The expectations hypothesis of the term structure: some empirical evidence for Portugal

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The Expectations Hypothesis of the Term Structure: Some Empirical Evidence for Portugal

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Abstract

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

The 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. In the single equation framework, the implications of the hypothesis in terms of the predictive ability of the spread are also clearly rejected.

Keywords: term structure of interest rates; expectations hypothesis; hypothesis testing; cointegration; Portugal.

JEL classification: E43, C22, C32.

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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 (1994), Jondeau and Ricart (1999), Lanne (2000), Sarno et al. (2005), 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, will be tested with Portuguese data for interbank money market (IMM) rates. To the best of our knowledge, this paper contains the first examination of the EH using Portuguese IMM data.

Some of the alternative ways of testing the hypothesis will be briefly reviewed, both in the framework of single and multiple equation models. We will focus particularly on cointegration analysis and on the predictive ability of the spread.

The results obtained support only a very weak version of the hypothesis and are in line with most of the conclusions in the literature. The empirical evidence supports the cointegration hypothesis of Portuguese rates and 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, in the single equation framework, the strict 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), Campbell and Shiller (1991), Drakos (2002), Evans and Lewis (1994), Hurn et al. (1995), Jondeau and Ricart (1999), Thornton (2006), Tzavalis (2003) and Tzavalis and Wickens (1997).

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. In section 3 we describe the data that we have used and in section 4 we present the main empirical evidence. Section 5 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 will be described.

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

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 ($\Phi^{(n)}$):

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

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 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 $H_0 : \delta_1 = 1 \quad vs \quad H_1 : \delta_1 \neq 1$. It should be noted that the error term of this equation is a MA$(n-m)$ process; see, e. g., Evans and Lewis (1994), Gerlash and Smets (1997) and Thornton (2005, 2006) for a closer look at this test.

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

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

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 ($\phi_h^{(m,n)}$). 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

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

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 τ_i, $i = 2, ..., l$:

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

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

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

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

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

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

Both implications can be tested in the context of a VAR model using the popular Johansen's approach (Johansen, 1995). 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 will use trace test statistics. The zero-sum restrictions will be 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. Moreover, assuming that the vector of interest rates is ordered in ascending order of maturity, the spread will predict in the direction indicated by the EH if the α_{ij} coefficients are negative for $i \neq l$ and $j = 1, ..., l-1$.

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 was already 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 (available from the authors) lead us to omit their presentation.

For the short end of the term structure, Treasury bills data are the most natural 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) an estimation procedure was performed. Several alternatives were analysed and 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 missing observations (20 for $r_t^{(6)}$), several alternative models were considered and a simple model in first differences minimizing MSFE was chosen.

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, 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.

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 unitary cointegration parameters. 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 $k_{MAX} = 18$.

Table 1 about here

On the other hand, assuming weak exogeneity (see below, table 7), table 2 reports the orders, (r, s), of the estimated bivariate ADL 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 we have not imposed the homogeneity restriction, this favourable evidence must be viewed with some caution. However, DOLS estimation and testing (whose results are not presented) provide clear evidence for unitary cointegration parameters.

Table 2 about here

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 [1]. 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.

Table 3 about here

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 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 4 about here

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;

[1] 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.

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 former are insufficient to discriminate against other hypotheses of the term structure, we may conclude that the EH appears to be valid only in some weak, "asymptotic" form [2].

4.2 In multiple equation models

Concerning the multiple equation approach, Johansen's ML method was implemented using PcGIVE 10.1 (Doornik and Hendry, 2001). 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. Obviously, augmenting the information set to cope with it is not an available option in the current context. 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 and SC 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

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

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}).

Table 5A about here

Table 5B about here

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 will be 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 6 about here

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 7 about here

5 Concluding Remarks

As far as we know, this is the first time that Portuguese IMM rates are used to test the EH. In general terms, mixed but very weak 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. When more detailed and demanding 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: although its forecasts of future short-term rates are in the correct direction, the spread between longer and shorter rates provides a forecast in the wrong direction for the behaviour of longer rates. Moreover, 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" version of the hypothesis, appears to hold for the Portuguese case.

A final remark refers to the robustness of these results to the sample period under analysis. Although our dataset covers a relatively short span of time, with no 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. After some deregulation in the monetary and financial markets, in the middle of 1994 interest rates began declining and a much more stable period initiated, both the spreads and the variation in interest rates exhibiting much less volatility.

All the inference procedures for the single equation setup were replicated for the two sub-samples but we could not find any contrasting difference between the results. A slight increase in the evidence favouring the EH is observed in the second sub-period but this concerns only the long-run properties of the data and may be attributed to the poor power performance of the tests 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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Table 1. P-values for augmented Engle-Granger tests for cointegration

	dependent variable: $r_t^{(n)}$				dependent variable: $r_t^{(m)}$			
rates	$k=6$	$k=12$	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

	dependent variable: $r_t^{(n)}$		dependent variable: $r_t^{(m)}$	
rates	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

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	$\widehat{\lambda}_0$	$\widehat{\lambda}_1$	HE 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

Table 5A. ***P*-values of trace tests for cointegration: fixed lag lengths**

		$p=6$		$p=12$		$p=18$	
rates	H_0	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**

		$\widehat{p}_{AIC}$			$\widehat{p}_{SC}$			$\widehat{p}_{LR}$		
rates	H_0	$\widehat{p}$	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

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

	$p=6$		$p=12$		$p=18$		$\hat{p}_{AIC}$		$\hat{p}_{SC}$		$\hat{p}_{LR}$	
vect.	$\widehat{\alpha}$	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