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ORIGINAL PAPER

Structural breaks and the expectations hypothesis of the term structure: evidence from Central European countries

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Abstract The expectations hypothesis of the term structure of interest rates in the Czech Republic, Hungary, Poland and Slovakia, which joined the EU on May 2004, is investigated in this paper. Using VAR and cointegration techniques in the presence of structural breaks, I examine several testable implications of the theory: (i) cointegration of interest rates, (ii) spread stationarity, (iii) validity of the cross-equation restrictions implied by the theory and (iv) no excess volatility of the actual spread relative to the theoretical spread. The results support the expectations hypothesis for the Czech Republic and Hungary and reject it for Poland and Slovakia.

Keywords Expectations hypothesis of the term structure · Structural breaks · Two-break LM unit root test · Cointegration with breaks · Theoretical spread

JEL Classification E43 · F15 · F42

1 Introduction

The term structure of interest rates, which gives the yield to maturity of different securities at a given point in time, has been the focus of monetary economists and policy makers for a long time, for several reasons: (i) the shape of the yield curve provides valuable information about the future movements of the long-term interest rates, and hence the long-term investment prospects of a country; (ii) the spread between the long and current short rate is a better predictor of a country's monetary policy stance than the rate of monetary growth allowed by a central bank; and (iii) empirical studies have suggested that the interest rate spread has good predictive

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power about the cyclical behaviour of an economy (Lahiri and Wang 1996; Estrella and Mishkin 1998; Estrella and Hardouvelis 1991).

The literature on the term structure of interest rates is large and growing. The expectations hypothesis of the term structure (EHTS) has been used extensively in many studies in order to explain the term structure of interest rates and the shape of the yield curve; see Shiller (1990) for an excellent survey of theory and empirical studies. According to the EHTS and for a given term premium, if future short rates are expected to rise, then the yield curve will be upward sloping. Conversely, if the future short rates are expected to fall, the yield curve will be downward slopping. In general, the empirical results are mixed regarding the validity of the EHTS of interest rates. Among others, Campbell and Shiller (1987) examined the economic and statistical significance of the EHTS for the USA and found evidence that provides partial support for the validity of the present value model of the term structure of interest rates. Hall et al. (1992) used yield series of US Treasury Bills and found cointegration among the interest rates, thereby interpreting the evidence as supportive of the EHTS. Hardouvelis (1994) investigated the EHTS for the G7 countries and rejected it only for the USA. Cuthbertson (1996) studied the EHTS for the UK interbank market and found evidence that supports the EHTS at shorter maturities but not at longer maturities. Jondeau and Ricart (1999) tested the EHTS for French, German, UK and US euro rates and found evidence in favour of the EHTS only for the French and UK rates. Cuthbertson and Bredin (2000) investigated the EHTS for Ireland and found evidence in favour of the EHTS. Bekaert and Hodrick (2001) studied the EHTS for Germany, the UK and the USA and found no evidence against the EHTS only for the UK. Lanne (2003) investigated the EHTS for the US Eurodollar deposit rates, allowing for potential regime shifts. He found evidence that supports the EHTS at the short end of the maturity spectrum once a potential regime shift was allowed for. Gravelle and Morley (2005) adopted the Kalman filter technique and strongly rejected the EHTS for Canada. Brüggemann and Lütkepohl (2005) analysed the relation between long and short rates for the euro area and the USA and found evidence in favour of the EHTS. Diebold et al. (2006) developed a yield curve model that incorporates yield factors and macroeconomic variables and they related it with the EHTS. They used US Treasury yields from early 1970s to 2000 and their evidence was in favour of the EHTS for certain periods, but not for the entire sample. Bekaert et al. (2007) examined the EHTS simultaneously with the uncovered interest rate parity, drawing data from Germany, Japan, the UK and the USA. In general, their evidence was against the EHTS but, economically, actual spreads and theoretical spreads do not behave very differently, especially at long horizons. Koukouritakis and Michelis (2008) studied the EHTS for the twelve newest EU countries and found evidence in favour of it for all countries except Malta.

The novelty of this paper lies on (i) the use of most recent data from the mid-1990s to the end of 2007 for studying the term structure of interest rates in four Central European (CE) countries that joined the EU on May 2004, namely the Czech Republic, Hungary, Poland and Slovakia; (ii) the use of recently developed Lagrange Multiplier (LM) unit root tests (Lee and Strazicich 2003), and cointegration tests (Lütkepohl and his associates in several papers noted below) for studying the EHTS in these four CE countries in the presence of structural shifts, which are likely to have been caused during the transition period of these countries from centrally planned economies to full EU members; and (iii) the use of the VAR approach proposed by Campbell and Shiller (1987, 1991) for testing the economic significance of the EHTS in these four CE countries. Briefly, the results provide support of the statistical and economic significance of the EHTS only for the Czech Republic and Hungary.

The paper is organised as follows. Section 2 describes briefly the EHTS of interest rates and discusses the testable implications of the theory. Section 3 outlines the unit root and cointegration tests in the presence of structural breaks, which are used in the subsequent analysis. Section 4 describes the data and analyses the empirical results. Section 5 concludes.

2 Testable implications of the EHTS

According to the EHTS, the relationship between an n-period bond yield $R_{n,t}$ and the average of the current and expected future rates on a set of m-period bond yields $r_{m,t}$, with $m < n$, can be written as

$$(1 + R_{n,t})^n = \varphi^*_{(n,m),t} \prod_{i=0}^{k-1} (1 + E_t r_{m,t+im}), \tag{1}$$

where $k = n/m$ is an integer and $\varphi^*_{(n,m),t}$ is a possible non-zero but stationary n-period term premium. The pure EHTS holds when there is no term premium of any kind, while the weaker version of the EHTS allows for a constant term premium in (1). Log-linearising (1) and subtracting $r_{m,t}$ from both sides I get

$$S_{(n,m),t} = \varphi_{n,m} + E_t \sum_{i=1}^{k-1} (1 - i/k)\Delta^m r_{m,t+im} = \varphi_{n,m} + E_t(S^*_{(m,n),t}), \tag{2}$$

where $S_{(n,m),t} \equiv (R_{n,t} - r_{m,t})$ is the actual yield spread, $\Delta^m r_{m,t+im} \equiv (r_{m,t+im} - r_{m,t})$ is the change in the short term (m-period) interest rates, $S^*_{(n,m),t} \equiv \sum_{i=1}^{k-1} (1 - i/k)\Delta^m r_{m,t+im}$ is the *perfect foresight spread*, which would obtain, under the EHTS, if economic agents had perfect foresight about future movements in interest rates, and $\varphi_{(n,m),t} = \ln(\varphi^*_{(n,m),t})$. It is clear from (2) that the actual spread S is an optimal forecast of the perfect foresight spread. Optimality implies that, given S, no other variable at time t can help predict future changes in short rates. An implication of this result is that S Granger causes changes in short rates. There are several other testable implications of the EHTS: (i) cointegration, (ii) the cointegrating vector linking long and short rates is (1, −1), (iii) cross-equation restrictions implied by the theory, and (iv) the variance ratio between the actual spread and that implied by the theory should be unity.

