Technical Paper

1/RT/00 **May 2000**

The Expectations Hypothesis of the Term Structure: The Case of Ireland

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The views expressed in this paper are the personal responsibility of the authors and are not necessarily held either by the Central Bank of Ireland or by the ESCB. All errors and omissions are the authors'. Comments and criticisms are welcome.

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Abstract

Using Irish money market rates (spot rates) with a term to maturity of 1, 3, and 6 months and monthly data, 1984-1997, we provide a number of tests of the expectations hypothesis (EH) of the term structure. The paper draws on co-integration techniques and the methodological approach of Campbell and Shiller (1987, 1991). On balance our results lend support to the EH and are broadly consistent with the recent findings for the UK, but are in sharp contrast to those for the US. It is encouraging that our results are consistent with those of recent studies at the short end of the maturity spectrum for the UK, (e.g. Cuthbertson, 1996).

1. Introduction

The expectations hypothesis (EH) of the term structure implies that the yield spread between the long rate and short rate is an optimal predictor of future changes in short rates over the life of the ‘long bond’.

There is a great deal of evidence on the EH for the US, based on the Campbell and Shiller (1991) VAR methodology using monthly data on spot rates (e.g. Hardouvelis, 1994). In general, for a wide variety of maturities from 1 to 12 months and for 2, 3, 4 ... 10-year maturities, Campbell and Shiller (1991) reject the EH. The (long-short) interest rate spread does not predict the direction of changes in the long-term interest rate that is consistent with the EH, and future changes in short-rates are not often closely correlated with the long-short spread (Campbell and Shiller, 1991).

Kugler (1988) using US, German and Swiss monthly data on one and three month Euromarket deposit rates found support for the EH only on German data (for the period of March 1974 to August 1986). Similarly, Engsted (1994) using Danish money market rates and (Engsted and Tanggaard, 1994) for longer maturity bonds find considerable support for the EH when the variation in interest rates is relatively large, such as in the post-1992 ERM ‘crisis period’. This is to be expected following the analysis of Mankiw and Miron (1986), for if interest rate stabilisation results in random walk behaviour for short rates, then the expected change in short rates is zero and the spread has no predictive power for future short rates (See also Rudebusch, 1995).

Using the Campbell-Shiller VAR methodology on data at the short end of the maturity spectrum (i.e. up to one year) Cuthbertson (1996) finds reasonable support for the EH on UK data. This is in contrast to Taylor (1992) who uses maturities of 5, 10 and 15 years and finds strongly against the EH (see also MacDonald and Speight, 1991). To our knowledge, the only related paper using Irish data, is that of Hurley (1990).[1] Using a number of interest rate combinations for the period 1979 to 1989, Hurley (1990) finds little evidence to support the EH. Although there is slight support using short rates, the predictive power is low.

However, as noted by the author, the study suffers a number of drawbacks. Firstly the data used in the study are yields taken from the Central Bank Bulletin and so are not properly compounded. Secondly, author also uses OLS estimation, which is inappropriate, given the inclusion of overlapping observations.

The main aim of this paper is to present evidence on the behaviour of the term structure of Irish interest rates at the short end of the maturity spectrum. The paper applies co-integration techniques and the methodological approach of Campbell and Shiller (1987, 1991).[2] To our knowledge the expectations hypothesis (EH) using the VAR approach has not been examined using Irish data. We test parameter restrictions on the VAR models using a high quality data set. By using spot rates based on quoted discount rates we avoid having to use an approximation to zero

[1] McGettigan (1995) drawing on the approach of McCulloch (1971, 1975) estimates the term structure of interest rates in Ireland indirectly using discount functions. As noted by McGettigan a similar approach was used by Breen and Keogh (1990). Breen (1991) in his comment on Joyce (1991) refers to the move towards the term structure, as opposed to the yield curve concept.

