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Short- and Long-Run Tests of the Expectations Hypothesis: The Portuguese Case

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Abstract

The purpose of this paper is to test both short- and long-run implications of the (rational) expectations hypothesis of the term structure of interest rates using Portuguese data for the interbank money market. The results support only a very weak, long-run or "asymptotic" version of the hypothesis, and broadly agree with previous (but separate) evidence for other countries.

Empirical evidence supports the cointegration of Portuguese rates and the "puzzle" well known in the literature: although its forecasts of future short-term rates are in the correct direction, the spread between longer and shorter rates fails to forecast future longer rates. Further short-run implications of the hypothesis in terms of the predictive ability of the spread are also clearly rejected, even for the more stable period which emerged in the middle nineties.

Keywords: Term Structure; Expectations Hypothesis; Hypothesis Testing; Structural Breaks; Portugal

JEL Classification: E43, C22, C32

1. Introduction

The expectations hypothesis (EH) of the term structure of interest rates, which states that the observed term structure can be used to infer market participants' expectations about future interest rates, has been at the origin of an extraordinary amount of econometric analysis; see, e.g., Campbell (1995), Campbell and Shiller (1987, 1991), Engsted and Tanggaard (1994a, b), Hall et al. (1992), Hardouvelis

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(1994), Jondeau and Ricart (1999), Lanne (2000), Sarno et al. (2007), Thornton (2006), and Tzavalis (2003).

Understanding the term structure of interest rates has always been viewed as crucial to assess the impact of monetary policy and its transmission mechanism, to predict interest rates, exchange rates and economic activity, and to provide information about expectations of participants in financial markets. In this paper, the EH of the term structure of interest rates, embedding the rational expectations hypothesis, is tested with Portuguese data for interbank money market (IMM) rates. Contrasting with most of the previous literature, we scrutinize both short- and long-run implications of the hypothesis. Also, to the best of our knowledge, this paper contains the first examination of the EH using Portuguese IMM data.

The results support only a very weak version of the hypothesis and are in line with most of the (separate) conclusions in the literature. The empirical evidence supports a somewhat stable long-run relation of Portuguese rates but also the "puzzle" well known in the literature of the EH: although forecasts of short-term rates changes based on the spread are in the correct direction, it fails in forecasting future longer rates because the forecasts are in the wrong direction. More importantly, the stricter short-run implications of the hypothesis on the predictive ability of the spread are clearly rejected by our data. Hence, our evidence closely agrees with most of the previous results in the literature; see, *inter alia*, Arshanapalli and Doukas (1994), Engsted and Tanggaard (1994a, b), Hurn et al. (1995) and Tillmann (2007) for long-run evidence, and Campbell and Shiller (1991), Evans and Lewis (1994), Jondeau and Ricart (1999), Thornton (2006), Tzavalis (2003) and Tzavalis and Wickens (1997) for results on the short-run.

To summarize all the evidence we propose that such a weak version of the hypothesis is called long-run or asymptotic, a term which we borrow from the rational expectations hypothesis literature. Moreover, contrasting with most empirical studies, we provide an historically and statistically based sample-split analysis, which confirms and reinforces the results for the whole sample.

Actually, this quest for robustness is also a trait of our study. Instead of relying on a single model / method, we diverge from previous empirical assessments of the EH in the range of methods and models that we use. For instance, although relatively standard, we employ two techniques rarely (if ever) used to test the EH: dynamic OLS (DOLS) estimation and testing and *t-ecm* tests for cointegration. As another example, we employ several vector autoregressive (VAR) and vector error-correction (VECM) models chosen with different criteria. Naturally, this allows us to robustify our inferences.

The remainder of this paper is organized as follows. Some of the most important implications and testing procedures of the EH are reviewed in the next section, both in the framework of single and multiple equation models. We focus particularly on cointegration analysis and on the predictive ability of the spread. In section 3 we describe the data that we have used and in section 4 we present the main empirical evidence. In section 5 we evaluate the robustness of the long-run relation, discuss the sample-split point and reassess the evidence considering a partition of the sample. Section 6 concludes the paper.

2. Some Implications and Testing Procedures of the EH

2.1 In Single Equation Models

In the single equation setup the focus is on pairs of interest rates. Some of the available tests regarding the spread between interest rates of different maturities are described below.

2.1.1 Cointegration

Since nominal interest rates are bounded below by zero, the I(1) property cannot be strictly justified on theoretical grounds. However, their typical high persistent behaviour in response to shocks has led to an almost universal consensus about the presence of a unit root. Hence, cointegration methods are applicable.

Assuming that interest rates correspond to I(1) processes, the EH requires cointegration between interest rates with different maturities. Denoting the long and the short rates with $r_t^{(n)}$ and $r_t^{(m)}$, respectively, the stationarity of the spread, $S_t^{(n,m)} = r_t^{(n)} - r_t^{(m)}$, is a necessary, although not sufficient, condition for the EH to hold, as it is an implication of several term structure models. In fact, as is sometimes pointed out, more traditional theories also demand this condition; see, e.g., Lanne (2000), Patterson (2000), and Taylor (1992).

If the spread is stationary, then the term / risk premium is also stationary and interest rates are driven by a common stochastic trend, preventing them from drifting too far apart from the equilibrium, so that profitable arbitrage opportunities do not persist. The rate of inflation is the most obvious candidate to represent this common trend (Domínguez and Novales 2000, Engsted and Tanggaard 1994b).

2.1.2 The Spread As a Predictor of Interest Rate Changes

Turning to the stricter implications, the fundamental equation characterizing the EH states that the long-term interest rate equals an average of current and expected short-term interest rates over the life of the long-term interest rate plus a constant term, representing the time invariant term / risk premium $(\Phi^{(n)})$:

$$r_t^{(n)} = \frac{1}{k}\sum_{i=0}^{k-1} E_t\left[r_{t+im}^{(m)}\right] + \Phi^{(n)}, \tag{1}$$

where k is an integer denoting n/m. Expectations formulated at time t for the future evolution of short-term interest rates drive the longer-term interest rate.

When short-term interest rates are expected to rise, longer-term interest rates will also rise.