2.1 Cointegration

Given that $R_{n,t}$ and $r_{m,t}$ are integrated of order one, (2) implies that the two rates should be cointegrated, as its right-hand side is a stationary process. Cointegration between long and short rates is consistent with the idea that market forces continuously adjust to correct any temporary disequilibrium, so that they do not allow for arbitrage opportunities.

2.2 The cointegrating vector is (1, −1)

If $R_{n,t}$ and $r_{m,t}$ are cointegrated, there exist constants β_R and β_r such that the linear combination $\beta_R R_{m,t} + \beta_r r_{m,t}$ is a stationary process. It is clear from the left hand side of (2) that the EHTS implies that the cointegrating vector $(\beta_R, \beta_r)' = (1, -1)'$.

2.3 Cross-equation restrictions

Campbell and Shiller (1987, 1991) proposed a VAR methodology evaluating the economic importance of deviation from the EHTS. They specify a VAR and derive a set of cross-equation restrictions that must hold under the EHTS. Using the VAR, they also compute the *theoretical spread*, an estimate of the perfect foresight spread, and then they compare it to the actual spread. Significant differences are interpreted as evidence against the EHTS.

Assuming that $x_t \equiv (\Delta r_{m,t}, S_{(n,m),t})'$ can be approximated by a stationary p-order VAR, one can write its companion form as a first-order VAR

$$z_t = Az_{t-1} + v_t, \tag{3}$$

where z_t is a $2p \times 1$ vector with elements, first $\Delta r_{m,t}$ and $p - 1$ lags and then $S_{(m,n),t}$ and $p - 1$ lags, A is the companion matrix of the VAR and v is a random error term.

Define the $2p \times 1$ vectors g and h such that $g'z_t = S_{(m,n),t}$ and $h'z_t = \Delta r_{m,t}$. The elements of g and h are all zero, except for the $p + 1$st element of g and the first element of h that are unity. Projecting both sides of (2) onto the information contained in z_t gives

$$S_{(m,n),t} = g'z_t = S'_{(m,n),t} \tag{4}$$

where

$$S'_{(m,n),t} = h'A\left[I - (m/n)(I - A^n)(I - A^m)^{-1}\right](I - A)^{-1}z_t \tag{5}$$

is the *theoretical spread*, computed from the VAR.

Since (3) holds for any general z_t, it must be the case that

$$g' = h'A\left[I - (m/n)(I - A^n)(I - A^m)^{-1}\right](I - A)^{-1}. \tag{6}$$

The set of restrictions in (6) are equivalent to the null hypothesis H_0: $S_{(n,m),t} = S'_{(n,m),t}$. This hypothesis can be tested using the Wald test, which is χ^2-distributed asymptotically under the null, with *2p* degrees of freedom. If H_0 is not rejected, then the EHTS holds. Otherwise, rejection of the H_0 is in favour of excess returns in the bonds market.

2.4 Variance ratio

Consider the variance ratio $\mathrm{VR} = \mathrm{var}(S_{(n,m),t})\big/\mathrm{var}(S'_{(n,m),t})$, together with the correlation between S_t and S_t'. If the EHTS holds, the correlation should be close to one, and the variances of the actual and the theoretical spreads should behave similarly over time. Thus, the *VR* should be close to unity. Campbell and Shiller (1991) note that this volatility test is preferable to formal tests of the *VAR* restrictions, because the latter may lead to rejection of the EHTS even though the deviations are quite small from an economic point of view.

As noted in the first section of the present paper, the empirical results regarding the validity of the EHTS are mixed. One main reason for the non validity of the EHTS can be found in the segmented markets theory of the term structure. According to this theory, the markets for different maturity bonds are completely separate and segmented, which means that the interest rate for each maturity bond is determined by the supply and demand for this bond and there are no effects from expected returns on bonds with different maturity. In other words, bonds with different maturities are not perfect substitutes, mainly due to uncertainty, since different maturities involve different risks. If there is relatively little shifting among bonds with different maturities, long rates may differ from the average of the current and expected future rates and thus, the EHTS is not valid. Another reason that affects the validity of the EHTS and is closely related with the countries that are examined in the present paper, is the degree of economic openness. The EHTS is more likely to be valid in countries with open and well-functioning markets and with developed secondary markets for securities. Also, as Jardet (2008) points out, the empirical rejection of the EHTS may be generated by a peso-problem (i.e. the effects of expected regime changes when this new regime is not observed in the data) or a time-varying term premium.

3 Unit roots and cointegration with structural breaks

During the transition period from the regime of centrally planned economies to the EU accession, the CE countries made important economic reforms that are likely to have caused structural breaks in their term structure of interest rates. These reforms are mainly associated with the implementation of several monetary policy regimes by the CE countries, in order to achieve price stability. For instance, the Czech Republic introduced a managed float of its currency against the German mark in mid-1990s, which was replaced by the euro in 1999, Poland introduced an "inflation targeting" monetary policy regime in 1998 in order to fight inflation, while Slovakia allowed the free floating of its currency in 1998, because it could not defend it against devaluation pressures. There were also international economic events that had a direct impact on the term structure of interest rates of these countries, such as the Russian financial crisis that affected the interest rates of Hungary.[1] Since the presence of structural breaks are known to have significant effects on the properties and interpretation of standard ADF-type unit root tests and Johansen-type cointegration tests, I employ recently developed tests that are valid in the presence of structural shifts in the data.

[1] These events are discussed analytically in Sect. 4.2 below and linked with the structural breaks that are determined endogenously by the unit root tests.

3.1 Unit root tests with structural breaks

I test for unit roots in the data using the two-break LM test developed by Lee and Strazicich (2003). This test has several desirable properties: (a) it determines the structural breaks endogenously from the data, (b) its null distributions is invariant to level shifts in a variable, and (c) it is easy to interpret; by including breaks under both the null and alternative hypotheses, a rejection of the unit root hypothesis implies unambiguously trend stationarity.

