Structural Breaks and the Term Structure of Interest Rates in the European Union

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Abstract

In this paper we test for the statistical and economic significance of the expectations hypothesis of the term structure of interest rates in the original 14 EU countries, in the presence of structural breaks in the data, using vector autoregression and cointegration techniques. We examine several testable implications of the theory including: (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 actual spreads relative to theoretical spreads. The empirical results are mixed and provide partial support of these implications for 12 out of 14 EU countries. Only for Greece and Portugal, the results strongly reject the expectations hypothesis of the term structure.

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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. There are several reasons for this. First, the shape of the term structure or yield curve provides valuable information about the future movements of the long-term interest rates, and hence the long-term investment prospects and economic growth of a country. Second, the spread between the long and current short rate is a better predictor of a country's monetary policy stance than the level of the short-term interest rates or the rate of monetary growth allowed by a central bank. Third, empirical studies have suggested that the interest rate spread has good predictive power about the future movement on economic activity, and hence the cyclical behaviour of an economy (Estrella and Hardouvelis (1991), Plosser and Rouwenhorst (1994), Lahiri and Wang (1996), Estrella and Mishkin (1998), Ang, Piazzesi and Wei (2006)). Also, as Jorion and Mishkin (1991) showed, the term structure has significant ability to forecast changes in inflation, particularly at long maturities

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, the interest rate on a long-term security is the average of the current short rate and the expected future rates on securities of shorter maturity, plus a possibly time varying term premium. Thus, 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.

The empirical literature has delivered mixed results regarding the validity of the EHTS of interest rates. Among others, Campbell and Shiller (1987, 1991) estimated bivariate vector autoregression (VAR) models for short rates and term spreads using US data, over the period 1959:1-1983:10. They showed that, although the expectations hypothesis was statistically rejected, long rates behaved in a very similar way to that implied by the expectations hypothesis. Hall, Anderson and Granger (1992) used monthly data from 1970:3 to 1988:12 for 12 yield series of US Treasury Bills and found cointegration among the interest rates, thereby interpreting the evidence as supportive of the EHTS. Hardouvelis (1994) used monthly data of different time spans for the G7 countries and found that the EHTS holds for all countries except the US. Cuthbertson (1996) studied the EHTS for the UK interbank market, using weekly data for 1-week, 4-week, 13-week, 26-week and 52-week discount rates, for the 1981-1992 period. Using the Johansen cointegration approach together with the Campbell-Shiller VAR methodology, he found evidence that gives some support for the EHTS at shorter maturities and a failure of the EHTS at longer maturities. Jondeau and Ricart (1999) used the Johansen approach to test the EHTS for French, German, the UK and the US euro rates. Using monthly data for 1-month, 3-month, 6-month and 12-month euro rates from 1975:1 to 1997:12, they could not reject the EHTS for French and UK rates, but they rejected it for German and US rates. Cuthbertson and Bredin (2000) investigated the EHTS for Ireland, using a sample consisted of monthly data for 1-month, 3-month and 6-month money market rates, for the 1984:1-1997:10 period. Using the Johansen cointegration approach together with the Campbell-Shiller VAR methodology, they found evidence that is consistent with the EHTS. Bekaert and Hodrick (2001) studied the EHTS for Germany, the UK and the USA, using 1-month and 12-month Eurocurrency interest rates, together with nominal exchange rates, for a sample that covers a period from 1975:1 to 1997:7. Using alternative statistical techniques, such as Wald, Lagrange Multiplier, and distance metric tests, and extensive Monte Carlo methods, they found no evidence against the EHTS only for the UK, weak evidence against the EHTS for the USA and stronger evidence against the EHTS for Germany. Lanne (2003) investigated the EHTS for the US Eurodollar 1-month, 3-month and 6-month deposit rates, allowing for potential regime shifts. His sample consisted of monthly observations, covering the period from 1983:1 to 1999:6, while his evidence supports the EHTS at the short end of the maturity spectrum once a potential regime shift was allowed for. Brüggemann and Lütkepohl (2005) analysed the relation between long and short rates for the Euro area and the USA, using monthly data from 1985:1 to 2004:12. Using cointegration and vector error-correction models (VECM), they found evidence that was consistent with the EHTS. Diebold, Rudebusch and Aruoba (2006) developed a yield curve model that incorporates yield factors and macroeconomic variables and they related it with the EHTS. They used monthly data for 17 US Treasury yields from 1972:1 to 2000:12 and their evidence was in favour of the EHTS for certain periods, but not for the entire sample. Bekaert, Wei and Xing (2007) examined the EHTS simultaneously with the uncovered interest rate parity, drawing data from Germany, Japan, the UK and the USA. Their sample consisted of monthly observations from 1972:01 through 1991:12 on implied zero-coupon yields with maturities of 3, 12, 24, 36, 48, and 60 months. Using VAR and Monte Carlo analysis, they found evidence that was, in general, against the EHTS but, economically, actual spreads and theoretical spreads did not behave very differently, especially at long horizons. Koukouritakis and

Michelis (2008) used cointegration and common trends techniques to study the EHTS among the twelve newest EU countries. Using monthly data for a long and a short rate from early 1990's to 2004:12, they found evidence in favour of the EHTS for all countries except Malta.

The present paper takes a fresh look at the study of the EHTS, by examining this issue in the context of the original 14 EU countries. The novelty of this study lies on the fact that we allow for structural breaks in the data. Our choice of the 14 EU countries for a sample that begins in the early 1980s and ends in 2009 provides a solid ground for this analysis. During this period, these countries moved to important policy reforms such as the German unification in 1990, and affected overall by events such as the financial crisis in 1992 and the subsequent collapse of the European Monetary System, the signing of the Maastricht Treaty and the introduction of the euro in 1999. These events are likely to have caused structural shifts in the term structure of interest rates of the original 14 EU countries. In order to capture these shifts, we use recently developed Lagrange Multiplier (LM) unit root tests (Lee and Strazicich (2003, 2004)) and cointegration tests (Johansen, Mosconi and Nielsen, 2000 and Lütkepohl and his associates in several recent papers noted below) in the present analysis. Finally, we use the VAR approach proposed by Campbell and Shiller (1987, 1991) in order to test for the economic significance of EHTS in the 14 EU countries. Briefly, our results are mixed and provide partial support of the statistical and economic significance of the EHTS for 12 out of 14 EU countries. The EHTS is strongly rejected only for Greece and Portugal.

The rest of the paper is organized 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 summarises and contains some concluding remarks.

2. The EHTS of Interest Rates and Testable Implications

According to the EHTS, the yield to maturity of an *n*-period bond $R_{n,t}$ is equal to the average of the current and expected future rates on a set of $m-$period short yields $r_{m,t}$, with $m<n$, plus the term premium. The relationship can be expressed in the following form

$$\left(1+R_{n,t}\right)^{n}=\varphi^{*}_{(n,m),t}\prod\nolimits_{i=0}^{k-1}\left(1+E_{t}r_{m,t+im}\right)\ , \tag{1}$$

where $k=n/m$ is an integer, $\varphi^{*}_{(n,m),t}$ is a possible non-zero but stationary *n*-period term premium and E_t is the expectations operator conditional on market information up to and including time t. The pure EHTS holds when there is no term premium of any kind $(\varphi^*_{(n,m),t} = 0)$. The weaker version of the EHTS allows for a constant term premium in equation (1).

The equality in equation (1) is established by the condition of no arbitrage opportunities to investors willing to hold both short-term and long-term bonds. Log-linearizing equation (1) we get

$$R_{n,t} = \varphi_{(n,m),t} + \left(1/k\right)\sum_{i=0}^{k-1} E_t r_{m,t+im} \quad , \tag{2}$$

where $\varphi_{(n,m),t} = \log\left(\varphi^*_{(n,m),t}\right)$. Equation (2) states that the yield of the $n-$period bond is equal to a term premium plus a simple arithmetic average of the current and expected future short rates. This representation is valid for pure discount bonds, (Shiller, 1979), which is the case with our data. The term premium may change with *m* and *n* but is assumed constant over time.

For the subsequent analysis it is convenient to subtract $r_{m,t}$ from both sides of equation (2) and write

$$S_{(n,m),t} = \varphi_{n,m} + E_t \sum_{i=1}^{k-1} (1 - i/k)\Delta^m r_{m,t+im} = \varphi_{n,m} + E_t\,(S^*_{(m,n),t}) \tag{3}$$

where $S_{(n,m),t} \equiv (R_{n,t} - r_{m,t})$ is the actual yield spread, $\Delta^m r_{m,t+im} \equiv \left(r_{m,t+im} - r_{m,t}\right)$ is the change in the short term (*m*-period) interest rates and $S^*_{(n,m),t} \equiv \sum_{i=1}^{k-1}(1-i/k)\Delta^m r_{m,t+im}$ is the *perfect foresight spread*, which would obtain, under the EHTS, if economic agents had perfect foresight about future movements in interest rates. It is clear from equation (3) that the actual spread *S* is an optimal forecast of the perfect foresight spread, a weighted average of changes in the short rates. 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. Here we concentrate on four testable predictions: (i) cointegration, (ii) the cointegrating vector linking long and short rates is $\left(1, -1\right)$, (iii) cross-equation restrictions implied by the theory, and (iv) the ratio of the variances of the actual spread to that implied by the theory should be unity.

(i) Cointegration

Given that $R_{n,t}$ and $r_{m,t}$ are integrated of order one, I(1), equation (3) implies that the two rates should be cointegrated, as its right hand side is a stationary or an I(0) process, provided that the term premium and changes in the short rates are stationary. Cointegration between long and short rates is consistent with the idea that market forces continuously adjust to correct any temporary disequilibrium, so that term-adjusted rates of return on different maturities do not drift apart permanently. Otherwise, this would give rise to arbitrage opportunities.

***(ii) The cointegrating vector is* (1,-1)**

In general, if $R_{n,t}$ and $r_{m,t}$ are cointegrated, there exist constants *a* and *b* such that the linear combination $aR_{m,t} + br_{m,t}$ is a stationary process, even though each of the rates is an $I(1)$ process. It is clear from the left hand side of equation (3) that the EHTS implies the cointegrating vector $(a,b)' = (1,-1)'$.

(iii) Cross-equation restrictions

Campbell and Shiller (1987, 1991) proposed a VAR methodology in order to evaluate the economic importance of deviation from the EHTS. For this, they specify a VAR and derive a set of cross-equation restrictions that must hold under the EHTS. Further, using the VAR, they compute the *theoretical spread*, an estimate of the perfect foresight spread, and then they compare it to the actual spread. Significant differences between the two measures of the spread are interpreted as evidence against the expectations hypothesis.

