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Heft Nr. 11 aus der Reihe
"Arbeitsberichte - Working Papers"
ISSN 1612-6483

Ingolstadt, im Dezember 2005

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Jörg Clostermann* and Franz Seitz+

December 2005

*) University of Applied Sciences
Ingolstadt
Esplanade 10
D- 85049 Ingolstadt
Germany
joerg.clostermann@fh-ingolstadt.de

+) University of Applied Sciences
Amberg-Weiden
Hetzenrichter Weg 15
D-92637 Weiden
Germany
f.seitz@fh-amberg-weiden.de

Abstract:
In the present paper we analyse whether fundamental macroeconomic factors, temporary influences or more structural factors have contributed to the recent decline in bond yields in the US. For that purpose, we start with a very general model of interest rate determination in which risk premia are captured via the macroeconomic (policy) environment. The empirical part consists of a cointegration analysis with an error correction mechanism from the mid 80s until 2005. We are able to establish a stable long-run relationship and find that the behaviour of bond rates in the last few years may well be explained by macroeconomic factors. These are driven by core price developments, monetary policy reflected in short-term interest rates and the business cycle. A changed structural demand for bonds does not seem to be at work. The existing overestimation of bond yields is not unusual historically. Finally, our bond yield equation outperforms a random walk model in different out-of-sample exercises.

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

JEL: C32, E43, E47

Are Bond Markets really Overpriced: The case of the US

1. Introduction

Long-term interest rates in Europe and in the US fell to an all-time low this year. Although they have increased gradually recently, long-term interest rates, especially in the US have been trading at historically low levels, although the economic environment there has been unfavorable for Treasuries: The US economy has so far been growing above trend, the Fed has raised its target rate several times, core inflation was until recently still heading higher and oil prices continue to rise.

In the monthly report of April 2005 the ECB 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, the speed of the decline in long-term yields which occurred from mid-2004 may suggest that other temporary market factors related to speculative behavior may also have played a role. The alleged widespread use of so-called carry trades, which generate interest income as they involve borrowing at low short-term interest rates and investing in longer-term maturities, may have amplified the trend of declining yields set in motion by more structural factors. As such trades appear to exploit market trends, they may have amplified the downturn in long-term interest rates. Speculative flows of this sort are likely to be 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 United States. There is, however, some indication that this factor played a less prominent role in most recent bond yield developments because the demand for US Treasuries from foreign official and institutional investors has leveled off in early 2005, as indicated by data published by the Federal Reserve.

To find out whether fundamental macroeconomic, temporary or more structural factors have been at work we choose the following way: First we discuss which fundamentals should theoretically determine bond yields. In a second step we estimate an interest rate model for the ten-year-US-Treasury notes. This model will be checked for parameter stability and whether there are hints of unexplained interest rate developments and of overestimations of the interest rate for the recent years. In doing that we also derive a "fair value" for the bond market. Furthermore, we undertake some out-of-sample forecasting exercises of our preferred model compared 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., Caporale and Williams, 2002; Brooke et al., 2000; 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., 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 deals 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. 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

On a general level, 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.[1] So far, however, such a model does not seem to have been developed with sufficient generality.

Therefore, and in line with other studies (see. e.g., Caporale and Williams, 2002) our analysis starts with a general model for the term structure of bond rates:

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

where r^l_t is the real long-term rate, r^s_t is the real short-term rate, l and s are the terms of the bonds, z_t is a set of variables that influences investors' risk perceptions and rp is the function defining that influence which gives us the term or risk premium on r^l_t.

To make this model operational, we need to define an explicit form for the function rp. Following Breedon, Henry and Williams (1999), Caporale and Williams (2002) and others, z_t is a capture-all variable for risks arising from macroeconomic policy developments. Specifically, we define

(2) $r_t^l = \beta r_t^s + \gamma rp(l, z_t)$

where $z=\{y, etc\}$ and y is a variable capturing the state of the business cycle. In "*etc*" we summarise three different sets of explanatory factors influencing the macroeconomic environment. First, and in line with Caporale and Williams (2002) we analyse the fiscal position d_t. Second, we ask whether the liquidity situation m_t helps to explain bond yields (see, e.g. ECB, 2005, 23; Jordá and Salyer, 2003). And third we investigate whether the stock prices p_t related variables are responsible for bond market developments (e.g. Durré and Giot, 2005). The variables included in "*etc*" will be considered in a second step. The pure expectations hypothesis implies that, if $l > h$ and $s \geq h$, where h is the holding period of the bond, the coefficient on r^l_t would be unity and γ would be 0.