[2] Earlier work may be found in Melino (1988) and Shiller (1989).

coupon yields and the par yield approximation which are required when analysing coupon paying bonds (Shiller, 1989). We also assess the results in comparison to the previous evidence.[3]

2. The Expectation Hypothesis

The Expectations hypothesis (EH) of the term structure posits that the return on an n-period bond $R_t^{(n)}$ is determined solely by expectations of (current and) future rates on a set of m-period assets $r_t^{(m)}$ where n > m. Using the continuously compounded spot rates the 'fundamental term structure' relationship is:

$$R_t^{(n)} = (1/k)\sum_{i=0}^{k-1} E_t r^{(m)}_{t+im} \tag{1}$$

where k = n/m is an integer. Reparameterising (1) enables the spread to be interpreted as the optimal predictor of future changes in short rates, $r_t^{(m)}$:

$$S_t^{(n,m)} = E_t \sum_{i=1}^{k-1} (1\text{-}i/k)\Delta^m r^{(m)}_{t+im} = E_t[PFS_t^{(n,m)}] \tag{2}$$

where $S_t^{(n,m)} = (R_t^{(n)} - r_t^{(m)})$ is the yield spread. For example, if m = 1 the above equation states that the spread between an n-period bond and a one-period bond equals the expectation of the change in one-period rates over

[3] Evans and Lewis (1994), Hardouvelis (1994), Campbell and Shiller (1991) and Shea (1992) use monthly data on a wide variety of yields for the United States. Engsted and Tanggaard (1994) and Engsted (1994) report results for short and long maturities using Danish data. Most of the above require an approximation to estimate some parts of the yield curve McCulloch (1987) and Shea (1984). Recent studies for the UK include Cuthbertson (1996) and Cuthbertson et al (1996).

the n-period horizon. The *Perfect foresight spread* is the spread that would be predicted by agents if they had perfect foresight about future movements in interest rates. A testable implication of equation 2 is that the spread, Granger causes future changes in short rates.[4] Assuming that ($R_t^{(n)}$, $r_t^{(m)}$) are I(1), then $\Delta r_t^{(m)}$ is I(0), which from equation 2 implies that the spread $S_t^{(n,m)}$ should also be I(0), and therefore ($R_t^{(n)}$, $r_t^{(m)}$) should be co-integrated with a co-integrating parameter of unity.[5]

If we now add the assumption of rational expectations (RE):

$$r_{t+im}^{(m)} = E_t r_{t+im}^{(m)} + \varepsilon_{t+im} \qquad (3)$$

we obtain the following single equation test of the null of the 'expectations hypothesis plus rational expectations', EH + RE:

$$PFS_t^{(n,m)} = \alpha + \beta S_t^{(n,m)} + \gamma\Omega_t + \varepsilon^*_t \qquad (4)$$

$$H_0: \alpha = \gamma = 0, \beta=1$$

where e^*_t is a moving average error of order (n-m-1) consisting of a weighted sum of future values of e_{t+im} . Under RE, e^*_t is independent of information at time t, W_t, and in particular is independent of the yield spread. If there is a constant term premia or if there are differential yet constant transactions costs (between investing 'long' and in a series of rolled-over short-term investments) then $\alpha \neq 0$. Under RE the right hand

[4] Strictly, failure of Granger causality does not constitute a rejection of the EH, but a failure to confirm it.

side variables in equation 4 are independent of ε^*_t and hence we do not require IV estimation. However a GMM estimator is employed to correct the covariance matrix for the moving-average error of order (n-m-1) and possible heteroscedasticity (Hansen, 1982; Newey and West, 1987).

3. VAR Methodology

If $Z_t = (S_t^{(n,m)}, \Delta r_t^{(m)})$ is stationary, then there exists a bivariate Wold representation (Hannan, 1970) which may be approximated by a vector autoregression (VAR) of order *p*, which in companion form is:

$$Z_t = AZ_{t-1} + v_t \qquad (5)$$

Using $\mathbf{e1'Z}_t = S_t^{(n,m)}$ and $\mathbf{e2'Z}_t = \Delta r_t^{(m)}$ in equation (2) (see Campbell and Shiller, 1991 and Taylor, 1992) we obtain the following restrictions:

$$e1' = e2'A\,[\,I - (m/n)(I - A^n)\,(I-A^m)^{-1}]\,(I - A)^{-1} \qquad (6)$$

We apply equation 6 on monthly data for (n,m) = (6,3), (6,1), (3,1). The restrictions in the parameters of the VAR in equation 6 are tested using a Wald test.[6] The VAR methodology suggests several approaches to testing the EH + RE under weakly rational expectations: (i) information at time t other than that contained in $S_t^{(n,m)}$ should not help to predict future changes in short rates (i.e. $S_t^{(n,m)}$ must be an optimal predictor of future changes in