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

$$z_t = Az_{t-1} + v_t \ , \tag{4}$$

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.

Next, define the $2p \times 1$ vectors g and h such that $g' z_t = S_{(m,n),t}$ and $h' z_t = \Delta r_{m,t}$, i.e., the elements of the vectors g and h are all zero, except for the $p+1st$ element of g and the first element of h, both of which are unity. Projecting both sides of equation (3) onto the information contained in z_t gives

$$S_{(m,n),t} = g' z_t = S'_{(m,n),t} \tag{5}$$

where

$$S'_{(m,n),t} = h' A\left[I - (m/n)(I - A^n)(I - A^m)^{-1}\right](I - A)^{-1} z_t \tag{6}$$

is the *theoretical spread*, computed from the VAR. In the present study m and n take different values, depending on data availability across the counties in our sample; see Table 1.

Since equation (4) holds for any general z_t, it must be the case that

$$g' = h'A\left[I - (m/n)(I - A^n)(I - A^m)^{-1}\right](I - A)^{-1} \quad (7)$$

It is clear from equations (5) and (6) that the set of restrictions in equation (7) are equivalent to the null hypothesis $H_0 : S_{(n,m),t} = S'_{(n,m),t}$. This hypothesis can be easily 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 provides an adequate description of the data. Otherwise it is rejected in favour of excess returns in the bonds market of a country.

(iv) Variance ratio

Equation (5) suggests an alternative way to test the empirical content of the EHTS. Consider the variance ratio $VR = \text{var}\left(S_{(n,m),t}\right) / \text{var}\left(S'_{(n,m),t}\right)$, together with the correlation between S_t and S'_t. If the EHTS provides an adequate description of the data, the correlation should be close to one, and the variances of the actual and the theoretical spreads should behave similarly over time. In this case, the variance ratio should be close unity. Otherwise, it will be significantly different from one. In particular, if $VR > 1$ there is "excess volatility" in the bonds market, in the sense that the actual spread is more volatile than the optimal predictor of future short rates.

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.

In section 3 below we use the two- and one-break LM unit root tests and system cointegration tests in the presence of structural breaks, in order to test the predictions (i) and (ii). Further, we adopt the VAR methodology of Campbell and Shiller in the presence of structural breaks, in order to test the predictions (iii) and (iv). In the next section we outline the unit root tests and cointegration models that will be used in the subsequent analysis.

3. Unit Roots and Cointegration with Structural Breaks

Table 1 lists the original 14 EU countries along with data sources and monthly observations on selected interest rates from the early 1980s to 2009:12, with different data spans for different countries, depending on data availability. During this period several, country specific and EU-wide, events have taken place in the EU, which are likely to have caused structural breaks in the term structure of interest rates across these countries. These events include, among others, a country's membership in the EU, the German unification in 1990, the European financial crisis in 1992, the ratification of the Maastricht Treaty in 1992, country specific reforms to satisfy the convergence criteria for participation in the European Monetary Union (EMU) and the introduction of the euro in 1999. 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, in the present study, we employ recently developed tests that are valid in the presence for structural shifts in the data.

3.1 Unit Root Tests with Structural Breaks

We test for unit roots in the data using the two-break and one-break LM tests developed by Lee and Strazicich (2003, 2004). These tests have several desirable properties: (a) they determine the structural breaks endogenously from the data, (b) their null distributions are invariant to level shifts in a variable, and (c) they are easy to interpret; by including breaks under both the null and alternative hypotheses, a rejection of the null hypothesis of a unit root implies unambiguously trend stationarity.

The LM tests are also easy to implement. Consider for example the two-break LM unit test for the process y_t generated by

$$y_t = \delta' Z_t + e_{t,} \qquad e_t = \beta e_{t-1} + A(L)\varepsilon_t, \qquad \varepsilon_t \sim iid\ N\left(0,\ \sigma^2\right) \tag{8}$$

where *A(L)* is a *k*-order polynomial in the lag operator *L* and Z_t is a vector of exogenous variables whose components are determined by the type of breaks 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, if T_{Bj} denotes the point in time the break occurs, 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, in addition, $DT_{jt} = t - T_{Bj}$ for $t \geq T_{Bj} + 1,\ j = 1,2$, and zero otherwise.

It is clear from (8) that y_t has a unit root if $\beta = 1$. Alternatively it is trend stationary if $\beta < 1$. According to the LM principle, a unit root test statistic can be obtained from the test regression

$$\Delta y_t = \delta' \Delta Z_t + \phi \tilde{S}_{t-1} + \sum\nolimits_{i=1}^{k} \theta_i \Delta \tilde{S}_{t-i} + u_t, \tag{9}$$

where $\tilde{S}_t = y_t - \tilde{\psi}_x - Z_t \tilde{\delta},\ t = 2,...,T$, 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 an iid error term with zero mean and finite variance. 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 equation (9) and can be tested by the LM test statistic:

$$\tilde{\tau} = t\text{-statistic for the hypothesis } \phi = 0. \tag{10}$$

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:

$$LM_\tau = \inf_\lambda \left\{ \tilde{\tau}(\lambda) \right\} \tag{11}$$

The critical values for the test are available in Lee and Strazicich (2003). They are invariant to the break locations$\left(\lambda_j\right)$ for Model A, but not for Model C.

In the present study, when the two-break LM test results showed that only one structural break is significant for some cases, we also compute the one-break LM test of Lee and Strazicich (2004). This was done not only because the one-break LM test appears more appropriate in this case, but also because we wanted to determine if including two breaks instead of one can adversely affect the power to reject the unit root hypothesis for these cases.

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 on cointegration in a VAR framework, there are two main approaches to testing for cointegration in the presence of structural breaks. One approach developed by Johansen, Mosconi and Nielsen (2000) (henceforth the JMN approach) 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 of a vector-valued stochastic process. JMN then derive the asymptotic distribution of the likelihood ratio (LR) or trace statistic for cointegration and obtain critical values or p-values, for the multivariate counterparts of models A and C above with q possible breaks, using the response surface methodology.

To illustrate the JMN approach, consider briefly the simple case of model A with only level shifts in the constant term μ of an observed $p-$dimensional time series Y_t , $t = 1,...,T$, of possibly $I(1)$ variables. JMN divide the sample observations into q sub-samples, according to the location of the break points, each of length $T_j - T_{j-1}$ for $j = 1,...,q$ and $0 = T_0 < T_1 < ... < T_q = T$, such that the last observation in the j^{th} sub-sample is T_j , while the first observation in the $(j+1)^{\text{th}}$ sub-sample is $T_j + 1$. They assume the following $VECM(k)$ for Y_t conditional on the first k observations of each sub-sample $Y_{T_{j-1}+1},...,Y_{T_{j-1}+k}$:

$$\Delta Y_t = \Pi Y_{t-1} + \mu D_t + \sum\nolimits_{i=1}^{k-1} \Gamma_i \Delta Y_{t-i} + \sum\nolimits_{i=1}^{k} \sum\nolimits_{j=2}^{q} g_{ji} D_{j,t-i} + \varepsilon_t, \quad \varepsilon_t \sim iidN(0,\Omega), \quad (12)$$

where $\mu = (\mu_{1,........,}\mu_q)$ and $D_t = (D_{1,t,........,}D_{q,t})'$ are of dimension $(p \times q)$ and $(q \times 1)$, respectively, and the $D_{j,t}$'s are dummy variables, such that $D_{j,t} = 1$ for $T_{j-1} + k + 1 \le t \le T_j$ and $D_{j,t} = 0$ otherwise, for $j = 1,....,q$.

As is well known, the hypothesis of at most r_0 cointegrating relations $(0 \le r_0 < p)$ among the components of Y_t can be stated in terms of the reduced rank of the $(p \times p)$ matrix Π in which case it can be written as $\Pi = \alpha\beta'$, where α and β are matrices of dimension $(p \times r)$. The cointegration hypothesis can then be tested by the likelihood ratio statistic

$$LR_{JMN} = -T \sum\nolimits_{i=r_0+1}^{p} \ln\left(1 - \hat{\lambda}_i\right) \quad (13)$$

where the eigenvalues $\hat{\lambda}_j$'s can be obtained by solving the related generalized eigenvalue problem, based on estimation of the $VECM(k)$ in equation (12) under the additional restrictions that $\mu_j = \alpha\rho_j^{'}$, $j = 1,.....,q$, where ρ_j is of dimension $1 \times r$.[1]

The second approach developed by Lütkepohl and his associates (henceforth the LST approach; see among others, Lütkepohl and Saikkonen (2000), Saikkonen and Lütkepohl (2000), Trenkler, Saikkonen and Lütkepohl (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. Thus, LST set up the data generation process (DGP) for Y_t by adding its

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[1] These restrictions are required in order to eliminate a linear trend in the level of the process Y_t. Using these restrictions in (12), estimation involves the reduced rank regression of ΔX_t on $(X_{t-1}^{'}, D_t^{'})^{'}$ each corrected for the regressors ΔX_{t-i} $(i = 1,.....,k-1)$ and $D_{j,t-i}$ $(i = 1,.....,k; j = 2,.....,q)$.

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 generalized 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 LM-type and 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. Lütkepohl Saikkonen and Trenkler (2003) study the statistical properties of their tests for the case of level shifts only, and compare them to the JMN test. They find that the LR-type tests perform better than the LM-type tests in finite samples. Further, their tests have better size and power properties than the JMN test in finite samples.

To illustrate the LST approach for LR-type tests, consider the case of a single shift in the level of Y_t. Assuming an exogenous break at time T_B in the level of μ_t, LST specify the following DGP for Y_t:

$$Y_t = \mu_t + X_t = \mu_0 + \mu_1 t + \delta d_t + X_t, \quad t = 1,....,T, \tag{14a}$$

where t is a linear time trend, $\mu_i \left(i = 0,1\right)$ and δ are unknown $\left(p \times 1\right)$ parameter vectors, d_t is a dummy variable defined as $d_t = 0$ for $t < T_B$ and $d_t = 1$ for $t \geq T_B$, and where the unobserved stochastic error X_t is assumed to follow a $VAR\left(k\right)$ process with VECM representation

$$\Delta X_t = \Pi X_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \varepsilon_t \ , \quad \varepsilon_t \ \sim iidN(0,\Omega), \quad t = 1,...,T \ . \tag{14b}$$

It is also assumed that the components of X_t are at most $I\left(1\right)$ variables and cointegrated $\left(i.e.\ \Pi = a\beta'\right)$ with cointegrating rank r_0, where $0 < r_0 \leq p$.