Consider this test for the process y_t generated by

$$y_t = \delta' Z_t + e_t, \qquad e_t = \beta e_{t-1} + A(L)\varepsilon_t, \qquad \varepsilon_t \sim iid\, N(0, \sigma^2) \tag{7}$$

where $A(L)$ is a k-order polynomial in the lag operator L and Z_t is a vector of exogenous variables of which components are determined by the type of breaks one wishes to examine in the process y_t. Lee and Strazicich (2003) extend Perron's (1989, 1993) single-break models to include two breaks in the level (Model A) and two breaks in both the level and trend (Model C) of y_t. Then for Model A, $Z_t = [1, t, D_{1t}, D_{2t}]'$ where $D_{jt} = 1$ for $t \geq T_{Bj} + 1$, $j = 1, 2$, and zero otherwise; and for Model C, $Z_t = [1, t, D_{1t}, D_{2t}, DT_{1t}, DT_{2t}]'$, where $DT_{jt} = t - T_{Bj}$ for $t \geq T_{Bj} + 1$, $j = 1, 2$, and zero otherwise. T_{Bj} denotes the break point in time.

Equation (7) denotes that y_t has a unit root if $\beta = 1$, while it is trend stationary if $\beta < 1$. According to the LM principle, a unit root test statistic is obtained from the test regression

$$\Delta y_t = \delta' \Delta Z_t + \phi \tilde{S}_{t-1} + \sum_{i=1}^{k} \theta_i \Delta \tilde{S}_{t-i} + u_t, \tag{8}$$

where $\tilde{S}_t = y_t - \tilde{\psi}_x - Z_t\tilde{\delta}$, $t = 2, \ldots, T$, in which $\tilde{\delta}$ is a vector of coefficients in the regression of Δy_t on ΔZ_t and $\tilde{\psi}_x = y_1 - Z_1\tilde{\delta}$, where y_1 and Z_1 are the first observations of y_t and Z_t, respectively, and u_t is a random error term. The lagged differences of $\tilde{S}_{t-i}$ are included as necessary to correct for serial correlation in u_t. The unit root null hypothesis is described by $\phi = 0$ in (8) and can be tested by the LM test statistic:

$$\tilde{\tau} = t\text{-statistic for the hypothesis } \phi = 0. \tag{9}$$

In order to endogenously determine the location of the two relative breaks $\lambda_j = T_{Bj}/T$, $j = 1, 2$, where T is the sample size, the two-break minimum LM test statistic is determined by a grid search over λ:

$$\mathrm{LM}_\tau = \inf_{\lambda}\{\tilde{\tau}(\lambda)\}. \tag{10}$$

The critical values for (10) are invariant to the break locations (λ_j) for Model A, but depend on the break locations for Model C, and are available in Lee and Strazicich (2003).

3.2 Cointegration tests with structural breaks

As in the case with unit root testing, structural breaks in the data can distort substantially standard inference procedures for cointegration. Thus, it is necessary to account for possible breaks in the data before inference on cointegration can be made.

In the recent literature, there are two main approaches to test for cointegration in the presence of structural breaks. One approach that was developed by Johansen et al. (2000) extends the standard VECM with a number of additional variables in order to account for q possible exogenous breaks in the levels and trends of the deterministic components. The most recent approach developed by Lütkepohl and his associates (henceforth the LST approach; see among others, Lütkepohl and Saikkonen (2000), Saikkonen and Lütkepohl (2000), Trenkler et al. (2008); and references therein) assumes that the structural breaks have occurred only in the deterministic part and do not affect the stochastic part of the process Y_t. Lütkepohl et al. (2003) studied the statistical properties of their tests for the case of level shifts only, and compared them to the Johansen et al. (2000) test. They found that their tests have better size and power properties than the Johansen et al. (2000) test in finite samples. For that reason, the LST approach has been used in the present paper.

LST set up the data generation process (DGP) for Y_t by adding its deterministic part μ_t to its stochastic part X_t, where the latter is an unobservable zero-mean purely stochastic VAR process, and use appropriate dummy variables to account for exogenous shifts in μ_t. Given this set up, LST propose a two-step procedure to test for cointegration. In the first step, they remove the deterministic part using a generalised least squares procedure under the hypothesis of r_0 cointegrating relations (GLS de-trending). In the second step, they test for cointegration in the de-trended series using their proposed LR-type test statistics. Several tests statistics can be derived depending on whether there are shifts only in the level of the process or shifts in both the level and the trend.

To illustrate the LST approach for LR-type tests, consider the case of a single shift in both the level and the trend of Y_t, at time T_B. LST specify the following DGP for Y_t:

$$Y_t = \mu_t + X_t = \mu_0 + \mu_1 t + \delta_0 d_t + \delta_1 b_t + X_t, \quad t = 1, \ldots, T, \tag{11a}$$

where t is a linear time trend, μ_i ($i = 0, 1$) and δ_i ($i = 0, 1$) are unknown ($p \times 1$) parameter vectors, d_t and b_t are dummy variables defined as $d_t = b_t = 0$ for $t < T_B$, and $d_t = 1$ and $b_t = t - T_B + 1$ for $t \geq T_B$. The unobserved stochastic error X_t is assumed to follow a VAR(k) process with VECM representation

$$\Delta X_t = \Pi X_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \varepsilon_t, \quad \varepsilon_t \sim iidN(0, \Omega), \quad t = 1, \ldots, T. \tag{11b}$$

It is also assumed that the components of X_t are at most integrated of order one processes and cointegrated (i.e. $\Pi = \alpha\beta'$) with cointegrating rank r_0.

Given the DGP in (11a), (11b) the first step of the LST approach involves obtaining estimates of the parameter vectors μ_0, μ_1, δ_0 and δ_1 in (11b) using a feasible GLS procedure under the null hypothesis $H_0(r_0)$: $rank(\Pi) = r_0$: versus $H_1(r_0)$: $rank(\Pi) > r_0$ (see Saikkonen and Lütkepohl (2000) for details). Having the estimated parameters, $\hat{\mu}_0$, $\hat{\mu}_1$, $\hat{\delta}_0$ and $\hat{\delta}_1$, one then computes the de-trended series $\hat{X}_t = Y_t - \hat{\mu}_0 - \hat{\mu}_1 t - \hat{\delta}_0 d_t - \hat{\delta}_1 b_t$. In the second step, an LR-type test for the null hypothesis of cointegration is applied to the de-trended series. This involves replacing X_t by $\hat{X}_t$ in the VECM (11b) and computing the LR or trace statistic:

$$\mathrm{LR}_{\mathrm{LST}} = -T \sum_{i=r_0+1}^{p} \ln(1 - \tilde{\lambda}_i) \tag{12}$$

where the eigenvalues $\tilde{\lambda}_i$'s can be obtained by solving a generalised eigenvalue problem, along the lines of Johansen (1988).

