The influence of short rate predictability and monetary policy on tests of the expectations hypothesis: some comparative evidence

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The influence of short rate predictability and monetary policy on tests of the expectations hypothesis: some comparative evidence

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Abstract

The aim of the present paper is to assess the role of short rate predictability and monetary policy in explaining different results from tests of the Expectations Hypothesis of the term structure. For this purpose McCallum (1994b) model for the interaction between the Expectations Hypothesis, a time-varying term premium and a policy reaction to the term spread is estimated using Eurorates for 8 countries in different subperiods between 1985 and 1995. The estimation is performed following Kugler (1997) modification of McCallum model. The results confirm previous findings by Kugler and suggest the important role played by monetary policy in explaining the empirical performance of the Expectation Hypothesis.

Keywords: expectation hypothesis, interest rates, monetary policy, term premium, term structure

JEL classification: C22, E43, E52

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1. INTRODUCTION

The validity of the Expectation Hypothesis (EH) of the term structure of interest rates (TS) has been long tested with alternate fortunes, whereby acceptance of the theory mainly realises for European countries (see inter alia Boero and Torricelli 1997, Engsted and Tanggaard, 1995, Gerlach and Smets, 1997) and refusal for the US (see Rudebusch, 1995, for a summary of different US studies). However, more 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 predictive content 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. Within this category falls Hardouvelis (1994) overreaction explanation, 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. Yet, recent empirical papers (e.g. Kugler, 1990, Gerlach and Smets, 1997) show that time-varying term premia are not as important as the scarce variability of short rates in diminishing the predictive content of the ET. The third explanation involves policy behaviour, and asserts that the limited variability of short rates is due to some particular monetary policy stances. The basic idea goes back to 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 so far: one merely empirical, the other theoretical. The former has been put forward by Dotsey and Otrok (1995) and Rudebusch (1995), who have empirically formalised Mankiw and Miron argument, by generating synthetic interest rate data from a Fed's interest rate targeting model which are then used to test the EH. Yet, the same 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 is therefore represented by a theoretical model of the type proposed by McCallum (1994b). The author develops a model, in a two- and in a N-period setting, for 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 has been successfully applied by Kugler (1997) and Hsu and Kugler (1997).

Aim of the present paper is to further explore the influence of monetary policy on tests of the expectations hypothesis and its ability to explain rejections compared to two alternative explanations: small sample bias and time-varying term premia. Differently from the presence of a time-varying term premium, the issue of biases in tests of the expectations hypothesis has been neglected in most of the empirical literature on tests of the EH, and only recently has received increased attention (see Bekaert, Hodrick and Marshall, 1997, and Schotman, 1997). Our analysis of the small sample bias effects draws on the empirical distributions derived in Bekaert et al. (1997). The empirical analysis is conducted with Euro-rates for eight different countries: USA, Japan, Germany, UK, France, Italy, Canada and Switzerland.

The plan of the paper is as follows. Section 2 outlines the main features of McCallum model and Kugler's exact solution and Section 3 presents the procedure used for the estimation of the model as described in Kugler (1997). In Section 4 we report new evidence on standard tests of the EH for eight countries, and conduct inference using both asymptotic and small sample distributions. In Section 5 we investigate the importance of the predictability of the short rate and of the term premium in tests of the EH. In Section 6 we present results from the estimation of the McCallum model and Section 7 concludes.

[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.

1. THE McCALLUM MODEL

In this section we describe the theoretical model tested in the present paper. The model was originally set up by McCallum (1994) 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's is an N-period model essentially characterised by an equation for the TS and an equation for the monetary policy rule. As for the former, it is represented by the EH modified by the existence of a time varying term premium of an autoregressive type, which implies that the return on a N period bond is given by:

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

$$\xi_t = \rho \xi_{t-1} + u_t \qquad (2)$$

where:

R_t^N is the return on a N period long bond,

N is time to maturity of the long bond,

r_t is the return on a one-period bond,

ξ_t is the term premium on the long bond with $|\rho| < 1$ and u_t white noise.[2]

Assuming that, for N large, the following hypothesis is reasonable:

$$E_t R_{t+1}^N = E_t R_{t+1}^{N-1} \qquad (3)$$

eq. (1) can be approximated as follows:

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

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

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

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[2] ξ_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 (1997b).

$$N(R_{t+1} - R_t) = (R_t - r_t) - \xi_t + N\varepsilon_t \qquad (5)$$

where $\varepsilon_t = R_{t+1} - E_t R_{t+1}$ 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 rates smoothing combined with a reaction to the term spread, i.e.:

$$\Delta r_t = r_t - r_{t-1} = \lambda(R_t - r_t) + \zeta_t \qquad (6)$$

where $\lambda \geq 0$ and ζ_t represents other components of policy behaviour and, for simplicity, is assumed to be white noise[3]. The rule is taken on the basis of «*the observation that actual policy behavior in the U.S. (and many other nations) involves manipulation of a short-term interest rate «instrument» or «operating» variable.*» (McCallum, 1994b).