(1) and (2) are specified in real terms. Two problems arise in this context. First, real rates are not directly observable but have to be proxied for empirical work. Second, the strength of the effect of expected inflation is ambiguous. Invoking the Fisher effect would allow to impose a one-to-one relationship between the nominal rate and expected inflation. Clearly, the Fisher relation holds in models in which the real interest rate does not depend on monetary variables and monetary neutrality holds. It

[1] See for a prototype model in this spirit Christiano et al. (2005).

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). The exact response of i^l to expected inflation is therefore an empirical matter. Rather than impose a coefficient of unity, we attempt to measure it directly by changing (2) to

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

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

(4) $$i_t^l = \alpha_o + \beta_1 i_t^s + \beta_2 \pi_t^e + \gamma_1 y_t + \gamma_{2,j} etc_t + \mu_t \quad for \;\; j = d, m, p$$

where α_0 is a constant and μ_t is a white noise error term.

Equations (3) and (4) represent a very general model with a number of testable economic implications. Moreover, our unrestricted approach allows to test alternative hypotheses within the same framework. For example, the pure expectations hypothesis implies $\alpha_0 = \gamma_i = 0 \;\; \nabla i$ and $\beta_1 = \beta_2 = 1$, in which case the Fisher effect also holds. For $\alpha_0 \neq 0$, we would have a constant term-premium model. For the more general cases, the coefficients would be less clearly determined a priori. For example, in a loanable funds framework, we might expect $\gamma_1 \neq 0, \gamma_{2,d} > 0, \gamma_{2,m} > 0, \gamma_{2,p} \neq 0, \beta_1, \beta_2 > 0$ with $\beta_2 = 1$ if the Fisher relation holds. Portfolio theory would imply $\gamma_{2,d} > 0$ if new debt issues are risky, and $\gamma_{2,d} < 0$ if they are of high quality, causing portfolio reallocation. If Ricardian equivalence holds, we would get $\gamma_{2,d} = 0$. If our stock market variables refer to risk considerations, we would expect $\gamma_{2,p} < 0$. If, on the other side, "*p*" stands for relative returns on the stock market, we would get $\gamma_{2,p} > 0$. Finally, the coefficient on *y* may be positive or negative depending on whether the supply of or the demand for bonds increases more.

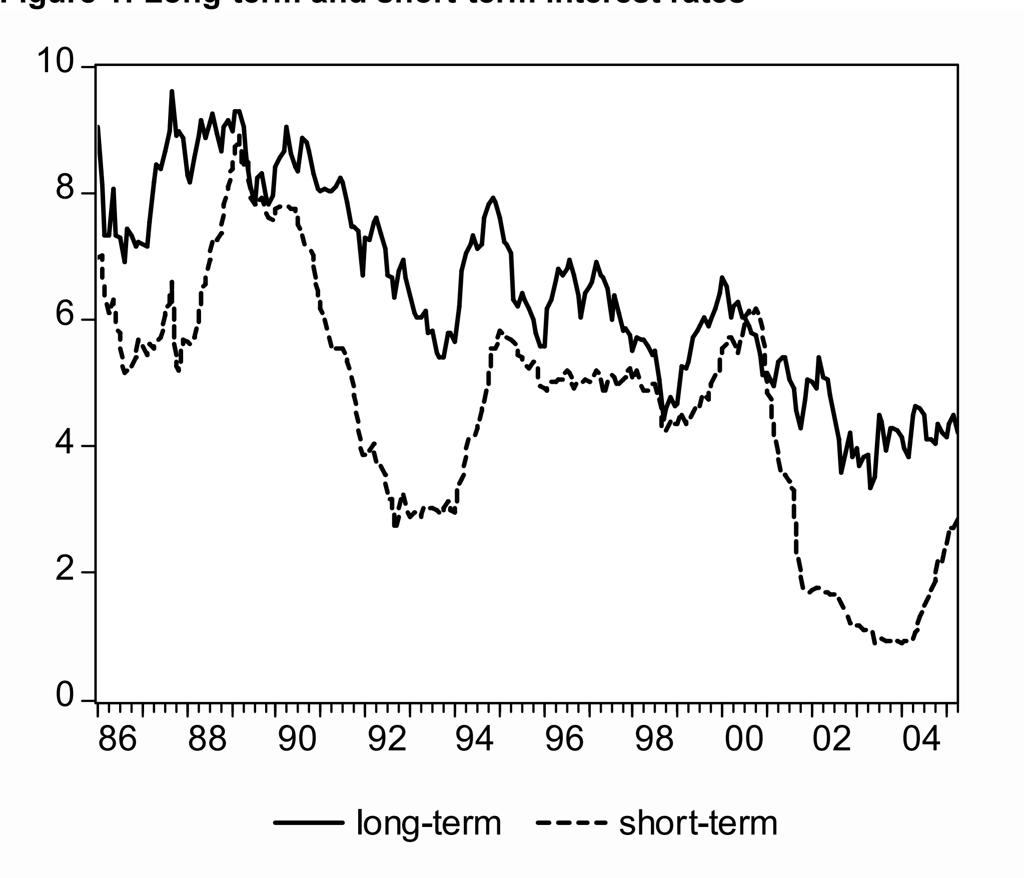
This framework therefore allows us to empirically test the general proposition that nominal long-term interest rates are determined solely by market participants, in which case the macroeconomic variables that we use will not be significant, against the alternative that macroeconomic performance is an important factor in interest rate determination. However, proper inference can only be drawn within and appropriate econometric framework, which we discuss in the next section.

