Tests of the Expectations Hypothesis and Policy Reaction to the Term Spread: some comparative evidence

Gianna Boero

University of Warwick and University of Cagliari

and

Costanza Torricelli

University of Modena

July 1998

Abstract

The aim of this paper is to evaluate the impact of monetary policy in tests of the Expectations Hypothesis of the term structure of interest rates. We apply the model developed by McCallum (1994b), in which the Expectations Hypothesis interacts with a policy reaction function and with a time-varying term premium, to eight countries with different monetary policy stances, within the period 1985 to 1995. The results suggest the importance of the treatment of monetary policy in explaining the empirical performance of the Expectations Hypothesis. Amongst other results, we also find that the model performs better for some countries than others depending upon the monetary policy stance adopted.

Keywords: Expectations Hypothesis, interest rates, monetary policy, term structure.

JEL classification: C22, E43, E52

The authors wish to thank Luisa Malaguti for helpful comments on a previous version of the present paper and acknowledge support from CNR97.01050.CT10 and CNR97.01211.CT10.

1. INTRODUCTION

There is an ongoing debate concerning the validity of the Expectations Hypothesis (EH) of the term structure of interest rates (TS). The hypothesis has been widely tested, for different countries and different time periods, with mixed results. Typically, the hypothesis has been accepted with data for European countries (see, inter alia, Boero and Torricelli 1997, Engsted and Tanggaard, 1995, Gerlach and Smets, 1997) and rejected for the US (see Rudebusch, 1995, for a summary of different US studies). However, the most recent findings for the US provide new evidence in favour of the EH (Hsu and Kugler, 1997).

In order to interpret this disparate evidence on the EH of the TS, three main explanations have been proposed in the literature. The first one rests on a departure from the assumption of Rational Expectations (RE), which is normally tested jointly with the EH. An example of this is the ‘overreaction’ explanation put forward by Hardouvelis (1994), according to which agents do not react rationally, but instead overreact to expected changes in the short rate signalled by the TS spread. A second possible explanation attributes the empirical failures of the EH to the existence of time-varying term premia. Due to the unobservable nature of the term premium, its influence on tests of the EH can be assessed only indirectly by including a proxy for the term premium in standard regression tests. Attempts along these lines can be found in Simon (1989), Boero, Madjlessi and Torricelli (1996) and Tzavalis and Wickens (1997). A number of papers (e.g. Kugler, 1990, Gerlach and Smets, 1997) have anyway shown that time-varying term premia are not as important as the scarce variability of short rates in diminishing the predictive power of the EH. The third explanation involves policy behaviour, and asserts that the limited variability of short rates is due to particular monetary policy stances. The basic idea stems from the argument suggested by Mankiw and Miron (1986) that the ability of the spread to predict future interest rate movements is enhanced in the presence of a money supply target policy and is diminished under interest rate stabilisation.

In order to investigate the issue further, two possible lines of investigation have been proposed in the literature: the first empirical, the second theoretical. The former has been put forward by Dotsey and Otrok (1995) and Rudebusch (1995), who have empirically formalised the argument of Mankiw and Miron, by generating synthetic interest rate data from a Federal Reserve interest rate targeting model, and then using these to test the EH. This type of empirical analysis cannot be replicated for countries where monetary policy is officially monetary targeting and public interest rate targets are not available.

An alternative line of research is represented by the theoretical model proposed by McCallum (1994b). The author develops a model of the interaction between the EH of the TS, a time-varying autoregressive term premium, and an interest rate smoothing monetary policy combined with a reaction to changes in the spread[1]. The model shows that results in support of the EH can be explained by a strong policy response to changes in the spread and a highly positively autocorrelated term premium. In a recent paper, Kugler (1997) extends the McCallum model, deriving an exact solution for the N-period long rate case. Applications of the model to US data (Hsu and Kugler, 1997) and to a number of different countries (Kugler, 1997) indicate the model is able to explain the results of standard regression tests of the EH.

The aim of the present paper is to explore further the influence of monetary policy on the outcome of tests of the expectations hypothesis by applying the McCallum model to a wider range of countries with different monetary policy stances, and to different sample periods. The countries considered are: USA, Japan, Germany, UK, France, Italy, Canada and Switzerland. We find that the McCallum model is better able to explain results for countries where the monetary policy involves interest rate smoothing and/or clearly responds to

[1] This model is closely related to the policy reaction model developed by McCallum (1994a) which explains failures of Uncovered Interest Parity as a consequence of systematic monetary policy behaviour.

changes in the long-short spread, than for countries operating with a more complex monetary policy. The empirical analysis is conducted with weekly Euro-rates, for different subperiods between 1985 and 1995, and rests on two standard regression tests involving term spreads: one predicting future short rates, the other future long rates.

We start the empirical analysis with the implementation of standard tests of the EH. The results from these tests are confronted with both asymptotic and small sample distributions. The latter are those newly derived by Bekaert, Hodrick and Marshall (BHM) (1997), where the main source of bias is represented by the high persistence in short interest rates. When we conduct inference with the BHM empirical distributions we find that small sample biases can seriously affect the interpretation of standard tests of the EH, particularly those based on the long rates.

Then, in the second step of the analysis, we compare standard regression tests with the results obtained with the estimation of the McCallum model and evaluate the ability of this model to explain deviations from the EH.

The plan of the paper is as follows. Section 2 outlines the main features of the McCallum model and the exact solution derived by Kugler (1997), and Section 3 presents the methodology followed for the implementation of the model. Sections 4 and 5 report evidence on tests of the EH for eight different countries: first we perfom standard regression tests ignoring monetary policy considerations and conduct inference using both asymptotic and small sample distributions (Section 4); then we estimate the McCallum model and evaluate its ability to explain conflicting evidence from standard tests of the EH (Section 5). Section 6 closes the paper with conclusions and further remarks.