Using equation (1) it is straightforward to get

$$S_t^{(n,m)} = E_t\left[S_t^{*(n,m)}\right] + \Phi^{(n)} = \sum_{i=1}^{k-1} \frac{k-i}{k} E_t\left[\Delta\, r_{t+im}^{(m)}\right] + \Phi^{(n)}\,, \tag{2}$$

where $S_t^{*(n,m)}$ denotes the perfect foresight spread, i.e., the spread that would obtain if there were perfect foresight about future interest rates, and $\Delta\, r_{t+im}^{(m)} = r_{t+im}^{(m)} - r_{t+(i-1)m}^{(m)}$. Hence, the spread is a weighted average (with declining weights) of expected changes of short-term interest rates plus the term / risk premium. Since the spread is such an optimal predictor, a test for the validity of the EH may be based on the equation

$$S_t^{*(n,m)} = \delta_0 + \delta_1 S_t^{(n,m)} + \xi_t\,, \tag{3}$$

where δ_0 represents $-\Phi^{(n)}$, testing $H_0 : \delta_1 = 1$ *vs* $H_1 : \delta_1 \neq 1$. It should be noted that the error term of this equation is a MA$(n-m)$ process; see, e.g., Evans and Lewis (1994), Gerlach and Smets (1997) and Thornton (2005, 2006) for a closer look at this test. It should be also noted that even when $r_t^{(n)}$ is a really *long-term* rate, the stationarity of the spread implies that equation (3) is testing for a short-run implication of the EH. A similar argument applies to the following tests.

Continuing to focus on the long-term behaviour of short-term rates, no other variable besides the spread should provide any help for predicting short-term interest rates changes. Therefore, in equation

$$S_t^{*(n,m)} = \delta_0 + \delta_1\, S_t^{(n,m)} + \delta_2'\, x_t + \eta_t\,, \tag{4}$$

where x_t denotes a vector of variables other than the spread, the EH demands that $\delta_1 = 1$ and $\delta_2 = 0$.

Changing the focus to the short-term behaviour of long-term interest rates, another important characterization of the EH is provided by

$$E_t\left[r_{t+m}^{(n-m)}\right] - r_t^{(n)} = \frac{m}{n-m}\left(S_t^{(n,m)} - \phi_h^{(n,m)}\right), \tag{5}$$

i.e., the expected (short-term) change of the long-term interest rate is defined as a proportion of the difference between the spread and the holding period term premium ($\phi_h^{(n,m)}$). When the long-term interest rate is expected to rise over the next m periods (in the short-term), potential capital losses are predictable. Therefore, the current long-term interest rate has to be higher than the short-term rate.

If the EH is true, the spread is also the optimal predictor of (short-term) changes of long-term interest rates. Based on equation (5), another EH test can be specified. As in equation (3), the simpler version tests whether $\lambda_1 = 1$ in

$$r_{t+m}^{(n-m)} - r_t^{(n)} = \lambda_0 + \lambda_1 \left[\frac{m}{n-m} S_t^{(n,m)} \right] + u_{t+m} \,, \tag{6}$$

and the augmented version is similar to the one of equation (4).

Concerning the predictive ability of the spread, the available empirical evidence tends to agree that:

a) the spread predicts the (long-term) changes in the short-term rates in the direction stated by the EH ($\hat{\delta}_1$ is generally positive, although sometimes statistically different from unity);

b) however, the spread does not predict the (short-term) changes in long-term rates in the direction required by the EH (usually $\hat{\lambda}_1$ is negative and significantly distinct from unity).

This is the "puzzle" well known in the EH literature, also known as the "Campbell-Shiller paradox". Besides providing a recent survey on previous attempts to solve this puzzle, Thornton (2006) demonstrates that it can emerge very often when the EH does not hold.

2.2 In Multiple Equation Models

In the multiple equation model framework the EH has two cointegration implications:

i) in a system of l interest rates with different maturities there should be one (and only one) common stochastic trend, which is responsible for the long-run movement of all interest rates, and

ii) in each of the $l-1$ cointegrating vectors the coefficients should sum zero.

While i) should be clear from the previous subsection, the restrictions of ii) deserve a closer look. Considering $m = 1$ and computing equation (1) for all maturities $\tau_i, i = 2, \ldots, l$:

$$r_t^{(\tau_i)} = \frac{1}{\tau_i} \sum_{i=0}^{\tau_i - 1} E_t \left[r_{t+i}^{(1)} \right] + \Phi^{(\tau_i)} = \frac{1}{\tau_i} \sum_{i=1}^{\tau_i - 1} E_t \left[r_{t+i}^{(1)} \right] + \frac{1}{\tau_i} r_t^{(1)} + \Phi^{(\tau_i)} \,.$$

But since $\frac{1}{\tau_i} \sum_{i=1}^{\tau_i - 1} E_t \left[-r_t^{(1)} \right] = -r_t^{(1)} + \frac{1}{\tau_i} r_t^{(1)}$, the previous equation may be written as

$$r_t^{(\tau_i)} = \frac{1}{\tau_i} \sum_{i=1}^{\tau_i - 1} E_t \left[r_{t+i}^{(1)} - r_t^{(1)} \right] + r_t^{(1)} + \Phi^{(\tau_i)} \,.$$

Taking a linear combination of all interest rates in the system, $\beta_1 r_t^{(1)} + \beta_2 r_t^{(\tau_2)} + \ldots + \beta_l r_t^{(\tau_l)}$ and using the previous equation for $r_t^{(\tau_i)}$, we get, apart from a constant term:

$$(\beta_1 + \beta_2 + \ldots + \beta_l) r_t^{(1)} + \frac{\beta_2}{\tau_2} \sum_{i=1}^{\tau_2 - 1} E_t \left[r_{t+i}^{(1)} - r_t^{(1)} \right] + \ldots + \frac{\beta_l}{\tau_l} \sum_{i=1}^{\tau_l - 1} E_t \left[r_{t+i}^{(1)} - r_t^{(1)} \right] . \tag{7}$$

Now, if interest rates correspond to I(1) processes the spreads will be I(0). Hence, the process of equation (7) will be I(0) iff $\beta_1 + \beta_2 + \ldots \beta_l = 0$.

Both implications can be tested in the context of a VAR model using the popular "Johansen's approach" (Johansen 1995; see also, e.g., Juselius 2006). Clearly, this is a case where the inclusion of a deterministic trend appears highly unreasonable. However, a constant term is required. But then, should a restricted or an unrestricted intercept be considered? While allowing for an unrestricted intercept appears implausible, there is a statistical justification for doing it: *"in vector error-correction models the cointegration rank test based on the unconstrained estimator has somewhat better local power than the test based on the constrained estimator"* (Lanne 2000).

For the cointegration rank analysis we have used trace test statistics. The zero-sum restrictions are tested employing likelihood ratio statistics. Besides these tests, Johansen's methodology also provides a test for the predictive ability of the spread concerning short-term interest rate changes. In order to do this, one must focus on the factor loadings (usually denoted with α_{ij}), which measure the influence of the error correction term in each equation. Under the EH, these coefficients should be statistically significant in all equations except in the one for the longer-term interest rate. In other words, the longer-term rate should be weakly exogenous for the cointegration vectors (see, e.g., Engsted and Tangaard 1994a, b).