3. Estimation

3.1 The Data

In what follows we estimate an equation for 10-years US Treasury notes from the mid 1980s until the end of 2005. 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 is done by economic reasoning and unit root tests (not shown, but available from the authors upon request). 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, compared to using monthly averages, do not introduce smoothness into the data which lends itself to autocorrelation in the residuals (Gujarati, 1995, 405). The rates are shown in figure 1.

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

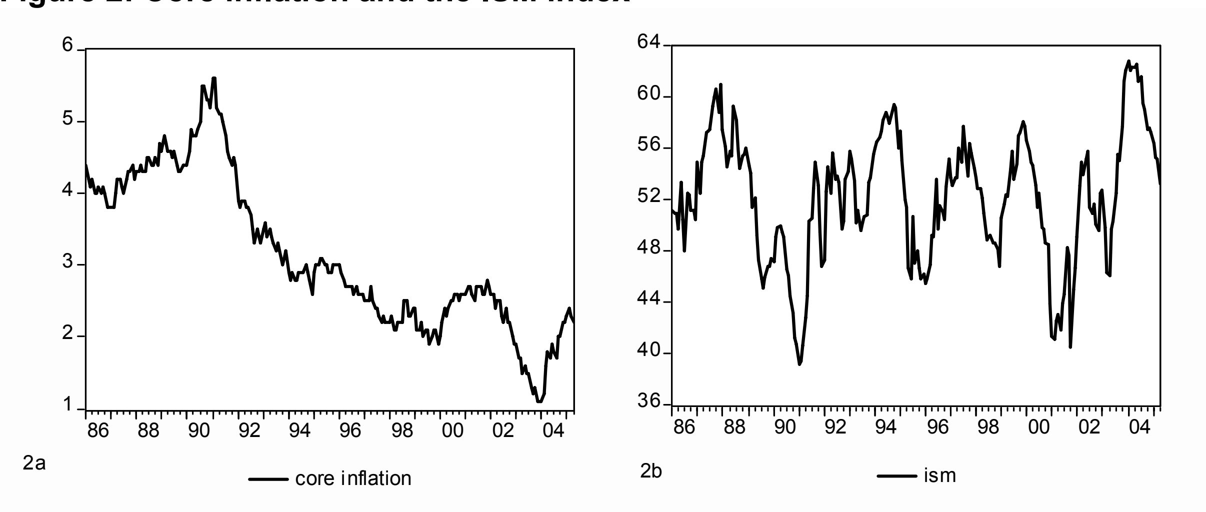


We measure expected inflation with core inflation, i.e. the annual change of headline CPI excluding food and energy prices to capture the general price trend (see figure 2a).[2] To capture the business cycle, we use the ISM Index for manufacturing (*ism*)

[2] We get slightly worse results with the overall headline measure.

from the Institute for Supply Management (see figure 2b). This variable has the advantage (and this is especially important for forecasting exercises) not to be revised and to be available with only a small publication lag. Our "*etc*"-variables comprise the public debt/deficit situation captured by total debt outstanding and total marketable debt (levels, changes, relative to GDP, per capita). As liquidity variables we consider the monetary aggregates M1, M2 and MZM and construct growth rates as well as money gaps as their difference between the actual development and trend. And finally we take into account the stock market as an alternative to investments in the bond market. We tried return (equity return) and risk (implied volatility on a broad stock market index and standard GARCH model derived volatility measures) variables. All the variables are available upon request and may alternatively be downloaded under HTTP://PEOPLE.FREENET.DE/CLOSTERMANN/DATA_US_BONDS.XLS.