2. THE McCALLUM MODEL

In this section we describe the theoretical model tested in the present paper. The model was originally set up by McCallum (1994b) and later developed by Kugler (1997). In the following, we present the McCallum model and the exact solution provided by Kugler to this model.

McCallum (1994b) develops an N-period model, characterised by an equation for the term structure and an equation for the monetary policy rule. The former is represented by the expectations hypothesis modified by the existence of an autoregressive time-varying term premium, which implies that the return on an N-period bond is given by:

$$R_t^N = \frac{1}{N}\left(r_t + \sum_{i=1}^{N-1} E_t r_{t+i} \right) + x_t \tag{1}$$

$$x_t = r x_{t-1} + u_t \tag{2}$$

where R_t^N is the return on an N-period long-term bond, N is time to maturity of the long-term bond, r_t is the return on a one-period bond, x_t is the term premium on the long bond with $|r| < 1$, and u_t is a white noise error.[2]

For N large, it is reasonable to assume the approximation:

$$E_t R_{t+1}^N = E_t R_{t+1}^{N-1} \tag{3}$$

Hence (1) can be approximated as follows:

$$R_t - N(E_t R_{t+1} - R_t) = r_t + x_t \tag{4}$$

where from now on we drop superscripts and R_t stands for the return on a N period bond.

For empirical tests, (4) can be more usefully rewritten as:

$$N(R_{t+1} - R_t) = (R_t - r_t) - x_t + Ne_t \tag{5}$$

where $e_t = R_{t+1} - E_t R_{t+1}$ is the expectational error, which under RE is uncorrelated with R_t and r_t.

The monetary policy rule is supposed to be aimed at interest rate smoothing combined with a reaction to the term spread, i.e.:

[2] x_t is not exactly the term-premium on an N-period bond, but instead a linear combination of term premia. A deeper discussion of this point and of other assumptions underlying McCallum's model can be found in Malaguti-Torricelli (1997).

$$\Delta r_t = r_t - r_{t-1} = l\left(R_t - r_t\right) + z_t \qquad (6)$$

where l ³$0$ is the policy parameter, and z_t represents other components of policy behaviour and, for simplicity, is assumed to be white noise[3]. The rule is based on the observation that actual monetary policy in many countries involves manipulation of a short-term interest rate instrument (McCallum, 1994b). Obviously, (6) represents a strong stylisation of actual monetary policy rules since Central Banks generally use a wider range of policy indicators other than the spread. Nevertheless, given the correlation between the spread and other indicators (e.g. real economic growth and inflation expectations), this simple rule can be thought of as also capturing policy stances of those Central Banks which officially do not use the spread as an indicator (e.g. the Bundesbank).

Combining (4) and (6) gives:

$$(1+\mathrm{N})\mathrm{R}_t = \mathrm{N}\mathrm{E}_t\mathrm{R}_{t+1} + (1+l)^{-1}\left[\mathrm{r}_{t-1} + l\mathrm{R}_t + z_t\right] + x_t \qquad (7)$$

which has to be solved for R_t.

The RE solution procedure is based on the minimum-state-variable (MSV) criterion discussed by McCallum (1983), whereby the solution is assumed to have the following form:

$$R_t = f_1 r_{t-1} + f_2 x_t + f_3 z_t \qquad (8)$$

and is given by:

$$R_t = r_{t-1} + \frac{1+l}{1+N-Nr(1+l)} x_t + z_t \qquad (9)$$

The relevant regressions, accordingly, become:

$$r_t - r_{t-1} = l r (R_{t-1} - r_{t-1}) + \frac{Nl}{1+N(1-r(1+l))} u_t + z_t \qquad (10)$$

and

$$R_t - R_{t-1} = (l r + r - 1)(R_{t-1} - r_{t-1}) + \frac{(1+l)}{1+N(1-r(1+l))} u_t + z_t \qquad (11)$$

[3] The analysis would not change if z_t was allowed to be autocorrelated (McCallum, 1994b, p. 5).

McCallum underlines that, except for very large values of ρ and/or λ, the coefficient of the spread in (11) will be negative, thus matching some empirical results for the US (e.g. Evans and Lewis, 1994, Campbell and Shiller, 1991) which cannot be reconciled with the constant term premium version of the EH.

Kugler (1997) offers an exact solution to the N-period case in McCallum's model. Specifically, the solution to the model hinges on the approximation assumed in equation (3), which allows one to eliminate the expected values for the short rate as far as period N. In order to avoid the use of the approximation in (3), Kugler calculates the (N-1) RE values of the short rate up to date N. The RE solutions are still attained according to the MSV criterion and by means of the method of undetermined coefficients. Kugler's regression equations for the spread and for the short rate are respectively the following:

$$(R_t - r_t) = r\,(R_{t-1} - r_{t-1}) + \frac{N}{N - l\sum_{j=1}^{N-1}(N-j)r^{\,j}}\,u_t \tag{12}$$

and

$$(r_t - r_{t-1}) = l\,r\,(R_{t-1} - r_{t-1}) + \frac{N l}{N - l\sum_{j=1}^{N-1}(N-j)r^{\,j}}\,u_t + z_t \tag{13}$$

The latter has the implication that the information content of the spread vanishes whenever λ or ρ tends to zero.[4] Kugler interprets this result by noting that the predictive content of the spread is based on predictable monetary policy reaction to the spread. If λ and/or ρ are zero, there is no predictable exogenous movement of the spread which determines the predictability of policy reaction.