[5] Strictly, for this to hold, forecast errors and any term premia must also be I(0). Shea (1992) examines the possibility of multiple co-integrating vectors, an issue not explored here since we concentrate on tests of bilateral relationships.

short rates), (ii) $S_t^{(n,m)}$ should Granger-cause (future) changes in short rates, (iii) if we define the 'theoretical spread' by $S_t^{(n,m)\prime}$ where,

$$S_t^{(n,m)\prime} = e2'A\,[\,I - (m/n)(I - A^n)\,(I-A^m)^{-1}]\,(I - A)^{-1}Z_t \qquad (7)$$

then the VAR restrictions test the hypothesis, H_0: $S_t^{(n,m)} = S_t^{(n,m)\prime}$. Campbell and Shiller (1991) note that formal tests of the VAR restrictions may lead to rejection of the EH even though the deviations from the null are quite small from an economic perspective. They suggest computing the theoretical spread (i.e. without imposing the VAR restrictions) and comparing the time series behaviour of the actual and theoretical spread using the standard deviation ratio SDR = $\sigma(S_t^{(n,m)\prime})/\ \sigma(S_t^{(n,m)})$, and the correlation coefficient, Corr($S_t^{(n,m)}$, $S_t^{(n,m)\prime}$).[7] If the EH plus weakly rational expectations holds then this should be reflected in SDR and Corr being close to unity (and hence the graphs of the 2 series moving together). If $\sigma(S_t^{(n,m)}) > \sigma(S_t^{(n,m)\prime})$ then there is 'excess volatility', that is the actual spread is more volatile than the optimal predictor of future short rates.

4. Empirical Results

4.1 The Data

The data used consist of monthly Irish money market rates (spot rates) with a term to maturity of less than 6 months. These rates were kindly provided by the Bank of Ireland from screen-quoted rates. The data set is from January 1984 to October 1997. The 1 month and the 6 month rates are graphed in figure 1.

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[6] In testing the VAR restrictions we use a GMM correction to the covariance matrix of the VAR system.

4.2 Unit Roots and Co-integration

Table 1 gives the results of unit root tests on the individual series R_t and $S_t^{(n,m)}$, which indicate that we cannot reject the null hypothesis that changes in short rates $\Delta r_t^{(m)}$ and the yield spread $S_t^{(n,m)}$ are I(0). Table 2 shows the OLS co-integration regression results and as can be seen the β in $R_t^{(n)} = \alpha + \beta r_t^{(m)}$ is 'banded ' by the 2 regressions (i.e. $R_t^{(n)}$ on $r_t^{(m)}$ and vice-versa)[8]. This result provides weak evidence in favour of the EH under the assumption of a constant or stationary term premium and any expectation scheme that yields I(0) forecast errors.

We now turn to tests based on the Johansen (1988) procedure. The Johansen results, shown in table 3, provide strong evidence that $R_t^{(n)}$ and $r_t^{(m)}$ are co-integrated and we cannot reject the hypothesis that the co-integrating vector is given by the theoretical value (-1,1). The (normalised) point estimates for the co-integrating vectors from the Johansen procedure are (-1, 0.99) for each case.

4.3 The Spread and the predictability of Changes in Short Rates

The regression of the perfect foresight spread, $PFS_t^{(n,m)}$ on the actual spread $S_t^{(n,m)}$ and the limited information set H_t (consisting of lags of $S_t^{(n,m)}$ and $\Delta r_t^{(m)}$) are shown in table 4. In all cases we do not reject the null of $H_0 : \beta=1$ or that information, available at time t or earlier does not incrementally add to the predictions of future interest rates. We also do not

[7] The standard errors of SDR = $\sigma(S_t^{(n,m)'})/ \sigma(S_t^{(n,m)})$ and Corr ($S_t^{(n,m)}, S_t^{(n,m)'}$) are non-linear functions of the estimated A matrix from the VAR and can be computed as $[f_\gamma(\gamma)' \Psi f_\gamma(\gamma)]$ where $f(\gamma)$ are the statistics of interest and Ψ is the (GMM) variance-covariance matrix of the parameters γ.