Given the DGP in equations (14a), (14b) the first step of the LST approach involves obtaining estimates of the parameter vectors μ_0, μ_1 and δ in (14a) using a feasible GLS procedure under the null hypothesis $H_0\left(r_0\right): rank\left(\Pi\right) = r_0:$ vs. $H_1\left(r_0\right): rank\left(\Pi\right) > r_0$ (e.g., see Saikkonen and Lütkepohl (2000) for details). Having the estimated parameters, $\hat{\mu}_0$, $\hat{\mu}_1$ and $\hat{\delta}$, one then computes the de-trended series $\hat{X}_t = Y_t - \hat{\mu}_0 - \hat{\mu}_1 t - \hat{\delta} d_t$. 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 (14b) and computing the LR or trace statistic:

$$LR_{LST} = -T\sum_{i=r_0+1}^{p} \ln\left(1-\tilde{\lambda}_i\right) \ , \tag{15}$$

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

Under the null hypothesis of cointegration, critical or $p-$values for a single level shift can be computed by the response surface techniques developed in Trenkler (2008). 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 equation (15) 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 (Trenkler et al., 2008, p. 338).

Since the JMN and LST approaches are designed to test the same null hypothesis of cointegration in the presence of structural breaks in the data, we employ both the LR_{JMN} and LR_{LST} test statistics in our empirical analysis of the EHTS. The break points that we use for the cointegration analysis are those determined from the data on the basis of the LM unit root tests discussed above.

4. Data and Empirical Results

4.1 Data

The data set consists of annualised monthly observations for 14 of the original 15 EU members, except for Luxembourg.[2] For most of the countries of our sample, we used data on four interest rates: one short-term treasury bill yield (either 3-month or 12-month rates), one medium-term government bond yield (either 2-year or 3-year rates) and two long-term government bond yields (5-year and 7-year or 10-year rates). For four countries data were available only for three interest rates: Finland and the UK (3-month, 5-year and 10-year rates) and Ireland and Portugal (2-year, 5-year and 10-year rates). For the Netherlands data were available only for the 3-month and the 10-year rates. Our sample consists of monthly observations of varying time spans for different countries, determined by data availability. For

[2] Luxembourg and Belgium have had a monetary union during 1922-2002, which was then subsumed into the European Monetary Union.

all the countries we used end-of-period data, except for Germany and Greece. For these two countries end-of-period data were not available and for that reason we used period average data. The details of the interest rate data are given in table 1.

4.2 Unit Root Results and Structural Breaks

Tables 2 and 3 report the unit root results from the two-break and one-break LM tests, respectively. We tested each interest rate series for a unit root using the two-break LM test at the 1-, 5- and 10 percent levels of significance. As noted above, when this test showed that only one structural break is significant we employed the one-break LM test at the same levels of significance. In order to determine the number of lags, k, in equation (9), we used a "general to specific" procedure at each combination of relative break points (λ_1, λ_2) for the two-break test, and at each single relative break point λ for the one-break test. Initially, we set the lag-length at $k = 12$, and examined the significance of the last lagged term, at the 10 percent level. The procedure was repeated until the last lagged term was found to be significantly different from zero, at which point the procedure stops.[3]

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 the term structures of all 14 EU countries. Also, the results in column 5 of table 2 indicate that all countries, except Finland, have experienced two breaks in their term structures. Table 3 reports similar results for Finland, which has experienced one break in its term structure.[4] As shown in the third column of tables 2 and 3, Model C fits the term structure data best in all cases, regardless of the presence of one or two structural breaks detected in the data. Hence, all countries have experienced shifts both in the deterministic levels and trends of their term structures, over the sample period. Note here that since the results in table 3 are consistent with the results of table 2 regarding the null hypothesis, there does not seem to be any detectable loss of power in using the two-break LM test to test the unit root hypothesis for the case of Finland.

The fifth column of tables 2 and 3 reports the structural breaks in each interest rate series, estimated from the data using the two-break and one-break LM tests, respectively, for each of the 14 EU countries. Not surprising, the estimated breaks correspond well to country-specific policy reforms or international events during the sample period.

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[3] We computed the one- and the two-break LM tests using the Gauss codes of J. Lee available at his website http://www.cba.ua.edu/~jlee/gauss .

[4] We also tested the interest rates of all countries for a second unit root. The null hypothesis was rejected in all cases. These results are available from the authors upon request.

For example, the two-break LM test detects one of the two breaks in three German interest rates in mid-1989, which is probably related with the preparation of the German reunification in 1990. The German unification probably affected the interest rates of two neighbour countries, the Netherlands and Sweden (2-, 5- and 10-year rates), which appear an estimated break during that period. The 10-year rate of Germany appears on break in late 1992. This date coincides with two important events: (a) the ratification of the Maastricht Treaty in early-1992, and its enforcement in late 1993, along with the nominal convergence criteria for participating in the European Monetary Union (EMU), and (b) the financial crisis in Europe in mid-1992, which was related to German reunification in 1990, and the subsequent widening, in mid-1993, of the exchange rate fluctuation margins of the ERM. France's term structure seems to have been affected by these two important events as well. Because of the historical proximity of the two events in (a) and (b), it is not easy to identify their effects separately, but it is clear from the results in tables 2 and 3 that, besides Germany and France, the events seems to have had a widespread impact on the term structures of Belgium, Denmark (3-month, 2- and 5-year rates), Ireland (2-year rate), Sweden (3-month rate) and the UK (3-month rate).

Other important events that had significant effects on the European term structures relate to the second stage of the EMU (January 1994 - December 1998), which included the creation of the European Monetary Institute, in January 1994, the adoption of the Stability and Growth Pact and the creation of ERM II, in June 1997, and the creation of the European Central Bank, in June 1998. As shown in tables 2 and 3, the interest rates of Austria (10-year rate), Denmark, (10-year rate), Finland, Greece (12-month, 3- and 5-year rates), Ireland (5- and 10-year rates), Italy (3-, 5- and 10-year rates), the Netherlands, Portugal, Spain, Sweden (5- and 10-year rates) and the United Kingdom (10-year rate) allow for a structural break during the second stage of the EMU. Additionally, the creation of the euro as an accounting currency and the introduction of the common monetary policy in third stage of the EMU (January 1999 and continuing) seem to have had an impact on the term structure of Austria (12-month, 2- and 5-year rates), Belgium, France, Greece (5-year rate), Italy (12-month, 3- and 5-year rates), Spain (10-year rate), Sweden (3-month and 2-year rates) and the UK (5-year rate).

Finally, country-specific policy reforms seem to have affected the term structure of interest rates in the 14 EU countries. For example, the Austrian interest rates have a break between late 1989 and early 1990, which is the year that country applied for full EU membership. The Danish term structure appears a second estimated break between mid-2006 and early 2007. During that period, the country's interest rates increased due to the expectations of monetary policy tightening. In Germany, the first estimated break varies across its term structure between late 1983 and early 1986, reflecting several important reforms in the money market procedures of the country. Greece's interest have estimated second breaks at the end of 2003 (12-month, 3- and 5-year rates) and in early 2006 (7-year rate). These breaks are probably related to the country's increasing fiscal deficits during those periods. A similar explanation can be given to the second estimated breaks in the interest rates of Portugal and Spain (3-month, 3- and 5-year rates) in 2006. The interest rates of Ireland have an estimated break in 1989, which is probably related to the country's reforms in the transaction system of securities that occurred during that year. The UK's interest rates were subject to a structural break between 1988 and 1989. This break can be associated with that country's worsening fiscal position in the late 1980s, and its preparations for participating in the ERM of the European Monetary System, in 1990.

4.3 Cointegration Results with Structural Breaks

In this section we examine the cointegration test results in the presence of structural breaks. We performed system cointegration tests using the JMN and LST tests described in Section 3.2. For this, we tested for cointegration different pairs of interest rates across different countries, by estimating several VECMs with $Y_t = \left(R_{n,t}, r_{m,t}\right)'$, where R_{nt} is the interest rate on a bond of longer maturity and r_{mt} is the interest rate on a bond of shorter maturity. To compute the JMN test, we estimated different VECMs for each EU country by including the corresponding level shifts and trend breaks of the bond of longer maturity, reported in tables 2 and 3. The LR_{JMN} test statistics and response surface p-values were computed using the JMulti software package, available at the website: http:///www.jmulti.de

To compute the LST test, for each of the 14 EU countries, except Finland, which were found to have two significant breaks in both the level and trend of their interest rates, we extended equation (14a) by adding a second step dummy and two linear trend dummies. For Finland, which was found to have one significant break in both the level and trend of its interest rates, we added a linear trend dummy to (14a). Then, for each pair of interest rates for each of the 14 EU countries we computed the LR_{LST} test statistic and the corresponding response surface p-value using GAUSS routines.[5]

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[5] We are grateful to Carsten Trenkler for kindly providing us with the Gauss codes to perform these estimations.

Table 4 reports the LR_{JMN} and LR_{LST} test results for each pair of interest rates for each of the 14 EU countries. The break points pertaining to each VECM are reproduced in the second column of the table.[6] The lag length, k, for each VECM, was selected using the Akaike information criterion (AIC).

Based on the results in table 4, we find evidence of at least one cointegrating vector in the term structures of all 14 EU countries. Also, for Denmark (2-year and 3-month rates), France (5-year and 3-month rates), Portugal (10- and 2-year rates), Sweden (5- and 2-year rates) and the UK (10-year and 3-month rates) the JMN test indicates a single cointegrating vector, while the LST test indicates no cointegration. Since the LST test has better size and power properties than the JMN test in finite samples, for these cases we conclude that there is no evidence of cointegration between short and long rates. Our results indicate that only for four countries, namely Finland, Greece, Ireland and Italy, there is evidence of cointegration between short and long rates for all maturities. For six countries, namely Belgium, Denmark, France, the Netherlands, Sweden and the United Kingdom, cointegration between short and long rates exists only when the very short rates (3-month) are included in the VECM. For Austria cointegration exists for the shorter (2-year and 12-month) and the longer (10- and 5-year) maturities. Further, the results for Germany indicate cointegration in three out of six pairs of interest rates, while for Portugal and Spain cointegration is established in two out of three and in five out of six cases, respectively.[7]

The JMN and LST tests for cointegration in the presence of structural breaks, assume that the ‘‘long-run’’ cointegration parameters remain constant over the sample period. Otherwise, the test results and inference would be invalid. For this reason, we first tested each VECM for parameter constancy, using the methodology developed by Hansen and Johansen (1999). These authors suggest a graphical procedure based on recursively-estimated 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 1.62, which is the 1 percent critical value for the Hansen and Johansen parameter constancy test. As shown in these plots, the null hypothesis of long run parameter constancy cannot be rejected in all

[6] The VECM and VAR results presented in tables 4, 5 and 6 were obtained using the break point(s) of the interest rate with the longest maturity for each country. Similar results are also obtained if we use the break point(s) of the interest rate with the shortest maturity. The latter results are available upon request.