Kugler does not present the regression equation for the long rates, which we have derived from his solutions as follows:

[4] This result extends McCallum's equivalent result for the 2-period case.

$$R_t - R_{t-1} = (l r + r - 1)(R_{t-1} - r_{t-1}) + \frac{(1+l)}{1 - l \sum_{i=1}^{N} (1 - \frac{j}{N+1}) r^{j}} u_t + z_t \quad (14)$$

By comparative inspection of (11) and (14), it is clear that the exactness of the solution calculated by Kugler is relevant only for the coefficient of the white noise term u_t. Since the coefficient of the spread is in all cases the same, the implications for the tests of the EH based on the value of the spread coefficient are as in McCallum.

3. MODELLING STRATEGY

In order to test whether the model presented in the previous section is able to explain empirical deviations from the expectations hypothesis, we follow the approach adopted by Kugler (1997) which consists in comparing the estimated value of the spread coefficient in standard regressions for tests of the EH with the value implied by the McCallum model. The type of regressions considered by Kugler use the spread to predict future short rates. In our empirical analysis we present evidence also for regressions which relate the spread to changes in the long rate.

First of all recall that the EH formulated in equation (1) implies that the slope coefficient β of the following regression equations should be equal to 1:

$$(1/N)\sum_{j=1}^{N-1}(N-j)\Delta r_{t+j} = a + b\,(R_t^N - r_t) + e_{t+N-1} \quad (15)$$

$$\mathrm{R}_{t+1}^{(N-1)} - \mathrm{R}_t^{N} = a + b\left(\frac{1}{(\mathrm{N}-1)}\right)(\mathrm{R}_t^{N} - \mathrm{r}_t) + e_{t+1} \quad (16)$$

where R^N_t and r_t are the N- and 1-period interest rates respectively, and Δr_{t+j} in (15) is equal to r_{t+j} - $r_{t+j-1.}$ Equation (15) uses the spread to predict (a weighted average of) changes in the short rate over an N-period horizon; in (16) the spread should predict the change in the N-period rate over the 1-period horizon.

Taking expected values of (12) and (13) and rearranging, it can be shown that Kugler's model implies:

$$\Delta r^{e}_{t+j} = l\, r^{\,j} (R_t - r_t)$$

Hence the counterpart of (15) in Kugler's model becomes:

$$(1/N)\sum_{j=1}^{N-1}(N-j)\Delta r^{e}_{t+j} = (1/N)\, l \sum_{j=1}^{N-1}(N-j)\, r^{\,j} (R^{N}_{t} - r_t) \qquad (17)$$

where the implied β is:

$$b_{MC} = (1/N)\, l \sum_{j=1}^{N-1}(N-j)\, r^{\,j} \qquad (18)$$

The counterpart of (16) is (14) where the implied β is:

$$\beta_{MC} = (N\text{-}1)(\lambda\rho+\rho\text{-}1) \qquad (19)$$

The suggestion from Kugler is to estimate β_{MC} by means of Indirect Least Squares in two stages: first, estimate the two reduced form equations (12) and (13) with OLS to obtain values for ρ and λρ, then divide the estimated value of λρ by ρ to obtain an estimate of l, and use (18) and (19) to obtain values of β_{MC}. These values are then compared with the estimates of β obtained from regressions (15) and (16).

Kugler (1997) applies quite successfully this methodology to regressions for the short rate, using the one- and three- month interest rates for the US, Japan, Germany and Switzerland in the period 1982-1992. He found that for the case of Japan, the good predictive power of the spread can be explained by both a high reaction of the monetary policy to the spread and a strong autocorrelation of the term premium. On the other hand, the low predictive power of the spread is explained by either low correlation of the term premium – which is the case for Germany and Switzerland – or low reaction of the monetary policy to the spread – as is the case for the US during that particular sample period. The latter case is further investigated in Hsu and Kugler (1997) where results in favour of the EH were found for the US for the more recent sample period 1987-95.

In order to assess further the empirical validity of the model proposed by McCallum, in the following sections we present new evidence on tests of the expectations hypothesis for a wider range of countries with different monetary policy rules, and for different sample periods. We start with standard tests ignoring policy behaviour and subsequently estimate the McCallum policy reaction model.

4. RESULTS FROM STANDARD REGRESSION TESTS

Different tests of the EH have been proposed in the literature. In this section we present evidence based on the implications of equation (1) that the term spread should predict future changes in the short rate and in the long rate.

The two regressions used to test these implications of the EH are equations (15) and (16). In these equations, ε_{t+N-1} and ε_{t+1} are forecast errors which under rational expectations are orthogonal to information at time t, and therefore uncorrelated with the regressor R^N_t-r_t, so OLS will give consistent estimates. However, the errors in (15) will be serially correlated following a MA(N-2) process, while the errors in (16) will follow a MA(m-1) process when the short rate has maturity m>1. So, standard errors are usually calculated with the Newey-West or Hansen and Hodrick corrections.

Tests of the predictive content of the spread imply testing for the significance of β (β=0), while tests of the EH with RE and constant term premium imply testing for β=1.

The data used in this study are weekly Euro-rates for the period 16-11-1985 to 11-11-1995 for USA, Japan, Germany, UK, France, Italy, Canada and Switzerland.

As we ultimately want to investigate the effects of monetary policy on tests of the EH, we only use the 1-month and 3-month rates, as these are more directly linked to monetary policy.[5]

In sub-section 4.1 we consider regressions for the short rate and in 4.2 regressions for the long rate.

