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Abstract

The expectations hypothesis of the term structure (EHT) implies cointegration between interest rates of different maturities and predicts certain values for adjustment speed. We estimate reduced-form vector error correction models of the US term structure. These are derived from a structural model combining the EHT, autocorrelated risk premia, interest rate smoothing and monetary policy feedback, which is able to capture a wide range of empirical outcomes. We explicitly test the necessary preconditions for the validity of the theoretical model. Premia persistence rises with longer-rate maturity, while the influence of the according spreads in the central bank reaction function diminishes.

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1 Introduction

The term structure, linking short- and long-term interest rates, carries evident importance for transmission of monetary policy. The expectations hypothesis (EHT) represents the most influential theoretical explanation for term structure relations. A lot of econometric research has been dedicated to examining empirical evidence of the EHT. Thereby, the sizeable predictive power of the spread, as implied by the pure EHT version, could regularly not be supported by the data (e.g., Campbell and Shiller 1991).

Among the numerous approaches dealing with this puzzle, persistent risk premia and interest rate smoothing are well established. For instance, Tzavalis and Wickens (1995) specify a GARCH-dependent term premium, while Caporale and Caporale (2008) explain shifts in premia by political risk. Mankiw and Miron (1986) argue that central bank behaviour is at the bottom of short-term rates being close to random walks. McCallum (2005) constructs a term structure model combining autocorrelated risk premia, interest rate smoothing and monetary policy feedback.

In the following, we analyse the whole range of maturities from one to 120 months, in order to deduce empirical regularities of the term-structure interaction between expectations, risk premia dynamics and monetary policy actions. We provide a solid empirical basis for the theoretical model by explicitly testing for cointegration along the term structure and checking the parameter restrictions implied by the use of the spread. Furthermore, we investigate the adjustment behaviour of both short and long rates by deriving and estimating the vector error correction (EC) form of the structural model.

2 Theoretical Considerations

The EHT states that arbitrage should equate the return of a single n-period investment to the expected return of a series of n successive one-period investments. In linearised form, the EHT can be written as

$$R_t^{(n)} = \frac{1}{n} \sum_{i=0}^{n-1} E_t r_{t+i} \ . \qquad (1)$$

In (1), $R_t^{(n)}$ denotes the interest rate of maturity $n > 1$ and r_t of maturity one. The operator E_t stands for the expectation given all available information at time t.

Taken into consideration that interest rates may be empirically approximated by integrated processes of order one (I(1)), Campbell and Shiller (1987) show that (1) implies cointegration between $R_t^{(n)}$ and r_t with the vector $(1, -1)$. Thereby, long-term rates will be weakly exogenous, approximately following random walks (see Pesando 1979). Assuming rational expectations, $E_t r_{t+1} = r_{t+1} + e_{t+1}$, with the expectation error e_t, the EC equation for the short-term rate in the two-period case ($n = 2$) results as

$$\Delta r_t = 2(R_{t-1}^{(2)} - r_{t-1}) - e_t \ . \tag{2}$$

However, empirical evidence (e.g., Campbell and Shiller 1991) has overwhelmingly shown that the considerable size of the adjustment coefficient is not supported by the data. One important modification of the pure EHT is given by autocorrelated risk premia. Adding a maturity-dependent AR(1) process[4] with constant term, (1) can be rewritten as

$$R_t^{(n)} = \frac{1}{n} \sum_{i=0}^{n-1} E_t r_{t+i} + v_t^{(n)} \quad \text{with} \quad v_t^{(n)} = a_n + \rho_n v_{t-1}^{(n)} + u_t^{(n)} \ , \ 0 < \rho_n < 1 \ . \tag{3}$$

The EC equation (2) now takes the form

$$\Delta r_t = 2(1 - \rho_2)(R_{t-1}^{(2)} - r_{t-1}) + rest \ , \tag{4}$$

where $rest$ contains a constant, short-run dynamics and errors. Besides the maturity n, the persistence ρ_n of the risk premium governs the adjustment coefficient, which can be rendered arbitrarily small.

A second explanation for the failure of the pure EHT is provided by interest rate smoothing. Mankiw and Miron (1986) argue that the predictive power of the spread has dramatically declined since the founding of the Federal Reserve System and attribute this fact to the interest rate stabilisation policy of the Fed. Approximately, the short rate may even follow a random walk, impairing predictability. In sum, the literature offers explanations for random-walk behaviour both of long and short rates. Notwithstanding that resolving this ambiguity is in the end an empirical task, we set up a theoretical framework that encompasses both hypotheses, in addition to intermediate solutions.