3. The Data

The most natural representation of the term structure of interest rates is with spot rates. But as zero coupon bonds are typically issued for maturities less than a year (short end of the maturity spectrum), spot rates have to be estimated from coupon bonds data for longer maturities. For long periods of time, this work has already been done for some countries but not for Portugal. Since this estimation is beyond our present purposes, a preliminary step of identifying alternative datasets was taken. This allowed us to get data for a 10-year government bond yield. Although we have used also this dataset at an initial stage, the rather limited scope of the results lead us to omit their presentation in the main body of the paper. In the Appendix we provide some results and some further comments.

For the short end of the term structure, Treasury bills data are the most common alternative. However, for the Portuguese case the number of missing observations is extremely high. At the end, interbank money market (IMM) rates were selected for several reasons. First, they represent the alternative providing the largest number of observations. Second, IMMs tend to be highly competitive, well integrated with other money markets, and internationally comparable. Finally, contrarily to the bond market, the IMM is much less influenced by large institutions aiming portfolio immunisation.

Monthly data for IMM rates for 1, 3 and 6 months—"value date of same day"—are available at the website of Banco de Portugal (section B.10). Our dataset covers the period from January 1989 to April 2004, i.e., $T = 184$. For the missing observations (2, 17 and 40 for $r_t^{(1)}$, $r_t^{(3)}$ and $r_t^{(6)}$, respectively) some alternatives were considered. Firstly, several univariate and multivariate models were applied to the first differences of the interest rates. However, we have not found the results satisfactory, particularly at the end of the sample. Moreover, when using multivariate models it was not possible to estimate all missing observations. Hence, using observations from the other two segments of the IMM appeared as an attractive and simple alternative[1]. Data for the three segments of the IMM are very highly correlated and it is common to observe that when there are no transactions in one of them the remaining two present some recorded operations. However, this estimation procedure did not allow us obtaining all the missing observations. Therefore we decided to adopt a two step procedure:

i) in the first stage, whenever possible, missing data were estimated with the monthly variation for rates with the same maturity but "value date deferred 1 or 2 days";

ii) for the remaining (20 for $r_t^{(6)}$ only) missing observations, several alternative models were considered and a simple multivariate model in first differences, relating $\Delta r_t^{(6)}$ with $\Delta r_t^{(1)}$ and $\Delta r_t^{(3)}$ and minimizing MSFE was chosen.

The first step is indeed innocuous because a graphical analysis shows that the behaviour of the series for the three segments is almost coincident. This behaviour manifests also in very high correlation coefficients: the lowest correlation is 0.97. On the other hand, since the second step may be viewed with some suspicion, we have made an additional robustness check, "dummying out" the estimated observations in most of the key estimations, i.e., considering them as potential additive outliers. The outcomes of this check are rather reassuring. Actually, in most of the cases the changes are barely perceptible.

Despite covering a relatively short span of time—less than 16 years—, our dataset refers to a period where the Portuguese economy made a long journey towards the full development of its financial markets. Having joined the European Union in 1986, Portugal began liberalizing its monetary and financial markets in the beginning of the nineties. Monetary policy was guided by the intermediate objective of exchange rate stability as a means to achieve price stability. Until the mid of 1994 Banco de Portugal had to intervene several times to protect the national currency (the "escudo") from speculative pressures. In particular, in the first six years of the sample period, it intervened in the foreign exchange markets while at the same time acting on the money market. Achieving exchange rate stability reflected

1 These two other segments are reported as "value date deferred one or two days", that is, the transactions are contracted in one day but the transference of funds occurs only one or two days later.

in a large volatility of interest rates until 1994. In section 5 below we take this into consideration.

Finally, in general terms the Portuguese IMM is closely integrated with other money markets and it has also a relatively high degree of competition. For instance, using the Herfindahl index as an indicator of competition, by the end of the sample period (in 2004) the Portuguese banking sector could be considered more competitive than those of Belgium, Denmark, Finland and the Netherlands (see European Central Bank, 2010, *Structure Indicators for the EU Banking Sector*). It is, however, much less competitive than the UK banking sector (whose IMM is analysed in Hurn et al. 1995).

4. Empirical Results

4.1 In Single Equation Models

Due to space constraints, some of the results for the single equation approach are only briefly presented. However, all the results are available from the authors. First, preliminary unit root testing, using ADF (augmented Dickey-Fuller), PP (Phillips-Perron) and WS (weighted symmetric, see Pantula et al. 1994) tests with several lag truncation parameters (k), provide overwhelming confirmation evidence for the I(1) hypothesis of interest rates. Tables A.1 and A.2 in the Appendix contain all these results, including those for the 10-year government bond yield.

Second, the same unit root tests, which may now be viewed as restricted cointegration tests, strongly support the stationarity of the spreads, i.e., cointegration with unit cointegration parameters (again, refer to tables A.1 and A.2 in the Appendix). Augmented Engle-Granger tests (see table 1) provide somewhat weaker evidence for cointegration but this appears to result only from the usual poor power behaviour of these tests. These were performed with fixed ($k = 6$ and 12) and estimated lag truncation parameters, using the general-to-specific t-sig procedure, denoted with GS t-sig, and the AIC + 2 rule (denoted with AIC + 2), as recommended by Pantula et al. (1994), using $k_{MAX} = 18$.

On the other hand, assuming weak exogeneity (see below, table 7), table 2 reports the orders, (r, s), of the estimated bivariate ADL (autoregressive distributed lag) models, chosen following the GS t-sig strategy and starting with $r_{MAX} = s_{MAX} = 12$, together with the *t-ecm* test statistics for cointegration. As the small sample 5% critical value is -3.232 (see Ericsson and MacKinnon 2002), these provide very strong evidence for cointegration when the dependent variables are the shorter-term rates. As the homogeneity restriction was not imposed, this favourable evidence must be viewed with some caution. However, DOLS estimation and testing (see table 8 below) provide clear evidence for unit cointegration parameters.

Table 1

***P*-values for Augmented Engle-Granger Tests for Cointegration**

rates	dependent variable: $r_t^{(n)}$				dependent variable: $r_t^{(m)}$			
	$k=6$	$k=12$	GS t-sig	AIC+2	$k=6$	$k=12$	GS t-sig	AIC+2
$r_t^{(3)}, r_t^{(1)}$	0.192	0.379	0.033	0.047	0.172	0.360	0.040	0.060
$r_t^{(6)}, r_t^{(3)}$	0.223	0.177	0.083	0.128	0.215	0.214	0.100	0.168
$r_t^{(6)}, r_t^{(1)}$	0.202	0.277	0.035	0.178	0.181	0.298	0.119	0.119

Table 2

***T-ecm* Test Statistics for Cointegration**

rates	dependent variable: $r_t^{(n)}$		dependent variable: $r_t^{(m)}$	
	ADL	*t-ecm*	ADL	*t-ecm*
$r_t^{(3)}, r_t^{(1)}$	ADL(10,3)	–2.086	ADL(8,9)	–5.431
$r_t^{(6)}, r_t^{(3)}$	ADL(7,5)	–4.757	ADL(6,7)	–11.15
$r_t^{(6)}, r_t^{(1)}$	ADL(10,4)	–2.509	ADL(4,7)	–5.313

Hence, in general terms, the analysis of the long-run properties of the data is strongly favourable to the EH. A much different picture is observed when more demanding implications are examined. Table 3 contains the results concerning equation (3), evaluating the predictive ability of the spread for short rate changes[2]. Although the sign of the estimates agrees with the EH, i.e., the predictions are in the correct direction, the restrictions it implies are very clearly and strongly rejected. Despite this evidence, the spread contains useful information about the future (long-run) behaviour of short-term interest rates, that is, $\widehat{\delta}_1$ is significant in all equations.