4.1 SHORT RATE REGRESSIONS

The equation estimated for the short rate is equation (15), where N=3-months. With monthly data the equation would be the following:

$$\left(\frac{1}{3}\right)\sum_{j=1}^{3-1}(3-j)\Delta r_{t+j} = a + b(R^3{}_t - r_t) + e_{t+3-1}$$

However, as we are using weekly data, we approximate the 1, 2 and 3-months horizons as 4, 9 and 13 weeks respectively, so the regression becomes:

$$\left(\frac{2}{3}\right)\left(\mathrm{r}_{t+4} - \mathrm{r}_t\right) + \left(\frac{1}{3}\right)\left(\mathrm{r}_{t+9} - \mathrm{r}_{t+4}\right) = a + b\left(\mathrm{R}_t^3 - \mathrm{r}_t\right) + e_{t+9} \qquad (20)$$

The errors in (20) follow an MA(h-1) process, where h is the forecast horizon (9 weeks in our regressions) and hence we computed Newey-West corrected standard errors with truncation lag equal to 8. The results are summarised in Table 1. The estimation period selected for each country varies within the interval 16/11/1985 and 11/11/1995, and, in each case, reflects the longest period for which the estimated β was found to be stable.

We first discuss evidence based on standard asymptotic distributions, and subsequently consider the effects of small sample bias.

Inference based on asymptotic distributions

Table 1 shows that all estimates for β are significantly different from zero, with the exception for Switzerland, thus confirming an overall information content of the spread for

---

[5] We thank Peter Kugler for kindly providing us with these data.

future short rates. However, tests of the EH (β=1), reported in the last column of Table 1, indicate that the EH is rejected for all countries at conventional significance levels, except for France, Italy and Canada. The evidence on France and Italy is in line with results in previous studies which mantain that the EH better describes weak-currency countries than strong-currency countries (see Gerlach and Smets, 1997). A possible rationale behind this explanation is that Central Banks of weak-currency countries often have to manage short rates in order to achieve mid-term intermediate exchange rate objectives. As these are usually well-known, short rates become more predictable, and this justifies higher values for β in those countries[6].

The effects of small sample bias

A common criticism of these regression-based tests of the EH is that they can be seriously biased in small samples. This is particularly true for the long rate regressions, as we discuss below, but in a recent paper Bekaert, Hodrick and Marshall (BHM) (1997) found that regressions for the short rate can also be affected by substantial positive bias. According to that study, the positive bias arises because under the assumption that the short rate is generated by an AR(1) process, the slope coefficient in these regressions can be shown to be a negative transformation of serial correlation coefficients. This transformation, combined with the negative bias in OLS estimates of autocorrelation coefficients for highly persistent data, generates a positive bias in the slope coefficients (see Bekaert et al., 1997, eqs. 7, 8 and 10).

In order to evaluate the effects of this kind of small sample bias in the tests presented so far, in what follows we conduct inference by using the 5% quantiles of the empirical distributions of the slope coefficient derived by BHM under two alternative data generating

[6] An empirical analysis of the importance of short rate predictability in tests of the EH can be found in a previous version of this paper (see Boero, Di Lorenzo and Torricelli, 1998).

processes (d.g.p.): an AR(1) for the short rate (see BHM, Panel C, Table 3) and a VAR-GARCH model for both the short rate and the spreads (see BHM, Panel C, Table 6). These empirical distributions are characterised by substantial positive bias (which would strengthen rejection of the EH) but also increased dispersion (which would weaken rejection). According to these distributions, to have a 5% rejection of the hypothesis β =1 in a one-tailed test, the slope coefficient should be smaller than 0.64 with the AR(1) d.g.p., and smaller than 0.62 with the VAR-GARCH d.g.p.. As evident from Table 1, by conducting inference with the small sample distributions, our results remain virtually unaffected, the only exceptions are USA and UK for which rejection would be weakened. These results are quite interesting, although they should be interpreted with some caution, as the critical values derived by BHM are not exactly applicable to our regressions for at least two reasons. First, we are using a very short term spread (3-1 month), while the closest spread considered in the BHM Monte Carlo experiment is 12-1 month. Second, BHM report critical values for only one sample size (524 observations) which is very close to the number of observations used in some of our regressions, but larger than that used in others.

4.2 LONG RATE REGRESSIONS

The equation estimated for the long rate is equation (16), with maturities N=3-months for the long rate, and 1-month for the short rate. With monthly data the equation would be the following:

$$R_{t+1}^{(3-1)} - R_t^3 = a + b\left(\frac{1}{(2)}\right)(R_t^3 - r_t\ ) + e_{t+1}$$

However, in our empirical tests this regression is modified for weekly data to obtain

$$R_{t+4}^3 - R_t^3 = a + b\left(\frac{1}{2}\right)(R_t^3 - r_t\ ) + e_{t+4} \tag{21}$$

In equation (21) we have used approximation (3), discussed in Section 2, $E_t R_{t+1}^N = E_t R_{t+1}^{N-1}$, which is commonly adopted in regressions of this type. However, while this approximation is irrelevant for large N, as we see shortly, it may have significant bias effects on the estimated value of β for small N.

Regressions for the long rate have been the focus of attention of many studies attempting to explain failures of the EH. In fact, while the EH implies that the slope coefficient should be equal to one, most of the empirical literature has reported very low values for the R^2, and estimated coefficients below unity, becoming negative as yields of longer-term bonds are used to form the dependent variable and the term spread. Negative values indicate that long rates move in the opposite direction to that implied by the theory.

Our estimates, reported in Table 2, are apparently at odds with previous findings, as they seem to support the EH in most cases. In what follows we first discuss the results using standard asymptotic distributions and then proceed to reinterpret the results using the BHM small sample distributions.

Inference based on asymptotic distributions

Table 2 reports our estimates of regression equation (21) and tests of the EH ($\beta = 1$) for the whole sample period 1985-1995 (518 observations), and for the sub-period 1991-95 (206 observations). The last two columns report the R^2 values. Contrary to previous findings in the literature, most point estimates are positive, with the exception of Switzerland, and some are close to one. However, some estimates are not significantly different from zero, and the very low R^2 values conform with previous results indicating that the spread between the long and short term interest rates has poor predictive content for changes in the longer rate.