Obviously, eq. (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. Yet, given the correlation between the spread and other indicators (e.g. real economic growth, inflation expectations) this simple rule can be thought of as capturing also policy instances of those Central Banks which officially do not use the spread as an indicator (e.g. the Bundesbank).

Combining (4) and (6) gives:

$$(1+N)R_t = NE_t R_{t+1} + (1+\lambda)^{-1}[r_{t-1} + \lambda R_t + \zeta_t] + \xi_t \qquad (7)$$

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:

$$R_t = \phi_1 r_{t-1} + \phi_2 \xi_t + \phi_3 \zeta_t \qquad (8)$$

and is given by:

$$R_t = r_{t-1} + \frac{1+\lambda}{N - \rho(N-1)(1+\lambda)}\xi_t + \zeta_t \qquad (9)$$

The relevant regressions accordingly become:

$$r_t - r_{t-1} = \lambda\rho(R_{t-1} - r_{t-1}) + \frac{\lambda}{N - \rho(N-1)(1+\lambda)} u_t + \zeta_t \qquad (10)$$

$$R_t - R_{t-1} = (\lambda\rho + \rho - 1)(R_{t-1} - r_{t-1}) + \frac{(1+\lambda)}{N - \rho(N-1)(1+\lambda)} u_t + \zeta_t \qquad (11)$$

McCallum underlines that, except for very large values of ρ and/or λ, the coefficient will be negative thus matching some empirical results for the U.S.(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) paper offers an exact solution to the N-period model just presented. Specifically, the solution to McCallum's model hinges on the approximation (3), which in fact allows one to get rid of the expected values for the short rate up to date N. In order to avoid the approximation above, Kugler has to look for 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. Accordingly, Kugler's regression equation for the spread and for the short rate are respectively the following:

$$(R_t - r_t) = \rho(R_{t-1} - r_{t-1}) + \frac{N}{N - \lambda\sum_{j=1}^{N-1}(N-j)\rho^j} u_t \qquad (12)$$

$$(r_t - r_{t-1}) = \lambda\rho(R_{t-1} - r_{t-1}) + \frac{N\lambda}{N - \lambda\sum_{j=1}^{N-1}(N-j)\rho^j} u_t + \zeta_t \qquad (13)$$

The latter has essentially the same implication as the corresponding equation in McCallum's two-period model, i.e. the information content of the spread vanishes whenever λ or ρ tends to zero. In fact Kugler concludes on the point: «*This finding can be interpreted as follows: the predictive power of the spread for the short rate is based on predictable policy reaction of the central bank to the spread. However, if ρ is zero there is no predictable exogenous movements of the spread which results in predictable policy reactions.*»

Kugler does not present the regression equation for the long rates, which we have derived from

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

his solutions:

$$R_t - R_{t-1} = (\lambda\rho + \rho - 1)(R_{t-1} - r_{t-1}) + \frac{(1+\lambda)}{1 - \lambda\sum_{i=1}^{N-1}(1 - \frac{j}{N})\rho^j} u_t + \zeta_t \qquad (14)$$

By comparative inspection of (11) and (14), it is clear that the exactness of the solution worked out 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.

2. KUGLER TEST OF THE MODEL

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 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. While Kugler carries out this test only with respect to regressions for the short rate we present evidence also for regressions for the longer rate.

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

$$(1/N)\sum_{j=1}^{N-1}(N-j)\Delta r_{t+j} = \alpha + \beta(R_t^N - r_t) + \varepsilon_{t+N-1} \qquad (15)$$

$$R_{t+1}^{(N-1)} - R_t^N = \alpha + \beta\left(\frac{1}{(N-1)}\right)(R_t^N - r_t) + \varepsilon_{t+1} \qquad (16)$$

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

Taking expected values of Eq.s (12) and (13) and rearranging, it can be shown that Kugler's

model implies:

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

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

$$(1/N)\sum_{j=1}^{N-1}(N-j)\Delta r^{e}_{t+j} = (1/N)\lambda\sum_{j=1}^{N-1}(N-j)\rho^{j}(R_t - r_t) \quad (17)$$

where the implied β is:

$$\beta_{MC} = (1/N)\lambda\sum_{j=1}^{N-1}(N-j)\rho^{j} \quad (18)$$

while the counterpart of (16) is Eq. (14) where the implied β is:

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

The suggestion from Kugler is to estimate β_{MC} by means of Indirect Least Squares in two stages:

i) Estimate using OLS the two reduced form equation (12) and (13) to obtain estimated value for ρ and $\lambda\rho$,

ii) Divide the estimated value of $\lambda\rho$ by ρ to obtain the estimated value of λ and use (18) and (19) to obtain values of β_{MC}.

These values are then compared with those obtained from estimation of 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 USA, 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 for the USA. The latter case is further investigated in Hsu and Kugler (1997).

In order to strengthen the empirical test of the model proposed by McCallum and further developed by Kugler, in the following sections we present empirical evidence on tests of the

expectations hypothesis for a wider range of countries with different monetary policy rules, and for different sample periods.