For that purpose, we recur to McCallum (2005), who proposes a term structure model

[4]Generalisations are straightforward. Note that constant premia would not change the adjustment parameter in (2).

that combines the EHT with a monetary policy rule: The central bank conducts interest rate smoothing, but may change the short rate in response to the yield spread. Thereby, the spread may provide an indicator for monetary policy expansiveness, future economic growth or expected inflation rates, approximating forward-looking counter-cyclical policy behaviour; see also Johnson (1988) in this context. The reaction function with interest rate smoothing, feedback intensity λ_n and policy shock $\varepsilon_t^{(n)}$ is given by

$$r_t = r_{t-1} + \lambda_n (R_t^{(n)} - r_t) + \varepsilon_t^{(n)} \,. \tag{5}$$

Importantly, equation (5) will only be balanced for I(1) interest rates if long and short rates are cointegrated with the vector $(1, -1)$, as follows from the EHT, too. According to the Granger representation theorem, cointegration implies an EC model. As solution of equations (3) and (5), McCallum (2005) derived for $n = 2$ the EC equation for the short rate.[5] Kugler (1997) generalised his result for any maturity, presenting according equations for the short rate and the spread. We develop the approach one step further, deriving the bivariate vector EC model (VECM):

$$\begin{pmatrix} \Delta R_t^{(n)} \\ \Delta r_t \end{pmatrix} = \begin{pmatrix} (1+\lambda_n)\theta_n a_n \\ \lambda_n \theta_n a_n \end{pmatrix} + \begin{pmatrix} \lambda_n \rho_n + \rho_n - 1 \\ \lambda_n \rho_n \end{pmatrix} (R_{t-1}^{(n)} - r_{t-1}) + \begin{pmatrix} 1 & (1+\lambda_n)\theta_n \\ 1 & \lambda_n \theta_n \end{pmatrix} \begin{pmatrix} \varepsilon_t^{(n)} \\ u_t^{(n)} \end{pmatrix} \,, \tag{6}$$

where $\theta_n = n/(n - \lambda_n \sum_{j=1}^{n-1} (n-j)\rho_n^j)$. Evidently, the spread has predictive power for the short rate as long as $\lambda_n \neq 0$ and $\rho_n \neq 0$. In (5) and (6), $\lambda_n = 0$ would render the short rate a random walk (up to possible short-run dynamics), triggering the Mankiw and Miron (1986) result of no predictability. The same holds for the limiting case $\rho_n = 0$, when the long rate would contain no distinct persistence in the risk premium.

The adjustment of $R_t^{(n)}$ reduces the spread for small λ_n and ρ_n ($\lambda_n \rho_n + \rho_n - 1 < 0$), but otherwise the long rate even departs further from equilibrium ($\lambda_n \rho_n + \rho_n - 1 > 0$). However, system stability is ensured for stationary term premia, as the characteristic equation has roots 0.5 and ρ_n. Long-rate predictability would be lost for $\lambda_n \rho_n + \rho_n = 1$. Notably, for $\lambda_n = 0$ and $\rho_n \to 1$, both interest rates would behave like random walks.

[5] As already provided in NBER Working Paper 4938 (1994).

3 Empirical Evidence

Our data set consists of monthly observations of certificate of deposit rates ($n = 1, 3, 6$) and constant maturity bond yields ($n = 12, 24, 36, 60, 84, 120$) obtained from the Fed. The sample begins in 1983(1) after the abolishment of the non-borrowed reserves targeting policy, and ends in 2008(9). We begin an additional sub-sample in 1988(1) after the start of the Greenspan era and the stock market crash. In this context, refer as well to Hsu and Kugler (1997), who argue that the spread gained importance as a policy indicator from 1988 on and estimate the Kugler (1997) model for the three-period case.

We check, whether the implications from EHT are fulfilled, within the VECM

$$\begin{pmatrix} \Delta R_t^{(n)} \\ \Delta r_t \end{pmatrix} = \begin{pmatrix} \alpha_1 \\ \alpha_2 \end{pmatrix} (R_{t-1}^{(n)} - \beta r_{t-1} - c) + \sum_{i=1}^{p} \Gamma_i \begin{pmatrix} \Delta R_{t-i}^{(n)} \\ \Delta r_{t-i} \end{pmatrix} + \begin{pmatrix} w_{1t}^{(n)} \\ w_{2t}^{(n)} \end{pmatrix} , \tag{7}$$

with adjustment parameters α, cointegration coefficient β, errors w and 2×2 parameter matrices Γ_i in the short-run dynamics. Moreover, estimates of ρ_n and λ_n can be recovered by comparing adjustment parameters from (6) and (7), see Table 1.