The finding of estimated slope coefficients close to 1 may depend on: (i) the particular nature of our regressions which only look at the very short end of the term structure, whereas regressions for the long rates are typically estimated in the longer end of the term structure, and (ii) the approximation used to construct the dependent variable. This point will be discussed further below.

The effects of small sample bias

We now return to the results in Table 2 and consider how they can be affected by small sample biases. In the study mentioned before by BHM, the small sample bias which affects all regression-based tests of the EH is shown to be particularly strong for the long rate regressions. BHM show that approximation (3) which we used in the estimation of equation (21) introduces a further error in the regression which exacerbates the small sample bias. In particular, they found that for maturity of the long rate N=12 and sample size T=524 the average of the OLS estimates of β is about 2, with similar value for the standard deviation. So, as already seen for the short rate regressions, the small sample distribution is biased upward and has an increased dispersion. However, the bias for the long rate regressions is much higher. Moreover, in contrast to the results obtained from regressions for the short rates, inference based on the small sample distributions for the long rates is not uniformly conclusive about rejection of the EH, as these distributions seem to be very sensitive to the data generating process used in the BHM Monte Carlo simulations. Specifically, according to the BHM empirical quantiles, the EH should be rejected at the 5% for values of $\beta<1.203$ when the d.g.p. for the short rate is an AR(1) model (BHM, Panel B, table 3), and for values of $\beta<0.131$ if inference is conducted under the assumption of a VAR-GARCH model for the short rate and the spreads (BHM, Panel B, Table 6). So, inference based on these critical values would indicate evidence against the EH for most countries (except France and marginally Canada) under the AR(1) d.g.p., while under the alternative d.g.p. the evidence would be generally in favour of the EH, with the sole exception for Switzerland.

To summarise, in this section we have conducted inference on regression-based tests of the EH with both asymptotic and small sample distributions. The empirical critical values that we used are those derived in Bekaert et al. (1997). Although these small sample distributions are affected by substantial positive bias (which depends on the persistence of the short rate), they are also characterised by increased dispersion which, in the case of regressions for the short rate, leads to results that are in general invariant to those based on asymptotic distributions. On the other hand, inference based on the empirical critical values for regressions for the long rate is inconclusive, due to the high sensitivity of the BHM small sample distributions to the data generating process.[7]

In the next section we estimate the McCallum model following the methodology introduced in Section 3, and evaluate its ability to explain deviations from the EH. We do this by comparing the values of the coefficients β implied by this model with those obtained from standard regression tests.

5. AN EVALUATION OF THE McCALLUM MODEL

In the present section we apply the McCallum model to the 8 countries considered in this study, and compare the implied values of β (β_{MC}) with those estimated from standard regression tests discussed in the previous sections. As described in Section 3, in the McCallum model the EH interacts with a policy reaction function, in the presence of a time-varying term premium, so the implied β in tests of the EH is a composite parameter reflecting policy behaviour (the λ coefficient in (6)) and the autoregressive component of the term premium (the ρ coefficient in (2)). The values of β_{MC} are obtained from equations (18) and (19) modified according to the weekly frequency of the data as follows:

$$\text{short rate regressions:}\quad \beta_{MC} = (1/3)\, l\,(2 \sum_{j=1}^{4} r^{j} + \sum_{j=5}^{8} r^{j}) \qquad (22)$$

[7] In a recent paper, Schotman (1997) assumes that interest rates (short and long) are generated by ARIMA (1,1,1) models and finds very large bias for the spread coefficient β. The bias is positive or negative depending on whether the sum of the autoregressive and moving average coefficients is negative or positive.

long rate regressions: $\beta_{MC} = 2(\lambda\rho+\rho-1)(1+\rho+\rho^2+\rho^3)$ (23)

Estimates of λ and ρ are obtained from the two Reduced Form (RF) equations (12) and (13). In particular, ρ is obtained by applying OLS to RF (12), while λ is obtained by Indirect Least Squares applied to RF (13) or, equivalently, by IVE applied to the policy reaction equation (6) with instrument $(R_{t-1}-r_{t-1})$. We first consider the ability of the McCallum model to explain results from standard regressions for the short rate. Then we turn to regressions for the long rate.

5.1 SHORT RATE REGRESSIONS

In this section we see whether the McCallum model is able to explain the results discussed in Section 4.1 and summarised in Table 1. For convenience, those results are reported again in Table 3 (columns 4 and 6) with estimates of λ and ρ, the implied β_{MC} computed as in equation (22), and a test for $\beta = \beta_{MC}$.

An important result is that the implied slope coefficients β_{MC} are in general consistent with the β estimates obtained from standard EH regressions: the Wald test never rejects the hypothesis $\beta = \beta_{MC}$. This results indicates that the McCallum model can rationalise different values for β, including low values as in the case of Switzerland, and suggests that explicit consideration of a monetary policy reaction function is important in providing an explanation of both failures and successes of the EH.

Table 3 also shows that differences in the values of β_{MC} are due more to different estimates of the monetary policy coefficient than to a different pattern of time variation in the term premium. In fact, while estimates of l range from a minimum of 0.194 for Germany to a maximum of 0.791 for France, those for r display a much lower variation (between 0.598 for Italy and 0.871 for Germany).

In particular, results point to a strong policy reaction to the spread for France (λ=0.791) and Italy (λ=0.751), a moderate reaction for the UK (λ=0.348), the US (λ=0.480) and Canada (λ=0.497), and a low reaction for Germany (λ=0.194), Switzerland (λ=0.272) and Japan (λ=0.301).