Figure 2: Core inflation and the ISM index



Our sample runs from 1986.1 until 2005.9. All data are monthly and all variables except interest rates and the inflation rate are in logarithms. The difference operator Δ refers to first (monthly) differences.

3.2 Econometric analysis

Standard unit root tests suggest that most of the variables are I(1) in levels and stationary in first differences.[3] The only exception is the "*ism*" index which (in line with theoretical considerations) is identified as a stationary variable. In a first step we

[3] Test results in detail are available from the authors upon request.

start with our 4 variable system (i^s, i^l, π^e, *ism*) which forms the basis of the theoretical model. It is only in a second step that we try to integrate the other variables (*d*, *m*, *s*).

Owing to the non-stationarity of the time series, the nominal long-term interest rate is estimated in 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 relationships (cointegration relationships) on which the theoretical considerations are based. The empirical analysis starts with an unrestricted vector error correction model which takes the following form:

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

where, in the first step, y_t represents the vector of the non-stationary variables i^l_t, i^s_t and π^e_t. ε_t denotes the vector of the independently and identically distributed residuals and η the vector of constants. The number of cointegration relationships corresponds to the rank of the matrix Π. If Π has reduced rank (0 < r < p), it can be separated into a (*p x r*)-dimensional matrix of the loading coefficients α and a (*p x r*)-dimensional matrix of the cointegration vectors β ($\Pi = \alpha\beta'$). The cointegration vectors represent the long-term equilibrium relationships of the system. The loading coefficients denote the importance of the cointegration relationships in the individual equations and the speed of adjustment following deviations from the long-term equilibrium. The lag order (*k*) of the system is determined by estimating an unrestricted VAR model in levels (based on the integrated variables i^l, i^s, and π^e) and using the information criteria suggested by Akaike (*AIC*) and, alternatively, by Schwarz (*SC*) and Hannan-Quinn (*HQ*), which are usually more restrictive with regard to the lag structure to be chosen. All criteria recommend a lag length of 1 (see table 1). However, as indicated by Likelihood-ratio-test, the residuals are auto-correlated in this case. For this reason, we choose a lag length of 3. This is the most parsimonious lag structure where no auto-correlation exists.

Table 1: Lag length tests

Lag	AIC	SC	HQ
0	8.36281	8.40671	8.38050
1	-1.05216	-0.87656	-0.98139
2	-1.03446	-0.72716	-0.91060
3	-1.02158	-0.58258	-0.84463
4	-1.00583	-0.43514	-0.77581
5	-0.98986	-0.28747	-0.70675
6	-0.99436	-0.16027	-0.65817
7	-0.95125	0.01454	-0.56198
8	-0.91974	0.17775	-0.47738

The number of cointegration vectors is verified by determining the cointegration rank with the trace-test and the lambda-max-test. Both tests suggest one cointegration relationship, i.e. one equilibrium relationship between the non-stationary variables i^l_t, i^s_t and π^e_t. (see table 2).

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

Eigenv.	L-max	Trace	r	p-r	L-max90	Trace90
0.0872	21.36	31.67	0	3	14.09	31.88
0.0364	8.68	10.31	1	2	10.29	17.79
0.0069	1.63	1.63	2	1	7.5	7.5

Therefore, it seems reasonable to restrict the *VECM* to one cointegration relationship and – as the unit root tests mentioned above suggest – to include the indicator for the expected stance of the business cycle "*ism*" as a stationary (exogenous) variable (with a lag length of 0 to 2) into the system. Hence, a *VECM* with the following structure is estimated:

(6) $$\begin{pmatrix} \mathrm{il}_{\mathrm{t}} \\ \mathrm{is}_{\mathrm{t}} \\ \pi^{\mathrm{e}}_{\mathrm{t}} \end{pmatrix} = \Gamma_1 \begin{pmatrix} \Delta\mathrm{il}_{\mathrm{t-1}} \\ \Delta\mathrm{is}_{\mathrm{t-1}} \\ \Delta\pi^{\mathrm{e}}_{\mathrm{t-1}} \end{pmatrix} + \begin{pmatrix} \alpha^{il} \\ \alpha^{is} \\ \alpha^{\pi} \end{pmatrix} \begin{pmatrix} 1 & \beta^{is} & \beta^{\pi} \end{pmatrix} \begin{pmatrix} \mathrm{il}_{\mathrm{t-1}} \\ \mathrm{is}_{\mathrm{t-1}} \\ \pi^{\mathrm{e}}_{\mathrm{t-1}} \end{pmatrix} + \psi \begin{pmatrix} \log(ism_t) \\ \log(ism_{t-1}) \\ \log(ism_{t-2}) \end{pmatrix} + \eta + \varepsilon_t \,.$$