Then, as expected, the EH is still strongly rejected when $\Delta r_t^{(n)}$, representing the short-run dynamics of the longer-term interest rate, is added as an additional regressor to equation (3) (cf. equation 4). However, the spread retains its statistical significance in all the regressions.

Turning to the predictive ability of the spread in respect to longer rate changes, we could not find a single trace of evidence for the validity of the EH. Table 4 contains the results for equation (6): all the estimates are in the incorrect predictive direction and all the p-values for the restrictions implied by the EH are equal to zero. Moreover, the spread does not seem to contain any relevant information about

2 Previously, Wu-Hausman exogeneity tests were performed, providing no evidence for the inconsistency of the OLS estimator. A similar preliminary analysis was performed also in relation to equation (6), providing the same type of results.

the future (short-run) behaviour of longer-term interest rates. Obviously, when $\Delta r_t^{(m)}$, which represents the short-run dynamics of the short-term interest rate, is included as an additional regressor, the evidence against the EH is confirmed.

Table 3

The Spread As a Predictor of Short Rate Changes (equation (3))

spread	$\widehat{\delta}_0$	$\widehat{\delta}_1$	EH *p*-val.
$S_t^{(3,1)}$	–0.077	0.238	0.000
$S_t^{(6,3)}$	–0.126	0.423	0.000
$S_t^{(6,1)}$	–0.235	0.491	0.000

Note: for the calculation of the Wald test statistics we have used a Newey-West correction with a Bartlett kernel and a bandwidth of $n - m$, but similar results arise when a (fixed) bandwidth equal to 12 is employed.

Table 4

The Spread As a Predictor of Long Rate Changes (equation (6))

spread	$\widehat{\lambda}_0$	$\widehat{\lambda}_1$	EH *p*-val.
$S_t^{(3,1)}$	–0.051	–0.220	0.000
$S_t^{(6,3)}$	–0.251	–0.153	0.000
$S_t^{(6,1)}$	–0.060	–0.087	0.000

Notes: a) we have also used a Newey-West correction with a bandwidth of $m - 1$ but similar results were obtained with a fixed bandwidth of 12; b) when $r_{t+m}^{(n-m)}$ is not available we have followed Hardouvelis (1994), using $r_{t+m}^{(n)}$ as a proxy.

To sum up, in single equation models the empirical evidence is mixed: on one hand, the long-run properties of the data are clearly supportive of the hypothesis; on the other hand, the "puzzle" well known in the literature is also observed for the Portuguese case and our data clearly fail to pass the tests on the predictive ability of the spread. Bearing in mind that the latter conditions are the ones which better characterize the EH and that the former are insufficient to discriminate against other hypotheses of the term structure, we may conclude that it appears to be valid only in some weak, "asymptotic" or long-run form[3].

[3] We borrow this term from the rational expectations hypothesis literature; see, e.g., Stein (1981) and Patterson (1987).

4.2 In Multiple Equation Models

Concerning the multiple equation approach, Johansen's maximum likelihood (ML) method was implemented using PcGIVE 10.1 (Doornik and Hendry 2001) and JMulTi 4.22 (Lütkepohl and Krätzig 2004). Results for systems with 2 and 3 IMM interest rates are presented below and, as previously mentioned, these were obtained including an unrestricted constant. However, all the procedures were also performed considering a restricted intercept, producing evidence which broadly agrees with the one which is presented.

In the modelling exercise we have faced two main problems: strong evidence for non-normality and for serial correlation of the disturbance vector. While non-Gaussianity is of no great concern (see, e.g., Gonzalo 1994, and Lütkepohl 2004), the latter problem may impart somewhat fragile estimates and inferences. The best way to cope with it is to enlarge the information set (at least considering the inflation rate). However, in the current testing framework this is not an admissible option. Instead, we employed a robustifying strategy, considering several dynamic specifications.

Basically, we obtained results for two rather different types of dynamic specifications, i.e., for fixed and for data dependent lag lengths (p). For the former, we used $p = 6$ and 12 for all systems and $p = 18$ only for bivariate systems. For the latter, besides resorting to the usual AIC (Akaike information) and SC (Schwarz) criteria, we have also employed a sequential general-to-specific (GS) strategy of eliminating insignificant lags based on likelihood ratio (LR) test statistics. When using the information criteria, we set $p_{max} = 18$ for bivariate systems and $p_{max} = 12$ for the trivariate case. For the GS-LR strategy, we used $p_{max} = 12$ and 6, respectively, and besides individual lag testing we have also used a joint confirmation test, testing all the restrictions imposed on the initial model.

Although maximum eigenvalue statistics were also computed for cointegration testing, we report only the evidence based on trace test statistics, which are more robust to non-Gaussianity. Besides the asymptotic p-values (denoted with λ_{trace}), tables 5A and 5B report also their finite sample corrected versions (λ^*_{trace}).

Considering bivariate systems, previous evidence for cointegration is generally confirmed but it appears weaker for the two longer-term rates. Strong evidence for cointegration is found in the trivariate system but, more importantly, there is only very weak support that the cointegration rank is equal to two. Actually, this condition seems to hold only when the SC criterion for lag selection is used. However, as is usually the case with SC, the chosen specification appears to be under-parameterized. As is well known, this tends to produce spurious finding for cointegration and for the number of cointegration vectors, and hence we give less weight to this evidence.

Taking these results into consideration, zero-sum restrictions regarding cointegration vectors were tested only in bivariate systems (see table 6). Now the evidence clearly tends to support the EH, confirming the one obtained with DOLS. Cointegrating vector estimates vary between $(1, -0.95)'$ and $(1, -0.99)'$.