4 TESTS OF THE EXPECTATIONS HYPOTHESIS

Different tests of the EH have been proposed in the literature. In this section we present evidence based on the implications of Eq. (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 Eq.s (15) and (16). In these equations ε_{t+N-1} and ε_{t+1} are forecast errors which under RE are orthogonal to information at time t, and therefore uncorrelated with the regressor R_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 (19) will follow a MA process of order m-1 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.

Data

The data set used in this study are weekly Eurorates for the period 16-11-1985 to 11-11-1995 for USA, Japan, Germany, U.K., 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.[4] On the other hand, as we will see shortly, this choice creates some problems to our tests based on regressions for the longer rates, as when these regressions are performed on the short end of the term structure they are subject to serious approximations errors in addition to a significant small sample bias (see Bekaert et al. 1997).

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

4.1 SHORT RATE REGRESSIONS

The equation estimated for the short rate is equation (15), which with N=3 months becomes:

$$(1/3)\sum_{j=1}^{3-1}(3-j)\Delta r_{t+j} = \alpha + \beta(R_t^3 - r_t) + \varepsilon_{t+3-1}$$

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

$$(2/3)(r_{t+4}-r_t)+(1/3)(r_{t+9}-r_{t+4}) = \alpha + \beta(R^3_t - r_t) + \varepsilon_{t+9} \qquad (20)$$

The errors in Eq. (20) follow an MA(h-1) process, where h is the forecast horizon, 9 weeks in our regressions, so we computed Newey-West corrected standard errors with truncation lag equal 8. The results are summarised in Table 1. The estimation period selected for each country varies between 16/11/1985 and 11/11/1995, and reflects the longest period for which the estimated β was found to be stable. We first discuss evidence based on standard distributions, then we will consider the effects of small sample bias.

TABLE 1 - Estimates of β in regressions for the short rate

Equation (20): $S^{*(3,1)}_t = \alpha + \beta(R^3_t - r_t) + \varepsilon_t$ where

$S^{*(3,1)}_t = (2/3)(r_{t+4}-r_t)+(1/3)(r_{t+9}-r_{t+4})$

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)	β = 0.77 (0.105)	4.36 *
Japan	16/11/90-11/11/95 (253)	β = 0.48 (0.098)	27.56 **
Germany	16/11/85-11/11/95 (513)	β = 0.60 (0.160)	6.15 *
U. K.	16/11/90-11/11/95 (253)	β = 0.64 (0.114)	9.67 **
France	16/11/90-11/11/95 (253)	β = 1.12 (0.223)	0.30
Italy	16/11/85-11/11/95 (513)	β= 0.80 (0.102)	3.57
Canada	16/11/90-11/11/95 (253)	β = 0.95 (0.259)	0.02
Switzerland	16/11/90-11/11/95 (253)	β = 0.27 ** (0.193)	13.69 **

Note. **(i)**:** indicates that the coefficient is not statistically different from zero at conventional levels;
(ii) the number in parentesis are heteroscedasticity and autocorrelation corrected standard errors; truncation lag 8.
(iii): * indicates rejection of H_0: β=1 at the 5%; ** indicates rejection at the 1%.

Inference based on asymptotic distributions

Table 1 shows that all estimates for β, but the one for Switzerland, are significantly different from zero, thus confirming an overall information content of the spread for 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 the results by Gerlach and Smets (1997), who mantain that the EH better describes weak-currency countries than strong-currency ones. The rationale behind this

explanation being that Central Banks of weak-currency countries often have to manage short rates in order to achieve mid-term intermediate exchange rate objectives, which are well-known and therefore make changes in the short rates more predictable. This justifies higher values for β in those countries.

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 will see below, but in a recent paper Bekaert et al. (1997) found that also regressions for the short rate can be affected by substantial positive bias.[5]

In order to evaluate the effects of small sample bias in the tests presented so far, we also conduct inference by using the 5% quantiles of the empirical distributions of the slope coefficient derived by Bekaert et al. under two alternative data generating processes: an AR(1) for the short rate (see Panel C, Table 3) and a VAR-GARCH model for the short rate and the spreads (see 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 the small sample distributions the slope coefficient should be smaller than 0.64 (with the AR(1) d.g.p.) and 0.62 (with the VAR-GARCH d.g.p.) to have a 5% rejection of the hypothesis $\beta=1$ in a one-tailed test.

So, overall, by conducting inference with the small sample distributions, our results remain virtually unaffected, the only exceptions are the U.S. and the U.K. for which rejection would be weakened. However, we must stress that these results are only indicative, and should be interpreted with caution, as the critical values derived by Bekaert et al. are not exactly applicable to our regressions with such a short term spread (the closest spread considered in the Monte Carlo experiment

[5] 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

conducted by Bekaert et al. is 12-1 month, whereas we are using a 3-1 month spread). Moreover they report results for only one sample size (524 observations) which is larger than the samples used in our tests.