Under the null hypothesis of cointegration, Trenkler et al. (2008) derive asymptotic results and p-values for the case of one level shift and one trend break in the Y_t process, and show that, in this case, the asymptotic distribution of the LR statistic in (12) depends on the location of the break point. They also discuss how the results can be extended to the general case of $q > 1$ break points. Also, critical or p-values for a single level shift can be computed by the response surface techniques developed by Trenkler (2008).

4 Data and empirical results

4.1 Data

I collected data for four CE countries that are EU members: the Czech Republic, Hungary, Poland and Slovakia. Due to lack of data availability, I collected data on two interest rates of the term structure for each country: treasury bill yields (short-term) and government bond yields (long-term). Unfortunately, the lack of data availability did not allow me to use securities with the same maturity for each country. For that reason, I had to use treasury bill rates with maturities of 3, 6 or 12 months as short term interest rates. For the long term I used securities of 5 or 10 years, except for Poland where I used yields of a medium term security (i.e. 2 years), because sufficient time series data for securities with longer maturity were not available. The sample consists of monthly data of varying time spans for different countries determined by data availability. Table 1 reports the data details and their sources.

4.2 Unit root results with structural breaks

Table 2 reports the unit root results from the two-break LM test. Each interest rate series was tested for a unit root at the 1, 5 and 10% levels of significance. In order to determine the number of lags, k, in (8), I used a "general to specific" procedure at each combination of break points (λ_1, λ_2). Initially, the lag-length was set at $k = 12$, and the significance of the last lagged term was examined at the 10% level. The procedure was repeated until the last lagged term was found to be significantly different than zero, at which point the procedure stops.[2]

[2] I computed the two-break LM test using the Gauss codes of J. Lee available at the Web site http://www.cba.ua.edu/~jlee/gauss.

Table 1 Description of data

Country	Time span	Variables	Source
Czech Republic	1993:08–2007:12	3-month treasury bill rates	IFS[a] (line 60c)
		5-year government bond yields[b]	IFS (line 61)
Hungary	1997:01–2007:12	3-month treasury bill rates	Central Bank of Hungary
		5-year government bond yields	Central Bank of Hungary
Poland	1994:02–2007:12	12-month treasury bill rates	Polish Ministry of Finance
		2-year government bond yields	Polish Ministry of Finance
Slovakia	1994:12–2007:12	6-month treasury bill rates	Central Bank of Slovakia
		10-year government bond yields[c]	IFS (line 61)

[a] International Financial Statistics CD-ROM of the International Monetary Fund

[b] For the period 1993:8–1999:12 the source is the Central Bank of the Czech Republic, because the IFS data series begins at 2000:1

[c] For the period 1994:12–2000:8 the source is the Central Bank of Slovakia, because the IFS data series begins at 2000:9

Table 2 Two-break minimum LM unit root test results

Country	Interest rate	Model	$\hat{k}$[a]	$\hat{T}_B$[b]	$\hat{\lambda}_1$, $\hat{\lambda}_2$[c]	LM-statistic
Czech Republic	$R_{n,t}$	C	9	1997:02, 1999:07	0.2, 0.4	−4.18
	$r_{m,t}$	C	9	1997:08, 1999:01	0.2, 0.4	−4.85
Hungary	$R_{n,t}$	C	3	1998:08, 2000:11	0.2, 0.4	−5.12
	$r_{m,t}$	C	12	2003:04, 2005:10	0.6, 0.8	−4.85
Poland	$R_{n,t}$	A	9	1996:01, 1999:11	NA	−3.32
	$r_{m,t}$	A	5	1996:01, 1998:01	NA	−2.39
Slovakia	$R_{n,t}$	A	11	1997:10, 1999:01	NA	−3.42
	$r_{m,t}$	A	11	1998:07, 1998:09	NA	−3.32

Model A				Model C		
Critical values			Break points	Critical values		
1%	5%	10%	$\lambda = (\lambda_1, \lambda_2)$	1%	5%	10%
−4.54	−3.84	−3.50	$\lambda = (0.2, 0.4)$	−6.16	−5.59	−5.27
			$\lambda = (0.6, 0.8)$	−6.32	−5.73	−5.32

$R_{n,t}$ and $r_{m,t}$ are the long-term and short-term interest rates respectively. *NA* not affected by the break points. The critical values for Models A and C are from Tables 1 and 2, respectively, of Lee and Strazicich (2003)

[a] $\hat{k}$ is the optimal number of lagged first-differenced terms included in the unit root test to correct for serial correlation

[b] $\hat{T}_B$ denotes the estimated break points

[c] $\hat{\lambda}_1$ and $\hat{\lambda}_2$ are the estimated critical value break points

As shown in the last column of Table 2, the unit root hypothesis with two structural breaks cannot be rejected at any of the three levels of significance for all interest rates.[3] As shown in the third column of Table 2, Model C fits the term structure data best for the Czech Republic and Hungary. Hence, these two countries have experienced two significant shifts both in the deterministic levels and trends of their term structures, over the sample period. Model A with only two significant level shifts fits the data best for Poland and Slovakia. The two breaks of each country were estimated endogenously by the two-break LM test, and are reported in column 5 of Table 2. Not surprisingly, the estimated breaks correspond closely to specific events that have taken place in the four CE countries over the sample period.

For the Czech Republic, both rates have two breaks in 1997 and in 1999. The first break coincides with the decision of the country's monetary authorities to introduce a managed float of the domestic currency against the German mark in that year. The second break is likely related to the replacement of the German mark with the euro as the benchmark currency and the subsequent considerable exchange rate fluctuations. For Hungary, the long rate breaks in 1998 and 2000 are probably related to the effects of the Russian financial crisis in 1998 and the significant appreciation of the domestic currency in 2000, which lowered imported inflationary pressures. The first short rate break in 2003 coincides with the speculative attack at a time when the domestic currency was close to the upper bound of its trading band and the subsequent interest rate cut by the country's central bank. The second short rate break in 2005 coincides with the significant decrease of inflation during that year.

Each of Poland's rates appears to have a first break in early 1996, which coincides with the significant widening of the domestic currency's fluctuation bands. The second long rate break in late 1999 is related to the implementation of a free floating exchange rate for the domestic currency and the removal of international capital flow restrictions. The second short rate break in early 1998 coincides with the introduction of an "inflation targeting" monetary policy regime, which led to a significant decrease of the inflation. Finally, Slovakia's rates have two structural breaks between late 1997 and early 1999, which coincide with a period of high and rising inflation. This led the Slovakian central bank to allow the free floating of the "koruna" in October 1998, because it could not defend it against devaluation pressures.