Table 5A

P-values of Trace Tests for Cointegration: Fixed Lag Lengths

rates	H_0	$p = 6$		$p = 12$		$p = 18$	
		λ_{trace}	λ^*_{trace}	λ_{trace}	λ^*_{trace}	λ_{trace}	λ^*_{trace}
$r_t^{(1)}, r_t^{(3)}$	$r = 0$	0.001	0.003	0.004	0.014	0.000	0.004
	$r = 1$	0.542	0.556	0.165	0.198	0.122	0.171
$r_t^{(1)}, r_t^{(6)}$	$r = 0$	0.024	0.037	0.003	0.012	0.001	0.010
	$r = 1$	0.545	0.559	0.198	0.233	0.172	0.227
$r_t^{(3)}, r_t^{(6)}$	$r = 0$	0.120	0.158	0.019	0.050	0.055	0.165
	$r = 1$	0.598	0.611	0.322	0.358	0.157	0.210
$r_t^{(1)}, r_t^{(3)}, r_t^{(6)}$	$r = 0$	0.007	0.021	0.012	0.096	—	—
	$r = 1$	0.193	0.270	0.046	0.141	—	—
	$r = 2$	0.530	0.552	0.234	0.290	—	—

Table 5B

P-values of Trace Tests for Cointegration: Estimated Lag Lengths

rates	H_0	p_{AIC}			$\widehat{p}_{SC}$			$\widehat{p}_{LR}$		
		$\widehat{p}$	λ_{trace}	λ^*_{trace}	$\widehat{p}$	λ_{trace}	λ^*_{trace}	$\widehat{p}$	λ_{trace}	λ^*_{trace}
$r_t^{(1)}, r_t^{(3)}$	$r = 0$	14	0.000	0.003	4	0.003	0.005	10	0.000	0.001
	$r = 1$		0.125	0.161		0.613	0.621		0.247	0.277
$r_t^{(1)}, r_t^{(6)}$	$r = 0$	18	0.001	0.010	1	0.011	0.013	11	0.005	0.016
	$r = 1$		0.172	0.227		0.664	0.667		0.214	0.245
$r_t^{(3)}, r_t^{(6)}$	$r = 0$	18	0.055	0.165	1	0.000	0.000	11	0.023	0.052
	$r = 1$		0.157	0.210		0.693	0.696		0.295	0.328
$r_t^{(1)}, r_t^{(3)}, r_t^{(6)}$	$r = 0$	11	0.004	0.040	1	0.000	0.000	6	0.007	0.021
	$r = 1$		0.084	0.197		0.000	0.000		0.193	0.270
	$r = 2$		0.186	0.234		0.785	0.787		0.530	0.552

Proceeding on the path of refining the restrictions required by the EH, table 7 contains the factor loading estimates (i.e., the $\widehat{\alpha}_{ij}$) and the p-values for weak exogeneity tests. The empirical evidence supports theory: at the usual 5 % significance level and with one exception only (in a case where SC is used), longer-term interest rates appear as weakly exogenous for the cointegration vectors. Moreover, confirming the evidence provided by the single equation approach, in every case the estimates present the required sign, i.e., the spread predicts short rate changes in the expected direction.

Table 6

***P*-values for Cointegrating Vector Restriction Tests**

rates	$p=6$	$p=12$	$p=18$	$\widehat{p}_{AIC}$	$\widehat{p}_{SC}$	$\widehat{p}_{LR}$
$r_t^{(1)}, r_t^{(3)}$	0.126	0.145	0.001	0.047	0.160	0.076
$r_t^{(1)}, r_t^{(6)}$	0.150	0.213	0.025	0.025	0.074	0.182
$r_t^{(3)}, r_t^{(6)}$	0.194	0.458	0.265	0.265	0.015	0.263

Table 7

Factor Loading Estimates and *P*-values for Weak Exogeneity Tests

vect.	$p=6$		$p=12$		$p=18$		$\widehat{p}_{AIC}$		$\widehat{p}_{SC}$		$\widehat{p}_{LR}$	
	$\widehat{\alpha}$	p-val.	$\widehat{\alpha}$	p-val.	$\widehat{\alpha}$	p-val.	$\widehat{\alpha}$	p-val.	$\widehat{\alpha}$	p-val.	$\widehat{\alpha}$	p-val.
$r_t^{(1)}$	–0.51	0.000	–0.60	0.000	–0.90	0.000	–0.75	0.000	–0.44	0.000	–0.57	0.000
$r_t^{(3)}$	–0.19	0.058	–0.17	0.218	–0.22	0.223	–0.22	0.114	–0.17	0.044	–0.10	0.419
$r_t^{(1)}$	–0.33	0.000	–0.44	0.000	–0.62	0.000	–0.62	0.000	–0.27	0.000	–0.43	0.000
$r_t^{(6)}$	–0.07	0.204	–0.04	0.576	–0.07	0.447	–0.07	0.447	0.02	0.724	–0.06	0.329
$r_t^{(3)}$	–0.61	0.001	–0.54	0.010	–0.65	0.026	–0.65	0.026	–0.62	0.000	–0.61	0.003
$r_t^{(6)}$	–0.27	0.068	–0.04	0.825	–0.09	0.698	–0.09	0.698	0.13	0.104	–0.14	0.423

5. Sample-split Analysis

Although our dataset covers a relatively short span of time, with no single sharp and abrupt change in monetary and financial conditions, it is possible to distinguish between two sub-periods, according to the degree of stability and deregulation in those markets. The first sub-period, roughly corresponding to the first third of the sample, is characterized by some instability, high interest rates and a rather volatile behaviour of the spreads. Much of this instability is explained by some external shocks, related to events in European foreign exchange markets occurring at the initial stages of the European Monetary System (EMS)[4]. After some deregulation in the monetary and financial markets and in the context of a smoother

4 Camarero and Tamarit (2002) find some evidence of a regime shift in the cointegration relation between long and short interest rates for the Spanish case occurring in 1994. They attribute this finding to reasons similar to ours.

EMS, in the middle of 1994[5] interest rates began declining and a much more stable period initiated, both the spreads and the variation in interest rates exhibiting much less volatility (see figure 1). Consequently, based on the historical description of the facts and on a simple graphical analysis, we decided to split the sample in two: the first sub-period ends in June 1994 and the second one, containing almost ten years of data, begins in July 1994 (which represents the hypothesized break date).