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:

$$R_{t+1}^{(3-1)} - R_t^3 = \alpha + \beta\left(\frac{1}{(3-1)}\right)(R_t^3 - r_t) + \varepsilon_{t+1}$$

This regression was modified for weekly data to obtain

$$R^3_{t+4} - R^3_t = \alpha + \beta(1/2)(R^3_t - r_t) + \varepsilon_{t+4} \qquad (21)$$

In Eq. (21) we have used approximation (3), discussed in Section 2, $E_tR^N_{t+1} = E_tR^{N-1}_{t+1}$, which is commonly adopted in regressions of this type. This approximation is irrelevant for large N, but may have significant bias effect on the estimated value of β for small N. These regressions 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, a vaste empirical literature has reported estimated coefficients below unity, and negative point estimates. These become more 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.

Inference based on asymptotic distributions

Our results reported in Table 2 and obtained for the whole sample period (1985-1995) and for a sub-period (1991-95) are not in line with previous findings: most point estimates are positive, except that for Switzerland, and some are close to one. This result may depend on the particular

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data, generates a positive bias in the slope coefficients (see Bekaert et al., 1997, eqs. 7, 8 and 10).

nature of our study which only looks 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. However, the very low R^2s (columns 5 and 6 in Table 2) confirm previous results whereby the spread between the long and short term interest rates has poor predictive content for changes in the longer rate.

TABLE 2 - Estimates of β in regressions for the long rate

Equation (21): $R^3_{t+4}-R^3_t = \alpha + \beta(1/2)(R^3_t-R^1_t) + \varepsilon_{t+4}$

Sample periods 16/11/85-11/11/95 (no. obs 518) and 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
U. K.	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** (.00)	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

Note. **(i)**:** indicates that the coefficient is not statistically different from zero at conventional significance levels;

(ii) the number in parentesis 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%.

The effects of small sample bias

Before we can draw any conclusion from these tests of the EH, we evaluate the results also in the light of the small sample distributions of the slope coefficients. In fact, in the study mentioned before by Bekaert et al. (1997), the small sample bias which affects all regression-based tests of the EH is shown to be particularly strong for the long rate regressions. Moreover, approximation (3) used in the estimation of Eq. (21) introduces a further error in the regression which exacerbates the small sample bias. Bekaert et al. (1997) found that for n=12 and sample size 524 the average of the OLS estimates of beta 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. Differently from regressions for the short rates, inference based on the small sample distributions is not uniformly conclusive about rejection (or lack of rejection) of the EH. In fact, according to the empirical quantiles tabulated in Bekaert et al. the EH should be rejected at the 5% for values of β <1.203 when the d.g.p. for the short rate is an AR(1) model (Panel B, table 3), and for values of β <0.131 if inference is conducted under the assumption of a VAR-GARCH model for the short rate and the spreads (Panel B, Table 6). So with these critical values our tests in Table 2 would find 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 only exception for Switzerland.

To summarise this section, inference on regression-based tests of the EH has been conducted with both asymptotic and small sample distributions. The empirical critical values are those derived in Bekaert et al. (1997). Although these tests are affected by substantial positive bias (which depends on the persistence of the short rate), the increased dispersion in the small sample distribution leads to results that are in general more favourable to the EH for regressions for the short rates, whereas results remain inconclusive for tests from regressions for the long rate. This

latter result depends on the sensitivity of the small sample distributions to the data generation process.[6]

The problems already underlined with the empirical distributions adopted in this section impose serious limitations to their use and inference based on them can be misleading. So, the important issue is to find explanations for the rejections of the EH. In the next section we will look at the effects of predictability of short rates and the importance of the term premium. Then we will look at a possible explanation in terms of policy behaviour.

5. PREDICTABILITY OF SHORT RATES AND IMPORTANCE OF THE TERM PREMIUM

So far we have presented standard tests of the EH with an attempt to evaluate the effects of small sample bias due to persistence in the short rate. We have seen that although these tests are affected by substantial positive bias, the increased dispersion in the small sample distribution leads to results that are more favourable to the EH for regressions for the short rates, whereas results remained inconclusive for tests from regressions for the long rate. This approach, however, is silent about the consequences of the presence of a time-varying term premium, in violation of the EH.

In this section we take a different perspective, and following Mankiw and Miron (1986) we examine the role of the predictability of short term rates for tests of the EH, in the presence of a time-varying term premium. For this analysis we focus only on regressions for the short rate. The conjecture put forth by Mankiw and Miron (1986) is that in the presence of a time-varying term premium, differences in the predictability of short rates can explain why the EH is supported in some countries (or in some periods) and not in others. Specifically, the EH should be better supported by the data the more predictable short term rates are, while the theory is more easily

[6] 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 beta. The bias is positive or negative depending on whether the sum of the autoregressive and moving average coefficients is negative or positive.

rejected for countries (or in periods) in which short term rates are difficult to predict.