The long run relationship of this system – after the cointegration coefficients have been normalised to the long-term interest rate i^l – is obtained from $il - \beta^{is} \cdot is - \beta^{\pi} \cdot \pi^e$, 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 must meet the condition of weak exogeneity, i.e. deviations from the long-term equilibrium are corrected solely through responses of i^l. As mentioned above, the extent to which the individual variables adjust to the long-term equilibrium is expressed in the α-values. In a formal test, the null-hypothesis of weak exogeneity of i^s and π^e ($\alpha^{is}=\alpha^{\pi}=0$) cannot be rejected at standard levels of significance ($\chi^2(2) = 1.85$, p-value = 0.40).[4] In contrast, the null of weak exogeneity of i^l has to be rejected ($\chi^2(1)= 5.93$, p-value = 0.01). Taking these results into account, the following regressions for the *VECM* ensue (see table 3):

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

Cointegrating Eq:	CointEq		
i^l_{-1}	1.00000		
i^s_{-1}	-0.33894		
	[-4.00423]		
π^e_{-1}	-0.94187		
	[-6.28504]		
Constant	-1.93210		
Error Correction:	Δi^l	Δi^s	$\Delta \pi^e$
α	-0.15442	0.00000	0.00000
	[-3.63747]	[NA]	[NA]
Δi^l_{-1}	0.13662	0.08942	0.00619
	[1.84560]	[1.56970]	[0.17619]
Δi^l_{-2}	-0.05845	0.05467	0.00365
	[-0.82399]	[1.00146]	[0.10838]
Δi^s_{-1}	-0.17303	0.02224	-0.01893
	[-1.80549]	[0.30152]	[-0.41645]
Δi^s_{-2}	-0.02669	0.00067	0.12034
	[-0.27777]	[0.00910]	[2.64033]
$\Delta \pi^e_{-1}$	0.08568	-0.07875	-0.04583
	[0.61807]	[-0.73825]	[-0.69696]
$\Delta \pi^e_{-2}$	0.08430	-0.02533	-0.15129
	[0.62115]	[-0.24257]	[-2.34990]
Constant	-4.53669	-4.30393	0.07834
	[-3.48175]	[-4.29249]	[0.12673]
log(ism)	2.48159	1.12518	-0.19300
	[4.97987]	[2.93424]	[-0.81641]
$\log(ism_{-1})$	-0.98932	-0.33458	-0.35354
	[-1.45583]	[-0.63984]	[-1.09669]
$\log(ism_{-2})$	-0.34950	0.29533	0.52435
	[-0.68733]	[0.75479]	[2.17376]
R-squared	0.15533	0.20323	0.09176
S.E. equation	0.28309	0.21784	0.13429
F-statistic	4.15606	5.76441	2.28326

Owing to the weak exogeneity of the fundamentals, switching to a single equation error correction model (*SEECM*; Engle et al., 1983, Johansen, 1992), may still improve the efficiency of the estimate. 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) where the error correction model

[4] When exogeneity is tested for each variable separately the conclusions do not change:

and the cointegration relation are estimated simultaneously.[5] Thus, we estimate the following equation

$$(7) \qquad \Delta i_t^l = \alpha \cdot (i_{t-1}^l - \beta \cdot x_{t-1}) + \sum_{i=1}^{k} \chi_i \cdot \Delta i_{t-i}^l + \sum_{j=-m}^{k} \varphi_j \cdot \Delta x_{t-j} + \sum_{k=0}^{k} \phi_k \cdot w_{t-k} + \mu_t$$

where x is the vector of *I(1)*-variables i^s and π^e entering the cointegration space, w is a vector of (stationary) regressors only entering short-run dynamics, α is the error correction term and μ 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.[6] To obtain the standard deviation and the t-statistics of the long-run coefficients *ß*, the Bewley transformation of the model has to be estimated (West, 1988).