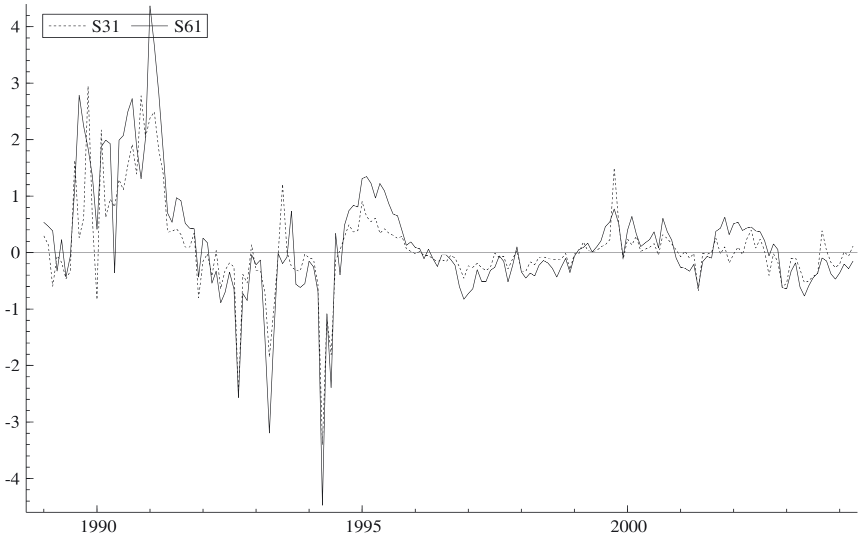

Figure 1: The Spreads $S_t^{(3,1)}$ (S31) and $S_t^{(6,1)}$ (S61)

A preliminary analysis provided strong support for our partition of the sample. In a first step, we have analysed the stability of the VECM models derived from the VAR models of table 5B, imposing $r = 1$ but no further restrictions. Using the assumed break date, both break-point (CH_{BP}) and sample-split (CH_{SS}) Chow statistics tend to reject the stability hypothesis, with bootstrap p-values of 0.0000. The only exceptions concern the CH_{SS} statistics for the models chosen using SC and for two of the models selected using LR tests (those for $(r_t^{(1)}, r_t^{(3)})$ and $(r_t^{(1)}, r_t^{(6)})$).

In a second step, we have searched for the data point where the Chow statistics are maximized. With the exception of the dates estimated using CH_{SS} for the models chosen with SC, which appear a bit earlier, all the remaining estimated break dates are located around 1994:7, ranging between 1992:11 and 1995:5.

[5] *"In the first half of 1994, . . ., faced with the emergence of downward pressure on the escudo, the Banco de Portugal intervened in the foreign exchange market, while at the same time acting in the money market to effect a significant hike in its intervention rates . . . The return to normal exchange conditions, . . ., allowed the Banco de Portugal to reintroduce its market intervention rates at the beginning of July . . .",* Banco de Portugal (1995, pp. 39 and 44).

Examining whether the cointegration relationships might have suffered some regime shift provided further partial evidence for our hypothesis about the sample-split point. Towards this purpose we have used the $\sup \tau$ (fluctuation) test of Hansen and Johansen (1999), testing the constancy of the largest eigenvalue and concentrating out the short-run parameters of the VECMs. In figures 2 and 3, where the dashed horizontal line represents the 5% critical value (1.36), we present two typical examples of this analysis: although the null hypothesis of constancy of the largest eigenvalue is not rejected at the 5% level, the time-path of the statistics exhibits some instability at the beginning of the sample and a more stable behaviour is observed only after the middle of 1994. The recursive estimates of the largest eigenvalues of the VAR models also show a rather similar behaviour: a common volatile pattern in the beginning of the sample which vanishes around 1994:7.

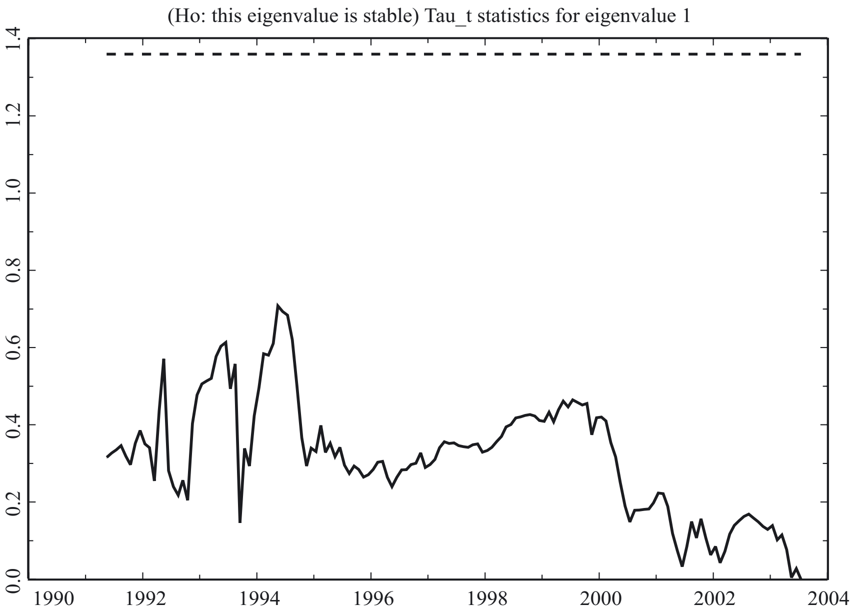

Figure 2: Fluctuation Statistic for the Largest Eigenvalue of the VAR Model for $(r_t^{(3)}, r_t^{(6)})$. With $\hat{p} = 11$.

On the other hand, when the short-run dynamics is not concentrated out, i.e., when all the parameters are estimated recursively, a tendency for an increase in the evidence for instability is observed and in some cases the null of stability is rejected[6]. Hence, in general terms and contrasting with the finding of Camarero and Tamarit (2002) for the Spanish case, we find much less evidence for a regime shift

6 Actually, this occurs only for the longer models chosen with AIC.

in the cointegration relations. Instead, it appears that the instability detected by the Chow statistics may be attributed mostly to the short-run dynamics. This lends support to the hypothesis recently stressed by Tillmann (2007), who argues that short-run dynamics are likely to shift across regimes, contrasting with the robustness of the long-run relation to regime shifts.

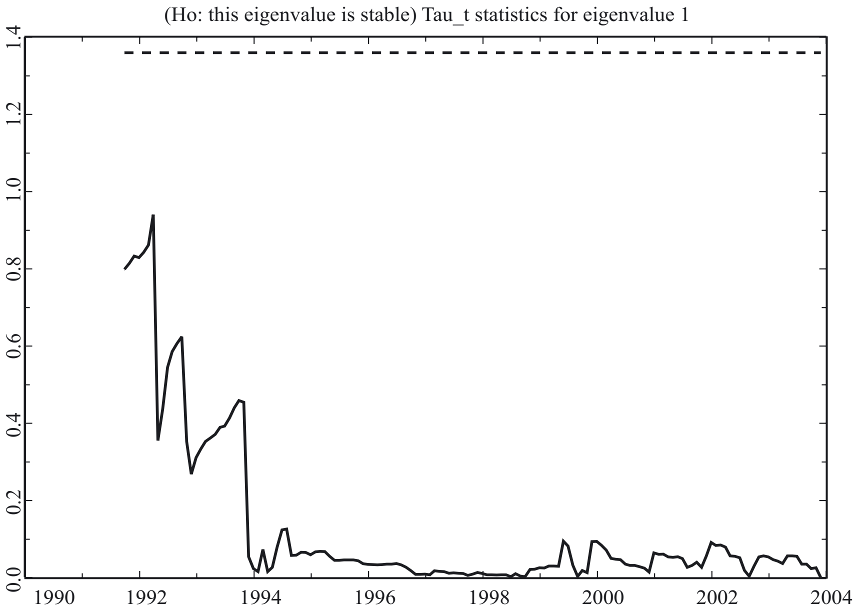

Figure 3: Fluctuation Statistic for the Largest Eigenvalue of the VAR Model for $(r_t^{(1)}, r_t^{(3)}, r_t^{(6)})$. With $\hat{p} = 6$.