[8] Hall (1986) suggests that the co-integration regressions of 'y_t on x_t' and 'x_t on y_t' should provide upper and lower limits for the co-integration parameter.

reject the null that the constant term premium is zero (i.e. $\alpha = 0$). The results therefore do not reject the EH + RE.

4.4 The Theoretical Spread and the VAR Results

Table 5 contains the results from the VAR models for $S_t^{(n,m)}$ and $\Delta r_t^{(m)}$. The lag length is chosen to minimise the Akaike Information Criterion (AIC), except for the rare occasions when additional lags are required to avoid any serial correlation in the residuals. A weak test of the EH is that the spread Granger-causes changes in short-term interest rates and this is not rejected for all maturities (table 5, column 3). There is also Granger-causality from $\Delta r_t^{(m)}$ to $S_t^{(n,m)}$ for the (6m, 3m) case, indicating feedback in the VAR regression.

For illustrative purposes the graph of the actual spread S_t and the theoretical spread S_t' are shown for (n,m) = (6, 1) and they move closely together, figure 2. In the regression of S_t on S_t' (table 6) the point estimate of the slope coefficients are very close to unity, for all 3 maturity combinations. The intercepts in these regressions are not statistically significantly different from zero in each case. Table 7 provides the metrics for the relationship between the actual spread S_t and the theoretical spread S_t'. The results indicate that the VAR restrictions are not rejected. For all maturities there is a strong correlation (column 4) between the actual spread $S_t^{(n,m)}$ and the predicted (theoretical) spread $S_t^{(n,m)\prime}$. The standard deviation ratio, SDR = $\sigma(S_t^{(n,m)\prime})/ \sigma(S_t^{(n,m)})$ yields estimates (column 3) which are all within two standard errors of unity. The Wald test of the VAR cross equation parameter restrictions (table 7, column 2), are not rejected for the (6m, 3m) and (6m, 1m) case, but are rejected for (3m, 1m).

On balance our results would appear to lend support to the EH, but how do we interpret the rejection of the VAR restrictions for the (3m, 1m) combination. Campbell and Shiller (1987) show that rejection implies either (1) information ($H_t \subset \Omega_t$) at time t or earlier (other than $S_t^{(n,m)}$) influences future changes in short-rates or (2) the influence of the current spread $S_t^{(n,m)}$ on future changes in interest rates via the chain rule of forecasting is less than required by the EH. However in contrast to (1) our single equation perfect foresight regressions in table 4 reject the null that H_t influences future changes in interest rates. Hence rejection of the VAR restrictions is probably due to the 'low weight' given to $S_t^{(n,m)}$ in the VAR regression. This may occur because over short forecasting horizons one might expect agents to utilise almost minute by minute observations of $S_t^{(n,m)}$ and $\Delta R_t^{(m)}$ and hence forecasts based on monthly data might not adequately mimic such behaviour for the (3m, 1m) maturities.

5. Interpretation of Results

In this section we analyse our results and compare them with those from other studies. On balance our results favour the EH. The perfect foresight regressions (table 4), the standard deviation ratios $\sigma(S_t^{(n,m)\prime})/\sigma(S_t^{(n,m)})$, and the coefficient of determination $Corr(S_t^{(n,m)}, S_t^{(n,m)\prime})$ (table 7), are consistent with the EH and this may be contrasted with the rejection of the VAR cross-equation restrictions for the (3m, 1m) combination, see table 7.[9]

First, the perfect foresight spread regressions implicitly allow potential future events (known to agents but not to the econometrician) to influence expectations, whereas the VAR approach requires the explicit information set known both to agents and the econometrician. Hence, if the econometrician erroneously excludes variables affecting traders' perceptions, then the estimated VAR coefficients may be biased, resulting in rejection of the VAR cross equation restrictions. Secondly, if agents actually do use the VAR methodology for forecasting, one would expect them to utilise almost minute by minute observations of ($S_t^{(n,m)}$, $\Delta r_t^{(m)}$) : hence forecasts based on monthly data seem unlikely to adequately mimic such behaviour. Thirdly, Campbell and Shiller (1991) have argued that rejection of the cross-equation parameter restrictions although statistically significant may not constitute a major departure form the EH on economic grounds, as long as the theoretical spread closely tracks the actual spread.[10]

Our perfect foresight spread results are broadly consistent with the results in Cuthbertson (1996); who found slope coefficients ranging from 0.73 to 1.3. The author could not reject the null, H_0:β=1, γ=0 which is consistent with the EH for the UK at the short end of the maturity spectrum. This is consistent with Hurn et al (1996) and Cuthbertson et al (1996).[11]

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[9] As has already been mentioned earlier, this is the first known study using the VAR methodology for Irish interest rates and as such comparisons will be made with similar approach's using US and UK data.

