwell, MA., 323-347.

Paesani, P., R. Strauch & M. Kremer (2006), Public Debt and Long-Term Interest Rates - The Case of Germany, Italy and the USA, ECB Working Paper No 656, July.

Pesavento, E. (2004), Analytical Evaluation of the Power of Tests for the Absence of Cointegration, Journal of Econometrics 122, 349-384.

Poole, W. (2005), Understanding the Term Structure of Interest Rates, Federal Reserve Bank of St. Louis Review 87, 589-595.

Rudebusch, G.D., E.T. Swanson & T. Wu (2006), The Bond Yield "Conundrum" from a Macro-Finance Perspective, Federal Reserve Bank of Dallas, Working Paper 2006-16.

Stock, J.H. (1987), Asymptotic Properties of Least Squares Estimators of Cointegrating Vectors, Econometrica 55, 1035-1056.

Tanzi, V. (1976), Inflation, Indexation and Interest Income Taxation, Banca Nazionale del Lavoro Quarterly Review 29, 54-76.

Tobin, J. (1965), Money and Economic Growth. Econometrica 33, 671–684.

Warnock, F.E. & V.C. Warnock (2005), International Capital Flows and U.S. Interest Rates, International Finance Discussion paper 840, September 2005.

West. K.D. (1988), Asymptotic Normality, When Regressors have a Unit Root, Econometrica 56, 1397-1417.

Wu, T. (2005), The Long-term Interest Rate Conundrum: Not Unraveled Yet, Federal Reserve Bank of San Francisco Economic Letter 2005-08.

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