[7] The pattern of these results is, in general, consistent with several studies, such as Fama (1984), Fama and Bliss (1987) and Campbell and Shiller (1987 and 1991), which find that the term structure of interest rates contains information for short and sometimes long maturities, but it is unreliable to predict movements in interest rates at intermediate maturities.

cases, as the time paths of the eigenvalues are always below the dotted line. As a result, the JMN and LST procedures have been applied correctly.

For all the cases in which there is evidence of cointegration, we also tested the null hypothesis $H_0 : a + b = 0$ where $(a, b)' = (1, -1)'$, that is, the unit vector belongs in the cointegration space as predicted by the EHTS. Under the null hypothesis, this likelihood ratio test is distributed as χ^2 with 1 degree of freedom asymptotically (Johansen, 1995, p. 104). As shown in the last column of table 4, for 12 out of 14 EU countries, except for Greece and Portugal, this hypothesis cannot be rejected at the 5 or 10 percent levels of significance. Thus, for these 12 countries, for all the cases in which the spreads are stationary, our results are also consistent with prediction (ii) of the EHTS. For Greece, the above null hypothesis cannot be rejected only for two out of six pairs of interest rates (7- and 5-years rates, 5- and 3-year rates), while for Portugal it is rejected for both pairs of interest rates in which there is spread stationarity.

Overall, our results give evidence that support the EHTS at short and sometimes long maturities, but not at intermediate maturities, for all 13 out of 14 EU countries. Only for Portugal, our results reject the EHTS of interest rates.

4.4 The Theoretical Spread and the VAR Results

In this section we present and analyse the results from the VAR models for $\Delta r_{m,t}$ and $S_{(n,m),t}$. For the rest of the analysis, we included only the pairs of interest rates for which table 4 shows evidence of spread stationarity and for which the unit vector belongs in the cointegration space. In each VAR we used the same structural breaks as we did in Section 4.3. The appropriate lag length, k, for each VAR was chosen using the likelihood ratio test (Johansen, 1995, p. 21). Also for each VAR, we performed a multivariate LM test for serial correlation.

Table 5 reports the VAR results for Granger causality. Column 4 gives the Wald test statistics (and p-values) for Granger non-causality, which, under the null hypothesis, 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 in almost all cases. The null hypothesis that the spread does not Granger cause short rate changes is rejected for any pair of interest rates for all countries, except for the 5- and 3-year rates of Greece, either at the 5 or at the 10 percent level of significance. Also, there is Granger causality from $\Delta r_{m,t}$ to the spread for almost half of the cases, indicating bi-directional causality in the VAR regressions (table 5, column 5). Further, as shown in the sixth column of table 5, the null hypothesis of no serial correlation in the VAR error term cannot be rejected in all cases, at the 5-percent level of significance, which strengthens the validity of our findings.

Table 6 (columns 3) reports the Wald test results for testing the VAR restrictions in equation (6).[8] As shown in table 6, these restrictions cannot be rejected for most of the cases, either at the 5- or at the 10-percent level of significance. Only in the cases of Finland (10- and 5-year rates), Germany (2-year and 12-month rates) and Greece, the restrictions in equation (6) are strongly rejected, as indicated by the very low p-values. However, rejection of the cross-equation restrictions in these four cases does not mean that the EHTS is devoid of any economic content. As pointed out 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, we also evaluated the economic significance of the EHTS by computing the variance ratio of the actual to theoretical spread and examining the correlations between $S_{(n,m),t}$ and $S'_{(n,m),t}$.

Columns 5 and 6 of table 6 show the results for the variance ratio VR and the correlation coefficient $corr\left(S_{(n,m),t}, S'_{(n,m),t}\right)$ between the actual and the theoretical spread, respectively. As shown in column 5, the variance ratios are not greater than two standard deviations from unity for all the cases, except for both cases of Greece. This implies that in 12 out of the 13 EU countries of table 6, the deviations from the EHTS are not economically or statistically significant. Also, for these 12 EU countries the correlation coefficient between $S_{(n,m),t}$ and $S'_{(n,m),t}$ is high and close to unity for most cases, except for Austria (2-year and 12-month rates) and Italy (10- and 3-year rates, 10-year and 12-month rates). Note that for Greece, for which both variance ratios are greater than two standard deviations from unity, the respective correlation coefficients are below 0.5. Figure 2, which plots $S_{(n,m),t}$ and $S'_{(n,m),t}$ for all pairs of yields of table 6, conveys similar information.

Overall, the results of this section provide strong support for the economic significance of the EHTS. The cross equation restrictions implied by the theory are rejected only in 4 out of 33 cases, while the variance ratios point to excess volatility in the actual spreads relative to the theoretical spreads only in 2 out of 33 cases. Thus, deviations from the theory are economically important only in 2 out of 33 cases of table 6; this is solid evidence in favour of the EHTS of interest rates.

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[8] All estimations of table 6 were performed by Gauss routines. We are grateful to John Y. Campbell, who kindly provided us with the Gauss codes. We modified these codes properly, in order to include the estimated structural breaks into the analysis.

Combining these results with the results of the previous section, they clearly provide support of the empirical adequacy of the EHTS in 12 out of the original 14 EU countries. For six countries (i.e. Belgium, Denmark, France, the Netherlands, Sweden and the United Kingdom) the EHTS holds only when the very short rate (3-month) is included in the analysis, while in the rest of the cases there is no evidence of cointegration. Only for three countries, namely Finland, Ireland and Italy, the EHTS holds for all maturities. For Austria there is evidence in favour of the EHTS in two cases (2-year and 12-month rates, 10- and 5-year rates), while in the other four cases there is no evidence of cointegration. For Germany, the results support the EHTS for three out of six cases, while in the other three cases, where the 10-year rate is included, there is no evidence of spread stationarity. For Spain we find evidence that supports the EHTS in five out of six cases, while in the other case (10-year and 3-month rates) the spread is not stationary.

For Greece and Portugal the results strongly reject the EHTS of interest rates. For Greece, the spreads are stationary in all six cases, but in four of them the unit vector does not belong in the cointegration space. In the other two cases (7- and 5-years rates, 5- and 3-year rates), even though the results are consistent with predictions (i) and (ii) of the EHTS, both variance ratios are greater than two standard deviations from unity and the respective correlation coefficients are below 0.5. Especially for these two cases, the results are also validated from Figure 2, where the actual and the theoretical spread do not seem to move together over time in both cases. Finally, for Portugal, the spreads are stationary in two cases but the unit vector does not span the cointegration space, while in the third case there is no evidence of cointegration.

In general, the above results imply that there is no evidence of excess volatility in 12 out of 14 EU countries, especially at shorter maturities, and deviations from the theory are not economically important.

5. Concluding Remarks

In this paper we investigated empirically the term structure of interest rates of the original 14 EU countries. Since the time span of our sample period covers more than two decades, the existence of structural breaks is quite possible and has been confirmed by the use of the two-break and the one-break minimum LM unit root tests. Our evidence shows that the interest rates of all 14 EU countries are non stationary and allow for two structural breaks or one structural break in the case of Finland. These breaks mainly appear during the period of the

ERM crisis and during the second stage of the EMU, where the European Monetary Institute established and the strengthening of monetary cooperation among potential EMU members were more than necessary. Since the interest rates follow random walks, we evaluated the expectations hypothesis of the term structure using the Johansen et al. and Lütkepohl et al. cointegration techniques in the presence of structural breaks along with the VAR approach of Campbell and Shiller.

In general, our empirical findings are providing support of the empirical adequacy of the EHTS for 12 out of 14 EU countries. For Belgium, Denmark, France, the Netherlands, Sweden and the United Kingdom, the results imply spread stationarity only when the very short rate is included in the analysis, while for Finland, Ireland and Italy the EHTS holds for all maturities. For Austria the EHTS holds for the shorter and the longer maturities, while for Germany the results support the EHTS only when the 10-year rate is excluded from the analysis. Also for Spain, the results show evidence in favour of the EHTS in almost all maturities. Finally, for Greece and the Netherlands, the EHTS is rejected at all maturities.

Overall, the above results are supportive of the economic significance of the expectations theory in all the original 14 EU countries, except Greece and Portugal. For these two countries our evidence implies predictability of excess returns and economically important deviations from the EHTS.

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Table 1
Description of data

Country	Time span	Treasury bill rates/ Government bond yields	Source
Austria	1981:04-2009:12	12-month, 2-year, 5-year, 10-year	Austrian Kontrollbank
Belgium	1987:06-2009:12	3-month, 3-year, 5-year, 10-year	National Bank of Belgium
Denmark	1988:06-2009:12	3-month, 2-year, 5-year, 10-year	National Bank of Denmark
Finland	1991:08-2009:12	3-month, 5-year, 10-year	Bank of Finland
France	1986:07-2009:12	3-month, 2-year, 5-year, 10-year	Bank of France
Germany	1980:01-2009:12	12-month, 2-year, 5-year, 10-year	Bundesbank
Greece	1993:04-2009:12	12-month, 3-year, 5-year, 7-year	Bank of Greece
Ireland	1985:01-2009:12	2-year, 5-year, 10-year	Central Bank and Financial Services Authority of Ireland
Italy	1992:10-2009:12	12-month, 3-year, 5-year, 10-year	Bank of Italy
The Netherlands	1986:04-2009:12	3-month, 10-year	Bank of The Netherlands
Portugal	1993:07-2009:12	2-year, 5-year, 10-year	Bank of Portugal
Spain	1989:07-2009:12	3-month, 3-year, 5-year, 10-year	Bank of Spain
Sweden	1987:01-2009:12	3-month, 2-year, 5-year, 10-year	Central Bank of Sweden
United Kingdom	1985:01-2009:12	3-month, 5-year, 10-year	Bank of England