[10] For example, suppose theory suggests an elasticity of unity between 2 variables and the estimated equation is lny = 0.99lnx with a standard error 0.001. While we strongly reject the null of a unit elasticity, the predicted values of lny_t will closely mirror the actual values.

[11] Campbell and Shiller (1991) use monthly US data from January 1952 to February 1987, and find little support for the EH at the short end of the maturity spectrum. The authors obtained slope coefficients β ranging between 0 and 0.5. They do find beta coefficients of around 1 for maturities of 4, 5 and 10 years.

Campbell and Shiller (1991) use monthly data on US government bonds for the period 1946 to 1986, and their results are broadly consistent with our reported results for the perfect foresight spread equations. However Campbell and Shiller find that the Corr($S_t^{(n,m)}$, $S_t^{(n,m)'}$) are relatively low being in the range 0-0.7 and the values the variance ratio are in the range 2-10 for maturities of less than 1 year. Campbell and Shiller do not directly test the VAR cross-equation restrictions but this has been done subsequently by Shea (1992) who in general finds they are rejected. Again our results are generally consistent with those of Cuthbertson (1996) using UK data at short maturities. Cuthbertson's results from the VAR models for $S_t^{(n,m)}$ and $\Delta R_t^{(m)}$ indicate that $S_t^{(n,m)}$ Granger causes $\Delta R_t^{(m)}$: a weak test of the pure expectations hypothesis (PEH).[12] The author also finds that for all maturities there are strong correlation's between the actual and theoretical spread, and that the variance ratio's are close to unity. Hurn et al (1996) using monthly LIBOR rates (1975-1991) find the VAR cross-equation restrictions are not rejected, while Cuthbertson (1996) rejects these.

6. Conclusions

Using a number of short-term maturities on monthly Irish money market rates from 1982 to 1997, we perform a number of tests of the EH of the term structure of interest rates for Ireland. On balance our results would appear to lend support to the EH. Based on the results presented here, it would appear that central banks targeting of day-to-day interest rates translates into control

[12] Consistent with the results found in our study, Cuthbertson (1996) finds Granger causality from $\Delta R_t^{(m)}$ to $S_t^{(n,m)}$ indicating substantial feedback in the VAR regressions.

of interest rates with a maturity of several months. The standard deviations ratio and the correlation coefficients give results in favour of the EH, while the VAR cross-equation restrictions are rejected for only 1 case out of 3 (for the (3m, 1m) combination) However, we do provide reasons why the cross-equation restrictions may be rejected.

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Table 1:
Unit Root Tests

Variable	Maturity	ADF(5)	PP-Stat
Interest Rate: $R_t^{(n)}$	1 month	-2.64	-2.62
	3 month	-2.35	-2.20
	6 month	-2.20	-1.82
Change in interest rate: $\Delta R_t^{(n)}$	1 month	-6.56	-11.49
	3 month	-6.42	-11.49
	6 month	-5.74	-12.07
Spread: $S_t^{(n,m)}$	(3,1) month	-4.49	-6.33
	(6,1) month	-4.40	-5.31
	(6,3) month	-4.32	-4.89

Notes:
The sample period is from January 1984 to October 1997. ADF(5) is the augmented Dickey-Fuller statistic with 5 lags, which ensures the regressions are free of serial correlation. PP is the Phillips-Perron (1988) statistic with correction for up to 5th order serial correlation. The critical value for both test statistics is -2.86 at the 5% significance level

Table 2:
OLS Cointegration Tests: $\mathbf{R_t}^{(n)} = a + b\mathbf{r_t}^{(m)} + e_t$
Sample Period: 1984.1 - 1997.10

Maturity of	Maturity of	β coeff.	ADF(1)
Dep. Variable	Expl. Variable		
6 month	1 month	0.84	-4.94
1 month	6 month	1.08	-5.30
6 month	3 month	0.93	-4.63
3 month	6 month	1.05	-4.75
3 month	1 month	0.93	-6.21
1 month	3 month	1.05	-6.44

Notes:
The augmented Dickey-Fuller (ADF) statistic for the residuals, ε_t, ensuring that enough lags are present to ensure no serial correlation remains. The critical value for the ADF statistic (at 5% significance) is -2.88 (MacKinnon, 1991). The cointegrating regressions are estimated for the period January 1984 to October 1997.