Subsequently we have proceeded considering July 1994 as the sample-split point and we have returned to the single equation approach analysis. In table 8 we present the results for DOLS estimation and testing that the cointegration parameter is unity, i.e., $H_0 : \beta_1 = 1$ *vs.* $H_1 : \beta_1 \neq 1$ in $r_t^{(n)} = \beta_0 + \beta_1 r_t^{(m)} + u_t$. Although these results should be viewed with great caution since the first subsample is very short, the results for the whole sample period appear to be masking somewhat distinct situations: only after 1994:6 did the two longer-term rates adjust more closely to the behaviour of the shorter-term rate.

In fact, for the pairs $(r_t^{(3)}, r_t^{(1)})$ and $(r_t^{(6)}, r_t^{(1)})$ the hypothesis is (strongly) rejected in the first period but not in the second, and a comparison of the estimates for β_1 is very helpful to understand this divergence. Also, the "crash" in the estimates for β_0 for these two pairs reflects the reduction in the level and in the volatility of the spread, and clearly confirms the visual impression obtained from figure 1.

Table 8

DOLS Estimates and Tests

rates	1989:1 – 2004:4			1989:1 – 1994:6			1994:7 – 2004:4		
	$\hat{\beta}_0$	$\hat{\beta}_1$	EH *p*-val.	$\hat{\beta}_0$	$\hat{\beta}_1$	EH *p*-val.	$\hat{\beta}_0$	$\hat{\beta}_1$	EH *p*-val.
$r_t^{(3)}, r_t^{(1)}$	0.045	1.032	0.129	3.968	0.785	0.000	–0.130	1.041	0.190
$r_t^{(6)}, r_t^{(3)}$	0.080	1.009	0.551	0.111	1.010	0.755	–0.079	1.029	0.744
$r_t^{(6)}, r_t^{(1)}$	0.070	1.046	0.236	5.475	0.706	0.001	–0.073	1.043	0.733

In tables 9 and 10 we revisit the predictive implications of the EH, now using a sample-split perspective. The testing results confirm and reinforce the previous evidence for the whole sample. In what concerns forecasting short rate changes (table 9), the forecasts for both sub-periods are in the correct direction but the EH is still firmly rejected. For the second sub-period this is somewhat surprising because a much more quiescent monetary and financial environment emerged in the second half of the nineties.

Table 9

The Spread As Predictor of Short Rate Changes (equation (3))

spread	1989:1 – 1994:6			1994:7 – 2004:4		
	$\hat{\delta}_0$	$\hat{\delta}_1$	EH *p*-val.	$\hat{\delta}_0$	$\hat{\delta}_1$	EH *p*-val.
$S_t^{(3,1)}$	–0.051	0.228	0.000	–0.092	0.289	0.000
$S_t^{(6,3)}$	–0.099	0.480	0.001	–0.133	0.152	0.000
$S_t^{(6,1)}$	–0.254	0.532	0.000	–0.217	0.251	0.000

Note: for the calculation of the Wald test statistics we have used a Newey-West correction with a Bartlett kernel and a bandwidth of $n - m$, but similar results arise when a (fixed) bandwidth equal to 12 is employed.

On the other hand, we still find again strong evidence against the EH when the predictive ability of longer rate changes is considered (table 10). Although we observe that the coefficient of $S_t^{(6,1)}$ now appears correctly signed in the second sub-period, it is far from statistically significant because its *p*-value is equal to 0.786.

Table 10

The Spread As a Predictor of Long Rate Changes (equation (6))

	1989:1 – 1994:6			1994:7 – 2004:4		
	$\hat{\lambda}_0$	$\hat{\lambda}_1$	EH *p*-val.	$\hat{\lambda}_0$	$\hat{\lambda}_1$	EH *p*-val.
$S_t^{(3,1)}$	0.020	–0.245	0.000	–0.090	–0.178	0.000
$S_t^{(6,3)}$	–0.198	–0.039	0.000	–0.265	–0.695	0.001
$S_t^{(6,1)}$	0.098	–0.160	0.000	–0.098	0.110	0.027

Notes: a) we have also used a Newey-West correction with a bandwidth of $m-1$ but similar results were obtained with a fixed bandwidth of 12; b) when $r_{t+m}^{(n-m)}$ is not available we have followed Hardouvelis (1994), using $r_{t+m}^{(n)}$ as a proxy.

6. Concluding Remarks

As far as we know, this is the first time that Portuguese IMM rates are used to test the EH. We have assessed both short- and long-run implications using single and multiple equation models. In general terms, mixed but very weak supporting evidence is provided by these data. Notwithstanding the mostly unfavourable evidence for the requirement that the cointegration rank equals two in trivariate systems, strong support is found only when general, long-run implications, are under scrutiny. Moreover, the long-run relations appear to be more stable than for the Spanish case, which has been sharing a common recent experience in terms of monetary and financial conditions. On the other hand, when more demanding short-run conditions are tested, the supporting evidence either becomes much weaker or vanishes completely.

In particular, the EH "puzzle" is also observed for the Portuguese case. Further still, all the test results concerning the predictive ability of the spread are totally at odds with the EH. Since our dataset covers only the short end of the maturity spectrum, these findings are consistent with most empirical evidence for other countries. Rather than viewing these results as "paradoxical", we follow Thornton (2006) and interpret them as invalidating the core of the EH. Hence, only a very weak, "asymptotic" and uncharacteristic version of the hypothesis, appears to hold for the Portuguese case.

Our results also suggest that, notwithstanding the fact that the Portuguese economy lags behind most advanced ones in many respects, its IMM does not possess any particular feature which might have produced a strikingly distinctive outcome. When more favourable evidence is reported—as in, for example, Hurn et al. (1995) for the London IMM—some deduction has to be made because some of the most stringent tests that we use here were not employed. On the other hand, when long-run implications are neglected, it is the negative evidence which most clearly emerges. Therefore, it appears that one of the main lessons to be drawn is that the same data are perfectly capable to produce both types of results. They are not novel but usually they appeared in separate studies, according to the main focus of the study. Moreover, as for other countries we may conjecture that the failure of the core of the EH is mainly due to the inadequacy of the rational expectations hypothesis. However, we cannot proceed further in this direction due to unavailability of data on interest rate expectations.