4.3 Cointegration results with structural breaks

This section examines the cointegration test results in the presence of structural breaks. To compute the LST test for the Czech Republic and Hungary that were found to have two structural breaks in both the level and trend of their interest rates, I extended the model described by (11a), (11b) by adding a second step dummy and a second linear trend dummy. For Poland and Slovakia which were found to have two breaks only in their level, I removed the linear trend dummy and added a second step dummy in the same model. The LR_{LST} test statistics and the corresponding

[3] Each interest rate was also tested for a second unit root. The null hypothesis was rejected in all cases. These results are available upon request.

Table 3 Testing for cointegration in the presence of structural breaks

Country	Breaks included in the VECM	$(p - r_0)$	$LR_{LST}(r_0)$	$\hat{k}$ [a]	H_0 [b]	
					LR	p-value
Czech Republic	Long rate:	2	27.48** (0.007)	1	1.41	0.234
	1997:02, 1999:07	1	5.56 (0.358)			
	Short rate:	2	22.35** (0.034)	3	0.18	0.672
	1997:08, 1999:01	1	3.0283 (0.674)			
Hungary	Long rate:	2	20.68* (0.063)	8	1.77	0.183
	1998:08, 2000:11	1	5.21 (0.361)			
	Short rate:	2	25.03** (0.017)	11	0.54	0.461
	2003:04, 2005:10	1	3.58 (0.647)			
Poland	Long rate:	2	5.51 (0.510)	5	NA	NA
	1996:01, 1999:11	1	0.15 (0.760)			
	Short rate:	2	5.69 (0.480)	5	NA	NA
	1996:01, 1998:01	1	0.34 (0.640)			
Slovakia	Long rate:	2	11.80* (0.061)	11	0.00	0.964
	1997:10, 1999:01	1	1.97 (0.185)			
	Short rate:	2	15.32** (0.015)	1	0.30	0.586
	1998:07, 1998:09	1	1.63 (0.240)			

LR is the likelihood ratio test statistic for H_0. Numbers in parentheses in column 4 are p-values

NA not applicable

** and * Rejection of the null hypothesis at the 0.05 and the 0.10 level of significance, respectively

[a] $\hat{k}$ denotes the estimated lag length in each VECM

[b] H_0 denotes the null hypothesis $(\beta_R, \beta_r)' = (1, -1)'$

response surface p-values were computed using GAUSS routines.[4] In order to check the robustness of the empirical results, I estimated two VECMs for each country: one that includes the long rate breaks and another with the short rate breaks.

Table 3 reports the LR_{LST} test results for each of the four CE countries. The break points included in each VECM are reproduced in the second column of the table. The lag length, k, for each VECM, was selected using the Akaike information criterion (AIC). The empirical results show evidence of cointegration between the long rate and the short rate for each of the Czech Republic, Hungary and Slovakia. In contrast, the hypothesis of cointegration is strongly rejected for Poland. These results hold for both VECMs of each country.

The LST cointegration test in the presence of structural breaks, assumes that the "long-run" cointegration parameters remain constant over time. Otherwise, the test results and inference would be invalid. For this reason, each VECM was first tested for parameter constancy, using the methodology developed by Hansen and Johansen (1999). They suggest a graphical procedure based on recursively estimated

[4] I am grateful to Carsten Trenkler for kindly providing me with the Gauss codes to perform these estimations.

eigenvalues. By inspecting the time paths of the eigenvalues, one can evaluate the constancy of the long-run parameters of the model. Figure 1 shows the time paths of the eigenvalues estimated for different VECMs. The dotted line in each plot corresponds to 1.36, which is the 5% critical value for the null hypothesis of long-run parameter constancy. As shown in these plots, this hypothesis cannot be rejected in all cases that there is evidence of cointegration, as the time paths of the eigenvalues are always below the dotted line. Thus, the LST procedure has been applied correctly.

For all the cases that there is evidence of cointegration, the null hypothesis H_0: $\beta_R + \beta_r = 0$ where $(\beta_R, \beta_r)' = (1, -1)'$ has been also tested. This hypothesis means that the unit vector belongs in the cointegration space as predicted by the EHTS. Under the null hypothesis, the likelihood ratio test is distributed as χ^2 with 1 degree of freedom asymptotically. As shown in the last column of Table 3, this hypothesis cannot be rejected for both VECMs of each of the Czech Republic, Hungary and Slovakia. Consequently, for these three countries in which the spreads are stationary, the empirical results are also consistent with prediction (ii) of the EHTS.

4.4 The theoretical spread and the VAR results

This section analyses the results from the VAR models for $\Delta r_{m,t}$ and $S_{(n,m),t}$. Since the VAR models require spread stationarity, the interest rates of Poland were excluded from the VAR analysis. For checking the robustness of the results again, I estimated two VARs for each country: one that includes the long rate breaks and another that includes the short rate breaks. The appropriate lag length, k, for each VAR was chosen using the likelihood ratio test (Johansen, 1995, p. 21). Also for each VAR, I performed a multivariate LM test for serial correlation.

Table 4 reports the VAR results. Column 4 presents the Wald test statistics and the respective p-values for Granger non-causality. Under the null hypothesis, these tests are χ^2-distributed asymptotically, with degrees of freedom equal to the number of lags in the VAR. As predicted by the EHTS, the actual spread Granger causes changes in the short rates of the Czech Republic, Hungary and Slovakia, since the null hypothesis that the spread does not Granger cause short rate changes is rejected in all cases. Also, there is Granger causality from $\Delta r_{m,t}$ to the spread for the Czech Republic and Slovakia, indicating bi-directional causality in the VAR regressions (Table 4, column 5). Further, as shown in the sixth column of Table 4, the null hypothesis of no serial correlation in the VAR error term cannot be rejected in all cases, even at the 10% level of significance, which strengthens the validity of the empirical findings.

Table 5 (column 3) reports the Wald test results for testing the VAR restrictions in Eq. (6).[5] For both VARs of the Czech Republic and Hungary, these restrictions cannot be rejected at the 5% level of significance. In contrast, they are strongly rejected for Slovakia, as indicated by the very low p-values. However, rejection of

[5] All estimations of Table 5 were performed by Gauss routines. I am grateful to John Y. Campbell, who kindly provided me with the Gauss codes. These codes were modified properly, in order to include the estimated structural breaks into the analysis.