Table 3:
Johansen procedure on $R_t^{(n)}$ and $r_t^{(m)}$

Interest rates (n,m)	Lag length		Cointegrating Vector	
			Normalised	LR test
(6,1) months	2		(-1, 0.99)	0.51
(6,3) months	3		(-1, 0.99)	0.04
(3,1) months	3		(-1, 0.99)	2.32

Notes:
In the Johansen procedure both the maximum eigenvalue test and the trace test do not reject the null of a unique co-integrating vector. The likelihood ratio (LR) statistic in column 4 tests the null that the co-integrating vector is (-1,1). Under the null, the reported test statistic has a critical value (at 5% significance level) of 3.84.

Table 4:
Does the Spread Predict Future Changes in Short-Rates?
Regression: $PFS_t^{(n,m)} = a + bS_t^{(n,m)} + gH_t$

Spread (n,m)	Coefficients		Wald Test		
	α s.e.(α)	β s.e.(β)	H_0:β=1 [p-value]	H_1:α=0, β=1 [p-value]	H_2:γ=0 [p-value]
(6,1)	-0.001 (0.001)	1.04 (0.15)	0.09 [0.77]	0.08 [0.78]	0.32 [0.58]
(6,3)	-0.0006 (0.0008)	1.02 (0.23)	0.007 [0.93]	0.007 [0.93]	2.88 [0.09]
(3,1)	-0.001 (0.001)	0.87 (0.21)	0.38 [0.54]	0.39 [0.53]	3.75 [0.06]

Notes:
The regression coefficients reported in columns 2 and 3 are from the regression with $\gamma = 0$ imposed. The method of estimation is GMM with a correction for heteroscedasticity and moving average errors using the Newey-West (1987) declining weights. The last 3 columns report Wald statistics and marginal significance levels for the null hypothesis stated. For $H_0 : \gamma = 0$ the reported results are for an information set which includes 4 lags of the change in the interest rates and the interest rate spread. The null H_0:β=1, is conditional on γ=0 while the null H_1:α=0, β=1 is also conditional on γ=0.

Table 5:
VAR model for ($S_t^{(n,m)}$, D$r_t^{(m)}$)

Spread (n,m)	Lag	Granger Tests	Causality	Ljung-	Box Q(26)	R2 -	Statistic
		S_t on $\Delta r_t^{(m)}$	$\Delta r_t^{(m)}$ on S_t	S_t - eqn.	Δr_t - eqn	S_t – eqn.	Δr_t - eqn
(6,1)	2	<0.01	0.48	9.36	17.4	0.52	0.22
(6,3)	2	<0.01	<0.01	13.3	27.7	0.62	0.19
(3,1)	2	<0.01	0.50	11.3	18.8	0.38	0.26

Notes:
'Lag' denotes the lag length that minimises the Akaike Information Criterion (AIC). Where the latter (occasionally) results in an equation system with serial correlation, the AIC is overridden and extra lags added (back) until any residual serial correlation is eliminated. The critical value for Q(26) is 38.89 (5% significance level). In columns 3 and 4 we report the marginal significance levels for the Granger-causality tests of $S_t^{(n,m)}$ on $\Delta r_t^{(m)}$ and vice versa (statistics are calculated after applying the GMM correction for heteroscedasticity used in Campbell and Shiller, 1991). The final 2 columns give the R^2 - statistic for each equation. The regressions are estimated for the whole sample period, January 1984 to October 1997.