In Table 3 below we list the countries according to a measure of the predictability of short rates and a measure of the relative importance of the risk premium. To compute these measures, we followed Gerlach and Smets (1997b) and Kugler (1990). The starting point is the expression for the probability limit (plim) of the OLS estimator of the slope coefficient in regressions for the short rate:

$$p\lim\hat{\beta} = \frac{\sigma^2(ES^*) + \rho\sigma(ES^*)\sigma(\xi)}{\sigma^2(ES^*) + \sigma^2(\xi) + 2\rho\sigma(ES^*)\sigma(\xi)} \qquad (22)$$

where S* is the dependent variable in regression (15) sometimes referred to as the roll-over spread, or the perfect foresight spread $S^* = (1/N)\sum_{j=1}^{N-1}(N-j)\Delta r_{t+j}$, and σ^2(ES*) is the variance of expected changes in the short rate ES*.

$\sigma^2(\xi)$ is the variance of the term premium ξ, and ρ the correlation between ES* and ξ. For simplicity we have omitted the time subscripts. From equation (22) it follows that when the variance of the term premium is zero, the plim of the estimate of β is 1, while in the presence of a time-varying term premium the coefficient estimate of β is biased, and the bias depends on the the variance of ES* and on ρ. If σ^2(ES*) goes to zero, the estimate of β tends to zero.

By division for the variance of changes in the short rate, σ^2(S*), the formula above can be rewritten as follows:

$$\text{plim } \beta = (R^2+\rho\Theta R)/(R^2+2\rho\Theta R+\Theta^2) \qquad (23)$$

where $R^2 = \sigma^2(ES^*)/\sigma^2(S^*)$ can be interpreted as a measure of the forecastability of changes in the short rates, and $\Theta^2 = \sigma^2(\xi)/\sigma^2(S^*)$ as a measure of the importance of the term premium relative to the variance of S*. Eq. (23) indicates that a low estimate for β can be attributed either to a low R^2 or to a high Θ^2 or both. As a measure of the forecastability of changes in the short rates we use the R^2 obtained from the standard regressions for the short rates. To compute Θ^2 we followed Kugler

(1990) and obtained $\sigma^2(\xi)$ as the variance of the fitted values from a regression of the ex post excess returns on the spread.[7] These two indices are presented in Table 3 with estimates of β and tests for the EH that $\beta=1$. The observation period is the same for all countries, 16/11/1991-11/11/95 (sample size 201), to facilitate comparison.

TABLE 3 Predictability of short rates (R^2) and importance of the term premium (Θ^2). Sample period 16/11/1991-11/11/95 (no. of obs. 201).

Country	**Estimates of $\beta^{(i)}$.**	**Wald test (ii) Chi-Sq for H_0: $\beta=1$**	**$R^{2(iii)}$**	**$\Theta^{2(iv)}$**
USA	$\beta = 0.77$ (0.105)	4.36*	$R^2 = 0.44$	$\Theta^2 = 0.29$
Japan	$\beta = 0.46$ (0.151)	12.74**	$R^2 = 0.15$	$\Theta^2 = 0.43$
Germany	$\beta = 0.55$ (0.135)	11.02**	$R^2 = 0.24$	$\Theta^2 = 0.60$
UK	$\beta = 0.68$ (0.098)	10.30**	$R^2 = 0.30$	$\Theta^2 = 0.31$
France	$\beta = 1.22$ (0.230)	0.96	$R^2 = 0.34$	$\Theta^2 = 0.07$
Italy	$\beta = 0.79$ (0.104)	3.84	$R^2 = 0.40$	$\Theta^2 = 0.24$
Canada	$\beta = 0.82$ (0.306)	0.30	$R^2 = 0.15$	$\Theta^2 = 0.09$
Switzerland	$\beta = 0.09$ ** (0.135)	20.34**	$R^2 = 0.01$	$\Theta^2 = 0.19$

Note. **(i)**: ** indicates that the coefficient is not significant at the 5% level. Corrected S.E.s in parentesis. **(ii)**: * indicates rejection of H_0: $\beta=1$ at the 5%; ** indicates rejection at the 1%.
(iii) $R^2 = \sigma^2(E(S^*))/\sigma^2(S^*)$ is the R^2 from regressions for the short rate; **(iv)** $\Theta^2 = \sigma^2(\xi)/\sigma^2(S^*)$ is a measure of the relative importance of the term premium, with $\sigma^2(\xi)$ computed as the variance of the fitted values from a regression of the ex post excess returns on the spread.

[7] The ex post excess returns are proxied by the following variable: $R^3_t - (1/3)(r_t+r_{t+1}+r_{t+2})$ which is regressed on a constant and the spread $R^3_t - r_t$.

From Table 3 it is clear that the predictability of the short rate and the importance of the term premium differ significantly amongst countries. The average value for R^2 is 0.24, and for Θ^2 is 0.28. The three cases where the EH is supported are France, Italy and Canada (see Table 1). These countries have either a high R^2 (above average) (France and Italy), or a very low Θ^2 (France and Canada) or both (France). Switzerland is the country with lowest β and this is explained with the very low R^2. Finally Japan and Canada form an interesting case: the same value of R^2 (below average) is associated in Canada with a very low Θ^2 (which explains the high β) and in Japan with a high Θ^2 (which explains the lower value of β).