Table 2
Two-break minimum LM unit root test results

Country	Interest rate	Model	$\hat{k}$	$\hat{T}_B$	$\hat{\lambda}_1$, $\hat{\lambda}_2$	$LM-$statistic
Austria	12-month	C	9	1990:01, 1999:04	0.2, 0.6	-4.7470
	2-year	C	3	1990:01, 1999:04	0.2, 0.6	-4.1976
	5-year	C	1	1989:10, 1999:01	0.2, 0.6	-3.9850
	10-year	C	1	1989:10, 1998:02	0.2, 0.6	-4.5350
Belgium	3-month	C	11	1993:07, 1999:04	0.2, 0.6	-4.7469
	3-year	C	11	1992:09, 1999:03	0.2, 0.6	-4.8741
	5-year	C	11	1992:07, 1999:06	0.2, 0.6	-4.9437
	10-year	C	10	1992:07, 1999:09	0.2, 0.6	-4.7718
Denmark	3-month	C	1	1993:09, 2006:12	0.2, 0.8	-4.2290
	2-year	C	5	1993:03, 2006:07	0.2, 0.8	-4.6987
	5-year	C	4	1993:01, 2006:06	0.2, 0.8	-5.1142
	10-year	C	4	1997:09, 2007:01	0.4, 0.8	-4.7462
Finland	3-month	C	4	1993:09^{n}, 2006:11	0.2, 0.8	-3.6010
	5-year	C	4	1997:08^{n}, 2008:02	0.4, 0.8	-4.6662
	10-year	C	4	1997:06^{n}, 2006:08	0.4, 0.8	-5.0620
France	3-month	C	10	1993:03, 1999:08	0.2, 0.6	-4.7396
	2-year	C	9	1992:06, 1999:03	0.2, 0.6	-4.5979
	5-year	C	9	1992:06, 1999:10	0.2, 0.6	-4.9877
	10-year	C	8	1992:06, 1999:10	0.2, 0.6	-4.9359
Germany	12-month	C	11	1983:12, 1989:07	0.2, 0.4	-4.1827
	2-year	C	11	1984:01, 1989:06	0.2, 0.4	-4.1044
	5-year	C	11	1985:02, 1989:05	0.2, 0.4	-4.2353
	10-year	C	11	1986:04, 1992:12	0.2, 0.4	-4.6400
Greece	12-month	C	2	1997:12, 2003:06	0.2, 0.6	-3.6408
	3-year	C	2	1997:10, 2003:05	0.2, 0.6	-5.2167
	5-year	C	0	1999:06, 2003:06	0.4, 0.6	-5.0379
	7-year	C	2	1998:11, 2006:04	0.4, 0.8	-5.1997
Ireland	2-year	C	9	1989:02, 1993:06	0.2, 0.4	-4.8355
	5-year	C	9	1989:06, 1998:01	0.2, 0.6	-5.1422
	10-year	C	9	1989:06, 1997:09	0.2, 0.6	-4.7635
Italy	12-month	C	6	2000:03, 2005:09	0.4, 0.8	-3.7911
	3-year	C	6	1994:12, 1999:04	0.2, 0.4	-4.8558
	5-year	C	6	1994:10, 1999:01	0.2, 0.4	-4.9470
	10-year	C	4	1994:10, 1998:01	0.2, 0.4	-5.0272
The Netherlands	3-month	C	10	1990:10, 1996:04	0.2, 0.4	-4.5210
	10-year	C	10	1989:11, 1998:03	0.2, 0.6	-5.1370
Portugal	2-year	C	5	1996:09, 2006:05	0.2, 0.8	-4.7518
	5-year	C	7	1996:12, 2006:08	0.2, 0.8	-4.6961
	10-year	C	7	1997:06, 2006:11	0.2, 0.8	-4.8292
Spain	3-month	C	5	1997:07, 2006:06	0.4, 0.8	-5.0746
	3-year	C	3	1997:03, 2006:05	0.4, 0.8	-4.6595
	5-year	C	4	1997:03, 2006:08	0.4, 0.8	-5.0329
	10-year	C	3	1996:07, 1999:10	0.4, 0.6	-5.0431

Table 2 *(continued)*

Country	Interest rate	Model	$\hat{k}$	$\hat{T}_B$	$\hat{\lambda}_1,\ \hat{\lambda}_2$	$LM-$statistic
Sweden	3-month	C	1	1992:06, 1999:09	0.2, 0.6	-4.1460
	2-year	C	6	1989:10, 1999:03	0.2, 0.6	-4.4153
	5-year	C	3	1989:07, 1996:11	0.2, 0.4	-4.6148
	10-year	C	4	1989:07, 1997:11	0.2, 0.4	-4.8919
United Kingdom	3-month	C	11	1988:04, 1993:03	0.2, 0.4	-5.1871
	5-year	C	11	1989:06, 1999:02	0.2, 0.6	-5.2403
	10-year	C	11	1989:06, 1998:05	0.2, 0.6	-5.2268

Break Points	Critical values for Model C		
$\lambda=(\lambda_1, \lambda_2)$	1%	5%	10%
λ=(0.2, 0.4)	-6.16	-5.59	-5.27
λ=(0.2, 0.6)	-6.41	-5.74	-5.32
λ=(0.2, 0.8)	-6.33	-5.71	-5.33
λ=(0.4, 0.6)	-6.45	-5.67	-5.31
λ=(0.4, 0.8)	-6.42	-5.65	-5.32

$\hat{k}$ is the estimated number of lags in the unit root test regression (9) to correct for serial correlation. $\hat{T}_B$ denotes the estimated break points. $\hat{\lambda}_1$ and $\hat{\lambda}_2$ are the estimated relative break points. The critical values are from table 2 of Lee and Strazicich (2003). ^{n} signifies that the relevant break is not significant at the 0.10 level of significance.

Table 3
One-break LM unit root test results

Country	Interest rate	Model	$\hat{k}$	$\hat{T}_B$	$\hat{\lambda}$	$LM-$statistic
Finland	3-month	C	5	1997:06	0.3	-3.4627
	5-year	C	1	1997:01	0.3	-3.8432
	10-year	C	2	1998:03	0.4	-3.9537

Break point	Critical values for Model C		
λ	1%	5%	10%
λ=0.3	-5.15	-4.45	-4.18
λ=0.4	-5.05	-4.50	-4.18

$\hat{k}$ is the estimated number of lags in the unit root test regression (9) to correct for serial correlation. $\hat{T}_B$ denotes the estimated break point. The critical values are from table 1 of Lee and Strazicich (2004).

Table 4
Testing for cointegration: JMN and LST tests

VECM for Y_t:	Break Points	$(p-r_0)$	$LR_{JMN}(r_0)$	$LR_{LST}(r_0)$	p-values		$\hat{k}$	$H_0:(a,b)'=(1,-1)'$	
					JMN	LST		LR	p-value
				Austria					
10y, 5y	1989:10	2	51.53**	21.19*	0.024	0.077	1	2.67	0.102
	1998:02	1	17.94	4.71	0.293	0.566			
10y, 2y	1989:10	2	34.86	12.69	0.476	0.581	1	NA	NA
	1998:02	1	8.72	4.55	0.932	0.591			
10y, 12m	1989:10	2	33.52	11.56	0.548	0.682	1	NA	NA
	1998:02	1	7.87	3.64	0.959	0.732			
5y, 2y	1989:10	2	43.02	14.07	0.152	0.462	12	NA	NA
	1999:01	1	13.02	1.58	0.678	0.971			
5y, 12m	1989:10	2	33.62	10.81	0.557	0.748	1	NA	NA
	1999:01	1	9.23	3.35	0.919	0.783			
2y, 12m	1990:01	2	51.44**	21.31*	0.026	0.076	1	2.39	0.122
	1999:04	1	14.17	3.12	0.588	0.817			
				Belgium					
10y, 5y	1992:07	2	34.77	16.67	0.445	0.256	6	NA	NA
	1999:09	1	12.77	3.32	0.647	0.758			
10y, 3y	1992:07	2	35.47	16.91	0.408	0.242	6	NA	NA
	1999:09	1	13.61	2.80	0.578	0.832			
10y, 3m	1992:07	2	31.97	15.26	0.591	0.352	6	NA	NA
	1999:09	1	14.35	3.98	0.519	0.656			
5y, 3y	1992:07	2	26.13	13.17	0.857	0.525	2	NA	NA
	1999:06	1	9.54	3.30	0.867	0.757			
5y, 3m	1992:07	2	45.93*	22.59**	0.070	0.047	11	0.01	0.916
	1999:06	1	12.37	3.22	0.670	0.768			
3y, 3m	1992:09	2	53.74**	23.57**	0.011	0.033	10	2.54	0.111
	1999:03	1	14.94	4.24	0.462	0.609			
				Denmark					
10y, 5y	1997:09	2	33.63	15.27	0.415	0.314	5	NA	NA
	2007:01	1	7.66	4.52	0.928	0.518			
10y, 2y	1997:09	2	32.61	15.88	0.467	0.273	5	NA	NA
	2007:01	1	8.01	4.26	0.913	0.557			
10y, 3m	1997:09	2	47.85**	22.12**	0.032	0.049	11	1.29	0.256
	2007:01	1	14.49	3.02	0.437	0.752			
5y, 2y	1993:01	2	36.63	17.62	0.213	0.165	4	NA	NA
	2006:06	1	10.65	0.86	0.664	0.976			
5y, 3m	1993:01	2	32.29	13.64	0.401	0.424	5	NA	NA
	2006:06	1	11.81	1.28	0.569	0.942			
2y, 3m	1993:03	2	40.99*	15.89	0.098	0.258	3	NA	NA
	2006:07	1	12.81	6.30	0.491	0.263			
				Finland					
10y, 5y	1998:03	2	38.01**	28.35**	0.037	0.001	5	0.22	0.636
		1	9.65	3.20	0.544	0.565			
10y, 3m	1998:03	2	46.15**	23.62**	0.004	0.009	9	1.27	0.259
		1	15.18	2.86	0.146	0.628			
5y, 3m	1997:01	2	38.74**	24.82**	0.024	0.005	4	0.23	0.633
		1	10.49	3.99	0.422	0.409			