Table 6:
Regression of the Actual Spread St on the Theoretical Spread St

Interest Rate Maturity	α		β		R^2 statistic
	coeff.	s.e.	coeff.	s.e.	
(6,1)	-0.01	0.02	0.90	0.06	0.93
(6,3)	0.02	0.01	0.81*	0.07	0.87
(3,1)	0.02	0.03	0.83	0.13	0.80

Notes:
The regressions are estimated for the whole sample period, January 1984 to October 1997. The estimated regression is $S_t = \alpha + \beta S_t' + \varepsilon_t$ which is estimated by GMM with heteroscedastic corrected errors. A star indicates the estimated coefficient is statistically different from that implied by the null hypothesis (at a 5% significance level), which for ' α ' is H_0: $\alpha = 0$ and for 'β ' is H_0 : $\beta = 1$. The theoretical spread S_t' is obtained from the predictions from the VAR using $z = [S_t, \Delta r_t]$.

Table 7:
Tests of the EH using weakly rational expectations

Spread (n,m)	Wald test W [.] = p-value	$\sigma(S_t^{(n,m)\prime})/\sigma(S_t^{(n,m)})$ (.) = std. Error	$Corr(S_t^{(n,m)}, S_t^{(n,m)\prime})$ (.) = std. Error
(6,1)	4.91 [0.30]	1.08(0.19)	0.96(0.05)
(6,3)	3.09 [0.54]	1.05(0.24)	0.93(0.10)
(3,1)	12.19 [0.02]	0.97(0.23)	0.89(0.12)

Notes:
The regressions are estimated for the whole sample period, January 1984 to October 1997.

Figure 1: 1 month and 6 month interest rate

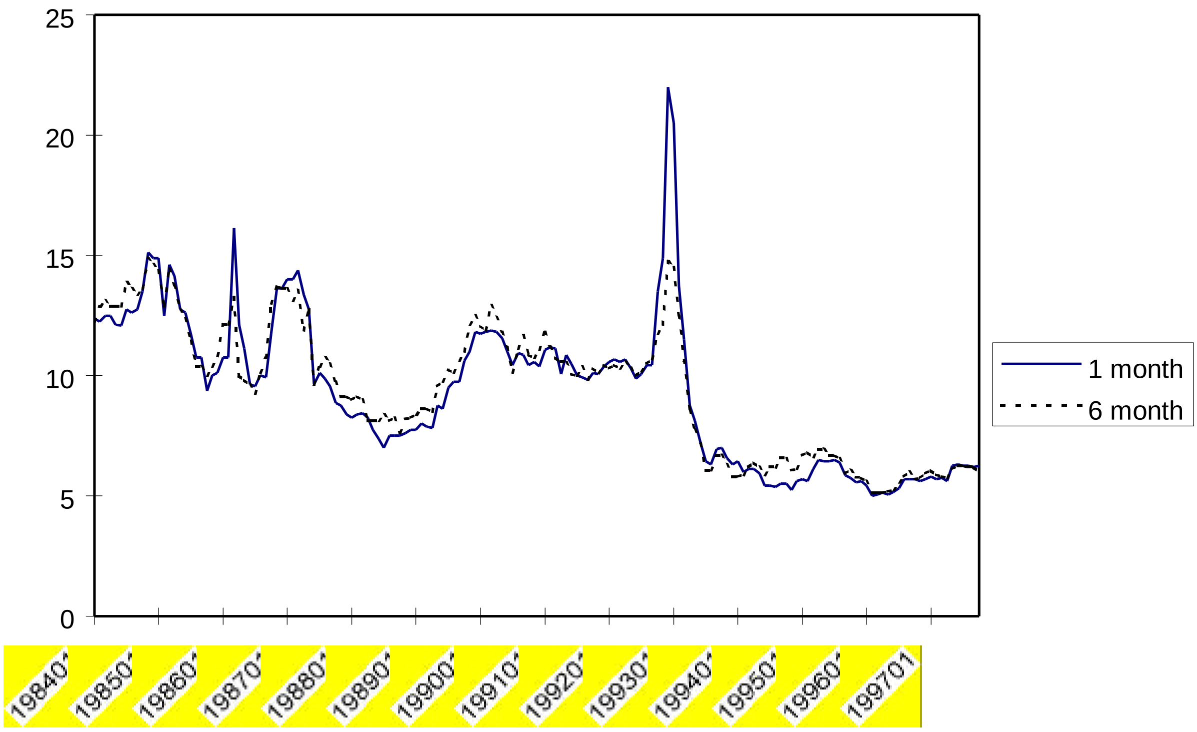

Figure 2: Actual and Theoretical Spread (6m - 1m)