To conclude, from this analysis it emerges that attempts to explain failures of standard tests of the EH must take into account differences in the predictability of changes in the short rate and the relative importance of a time-varying term premium. However, the analysis so far does not provide an explanation of why interest rates are more predictable in some countries than in others, or why term premia are more variable in some countries than in others, and therefore it can not rationalise differences in the results of tests of the EH.

In the next section we present an attempt to explain different evidence obtained from standard tests of the EH which involves considerations about policy behaviour. We do this by estimating the McCallum model introduced in Section 2, and comparing the value of the coefficient β implied by this model with that obtained from standard test of the EH discussed in Section 4.

6. AN APPLICATION OF 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. 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 (λ) and the autoregressive component of the term premium (ρ). The values of β_{MC} are obtained from Eq.s (18) and (19) modified according to the weekly frequency of the data as follows:

β_{MC} for short rate regressions:

$$\beta_{MC} = (1/3)\lambda(2\sum_{j=1}^{4}\rho^{j} + \sum_{j=5}^{8}\rho^{j}) \qquad (24)$$

β_{MC} for long rate regressions:

$$\beta_{MC} = 2(\lambda\rho+\rho-1)(1+\rho+\rho^{2}+\rho^{3}) \qquad (25)$$

Estimates of λ and ρ are obtained from the two Reduced Form 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 $(R_{t-1}-r_{t-1})$.

6.1 SHORT RATE REGRESSIONS

In Table 4 we report estimates of λ, ρ, the implied β_{MC}, for regressions for the short rate (Eq. 24) and a test for $\beta_{HAT}=\beta_{MC}$, where β_{HAT} are the same estimates of β reported in Table 1 and Table 3.

A general result is that the implied slope coefficient β_{MC} is consistent with the β estimates obtained from standard EH regressions, indicating that the McCallum model can rationalise very different values for β, including very low values like in the case of Switzerland. Values statistically different from one are consistent with EH and reflect the way in which monetary policy responds to changes in TS.

TABLE 4- Estimation of McCallum model:
comparison with results from regressions for the short rate

Country and sample period	**Estimates of λ and ρ[(i)]**	β_{MC}[(ii)]	**Estimated β from standard regressions**	**Wald test[(iii)] H_0: βhat=β_{MC}**	**Wald test[(iii)] H_0: βhat =1**
USA 11/91-11/95	λ = 0.480 (.079) ρ = 0.810 (.040)	β_{MC} = 0.97	β = 0.77 (0.105)	Chi-Sq.=3.27	4.36 *
Japan 11/90-11/95	λ = 0.301 (.074) ρ = 0.801 (.035)	β_{MC} =0.58	β = 0.48 (0.098)	Chi-Sq.=1.04	27.56 **
Japan 11/91-11/95	λ = 0.272 (.090) ρ = 0.797 (.041)	β_{MC} =0.52	β = 0.46 (0.151)	Chi-Sq.=0.18	12.74**
Germany 11/85-11/95	λ = 0.194 (.042) ρ = 0.871 (.021)	β_{MC} =0.49	β = 0.60 (0.160)	Chi-Sq.=0.42	6.15 *
Germany 11/91-11/95	λ = 0.126 (.059) ρ = 0.914 (.026)	β_{MC} =0.38	β = 0.55 (0.135)	Chi-Sq.=1.48	11.02**
UK 11/90-11/95	λ = 0.348 (.059) ρ = 0.823 (.034)	β_{MC} =0.74	β = 0.64 (0.114)	Chi-Sq.=0.72	9.67 **
UK 11/91-11/95	λ = 0.446 (.080) ρ = 0.783 (.043)	β_{MC} =0.81	β = 0.68 (0.098)	Chi-Sq.=1.71	10.30**
France 11/90-11/95	λ = 0.791 (.207) ρ = 0.665 (.046)	β_{MC} =0.93	β = 1.12 (0.223)	Chi-Sq.=0.73	0.30

Table 4 (continued)

France 11/91-11/95	$\lambda = 0.920$ (.263) $\rho = 0.631$ (.053)	$\beta_{MC} = 0.95$	$\beta = 1.22$ (0.230)	Chi-Sq.=1.41	0.96
Italy 11/85-11/95	$\lambda = 0.751$ (.124) $\rho = 0.598$ (.035)	$\beta_{MC} = 0.69$	$\beta = 0.80$ (0.102)	Chi-Sq.=1.18	3.57
Italy 11/91-11/95	$\lambda = 0.567$ (.150) $\rho = 0.643$ (.053)	$\beta_{MC} = 0.61$	$\beta = 0.79$ (0.104)	Chi-Sq.=2.85	3.84
Canada 11/90-11/95	$\lambda = 0.497$ (.101) $\rho = 0.771$ (.039)	$\beta_{MC} = 0.86$	$\beta = 0.95$ (0.259)	Chi-Sq.=0.11	0.02
Canada 11/91-11/95	$\lambda = 0.494$ (.121) $\rho = 0.752$ (.045)	$\beta_{MC} = 0.80$	$\beta = 0.82$ (0.306)	Chi-Sq.=0.01	0.30
Switzerland 11/90-11/95	$\lambda = 0.272$ (.112) $\rho = 0.779$ (.038)	$\beta_{MC} = 0.49$	$\beta = 0.27$ ** (0.193)	Chi-Sq.=1.21	13.69 **
Switzerland 11/91-11/95	$\lambda = 0.192$* (.110) $\rho = 0.809$ (.038)	$\beta_{MC} = 0.38$	$\beta = 0.09$ ** (0.135)	Chi-Sq.=2.11	20.34**

Note. **(i)**: * indicates that λ is not significant at the 1% level.