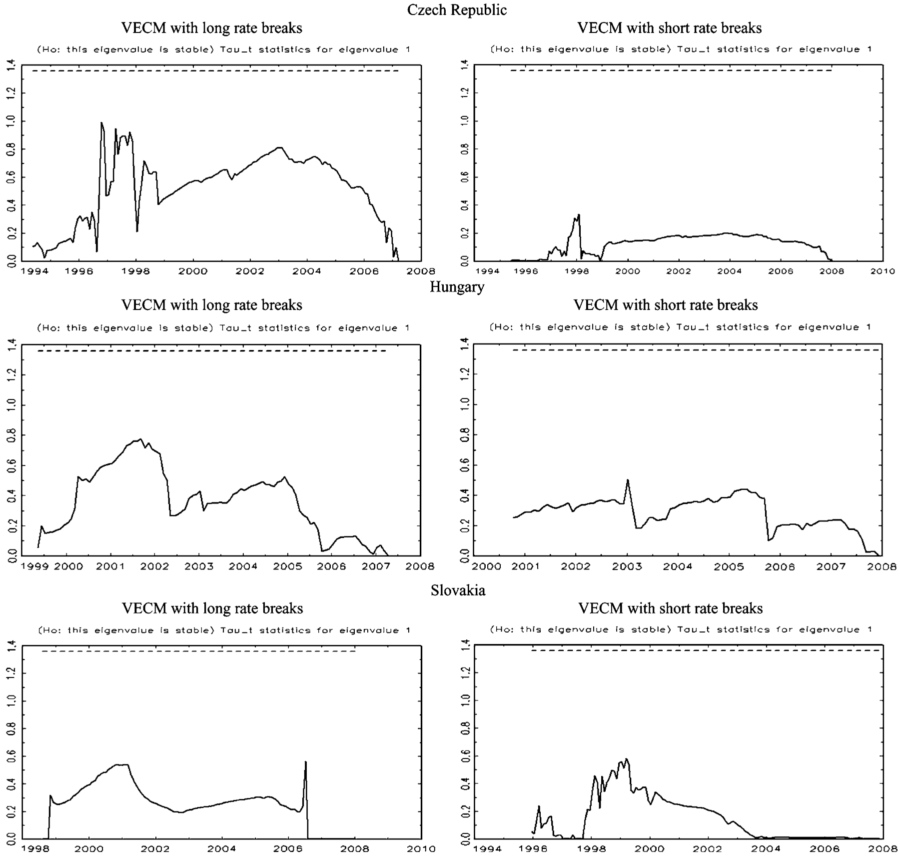

Fig. 1 Parameter constancy tests

the cross-equation restrictions in the case of Slovakia does not mean that the EHTS is devoid of any economic content. As indicated by Campbell and Shiller (1991), it is quite possible that minor deviations from the EHTS may lead to statistical rejection of theory. For this reason, the economic significance of the EHTS has been also evaluated by computing the variance ratio of the $S_{(n,m),t}$ to the $S'_{(n,m),t}$ and examining the correlations between them.

Columns 5 and 6 of Table 5 show the results for the variance ratio and the correlation coefficient corr($S_{(n,m),t}$, $S'_{(n,m),t}$) between the actual and the theoretical spread, respectively. Column 5 indicates that for both VARs of each of the Czech Republic and Hungary the variance ratios are not greater than two SDs from unity. Also, column 6 shows that the correlation coefficient between $S_{(n,m),t}$ and $S'_{(n,m),t}$ is high and close to unity. This implies that for these two countries, the deviations from the EHTS are not economically or statistically significant. In contrast, there is evidence against the EHTS for Slovakia. For this country, the variance ratio is very low and greater than two SDs from unity, while the correlation coefficient between the actual and the theoretical spread is quite low and far from unity. Figure 2, which

Table 4 VAR model for ($S_{(n,m),t}$, $\Delta r_{m,t}$)

Country	Breaks included in the VAR	$\hat{k}$ [a]	Granger causality tests		LM test	R^2	
			$S_{(n,m),t}$ to $\Delta r_{m,t}$	$\Delta r_{m,t}$ to $S_{(n,m),t}$		$\Delta r_{m,t}$-eqn	$S_{(n,m),t}$-eqn
Czech Republic	Long rate:	2	40.59*	15.91*	7.51	0.602	0.782
	1997:02, 1999:07		(0.000)	(0.000)	(0.111)		
	Short rate:	2	35.88*	14.89*	7.02	0.604	0.777
	1997:08, 1999:01		(0.000)	(0.001)	(0.135)		
Hungary	Long rate:	3	11.75*	5.78	2.53	0.105	0.848
	1998:08, 2000:11		(0.008)	(0.123)	(0.638)		
	Short rate:	3	11.38*	5.38	2.16	0.106	0.847
	2003:04, 2005:10		(0.010)	(0.146)	(0.706)		
Slovakia	Long rate:	3	25.89*	193.53*	2.68	0.492	0.709
	1997:10, 1999:01		(0.000)	(0.000)	(0.613)		
	Short rate:	3	24.28*	145.79*	1.13	0.441	0.643
	1998:07, 1998:09		(0.000)	(0.000)	(0.889)		

Numbers in the LM test column are multivariate LM test statistics, which under the null hypothesis of no autocorrelation, are distributed as χ^2 asymptotically, with degrees of freedom d^2, where $d = 2$ is the dimension of the VAR. Numbers in parentheses are p-values

* Rejection of the null hypothesis at the 0.05 level of significance

[a] $\hat{k}$ denotes the estimated lag length in each VAR

Table 5 Testing the EHTS of interest rates

Country	Breaks included in the VAR	Wald tests		VR	Corr ($S_{(n,m),t}$, $S'_{(n,m),t}$)
		Test statistic	*df*		
Czech Republic	Long rate:	5.03	4	0.952	0.996
	1997:02, 1999:07	(0.284)		(0.703)	(0.024)
	Short rate:	4.80	4	1.050	0.997
	1997:08, 1999:01	(0.309)		(1.120)	(0.020)
Hungary	Long rate:	4.71	6	0.821	0.996
	1998:08, 2000:11	(0.582)		(1.180)	(0.007)
	Short rate:	4.63	6	0.761	0.996
	2003:04, 2005:10	(0.592)		(1.240)	(0.010)
Slovakia	Long rate:	378.00*	6	0.405→	0.451
	1997:10, 1999:01	(0.000)		(0.127)	(0.234)
	Short rate:	611.00*	6	0.382→	0.487
	1998:07, 1998:09	(0.000)		(0.120)	(0.271)

Under the null hypothesis H_0: $S_{(n,m),t} = S'_{(n,m),t}$, the Wald test statistics are χ^2-distributed, asymptotically with $2p$ degrees of freedom (*df*), where p is the VAR order for $x_t \equiv (\Delta r_{m,t}, S_{(n,m),t})'$. Numbers in parentheses in column 3 are p-values. Numbers in parentheses in columns 5 and 6 are standard errors