Table 4 *(continued)*

VECM for Y_t:	Break Points	$(p-r_0)$	$LR_{JMN}(r_0)$	$LR_{LST}(r_0)$	p-values		$\hat{k}$	$H_0:(a,b)'=(1,-1)'$	
					JMN	LST		*LR*	p-value
				France					
10y, 5y	1992:06	2	40.66	15.21	0.208	0.362	4	NA	NA
	1999:10	1	12.23	0.90	0.709	0.994			
10y, 2y	1992:06	2	37.59	13.32	0.329	0.520	3	NA	NA
	1999:10	1	11.65	1.70	0.753	0.958			
10y, 3m	1992:06	2	41.90	13.86	0.169	0.472	6	NA	NA
	1999:10	1	15.36	5.32	0.462	0.468			
5y, 2y	1992:06	2	36.71	12.16	0.370	0.624	4	NA	NA
	1999:10	1	9.74	3.52	0.875	0.740			
5y, 3m	1992:06	2	45.85*	13.86	0.081	0.472	4	NA	NA
	1999:10	1	13.66	0.93	0.596	0.993			
2y, 3m	1992:06	2	53.75**	22.90**	0.012	0.043	4	0.31	0.577
	1999:03	1	18.25	0.59	0.257	0.998			
				Germany					
10y, 5y	1986:04	2	33.70	14.16	0.399	0.411	3	NA	NA
	1992:12	1	14.96	2.32	0.383	0.846			
10y, 2y	1986:04	2	31.88	13.34	0.494	0.481	3	NA	NA
	1992:12	1	13.86	2.08	0.464	0.877			
10y, 12m	1986:04	2	33.12	12.24	0.428	0.579	3	NA	NA
	1992:12	1	15.13	3.88	0.372	0.607			
5y, 2y	1985:02	2	44.22**	21.21**	0.031	0.047	5	0.46	0.499
	1989:05	1	14.83	7.85	0.276	0.128			
5y, 12m	1985:02	2	45.12**	21.23**	0.024	0.047	3	0.00	0.988
	1989:05	1	14.16	3.22	0.314	0.608			
2y, 12m	1984:01	2	52.94**	23.18**	0.003	0.024	8	1.76	0.185
	1989:06	1	16.56	4.84	0.187	0.377			
				Greece					
7y, 5y	1998:11	2	48.68**	22.16*	0.042	0.054	6	0.39	0.531
	2006:04	1	20.64	4.63	0.151	0.556			
7y, 3y	1998:11	2	67.64**	20.48*	0.000	0.091	3	8.95**	0.003
	2006:04	1	14.88	2.22	0.485	0.904			
7y, 12m	1998:11	2	68.59**	20.36*	0.000	0.094	6	18.40**	0.000
	2006:04	1	15.63	2.04	0.247	0.923			
5y, 3y	1999:06	2	67.73**	43.17**	0.000	0.000	3	0.65	0.421
	2003:06	1	15.27	6.10	0.476	0.367			
5y, 12m	1999:06	2	60.81**	27.89**	0.002	0.008	6	20.96**	0.000
	2003:06	1	12.12	2.87	0.725	0.837			
3y, 12m	1997:10	2	51.32**	39.17**	0.026	0.000	2	7.88**	0.005
	2003:05	1	11.20	6.27	0.803	0.354			

Table 4 *(continued)*

VECM for Y_t:	Break Points	$(p-r_0)$	$LR_{JMN}(r_0)$	$LR_{LST}(r_0)$	p-values JMN	p-values LST	$\hat{k}$	$H_0:(a,b)'=(1,-1)'$ LR	$H_0:(a,b)'=(1,-1)'$ p-value
				Ireland					
10y, 5y	1989:06	2	51.01**	21.80*	0.018	0.057	3	1.73	0.188
	1997:09	1	13.48	1.54	0.545	0.952			
10y, 2y	1989:06	2	46.53*	20.21*	0.051	0.093	2	0.36	0.550
	1997:09	1	10.59	5.83	0.771	0.367			
5y, 2y	1989:06	2	48.00**	20.16*	0.038	0.094	2	0.32	0.574
	1998:01	1	10.28	3.08	0.798	0.768			
				Italy					
10y, 5y	1994:10	2	44.39**	26.26**	0.026	0.007	10	2.42	0.120
	1998:01	1	12.37	0.39	0.433	0.992			
10y, 3y	1994:10	2	54.58**	26.78**	0.002	0.006	12	0.83	0.364
	1998:01	1	17.18	0.01	0.153	0.999			
10y, 12m	1994:10	2	47.87**	23.11**	0.011	0.024	12	0.68	0.410
	1998:01	1	15.47	0.01	0.228	0.999			
5y, 3y	1994:10	2	47.32**	21.12*	0.018	0.053	10	2.40	0.122
	1999:01	1	10.46	0.27	0.645	0.998			
5y, 12m	1994:10	2	50.72**	20.22*	0.008	0.071	3	0.86	0.355
	1999:01	1	9.79	1.88	0.705	0.857			
3y, 12m	1994:12	2	53.07**	21.20*	0.005	0.054	3	0.19	0.659
	1999:04	1	9.16	2.41	0.770	0.788			
				The Netherlands					
10y, 3m	1989:11	2	50.10**	21.73*	0.019	0.055	10	0.02	0.898
	1998:03	1	15.48	4.25	0.371	0.566			
				Portugal					
10y, 5y	1997:06	2	43.95*	22.31**	0.070	0.044	12	26.17**	0.000
	2006:11	1	13.42	2.85	0.497	0.763			
10y, 2y	1997:06	2	43.08*	10.00	0.082	0.771	10	NA	NA
	2006:11	1	11.79	0.98	0.624	0.978			
5y, 2y	1996:12	2	55.00**	20.03*	0.004	0.089	1	6.32**	0.012
	2006:08	1	20.13	2.20	0.116	0.850			
				Spain					
10y, 5y	1996:07	2	52.30**	23.32**	0.012	0.033	8	1.69	0.194
	1999:10	1	20.51	3.18	0.124	0.736			
10y, 3y	1996:07	2	44.61*	21.18*	0.070	0.066	8	2.43	0.119
	1999:10	1	14.04	2.69	0.482	0.811			
10y, 3m	1996:07	2	31.48	18.46	0.543	0.147	3	NA	NA
	1999:10	1	15.54	3.27	0.372	0.721			
5y, 3y	1997:03	2	58.12**	31.17**	0.003	0.002	1	0.69	0.406
	2006:08	1	11.97	3.16	0.670	0.749			
5y, 3m	1997:03	2	54.92**	23.64**	0.007	0.031	5	2.51	0.113
	2006:08	1	19.57	0.54	0.166	0.998			
3y, 3m	1997:03	2	53.10**	22.44**	0.012	0.047	7	2.24	0.134
	2006:05	1	18.98	0.53	0.199	0.998			

Table 4 *(continued)*

VECM for Y_t:	Break Points	$(p-r_0)$	$LR_{JMN}(r_0)$	$LR_{LST}(r_0)$	p-values		$\hat{k}$	$H_0:(a,b)'=(1,-1)'$	
					JMN	LST		LR	p-value
				Sweden					
10y, 5y	1989:07	2	34.74	18.62	0.305	0.123	2	NA	NA
	1997:11	1	13.63	6.24	0.457	0.278			
10y, 2y	1989:07	2	36.59	15.83	0.230	0.256	3	NA	NA
	1997:11	1	15.94	4.76	0.294	0.455			
10y, 3m	1989:07	2	52.37**	22.17**	0.007	0.042	9	0.84	0.358
	1997:11	1	19.44	6.68	0.131	0.237			
5y, 2y	1989:07	2	41.91*	17.09	0.079	0.183	3	NA	NA
	1996:11	1	17.66	2.12	0.190	0.848			
5y, 3m	1989:07	2	44.77**	24.89**	0.043	0.016	5	0.01	0.933
	1996:11	1	15.81	2.42	0.287	0.804			
2y, 3m	1989:10	2	57.85**	19.85*	0.002	0.092	8	1.97	0.160
	1999:03	1	18.80	2.00	0.166	0.890			
				United Kingdom					
10y, 5y	1989:06	2	42.83	17.36	0.114	0.205	2	NA	NA
	1998:05	1	12.22	3.42	0.66	0.719			
10y, 3m	1989:06	2	43.88*	13.02	0.093	0.519	2	NA	NA
	1998:05	1	12.51	2.04	0.633	0.909			
5y, 3m	1989:06	2	58.51**	22.08*	0.003	0.054	1	1.15	0.284
	1999:02	1	16.52	1.46	0.334	0.964			

y stands for years and m stands for months. $\hat{k}$ is the estimated lag length in each VECM. LR is likelihood ratio test statistic for H_0 . ** and * denote rejection of the null hypothesis at the 0.05 and the 0.10 level of significance, respectively. NA stands for "Not Applicable".

Table 5

VAR Model for $(S_{(n,m),t}, \Delta r_{m,t})$: Granger causality tests

VAR for Y_t:	Break Points	$\hat{k}$	Wald statistics for Granger non-causality		*LM* test
			$S_{(n,m),t}$ to $\Delta r_{m,t}$	$\Delta r_{m,t}$ to $S_{(n,m),t}$	
			Austria		
10y, 5y	1989:10, 1998:02	2	3.41* (0.065)	0.90 (0.639)	3.07 (0.547)
2y, 12m	1990:01, 1999:04	9	32.51** (0.000)	7.04 (0.633)	4.19 (0.380)
			Belgium		
5y, 3m	1992:07, 1999:06	6	34.18** (0.000)	12.77** (0.047)	3.51 (0.477)
3y, 3m	1992:09, 1999:03	6	38.44** (0.000)	13.17** (0.040)	3.11 (0.540)
			Denmark		
10y, 3m	1997:09, 2007:01	11	30.67** (0.001)	28.31** (0.003)	4.19 (0.380)
			Finland		
10y, 5y	1998:03	12	21.10** (0.049)	30.94** (0.002)	1.05 (0.902)
10y, 3m	1998:03	10	19.64** (0.033)	37.97** (0.000)	5.38 (0.250)
5y, 3m	1997:01	8	15.34* (0.053)	10.94 (0.205)	3.18 (0.529)
			France		
2y, 3m	1992:06, 1999:03	6	26.71** (0.000)	31.12** (0.000)	2.90 (0.575)
			Germany		
5y, 2y	1985:02, 1989:05	11	35.58** (0.000)	33.43** (0.000)	1.68 (0.795)
5y, 12m	1985:02, 1989:05	12	76.95** (0.000)	26.76** (0.008)	3.42 (0.491)
2y, 12m	1984:01, 1989:06	11	108.68*** (0.000)	23.54** (0.015)	1.91 (0.753)
			Greece		
7y, 5y	1998:11, 2006:04	11	56.37** (0.000)	11.93 (0.369)	4.42 (0.352)
5y, 3y	1999:06, 2003:06	9	9.83 (0.365)	28.93** (0.001)	7.03 (0.134)
			Ireland		
10y, 5y	1989:06, 1997:09	2	15.62** (0.000)	22.68** (0.000)	0.40 (0.982)
10y, 2y	1989:06, 1997:09	2	23.84** (0.000)	18.13** (0.000)	4.95 (0.292)
5y, 2y	1989:06, 1998:01	2	22.22** (0.000)	10.85** (0.004)	6.58 (0.160)
			Italy		
10y, 5y	1994:10, 1998:01	7	15.40** (0.031)	11.88 (0.105)	1.51 (0.824)
10y, 3y	1994:10, 1998:01	9	15.50* (0.077)	8.06 (0.528)	2.86 (0.582)
10y, 12m	1994:10, 1998:01	5	29.76** (0.000)	5.93 (0.313)	3.79 (0.435)
5y, 3y	1994:10, 1999:01	3	9.26** (0.026)	2.25 (0.523)	6.39 (0.172)
5y, 12m	1994:10, 1999:01	2	36.20** (0.000)	4.50 (0.105)	4.65 (0.325)
3y, 12m	1994:12, 1999:04	2	44.30** (0.000)	1.92 (0.384)	6.53 (0.163)
			The Netherlands		
10y, 3m	1989:11, 1998:03	7	13.11* (0.069)	20.93** (0.004)	2.60 (0.628)
			Spain		
10y, 5y	1996:07, 1999:10	9	18.73** (0.028)	8.16 (0.518)	4.22 (0.377)
10y, 3y	1996:07, 1999:10	9	14.90* (0.094)	12.09 (0.208)	5.87 (0.209)
5y, 3y	1997:03, 2006:08	9	16.10* (0.065)	25.91** (0.002)	2.45 (0.653)
5y, 3m	1997:03, 2006:08	10	66.80** (0.000)	8.72 (0.559)	4.18 (0.383)
3y, 3m	1997:03, 2006:05	10	96.52** (0.000)	12.04 (0.282)	3.96 (0.411)
			Sweden		
10y, 3m	1989:07, 1997:11	10	20.85** (0.022)	12.85 (0.232)	3.87 (0.424)
5y, 3m	1989:07, 1996:11	10	39.57** (0.000)	12.74 (0.238)	1.86 (0.761)
2y, 3m	1989:10, 1999:03	10	66.76** (0.000)	11.72 (0.304)	0.80 (0.939)