(ii): computed as $\beta^{MC} = (1/3)\lambda(2\sum_{j=1}^{4}\rho^j + \sum_{j=5}^{8}\rho^j)$

(iii): Wald test computed with Newey-West corrected standard errors, consistent in the presence of MA(8) errors, and heteroscedastic.

So, these results seem to suggest that explicit consideration of a monetary policy reaction function is important in providing an explanation of both failures and successes of the EH. Moreover, the table shows that the differences in the values of the theoretical βs are due more to different monetary policy instances than to different pattern of time variation in the term premium. In fact, while estimated values for λ range from a minimum of 0.126 for Germany to a maximum of 0.92 for France, those for ρ display a much lower variation. In particular, results point to a strong policy reaction to the spread for France (λ=0.92) and Italy (λ=0.57) (two countries for which the EH could not be rejected), a low policy reaction to the spread for Germany (λ=0.126), Switzerland (λ=0.192) and Japan (λ=0.272), and a moderate reaction for UK (λ=0.446), U.S. (λ=0.48) and Canada (λ=0.494).

Table 5 casts further light on this point, showing that the ranking between the countries is similar according to the three parameters λ, β_{HAT} and β_{MC}. The four countries with highest β coincide with those with higher λ, therefore confirming the importance of monetary policy in explaining differences between the estimated values of β.

The high values of λ obtained for France and Italy can be interpreted in line with Gerlach and Smets (1997) as follows: both countries operate with intermediate exchange rate targets, so policy is more predictable. In the period under investigation both countries had to defend their exchange rate by raising the short rate in response to a high spread signalling a future depreciation in the exchange rate. Within McCallum's model, depreciation fears can be thought of as the exogenous shock u_t in Eq. (2) that increases the risk premium ξ_t (and hence the spread) requested to invest in a currency under devaluation risk.

The moderately high value of λ for the US is in line with recent findings for the US (Hsu and Kugler, 1997) reflecting the increased reliance on the spread as a policy indicator rather than monetary aggregates. This estimate of λ combined with a high value for ρ (0.81) implies a value for the slope coefficient β_{MC} close to 1 (0.97).

Finally the case of Germany, where an extremely low value of λ is displayed, corresponds to official statements of the Bundesbank which clearly indicate that the spread is not used as a policy indicator. Yet the model cannot be rejected in this case too, since the low value for the estimated β indicates that either monetary policy was in the period little predictable or the term premia were highly variable or both (see also high value for Θ^2 in Table 3).

TABLE 5- Countries ranked according to values of λ, β_{HAT} and β_{MC}. Sample period 16/11/91-11/11/95.

λ		β_{HAT}		β_{MC}	
France	$\lambda = 0.92$	France	β=1.22	USA	β_{MC} =0.97
Italy	$\lambda = 0.56$	Canada	β=0.82	France	β_{MC}=0.95
Canada	$\lambda = 0.49$	Italy	β=0.79	UK	β_{MC}=0.81
USA	$\lambda = 0.48$	USA	β=0.77	Canada	β_{MC}=0.80
UK	$\lambda = 0.44$	UK	β=0.68	Italy	β_{MC}=0.61
Japan	$\lambda = 0.21$	Japan	β=0.46	Japan	β_{MC}=0.52
Switzerland	$\lambda = 0.19$	Germany	β=0.55	Germany	β_{MC}=0.38
Germany	$\lambda = 0.12$	Switzerland	β=0.09	Switzerland	β_{MC}=0.38

The main problem with the implementation of the McCallum model consists in the estimates of λ. In fact, these are based on an extremely simplified policy reaction function, where policy responds only to the spread (reflecting 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, in the remainder of this section we investigate the possibility that Central Banks react to lagged short rate changes, and reestimate 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}). These estimates of λ are very close to those reported in Table 4 for most countries. The only exceptions are France and Italy. For example, the value of λ for France changed from 0.92 to 0.44 when the AR(1) procedure was employed, implying significant changes in the value of β_{MC}. So, the stylised reaction function of the McCallum model is adequate to describe the monetary policy in countries like the U.S. where the spread is clearly used as an important indicator, while for countries like France and Italy the econometric estimation of the model requires consideration of a more complex set of policy indicators.

6.2 LONG RATE REGRESSIONS

Finally in Table 6 we compare estimates of the β coefficients obtained from regressions for the long rate (same as in Table 2) with the value for β implied by the policy reaction model (Eq. 25). 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 (β_{HAT} in the Table) cannot reject this hypothesis. Although these results are somehow suggestive of the ability of the McCallum model to 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 above, but also because of the very low R^2s.