	1983(1)-2008(9)						1988(1)-2008(9)					
n	lags	$\hat{\beta}$	$\hat{\alpha}_1$	$\hat{\alpha}_2$	$\hat{\rho}_n$	$\hat{\lambda}_n$	lags	$\hat{\beta}$	$\hat{\alpha}_1$	$\hat{\alpha}_2$	$\hat{\rho}_n$	$\hat{\lambda}_n$
3	1-2, 12	1.005	0.306	0.652	0.654	1.000	1-2, 12	1.001	0.447	0.844	0.703	1.207
		(0.006)	(0.15)	(0.14)				(0.01)	(0.16)	(0.14)		
6	1	1.010	0.112	0.419	0.693	0.605	1, 12	1.004	0.239	0.520	0.719	0.723
		(0.01)	(0.07)	(0.07)				(0.01)	(0.07)	(0.06)		
12	1	1.019	0.040	0.152	0.887	0.172	1, 12	0.978	0.066	0.233	0.833	0.280
		(0.04)	(0.03)	(0.03)				(0.03)	(0.03)	(0.03)		
24	1, 12	1.047	0.004	0.094	0.911	0.103	1, 12	0.983	0.011	0.143	0.868	0.165
		(0.06)	(0.02)	(0.02)				(0.04)	(0.02)	(0.02)		
36	1, 12	1.035	−0.007	0.074	0.919	0.081	1-3, 12	1.030	0.003	0.085	0.918	0.093
		(0.07)	(0.02)	(0.02)				(0.07)	(0.02)	(0.02)		
60	1, 12	1.002	−0.014	0.051	0.935	0.055	1-3, 12	1.007	−0.008	0.065	0.927	0.070
		(0.10)	(0.01)	(0.01)				(0.09)	(0.02)	(0.01)		
84	1-3, 12	1.040	−0.011	0.043	0.946	0.045	1-3, 12	0.993	−0.008	0.060	0.933	0.064
		(0.12)	(0.01)	(0.01)				(0.10)	(0.01)	(0.01)		
120	1-3, 12	1.012	−0.013	0.039	0.948	0.042	1-3, 12	0.966	−0.010	0.055	0.935	0.059
		(0.13)	(0.01)	(0.01)				(0.11)	(0.01)	(0.01)		

Notes: n: maturity in months, β: cointegration coefficient, α: adjustment coefficient, ρ: risk premium AR(1) coefficient, λ: feedback coefficient, standard errors in parentheses

Table 1: VECM results

Johansen Trace tests uniformly reject the null hypotheses that the cointegration rank is zero at the 1% significance level, except for the three longest maturities in the first period, where the 5% level applies. LR tests in no case reject the null hypothesis $\beta = 1$. The presence of cointegration together with the parameter restriction ensure stationarity of the spreads. Importantly, this provides the necessary condition for the validity of the structural specification (3), (5).

Adjustment of the one-period rate is faster than that of the longer-term rates ($\hat{\alpha}_2 > |\hat{\alpha}_1|$), but reduces with increasing maturity. The long rate adjustment coefficients are "wrong"-directed for maturities until two (three) years and become slightly negative thereafter. While $\hat{\alpha}_2$ is significant in all estimated VECMs, this holds for $\hat{\alpha}_1$ only until the maturity of six (twelve) months. In our model, this reduced-form outcome results from underlying structural processes, namely risk premia persistence and policy feedback. The former is of substantial strength even for relatively short maturities and rises to nearly 1 for the ten-year bond. Intuition is straightforward: When one month passes, the fraction ρ of the risk premium survives until the next period. This fraction will be the higher the less a single period counts for the whole premium (for example only 1/120 for the ten-year bond) and the less news from this period count for the overall outlook until maturity. In contrast, the impacts λ of the respective spreads in the central bank reaction function start at about one for the three-month rate and approach zero for long maturities. Again, this pattern is in line with a priori expectations, as monetary policy is likely to concentrate its forward-looking behaviour on near-future outlooks. In this sense, the spreads against relatively short rates provide immediately relevant information.

Risk premia persistence stays about the same in both estimation periods. This is not surprising, because the sample partition was largely guided by criteria connected to monetary policy. Indeed, the feedback coefficients are uniformly higher from 1988 on, confirming the presumption of increased significance of the spread for central bank decisions. Furthermore, as far as the spread incorporates inflation expectations, the result can generally be interpreted in the sense of higher inflation sensitivity often supposed to characterise the Greenspan era.

4 Conclusion

We adopt a structural framework of McCallum (2005) and Kugler (1997) in order to model simultaneously bond market arbitrage and central bank behaviour. We derive the vector EC form from the structural equations, test whether essential preconditions are fulfilled and estimate the VECM for a wide range of the US term structure.

We find that persistence of risk premia is the higher the larger the difference in maturity between the considered interest rates. This result probably reflects the fact that the weight of news from a single period, relative to the incumbent premium, is naturally lower for long-horizon risk assessments. Following the stylised policy rule, the Fed reacts strongly to spreads against money-market rates, but much less against long-term bond rates. Since it is common sense that central banks align operative policy to prospects of the near future, again the potential of the model to grasp real-world phenomena is underlined.

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