* Rejection of the null hypothesis at the 0.05 level of significance

→ A variance ratio that is greater than two SDs from unity

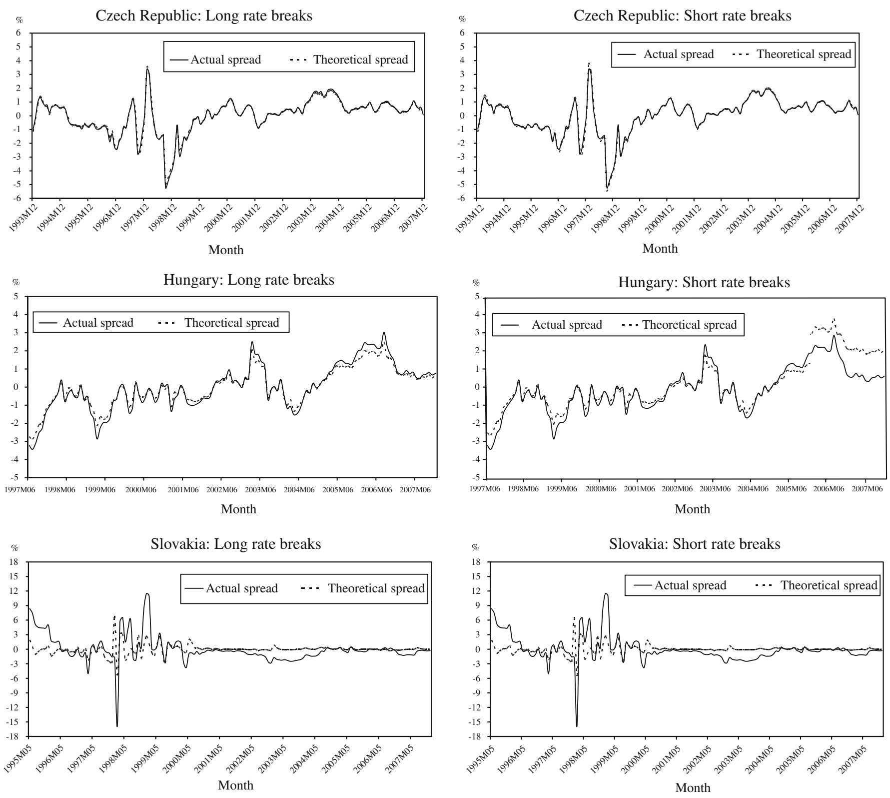

Fig. 2 Term structure: deviations from means of actual spread ($S_{(n,m),t}$) and theoretical spread ($S'_{(n,m),t}$)

plots the $S_{(n,m),t}$ and the $S'_{(n,m),t}$ between the interest rates of each of the above three CE countries, conveys similar information. For the Czech Republic and Hungary the actual and the theoretical spread seem to move together over time, while for Slovakia the low correlation between the actual and the theoretical spread is clear.

Combining these results with the results of the previous section, they clearly provide support of the empirical adequacy of the EHTS only in the cases of the Czech Republic and Hungary. On the contrary, for Poland the EHTS is strongly rejected and for Slovakia the actual spread is less volatile than the theoretical spread. The above results are somehow different than those provided by Koukouritakis and Michelis (2008) (KM) and they more accurate for three reasons: (a) the use of unit root and cointegration testing in the presence of structural breaks, which allows to avoid distortions in standard inference procedures, (b) the implementation of cross-equation restrictions tests and variance ratio estimations, which are not used in the KM study, and (c) the use of a larger time span for the countries of the sample. The implementation of these techniques in the present study led to no rejection of the $(1, -1)'$ null hypothesis for the cointegrating vector, in the cases of the Czech Republic and Hungary, while it gave a clear-cut result for Poland, which is against the EHTS. For the case of Slovakia, the present paper and the KM study both indicate cointegration with cointegrating vector $(1, -1)'$. However, the results of cross-equation restrictions tests and variance ratio estimations in the present study indicate rejection of the EHTS for this country, since the cross-equation restrictions are strongly rejected and the variance ratio is very low and greater than two SDs from unity.

The results for the Czech Republic and Hungary can be explained by the fact that during their transition period these two countries had accelerated the process of financial market liberalization. Additionally, according to the EU Commission Report they were considered as market economies since late 1990s. On the contrary, for Poland and Slovakia the overall transition process was impeded by political instability and the lack of social consensus during the 1990s. As a result, the financial markets of these two countries became fully market-determined only in the first years of the present decade, and this may explain the rejection of the empirical adequacy of the EHTS.

Especially for Slovakia, the above results are of great interest because the country became the sixteenth member of the euro area since January 2009. As Angeloni and Ehrmann (2003) indicate, the transmission process of the European Central Bank's (ECB) monetary policy through the interest-rate channel is quite similar across the EMU members. Thus, the rejection of EHTS for Slovakia raises some doubts about the effectiveness of the euro area's monetary policy on the country, since the short rate, which is directly affected by the decisions of the ECB, has a limited impact on the long rate. However, the empirical results indicate that the rejection of the EHTS for Slovakia is a result of significant deviations of the $S_{(n,m),t}$ from the $S'_{(n,m),t}$, even though the short and the long rate are cointegrated. As shown in the third panel of Fig. 2, these deviations took place mainly in the 1990s, where the country was facing high and volatile inflation that had a direct impact on its term structure. After the country's accession to the EU in May 2004, which signalised the end of the transition process, the $S_{(n,m),t}$ and the $S'_{(n,m),t}$ seem to move together and thus, the deviations from the EHTS may not be economically significant.

5 Concluding remarks

This paper investigated empirically the term structure of interest rates in four CE countries that are EU members: the Czech Republic, Hungary, Poland and Slovakia. Since the time span of the sample period covers more than one decade, the existence of structural breaks was quite possible and confirmed by the use of the two-break minimum LM unit root tests. The empirical evidence shows that all interest rates are non stationary and allow for two structural breaks. These breaks occurred during the transition period of these countries from centrally planned economies to full EU members. Since the interest rates follow random walks, the expectations hypothesis of the term structure has been evaluated using the VAR approach of Campbell and Shiller along with the Lütkepohl et al. cointegration approach in the presence of structural breaks. The empirical findings of the present paper are providing support of the empirical adequacy of the EHTS for the Czech Republic and Hungary. For Poland and Slovakia, the empirical evidence implies economically important deviations from the EHTS. Especially for Slovakia, these deviations took place mainly in the 1990s due to high and volatile inflation, which had a direct effect on the country's interest rates.