Table 4 casts further light on the ability of the McCallum model to explain different results from tests of the EH, showing that the ranking between the countries is similar according to the three parameters λ, estimated β and β_{MC}.

However, at this point, a note of caution is necessary, particularly with respect to the estimates of λ. In fact, these are based on an extremely simplified policy reaction function, where policy responds only to the spread (reflecting the Central Bank reaction to changes in expected future inflation), and may therefore suffer from omitted variable bias. Moreover, the estimates of λ are based on the assumption that the error term ζ_t is not autocorrelated. To improve on the empirical estimation of λ, ideally one would include other potentially important policy indicators (recent inflation, exchange rate, output), but this route would require specification of an expanded macroeconometric model which endogenously explains the added variables. Instead, following Kugler, we have considered the possibility that Central Banks react to lagged short rate changes, as well as the spread, and reestimated the policy reaction function with instrumental variables plus an AR(1) error, using as instruments the lagged change in the short rate (r_{t-1}-r_{t-2}) and the lagged spread (R_{t-1}-r_{t-1}). With this estimation procedure we found, in general, values of λ very close to those reported in Table 3, which we have chosen not to report here for reasons of space.[8] However, there were two exceptions, France and Italy, for which the estimated λ did not seem to be robust. The value for France, for example, changed from 0.791 to 0.44 when the AR(1) procedure was employed, implying significant changes in the value of β_{MC}, and suggesting that the highly

[8] Results are available from the authors on request.

stylised policy reaction function cannot adequately describe monetary policy in that country. Similar findings obtained for Italy. Both these countries operate with intermediate exchange rate targets, so the simple policy reaction function which combines interest rate smoothing with a reaction to the spread may need to be expanded with the inclusion of a more complex set of policy indicators, to obtain better estimates for λ.

On the other hand, the stylised reaction function of the McCallum model seems particularly adequate to describe the monetary policy in countries where the spread is clearly used as an important indicator. For example, with regard to the US, the moderately high value of l is in line with recent findings (Hsu and Kugler, 1997) and reflects the increased reliance of the Federal Reserve on the spread as a policy indicator rather than on monetary aggregates. This estimate of l combined with a high value for ρ (0.81) implies a value for the slope coefficient β_{MC} close to 1 (0.97).

Another interesting result is that for Germany, where the low value of l corresponds to official statements of the Bundesbank that the spread is not used as a policy indicator. Moreover, while the Bundesbank is officially monetary targeting, the estimation of the McCallum policy reaction model suggests that there may be elements of interest rate smoothing in the German monetary policy. The low estimate of λ for Germany combined with a high persistence in the spread (ρ=0.871) imply a value for β_{MC} of 0.49 which is not too far away from the estimated β (0.60).

5.2 LONG RATE REGRESSIONS

Finally, in Table 5 we compare estimates of the β coefficients obtained from regressions for the long rate (repeated from Table 2) with the value for β implied by the policy reaction model (equation (23)). As we can see, this model can again rationalise different results for β, including negative values, and, in most cases, the Wald test for the equality of the β_{MC} and the β estimated from regressions for the long rate cannot reject this hypothesis. Although these results imply that the McCallum model can rationalise different findings from tests of the EH, it is important to emphasise that results from these regressions for the long rates are to be interpreted with caution not only because of the small sample bias discussed in Section 4.2, but also because of the very low R^2 values.

6. CONCLUSIONS

In this paper we have tested the Expectations Hypothesis of the term structure across eight countries with different monetary policy rules and have examined the ability of the McCallum model (1994b) to explain conflicting evidence from standard tests. The empirical analysis was conducted in two steps. First, we performed standard regression tests of the EH using the term spread to predict future changes in the short rates and in the long rates. Second, we estimated the McCallum model, in which the expectations hypothesis interacts with a policy reaction function and with an autoregressive time-varying term premium, and compared the estimates of the term structure coefficients implied by this model with those obtained in standard regressions.

With respect to the first step, we found that in the case of regressions for the short rates the coefficient of the term spread was significantly different from zero in most cases, thus confirming an overall information content of the spread for future short rates. However, tests of the Expectations Hypothesis (β=1) indicated rejection for five countries out of eight. In the case of regressions for the long rates, contrary to previous findings in the literature, we found that most point estimates were positive, and some were close to one. However, for these regressions we also found very low R^2 values, which conforms with previous results, indicating that the spread between the long and short term interest rates has poor predictive content for changes in the longer rate. To complete the first step of the analysis, we also confronted the results of tests of the EH with the empirical critical values recently derived in Bekaert *et al*. (1997). These take into account a particular kind of bias due to persistence in the short rates. We found that this kind of bias did not affect substantially the results of tests from the short rate regressions, whereas inference based on the small sample distributions for the long rate regressions produced conflicting results, depending on the assumptions underlying the data generating process.

With respect to the second step, we found that the McCallum model was in general able to rationalise different values for β. This result suggests that values of the slope coefficients statistically different from one are consistent with the EH and deviations from the EH reflect the way in which monetary policy responds to changes in the term spread. Amongst other results we also found that the model performs better for some countries than others depending upon the monetary policy stance adopted. Specifically, the model is better able to explain results for countries where the monetary policy involves interest rate smoothing and/or clearly responds to changes in the long-short spread, than for countries operating with a more complex monetary policy.

REFERENCES

Bekaert, Hodrick and Marshall, 1997, On biases in tests of the expectations hypothesis of the term structure of interest rates, Journal of Financial Economics, 44, 309-348.

Boero G., F. Madjlessi F. and C. Torricelli, 1996, The information in the term structure of German interest rates, *Proceedings of the VI International AFIR Colloquium*, Nürnberg, October 1996, VVW Karlsruhe.