Table 5 *(continued)*

VAR for Y_t:	Break Points	$\hat{k}$	Wald statistics for Granger non-causality		LM test
			$S_{(n,m),t}$ to $\Delta r_{m,t}$	$\Delta r_{m,t}$ to $S_{(n,m),t}$	
			United Kingdom		
5y, 3m	1989:06, 1999:02	11	23.47** (0.015)	42.79** (0.000)	4.57 (0.334)

$\hat{k}$ is the estimated lag length in each VAR. 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. **, * denote rejection of the null hypothesis at the 0.05 and 0.10 level of significance, respectively.

Table 6
VAR Results: Wald tests for cross-equation restrictions, variance ratios and correlation coefficients of actual and theoretical spreads

VAR for Y_t:	Break Points	Wald tests: Test statistic	df	VR	$corr\left(S_{(n,m),t}, S'_{(n,m),t}\right)$
		Austria			
10y, 5y	1989:10, 1998:02	4.56 (0.336)	4	0.456 (0.623)	0.974 (0.084)
2y, 12m	1990:01, 1999:04	25.30 (0.117)	18	0.597 (0.254)	0.448 (0.655)
		Belgium			
5y, 3m	1992:07, 1999:06	9.15 (0.690)	12	0.551 (1.480)	0.956 (0.433)
3y, 3m	1992:09, 1999:03	11.60 (0.476)	12	0.640 (1.110)	0.952 (0.287)
		Denmark			
10y, 3m	1997:09, 2007:01	3.29 (0.999)	22	0.728 (4.070)	0.976 (0.610)
		Finland			
10y, 5y	1998:03	80.20** (0.000)	24	1.010 (0.504)	0.816 (0.186)
10y, 3m	1998:03	12.50 (0.896)	20	1.010 (2.490)	0.973 (0.292)
5y, 3m	1997:01	14.10 (0.592)	16	0.828 (1.420)	0.971 (0.097)
		France			
2y, 3m	1992:06, 1999:03	6.92 (0.863)	12	0.905 (0.922)	0.995 (0.020)
		Germany			
5y, 2y	1985:02, 1989:05	22.00 (0.460)	22	0.916 (1.400)	0.954 (0.213)
5y, 12m	1985:02, 1989:05	29.10 (0.215)	24	0.837 (1.840)	0.955 (0.435)
2y, 12m	1984:01, 1989:06	58.20** (0.000)	22	0.595 (0.379)	0.753 (0.628)
		Greece			
7y, 5y	1998:11, 2006:04	180.00** (0.000)	22	0.514† (0.207)	0.478 (1.020)
5y, 3y	1999:06, 2003:06	3590.00** (0.000)	18	0.194† (0.353)	0.479 (0.720)
		Ireland			
10y, 5y	1989:06, 1997:09	1.25 (0.871)	4	0.928 (0.801)	0.999 (0.006)
10y, 2y	1989:06, 1997:09	3.13 (0.537)	4	0.950 (1.230)	0.998 (0.011)
5y, 2y	1989:06, 1998:01	2.82 (0.588)	4	0.722 (0.578)	0.993 (0.028)
		Italy			
10y, 5y	1994:10, 1998:01	4.12 (0.995)	14	2.160 (2.650)	0.766 (0.415)
10y, 3y	1994:10, 1998:01	10.80 (0.905)	18	1.010 (1.500)	0.622 (1.120)
10y, 12m	1994:10, 1998:01	3.35 (0.972)	10	0.839 (1.720)	0.556 (3.740)
5y, 3y	1994:10, 1999:01	6.35 (0.385)	6	1.260 (1.350)	0.845 (0.249)
5y, 12m	1994:10, 1999:01	2.16 (0.706)	4	1.190 (1.280)	0.997 (0.014)
3y, 12m	1994:12, 1999:04	0.71 (0.950)	4	1.070 (0.680)	0.998 (0.006)
		The Netherlands			
10y, 3m	1989:11, 1998:03	6.56 (0.950)	14	0.824 (1.810)	0.986 (0.222)
		Spain			
10y, 5y	1996:07, 1999:10	22.70 (0.203)	18	1.040 (1.420)	0.935 (0.123)
10y, 3y	1996:07, 1999:10	18.50 (0.421)	18	1.280 (1.930)	0.956 (0.107)
5y, 3y	1997:03, 2006:08	19.10 (0.385)	18	1.000 (1.360)	0.872 (0.356)
5y, 3m	1997:03, 2006:08	24.60 (0.219)	20	0.667 (0.948)	0.785 (0.816)
3y, 3m	1997:03, 2006:05	22.90 (0.293)	20	0.714 (0.645)	0.700 (0.749)
		Sweden			
10y, 3m	1989:07, 1997:11	10.20 (0.965)	20	0.882 (2.040)	0.985 (0.077)
5y, 3m	1989:07, 1996:11	11.80 (0.923)	20	0.932 (1.520)	0.978 (0.109)
2y, 3m	1989:10, 1999:03	18.10 (0.583)	20	1.100 (0.809)	0.952 (0.185)

Table 6 *(continued)*

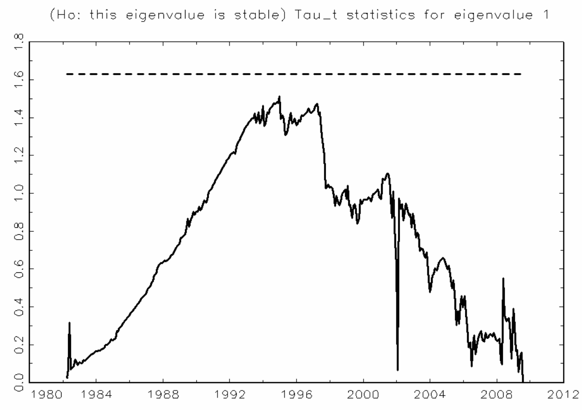
VAR for Y_t:	Break Points	Wald tests		VR	$corr\left(S_{(n,m),t}, S'_{(n,m),t}\right)$
		Test statistic	df		
		United Kingdom			
5y, 3m	1989:06, 1999:02	12.60 (0.943)	22	0.952 (1.650)	0.968 (0.203)

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 $x_t \equiv \left(\Delta r_{m,t}, S_{(n,m),t}\right)'$. Numbers in parentheses in column 3 are p-values. Numbers in parentheses in columns 5 and 6 are standard errors. ** and * denote rejection of the null hypothesis at the 0.05 and 0.10 level of significance, respectively. † indicates a variance ratio that is greater than two standard deviations from unity.