The first part of (7) with the variables in levels describes the cointegration relationship that has been normalised to the long-term interest rate. The *SEECM* differs from the *VECM* in that contemporary variables as well as leads of the (weakly) exogenous variables are added in order to improve the asymptotical properties of the estimates.[7] The lead length (*m*) is restricted to a maximum of two as recommended by Banerjee et al. (1998, 275); the lag length (*k*) is restricted to a maximum of four. The regression has been run with the so-called backward procedure, i. e. insignificant coefficients (error probability > 5 %) have been successively omitted. The final regression reads as (absolute t-values in parentheses)

$$(8) \qquad \begin{aligned} \Delta i_t^l = & -\underset{(4.6)}{0.19} \cdot (i_{t-1}^l - \underset{(5.1)}{0.35}\, i_{t-1}^s - \underset{(7.6)}{0.95}\, \pi^e + \underset{(2.7)}{13.50}) + \underset{(1.9)}{0.13}\, \Delta i_{t-1}^l + \underset{(6.9)}{0.54}\, \Delta i_t^s - \underset{(2.4)}{0.21}\, \Delta i_{t-1}^s \\ & + \underset{(4.0)}{1.89}\, ism_t - \underset{(2.6)}{1.17}\, ism_{t-1} + \mu_t \end{aligned}$$

sample: 1986:01-2005.09; R^2 = 0.30; SE = 0.26; LM(1) = 0.04; LM(4) = 1.03; ARCH(1) = 0.20; ARCH(4) = 1.65; JB = 0.73; CUSUM: stable; CUSUM square: stable.

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i^s: $\chi^2(1)$= 0.20, π: $\chi^2(1)$= 1.24.

[5] 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.

[6] 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.

[7] These approaches are based on Phillips and Lorethan (1991) as well as Saikkonen (1991). Chinn and Johnston (1997) apply this approach to empirical exchange rate modelling.

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. This result confirms the existence of the Fisher effect and is in line with Keeley and Hutchison (1986) who emphasise that this result is due to monetary regime stability. The Greenspan era on which we concentrate in this paper obviously was characterised 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 rates. The coefficient of -0.35 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 35 basis points. Accordingly, the term structure is going to flatten (see also Diebold et al., 2006). The less than proportional response of i^l to is i^s also found in Ducoudré (2005) in the case of the US. The overall effect 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 Diebod et al., 2006). The (contemporarily estimated) reaction of i^l to *ism* is positive and highly significant. In the short run, a contemporaneous increase of 1% in *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 also being driven, in the short run, by non-fundamental factors. This could be due to the market behaviour of chartists (Nagayasu 1999) whose interest rate forecasts are customarily based to some extent on past interest rate movements.

To explain the behaviour of bond yields, we do not need any of our *etc*-variables. After having taken into account the three factors discussed above, they have no significant influence. Concerning the fiscal variables, this is not an uncommon result (Ducoudré, 2005), but, e.g., contrasts with Caporale and Williams (2002) who use a different definition of the variables and a different frequency. The lacking effect of liquidity may be rationalised by the inclusion of expected inflation and a short-term interest rate which may dominate money or incorporate the effects of money on long-term interest rates (Smirlock, 1986). And finally, both for the long-run relation but also for short-run dynamics, stock market developments do not seem to matter much for bond markets. This is generally not true in models which use high frequency data (see, e.g. Ehrmann et al. 2005)