Boero G. and C.Torricelli, 1997, The Expectations Hypothesis of the term structure of interest rates: evidence for Germany, *Contributi di ricerca CRENOS*, n. 4, Università di Cagliari, settembre.

Campbell J.Y. and R.J. Shiller, 1991, Yield spread and interest rate movements: a bird's eyes view, *Review of Economic Studies*, 58, 495-514.

Dotsey M. and C. Otrok, 1995, The rational expectation hypothesis of the term structure, monetary policy and time-varying term premia, *Economic Quarterly, Federal Reserve Bank of Richmond*, 81, 65-81.

Engsted, T. and C. Tanggaard, 1995, The predictive power of yield spreads for future interest rates: Evidence from the Danish term structure, *Scandinavian Journal of Economics*, 97, 145-159.

Evans M.D.D. and K.K. Lewis, 1994, Do stationary risk premia explain it all? - Evidence from the term structure, *Journal of Monetary Economics*, 33, 285-318.

Gerlach, S. and F. Smets, 1997, The Term Structure of Euro-rates: some evidence in support of the expectations hypothesis, *Journal of International Money and Finance*, Vol. 16, no. 2, 285-303.

Hsu C. and P.Kugler, 1997, The revival of the Expectations Hypothesis of the US term structure of interest rates, *Economics Letters*, 55, 115-120.

Kugler, P., 1990, The term structure of interest rates and rational expectations, *Journal of International Money and Finance,* Vol. 9, pp. 234-244.

Kugler P., 1997, Central Bank policy reaction and the Expectation Hypothesis of term structure, *International Journal of Finance and Economics*,2, 217-224.

Malaguti L. and C. Torricelli, 1997, The interaction between monetary policy and the Expectation Hypothesis of the term structure of interest rates in a N-period Rational Expectation model, *Materiali di Discussione*, n.181, Dipartimento di Economia Politica, Università di Modena, luglio.

Mankiw N.G. and J.A..Miron, 1986, The changing behaviour of the term structure of interest rates, *Quarterly Journal of Economics*, 101, 211-228.

McCallum B.T., 1983, On non-uniqueness in rational expectations models: an attempt at perspective, *Journal of Monetary Economics*, 11, 139-168.

McCallum B.T., 1994a, A reconsideration of the Uncovered Interest Rate Parity Relationship, *Journal of Monetary Economics*, 33, 105-132.

McCallum B.T., 1994b, Monetary policy and the term structure of interest rates, NBER Working Paper Series, N. 4938.

Rudebusch G.D., 1995, Federal Reserve interest rate targeting, rational expectations, and the term structure, *Journal of Monetary Economics*, 35, 245-274.

Schotman P.C., 1997, Small sample properties of the regression test of the expectations model of the term structure, *Economics Letters*, 54, 129-134

Simon, D.P., 1989, Expectations and risk in the Treasury bill market: An instrumental variables approach, *Journal of Financial and Quantitative Analysis*, 24, 357-365.

Tzavalis E. and M.R. Wickens, 1997, Explaining the failures of the Term Spread Models of the Rational Expectations Hypothesis of the Term Structure, *Journal of Money, Credit and Banking*, Vol. 29, No. 3, 364-380.

TABLE 1 - Estimates of β in regressions for the short rate (equation 20)

Country and sample period (no. of obs.)		**Estimates of $\beta^{(i)}$. (corrected SEs)$^{(ii)}$**	**Wald test $^{(iii)}$ Chi-Sq for H_0: β =1**
USA	16/11/91-11/11/95 (201)	0.77 (0.105)	4.36 *
Japan	16/11/90-11/11/95 (253)	0.48 (0.098)	27.56 **
Germany	16/11/85-11/11/95 (513)	0.60 (0.160)	6.15 *
UK	16/11/90-11/11/95 (253)	0.64 (0.114)	9.67 **
France	16/11/90-11/11/95 (253)	1.12 (0.223)	0.30
Italy	16/11/85-11/11/95 (513)	0.80 (0.102)	3.57
Canada	16/11/90-11/11/95 (253)	0.95 (0.259)	0.02
Switzerland	16/11/90-11/11/95 (253)	0.27 ** (0.193)	13.69 **

Notes: **(i)**** indicates that the coefficient is not statistically different from zero at the 5% significance level; **(ii)** the number in parenthesis are heteroscedasticity and autocorrelation corrected standard errors with truncation lag 8; **(iii)** * indicates rejection of H_0: $\beta = 1$ at the 5%; ** indicates rejection at the 1%.

TABLE 2 - Estimates of β in regressions for the long rate (equation 21)

Sample periods: 16/11/85-11/11/95 (no. obs 518);

16/11/91-11/11/95 (no. obs. 206)

Country	**Estimates of β(i). (corrected SEs)(ii)**		**Wald test Chi-Sq for H_0: β =1 (prob. of rejection)(iii)**		**R^2**	
	85-95	91-95	85-95	91-95	85-95	91-95
USA	0.54** (0.33)	0.92 (0.33)	1.92 (.17)	0.05 (.81)	0.03	0.10
Japan	0.92 (0.25)	0.50 (0.22)	0.09 (.77)	1.78 (.18)	0.13	0.02
Germany	0.54** (0.39)	0.64 (0.30)	1.42 (.23)	1.50 (.22)	0.03	0.05
UK	1.07 (0.31)	0.84 (0.38)	0.05 (.82)	0.18 (.67)	0.08	0.07
France	1.30 (0.61)	1.86 (0.51)	0.24 (.62)	2.8 (.09)	0.11	0.14
Italy	0.58 (0.20)	0.51 (0.14)	4.4* (.04)	12.2** (.01)	0.04	0.05
Canada	1.23 (0.38)	0.89** (0.82)	0.38 (.54)	0.02 (.89)	0.09	0.03
Switzerland	0.18** (0.24)	-0.18 ** (0.38)	11.9** (.001)	9.4** (.002)	0.003	0.002

Notes: **(i)**:** indicates that the coefficient is not statistically different from zero at the 5% level; **(ii)** the number in parenthesis are heteroscedasticity and autocorrelation corrected standard errors, with truncation lag 3; **(iii)** * indicates rejection of H_0: β = 1 at 5%, ** indicates rejection at 1%.