It is true that unanticipated events related to exchange rate pressures might have lead to large deviations between the actual and the perfect foresight spread. However, these events occurred only until mid 1994. Actually, a final remark refers to the robustness of these results to the sample period under scrutiny. The sample-split analysis allowed us to confirm and to strengthen them. All the inference procedures for the single equation setup were replicated for the two subsamples but we could not find any sharply contrasting difference between the results. A slight increase in the evidence favouring the EH is observed in the calmer second sub-period but this concerns only the long-run properties of the data and may be attributed to the poor performance of the methods used on the (very) small sample of the first sub-period. In other words, our evidence concerning the EH appears to be robust to the sample period.

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Appendix

A. Unit Root Tests

Table A.1

Augmented Dickey-Fuller (ADF) and Philips-Perron (PP) Unit Root Tests: Levels, First Differences and Spreads

	ADF						PP					
	GS *t-sig*		$k=6$	$k=12$	AIC+2		GS *t-sig*		$k=6$	$k=12$	AIC+2	
	k	*p*-val.	*p*-val.	*p*-val.	k	*p*-val.	k	*p*-val.	*p*-val.	*p*-val.	k	*p*-val.
$r_t^{(1)}$	16	0.830	0.961	0.929	6	0.961	16	0.908	0.919	0.918	6	0.919
$r_t^{(3)}$	9	0.884	0.963	0.895	12	0.895	9	0.926	0.933	0.924	12	0.924
$r_t^{(6)}$	18	0.449	0.948	0.852	12	0.852	18	0.916	0.940	0.928	12	0.928
$r_t^{(120)}$	18	0.687	0.630	0.725	18	0.687	18	0.876	0.878	0.872	18	0.876
$\Delta r_t^{(1)}$	18	0.085	0.000	0.002	4	0.000	18	0.000	0.000	0.000	4	0.000
$\Delta r_t^{(3)}$	8	0.003	0.000	0.003	11	0.006	8	0.000	0.000	0.000	11	0.000
$\Delta r_t^{(6)}$	17	0.068	0.000	0.006	10	0.007	17	0.000	0.000	0.000	10	0.000
$\Delta r_t^{(120)}$	17	0.003	0.000	0.002	16	0.007	17	0.000	0.000	0.000	16	0.000
$S_t^{(3,1)}$	17	0.001	0.010	0.023	18	0.001	17	0.000	0.000	0.000	18	0.000
$S_t^{(6,3)}$	13	0.003	0.012	0.005	15	0.003	13	0.000	0.000	0,000	15	0,000
$S_t^{(6,1)}$	18	0.002	0.011	0.012	9	0.006	18	0.000	0.000	0.000	9	0.000
$S_t^{(120,1)}$	7	0.048	0.095	0.067	5	0.128	7	0.007	0.010	0.004	5	0.012
$S_t^{(120,3)}$	9	0.070	0.197	0.060	14	0.095	9	0.025	0.039	0.017	14	0.013
$S_t^{(120,6)}$	14	0.126	0.232	0.081	11	0.083	14	0.070	0.113	0.074	11	0.080

Note: ADF denotes the usual τ_c statistic and PP denotes the modified z_c version (z_c^*).

Table A.2

Weighted Symmetric Unit Root Tests (WS): Levels, First Differences and Spreads

	GS *t-sig*		$k=6$	$k=12$	AIC+2	
	k	WS_c	WS_c	WS_c	k	WS_c
$r_t^{(1)}$	16	–1.054	–0.571	-0.778	5	–0.537
$r_t^{(3)}$	9	–1.042	–0.640	–0.969	12	–0.969
$r_t^{(6)}$	18	–1.637	–0.738	–1.055	11	–1.153
$r_t^{(120)}$	18	–0.422	–0.820	–0.735	17	–0.661
$\Delta r_t^{(1)}$	18	–1.713	–5.070*	–3.217*	4	–7.948*
$\Delta r_t^{(3)}$	8	–2.990*	–4.611*	–2.582*	11	–2.792*
$\Delta r_t^{(6)}$	17	–1.538	–3.785*	–2.458	10	–2.513
$\Delta r_t^{(120)}$	17	–3.148*	–3.771*	–3.209*	16	–2.753*
$S_t^{(3,1)}$	17	–3.249*	–2.742*	–2.391	18	–3.021*
$S_t^{(6,3)}$	13	–2.775*	–2.486	–2.218	7	–2.624*
$S_t^{(6,1)}$	18	–2.427	–2.704*	–2.432	9	–2.540
$S_t^{(120,1)}$	7	–2.091	–1.834	–1.951	5	–1.722
$S_t^{(120,3)}$	9	–1.959	–1.499	–2.036	16	–1.867
$S_t^{(120,6)}$	14	–1.699	–1.452	–1.946	11	–1.936

Note: "*" denotes a rejection at the 5 % asymptotic level (the critical value is –2.55).

B. Some Further Results for 10-year Government Bond Yield

The empirical analysis described in section 4 included also a 10-year government bond. Despite the evidence of a unit root for this rate (see tables A.1 and A.2), no strong evidence for cointegration emerged from these data (see tables A.1 and A.2, and B.1 and B.2 below). Therefore, the predictive analysis of the spread cannot be considered applicable. Moreover, since these data refer to yields to maturity, they are not strictly comparable with spot rates. While this is not a severe problem for the analysis of long-run properties, it invalidates the usual interpretation of the predictive analysis. Results for this rate using the techniques of subsection 4.2 are also available from the authors.

Table B.1

***P*-values for Augmented Engle-Granger Tests for Cointegration**

rates	dependent variable: $r_t^{(n)}$				dependent variable: $r_t^{(m)}$			
	$k=6$	$k=12$	GS *t*-sig	AIC+2	$k=6$	$k=12$	GS *t*-sig	AIC+2
$r_t^{(120)}, r_t^{(1)}$	0.173	0.065	0.046	0.045	0.286	0.143	0.101	0.098
$r_t^{(120)}, r_t^{(3)}$	0.094	0.009	0.061	0.082	0.187	0.025	0.025	0.157
$r_t^{(120)}, r_t^{(6)}$	0.043	0.018	0.183	0.019	0.093	0.041	0.263	0.041

Table B.2

***T-ecm* Test Statistics for Cointegration**

rates	dependent variable: $r_t^{(n)}$		dependent variable: $r_t^{(m)}$	
	ADL	*t-ecm*	ADL	*t-ecm*
$r_t^{(120)}, r_t^{(1)}$	ADL(6,11)	–2.814	ADL(4,11)	–1.969
$r_t^{(120)}, r_t^{(3)}$	ADL(9,3)	–1.963	ADL(10,1)	–3.059
$r_t^{(120)}, r_t^{(6)}$	ADL(9,3)	–1.629	ADL(12,10)	–2.223