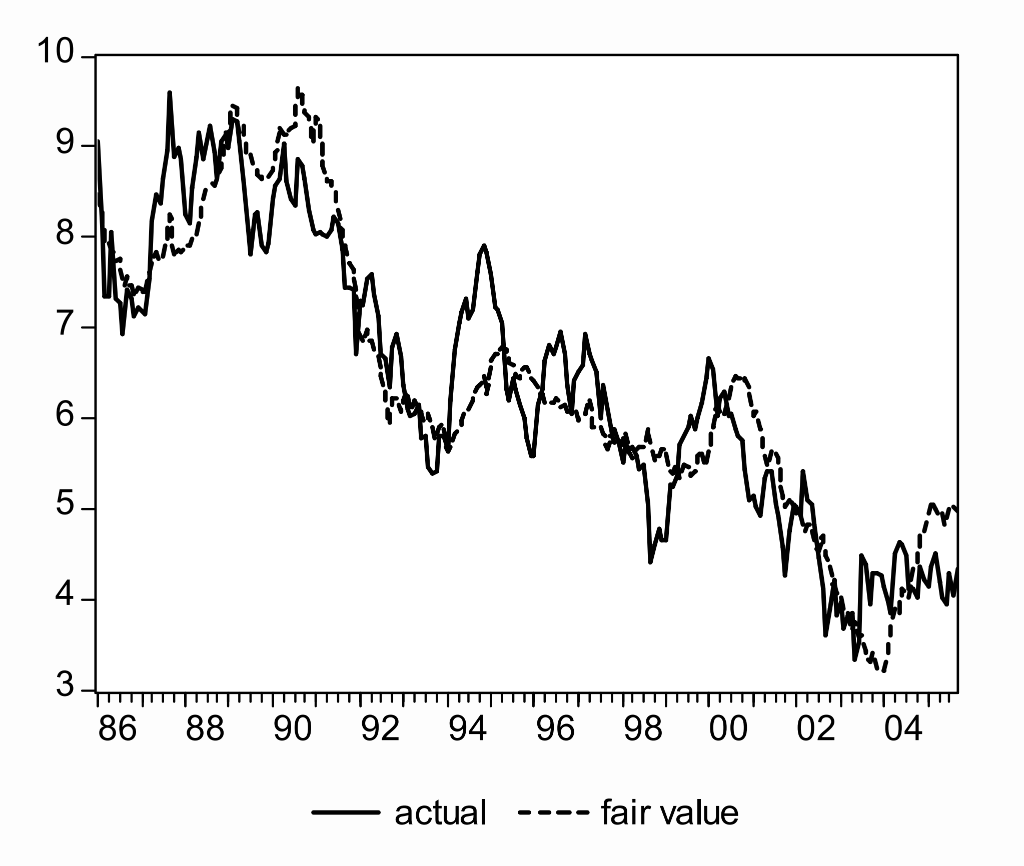
The coefficient of the error correction term is negative and, measured against the critical values of Banerjee et al. (1998), highly significant. Thus, the condition for a long-term stable equilibrium is satisfied. The parameter value of 0.185 suggests a half-life period of shocks of round about 3 months. In other words, the gap between the long-term nominal interest rate and its equilibrium value on basis of the cointegration relationship is closed by half in one quarter after an exogenous shock or is reduced by over 90% in a period of one year, respectively.[8]

Breusch-Godfrey Lagrange multiplier tests (LM) give no indication of autocorrelation in the residuals (1st and 4th order). Nor can the Lagrange multiplier (ARCH) test for autoregressive conditional heteroscedasticity (1st and 4th order) identify any violation of the basic assumptions of the regression. 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 reveals the stability of the found 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 this issue, we use our preferred model to calculate a "fair value" of bond yields. To be specifically, it is derived from our cointegration relation. Figure 3 shows that bond markets were overvalued in the course of 2004/2005. But obviously the "disequilibria" is not unusual in historical perspective. Our three macroeconomic factors seem to capture the evolution of bond yields quite well. Therefore, it is not necessary to recur to a changing structural demand for bonds. The fair value in September 2005 is 4.97 %, compared to an actual value of 4.33 %.

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

Figure 3: The fair value of bonds



3.3 Forecast evaluation

In order to check 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), this model has become a very popular benchmark in forecast evaluation. In line with the unit root tests the *RWM* is specified without a constant and a 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. In doing this, we assume that the forecaster has no idea as how the exogenous variables evolve and bases his predictions of them on completely endogenous structures. Thus, the forecasts include only information that had actually been available at the time it was carried out. In the case of our *SEECM*, forecast values of the exogenous variables are generated on the basis of univariate ARIMA-models or in case of "*ism"* on the basis of an ARMA-model. The AR-/MA structure is chosen by AICC, a bias-corrected version of the Akaike information criterion (Brockwell and Davis, 2002, 171). In contrast to this narrow information set the second approach assumes that the exogenous variables for the forecasting horizon are known to the forecaster. Realistically, the actual forecasting environment should be somewhere between these two cases.

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

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

The forecasts are carried out recursively, the "first" forecast period runs from 1994:11 to 1995:10. The forecast "window" is successively extended month by month. Consequently, the next forecast period is from 1994:12 to 1995:11. And the last forecast period is from 2004:10 to 2005:9. Doing so, we get 120 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) $$RMSE_h = \sqrt{\frac{1}{T}\sum_{t=1}^{T} e^2_{t+h,t}}$$

A smaller *RMSE* implies better forecast performance. A formal test based on the so-called loss differential (Diebold and Mariano, 1995) provides information on the significance of the relative forecasts. Assuming a quadratic loss function, the loss differential is defined as the squared forecast error of the benchmark model (the *RWM* in our case) minus the squared forecast error of the *SEECM*. The test is pursued by regressing the loss differential solely on a constant term. If the constant term is significantly positive the *SEECM* provides significantly better forecasts than the *RWM*.