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References

Angeloni, I., & Ehrmann, M. (2003). *Monetary policy transmission in the euro area: any changes after EMU?* (European Central Bank Working Paper, No. 240). Frankfurt a.M.

Bekaert, G., & Hodrick, R. J. (2001). Expectations hypotheses tests. *Journal of Finance, 56*(4), 1357–1394. doi:10.1111/0022-1082.00371.

Bekaert, G., Wei, M., & Xing, Y. (2007). Uncovered interest rate parity and the term structure. *Journal of International Money and Finance, 26*(6), 1038–1069. doi:10.1016/j.jimonfin.2007.05.004.

Brüggemann, R., & Lütkepohl, H. (2005). Uncovered interest rate parity and the expectations hypothesis of the term structure: Empirical results for the US and Europe. *Applied Economics Quarterly, 51*(2), 143–154.

Campbell, J. Y., & Shiller, R. J. (1987). Cointegration and tests of present value models. *Journal of Political Economy, 95*(5), 1062–1088. doi:10.1086/261502.

Campbell, J. Y., & Shiller, R. J. (1991). Yield spreads and interest rate movements: A bird's eye view. *Review of Economic Studies, 58*(3), 495–514. doi:10.2307/2298008.

Cuthbertson, K. (1996). The expectations hypothesis of the term structure: The UK interbank market. *Economic Journal, 106*(3), 578–592. doi:10.2307/2235564.

Cuthbertson, K., & Bredin, D. (2000). The expectations hypothesis of the term structure: The case of Ireland. *The Economic and Social Review, 31*(3), 267–281.

Diebold, F. X., Rudebusch, G. D., & Aruoba, S. B. (2006). The macroeconomy and the yield curve: A dynamic latent factor approach. *Journal of Econometrics, 131*(1–2), 309–338. doi:10.1016/j.jeconom.2005.01.011.

Estrella, A., & Hardouvelis, G. A. (1991). The term structure as a predictor of real economic activity. *Journal of Finance, 46*(2), 555–576. doi:10.2307/2328836.

Estrella, A., & Mishkin, F. (1998). Predicting U.S. recessions: Financial variables as leading indicators. *Review of Economics and Statistics, 80*(1), 45–61. doi:10.1162/003465398557320.

Gravelle, T., & Morley, J. C. (2005). A kalman filter approach to characterizing the Canadian term structure of interest rates. *Applied Financial Economics, 15*(10), 691–705. doi:10.1080/09603100500107917.

Hall, A. D., Anderson, H. M., & Granger, C. W. J. (1992). A cointegration analysis of treasury bill yields. *Review of Economics and Statistics, 74*(1), 116–126. doi:10.2307/2109549.

Hansen, H., & Johansen, S. (1999). Some tests for parameter constancy in cointegrated VAR-models. *Econometrics Journal, 2*(2), 306–333. doi:10.1111/1368-423X.00035.

Hardouvelis, G. A. (1994). The term structure spread and future changes in long and short rates in the G7 countries. *Journal of Monetary Economics, 33*(2), 255–283. doi:10.1016/0304-3932(94)90003-5.

Jardet, C. (2008). Term structure anomalies: Term premium or peso-problem? *Journal of International Money and Finance, 27*(4), 592–608. doi:10.1016/j.jimonfin.2008.01.011.

Johansen, S. (1988). Statistical analysis of cointegration vectors. *Journal of Economic Dynamics and Control, 12*(2), 231–254. doi:10.1016/0165-1889(88)90041-3.

Johansen, S. (1995). *Likelihood-based inference in cointegrated vector autoregressive models*. Oxford: Oxford University Press.

Johansen, S., Mosconi, R., & Nielsen, B. (2000). Cointegration analysis in the presence of structural breaks in the deterministic trend. *Econometrics Journal, 3*(3), 216–249. doi:10.1111/1368-423X.00047.

Jondeau, E., & Ricart, R. (1999). The expectations hypothesis of the term structure: Tests on US, German, French, and UK euro-rates. *Journal of International Money and Finance, 18*(5), 725–750. doi:10.1016/S0261-5606(99)00029-7.

Koukouritakis, M., & Michelis, L. (2008). The term structure of interest rates in the 12 newest EU countries. *Applied Economics, 40*(4), 479–490. doi:10.1080/00036840600690249.

Lahiri, K., & Wang, J. (1996). Interest rate spreads as predictors of business cycles. In G. Maddala, & C. Rao (Eds.), *Handbook of statistics* (Vol. 14, pp. 297–315). Amsterdam: Elsevier.

Lanne, M. (2003). Testing the expectations hypothesis of the term structure of interest rates in the presence of a potential regime shift. *The Manchester School, 71*(Supplement), 54–67. doi:10.1111/1467-9957.71.s.4.

Lee, J., & Strazicich, M. C. (2003). Minimum Lagrange multiplier unit root test with two structural breaks. *Review of Economics and Statistics, 85*(4), 1082–1089. doi:10.1162/003465303772815961.

Lütkepohl, H., & Saikkonen, P. (2000). Testing for the cointegrating rank of a VAR process with a time trend. *Journal of Econometrics, 95*(1), 177–198. doi:10.1016/S0304-4076(99)00035-4.

Lütkepohl, H., Saikkonen, P., & Trenkler, C. (2003). Comparison of tests for the cointegrating rank of a VAR process with a structural shift. *Journal of Econometrics, 113*(2), 201–229. doi:10.1016/S0304-4076(02)00200-2.

Perron, P. (1989). The great crash, the oil price shock, and the unit root hypothesis. *Econometrica, 57*(6), 1361–1401. doi:10.2307/1913712.

Perron, P. (1993). Erratum: The great crash, the oil price shock, and the unit root hypothesis. *Econometrica, 61*(1), 248–249. doi:10.2307/2951792.

Saikkonen, P., & Lütkepohl, H. (2000). Testing for the cointegrating rank of a VAR process with structural shifts. *Journal of Business and Economic Statistics, 18*(4), 451–464. doi:10.2307/1392226.

Shiller, R. J. (1990). The term structure of interest rates. In B. M. Friedman, & F. H. Hahn (Eds.), *Handbook of monetary economics* (Vol. 1, pp. 627–722). Amsterdam: North-Holland.

Trenkler, C. (2008). Determining p-values for systems cointegration tests with a prior adjustment for deterministic terms. *Computational Statistics, 23*(1), 19–39. doi:10.1007/s00180-007-0066-8.

Trenkler, C., Saikkonen, P., & Lütkepohl, H. (2008). Testing for the cointegrating rank of a VAR process with level shift and trend break. *Journal of Time Series Analysis, 29*(2), 331–358. doi:10.1111/j.1467-9892.2007.00558.x.