TABLE 6 Estimation of McCallum model: comparison with results from regressions for the long rates. Sample period: 16/11/1991-11/11/95 (no. obs 206).

Country	Estimates of λ and ρ[(i)].	β_{MC}[(ii)]	Estimated β[(iii)] (β_{HAT})	Wald test [(iv)] H_0: $\beta_{HAT} = \beta_{MC}$ (prob. of rejection)	Wald test[(iv)] H_0: β=1 (prob. of rejection)
USA	λ = 0.480 (.079) ρ = 0.810 (.040)	β_{MC} = 1.19	β_{HAT} = 0.92	0.65(.42)	0.05(0.81)
Japan	λ = 0.272 (.090) ρ = 0.797 (.041)	β_{MC} =0.09	β_{HAT} = 0.50**	1.19(.28)	1.78(0.18)
Germany	λ = 0.126 (.059) ρ = 0.914 (.026)	β_{MC} =0.20	β_{HAT} = 0.63	2.16(.14)	1.50(.22)
UK	λ = 0.446 (.080) ρ = 0.783 (.043)	β_{MC} =0.76	β_{HAT} = 0.84	0.05(.82)	0.18(0.67)
France	λ = 0.920 (.263) ρ = 0.631 (.053)	β_{MC} =0.95	β_{HAT}= 1.86	3.18(.08)	2.8(0.09)
Italy	λ = 0.567 (.150) ρ = 0.643 (.053)	β_{MC} =0.02	β_{HAT}= 0.51	12.15(.00)**	12.2(0.01)**
Canada	λ = 0.494 (.121) ρ = 0.752 (.045)	β_{MC} =0.68	β_{HAT}= 0.89**	0.07(.80)	0.02(0.89)
Switzerland	λ = 0.192* (.110) ρ = 0.809 (.038)	β_{MC} =-.28	β_{HAT}= -.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) β_{MC} is computed from the formula: $2(\lambda\rho+\rho-1)(1+\rho+\rho^2+\rho^3)$

(iii) these values are taken from Table 2; ** indicates that the coefficient is not statistically different from zero at conventional levels;

(iv): Wald test computed with Newey-West corrected standard errors, consistent in the presence of MA(3) errors, and heteroscedastic. ** indicates rejection at 1%.

7. CONCLUSIONS

In this paper we tested the EH and possible explanations for its failures across eight countries in the period 1985-1995. To this end we performed our empirical analysis in three steps. First, we simply tested the EH obtaining disparate evidence across the countries considered. Secondly, we tried to account for these latter results by examining the predictability of the short rate in the presence of a time-varying term premium. Finally, we tested McCallum model to assess how the EH interacts with a policy reaction function in the presence of a time-varying term premium. Our results can be summarised as follows.

Inference on regression-based tests of the EH was conducted with both asymptotic and small sample distributions on the basis of the empirical critical values derived in Bekaert et al. (1997). Although these tests are affected by substantial positive bias, the increased dispersion in the small sample distribution lead to more favourable results to the EH for regressions for the short rate, whereas results remained inconclusive for tests on the long rate.

The analysis conducted in section 4, however, could not throw any light on the possible effects of a time varying term premium. One root commonly followed by researcher in this area, is to explain different results in tests of the EH with differences in the forecastability of the short rate, in the presence of a time-varying term premium. In section 5 we attempted to establish a link between the forecastability of the short rate and the outcome of the tests of the EH for the different countries. From this analysis it was clear that the predictability of the short rate and the importance of the term premium differ significantly amongst countries. There was also strong evidence that high values of β are associated with high predictability of the short rate and a low relevance of the term premium.

Finally, in section 6, we investigated the role of a policy reaction function in explaining the empirical evidence obtained in the previous sections. A general result was that the slope coefficient implied by the model, β_{MC}, is consistent with the β estimates obtained from standard EH regressions, indicating that the McCallum model can rationalise very different values for β,

including very low values like in the case of Switzerland. Moreover, the differences in the values of the theoretical βs are due more to the different monetary policy instances than to the different pattern of time variation in the term premium. In fact, while estimated values for λ range from a minimum of 0.12 for Germany to a maximum of 0.92 for France, those for ρ display a much lower variation. In order to account for the extremely simplified reaction function assumed in the model, we also reestimated the policy reaction function with IV and an AR(1) error term. Estimates did not significantly change, but for the case of France and Italy, which suggests that – at least for some countries - a richer set of monetary policy indicators should be considered.

In conclusion, the tests presented in the present paper stress that, in order to explain the differential performance of the EH across the countries considered in this study, the relevance of monetary policy consideration appears to be bigger than that of a time-varying term premium. Yet, since McCallum model takes rather simple assumptions on both the policy reaction function and the time-varying term premium, the present research can be extended as to consider, on one hand, a different time-pattern of variation in the term premium and, on the other, a wider set of monetary policy indicators in the policy reaction function.

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