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

$$if\ \ sign\left(i^l_{t+h} - i^l_t\right) = sign\left(i^l_{t+h,t} - i^l_t\right) \Leftrightarrow J = 1$$
$$if\ \ sign\left(i^l_{t+h} - i^l_t\right) \neq sign\left(i^l_{t+h,t} - i^l_t\right) \Leftrightarrow J = 0$$

Therefore, our hit ratio *HR* is defined as

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

The higher the *HR*, the more often the forecast value signals the correct direction of interest rate changes.[9] 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 correct direction of change - 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 for the two longest forecast horizons (h=11, 12). Overall, the results underpin the superiority of the *SEECMs*, especially the longer the forecast horizon is. Moreover, it is obvious that the *SEECM* does a better job the better are the forecaster's predicting abilities of 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	27.22	0.38	56.67	0.19	25.01	0.02	61.67	0.05
2	39.52	0.28	58.33	0.08	33.64	0.00	70.00	0.00
3	46.48	0.40	63.33	0.03	35.58	0.00	76.67	0.00
4	53.50	0.33	67.50	0.00	38.41	0.00	79.17	0.00
5	58.06	0.22	60.00	0.07	40.42	0.00	80.00	0.00
6	61.09	0.15	60.83	0.07	41.70	0.00	83.33	0.00
7	64.07	0.14	64.17	0.02	42.61	0.00	80.83	0.00
8	66.44	0.17	62.50	0.04	43.17	0.00	83.33	0.00
9	68.20	0.15	60.83	0.10	44.51	0.00	77.50	0.00
10	70.51	0.12	64.17	0.02	45.44	0.00	73.33	0.00
11	72.15	0.08	67.50	0.00	46.51	0.00	76.67	0.00
12	73.39	0.05	67.50	0.00	47.39	0.00	74.17	0.00

4. Summary and conclusions

Our results reveal that the development of long-term bond yields in the US may very well be explained by three standard macroeconomic factors which are widely considered to be the minimum set of fundamentals needed to capture basic

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

macroeconomic dynamics (see also Diebold et al., 2006): Monetary policy, the business cycle and inflation expectations. These three variables are able to capture the development of bond yields in a stable manner and outperform a random-walk model in different forecasting exercises. Further macro variables are not needed or do not exert a significant influence. We also need not recur to changed structural factors, e.g. a changed demand by institutional investors, in capturing the evolution of bond yields in 2004/2005.

Our forecasting exercises show that we are able to outperform a simple random walk model with our two evaluation criteria. In these tests the fully-dynamic approach assumes an information set of the forecaster with regard to the exogenous variables which, in real world applications, is obviously conservative. On the other side, the perfect foresight case neglects informational deficiencies. The simple random walk model which we use may be criticised to be too "naive" in that it may 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 in further evaluation metrics, e.g. a consistency criterion, to check the robustness of our results. This is left to future research.

Compared to our equilibrium or fair-value concept, bond yields since fall 2004 were lower than expected by our macroeconomic factors. But the disequilibria we get are not unusual in a historical context. However, there are hints of some instabilities in the last years indicated by a Chow breakpoint test. This suggests to be cautious in drawing too far-reaching conclusions.

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Biographie

Prof. Dr. Jörg Clostermann ist Professor für Volkswirtschaftslehre, Quantitative Methoden und Finanzmarktanalyse an der Fachhochschule Ingolstadt. Vor seiner Berufung war er in der Hauptabteilung Volkswirtschaft der Deutschen Bundesbank, Abteilung Außenwirtschaft, in Frankfurt tätig. Seine Interessens- und Beratungsschwerpunkte sind Prognosemodelle und Finanzmarktanalysemodelle.

Contact: joerg.clostermann@fh-ingolstadt.de

Prof. Dr. Franz Seitz lehrt Volkswirtschaftslehre an der Fachhochschule Amberg-Weiden. Vor seiner Berufung war er in der Hauptabteilung Volkswirtschaft der Deutschen Bundesbank, Abteilung Geld, Kredit und Kapitalmarkt, in Frankfurt tätig. Er fungiert inzwischen auch als Consultant der Europäischen Zentralbank und der deutschen Bundesbank.

Contact: f.seitz@fh-amberg-weiden.de

Impressum

Herausgeber
Der Präsident der
Fachhochschule Ingolstadt

Esplanade 10
85049 Ingolstadt
Telefon: 08 41 / 93 48 - 0
Fax: 08 41 / 93 48 - 200
E-Mail: info@fh-ingolstadt.de

Druck
Hausdruck

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ISSN 1612-6483