TABLE 3- Estimation of the McCallum model: comparison with results from regressions for the short rate					
Country and sample period (1)	**Estimates of λ and ρ** (2)	β_{MC} (i) (3)	**Estimates of β** (4)(ii)	**Wald test(iii) H_0: $\beta = \beta_{MC}$** (5)	**Wald test(iii) H_0: $\beta = 1$** (6)(ii)
USA 11/91-11/95	λ = 0.480 (.079) ρ = 0.810 (.040)	0.97	0.77 (0.105)	3.27	4.36 *
Japan 11/90-11/95	λ = 0.301 (.074) ρ = 0.801 (.035)	0.58	0.48 (.098)	1.04	27.56 **
Germany 11/85-11/95	λ = 0.194 (.042) ρ = 0.871 (.021)	0.49	0.60 (.160)	0.42	6.15 *
UK 11/90-11/95	λ = 0.348 (.059) ρ = 0.823 (.034)	0.74	0.64 (.114)	0.72	9.67 **
France 11/90-11/95	λ = 0.791 (.207) ρ = 0.665 (.046)	0.93	1.12 (.223)	0.73	0.30
Italy 11/85-11/95	λ = 0.751 (.124) ρ = 0.598 (.035)	0.69	0.80 (.102)	1.18	3.57
Canada 11/90-11/95	λ = 0.497 (.101) ρ = 0.771 (.039)	0.86	0.95 (.259)	0.11	0.02
Switzerland 11/90-11/95	λ = 0.272 (.112) ρ = 0.779 (.038)	0.49	0.27 ** (.193)	1.21	13.69 **

Note: **(i)** computed as in equation (22) : $\beta_{MC} = (1/3)\, l\,(2 \sum_{j=1}^{4} r^j + \sum_{j=5}^{8} r^j)$; **(ii)** The values in columns 4 and 6 are taken from Table 1; ** in column 4 indicates that the coefficient is not statistically different from zero at the 5%; * and ** in column 6 indicates rejection at the 5% and 1% respectively; **(iii)** Wald test (chi-square values) computed with Newey-West corrected standard errors, consistent in the presence of MA(8) errors and heteroscedastic.

TABLE 4- Countries ranked according to values of λ, estimated β and $\beta_{MC.}$

λ		**estimated β**		β_{MC}	
France	0.791	France	1.12	USA	0.97
Italy	0.751	Canada	0.95	France	0.93
Canada	0.497	Italy	0.80	Canada	0.86
USA	0.480	USA	0.77	UK	0.74
UK	0.348	UK	0.64	Italy	0.69
Japan	0.301	Germany	0.60	Japan	0.58
Switzerland	0.272	Japan	0.48	Switzerland	0.49
Germany	0.194	Switzerland	0.27	Germany	0.49

TABLE 5 Estimation of the McCallum model: comparison with results from regressions for the long rates.
Sample period: 16/11/1991-11/11/95 (no. obs 206).

Country (1)	**Estimates of λ and $\rho^{(i)}$** (2)	$\beta_{MC}^{(ii)}$ (3)	**Estimates of β** (4) $^{(iii)}$	**Wald test $^{(iv)}$ H_0: $\beta = \beta_{MC}$ (prob. of rejection)** (5)	**Wald test$^{(iv)}$ H_0: β=1 (prob. of rejection)** (6) $^{(iii)}$
USA	λ = 0.480 (.079) ρ = 0.810 (.040)	1.19	0.92	0.65 (.42)	0.05 (.81)
Japan	λ = 0.272 (.090) ρ = 0.797 (.041)	0.09	0.50	1.19 (.28)	1.78 (.18)
Germany	λ = 0.126 (.059) ρ = 0.914 (.026)	0.20	0.64	2.16 (.14)	1.50 (.22)
UK	λ = 0.446 (.080) ρ = 0.783 (.043)	0.76	0.84	0.05 (.82)	0.18 (.67)
France	λ = 0.920 (.263) ρ = 0.631 (.053)	0.95	1.86	3.18 (.08)	2.8 (.09)
Italy	λ = 0.567 (.150) ρ = 0.643 (.053)	0.02	0.51	12.15** (.00)	12.2** (.01)
Canada	λ = 0.494 (.121) ρ = 0.752 (.045)	0.68	0.89**	0.07 (.80)	0.02 (.89)
Switzerland	λ = 0.192* (.110) ρ = 0.809 (.038)	-.28	-.18**	0.07 (.78)	9.4** (.002)

Note. **(i)**: λ and ρ are estimated as explained in the text; * indicates that λ is not significantly different from zero at the 1% level; **(ii)** computed as in equation (23): $\beta_{MC}=2(\lambda\rho+\rho-1)(1+\rho+\rho^2+\rho^3)$; **(iii)** the values in columns 4 and 6 are taken from Table 2; ** in column 4 indicates that the coefficient is not statistically different from zero at the 5%; ** in column 6 indicates rejection at the 1%; **(iv)**: Wald test (chi-square values) computed with Newey-West corrected standard errors, consistent in the presence of MA(3) errors, and heteroscedastic.