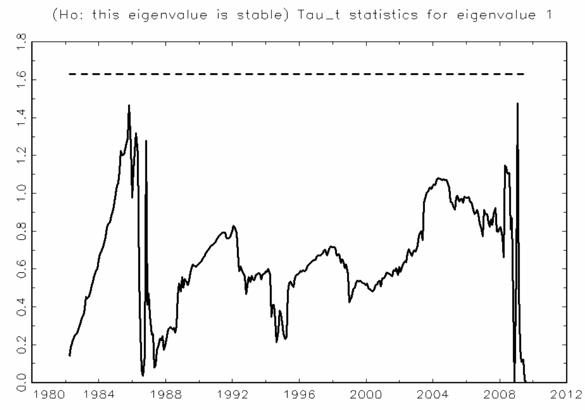
Figure 1

Parameter constancy tests

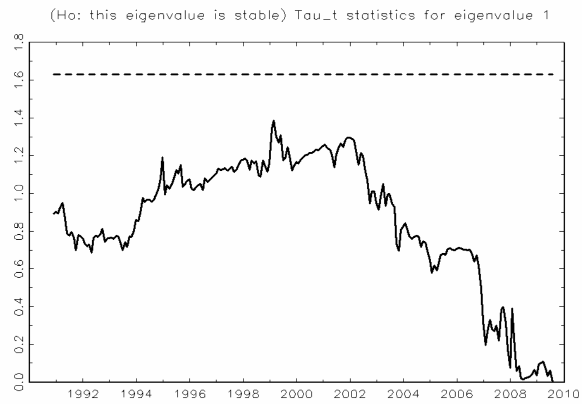
Austria: 10 year – 5 year VECM



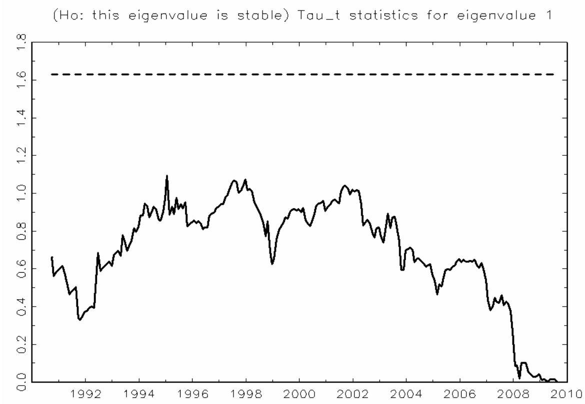
Austria: 2 year – 12 month VECM



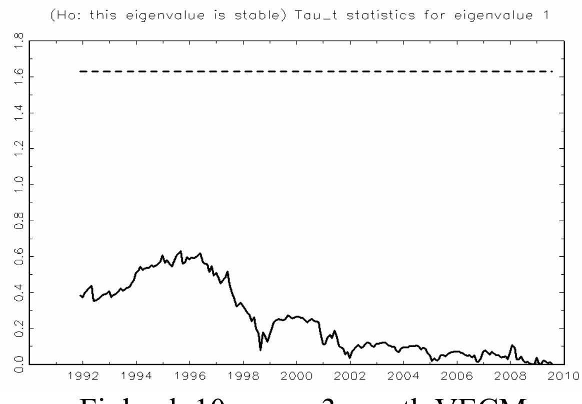
Belgium: 5 year – 3 month VECM



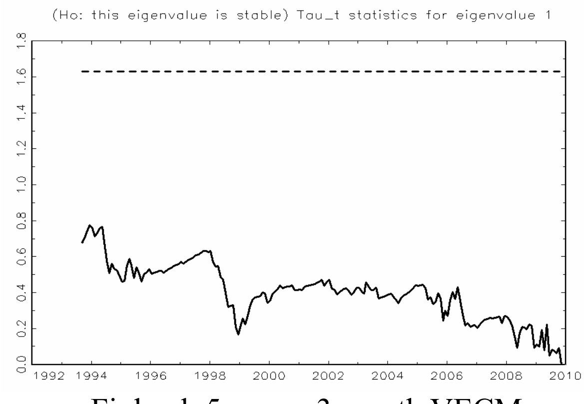
Belgium: 3 year – 3 month VECM



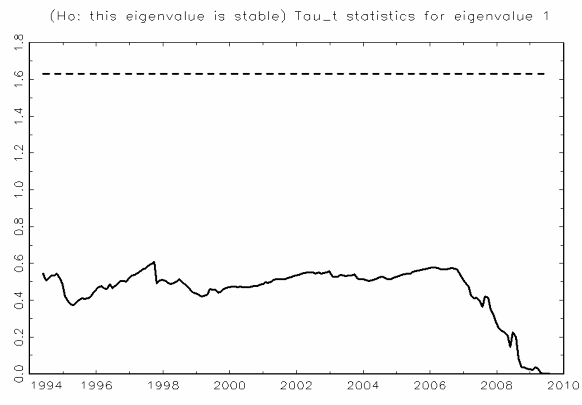
Denmark: 10 year – 3 month VECM



Finland: 10 year – 5 year VECM



Finland: 10 year – 3 month VECM



Finland: 5 year – 3 month VECM

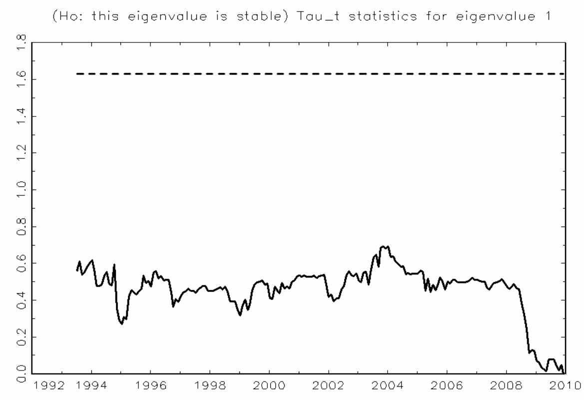
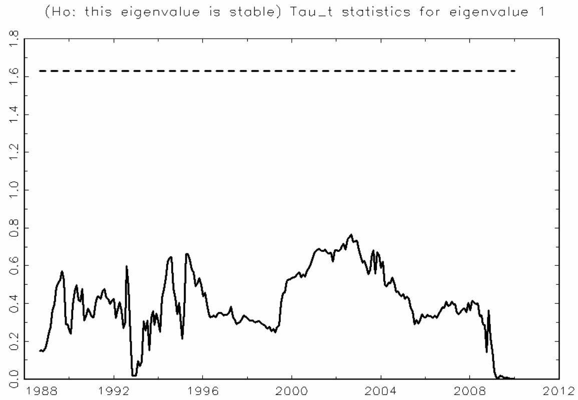
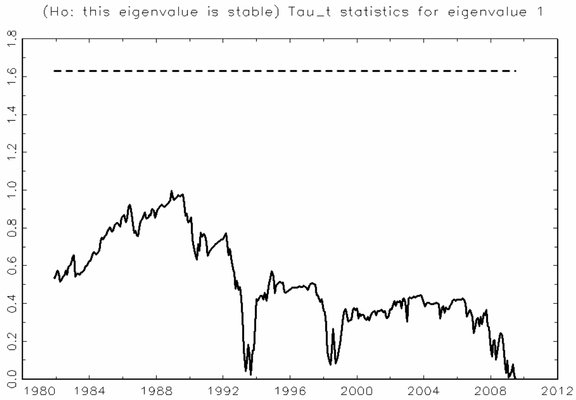


Figure 1 *(continued)*

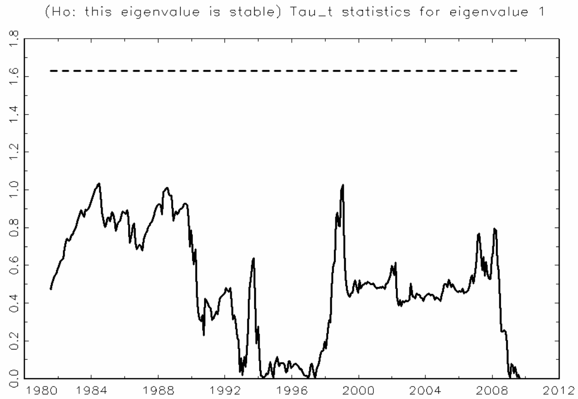
France: 2 year – 3 month VECM



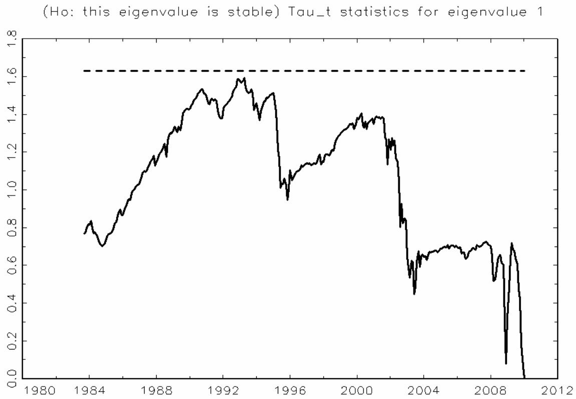
Germany: 5 year – 2 year VECM



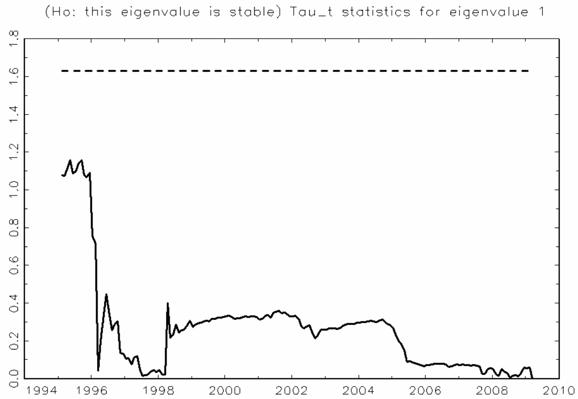
Germany: 5 year – 12 month VECM



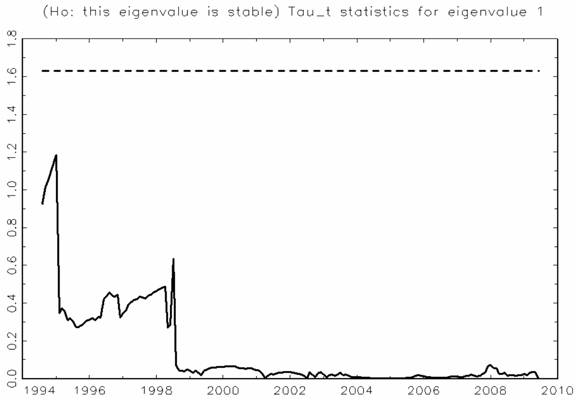
Germany: 2 year – 12 month VECM



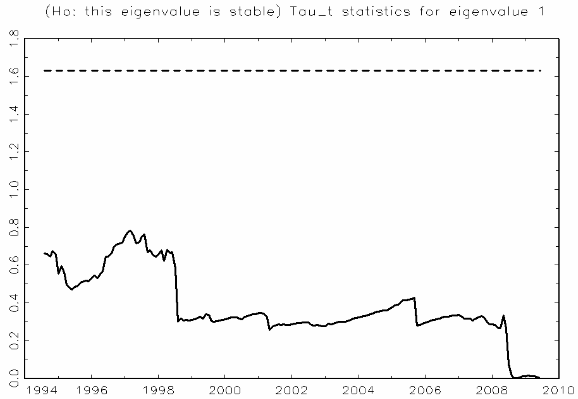
Greece: 7 year – 5 year VECM



Greece: 7 year – 3 year VECM



Greece: 7 year – 12 month VECM



Greece: 5 year – 3 year VECM

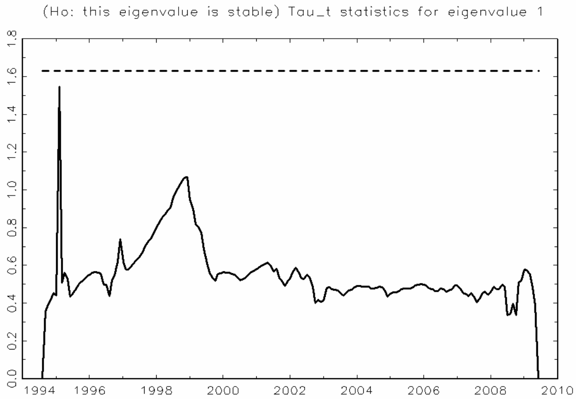
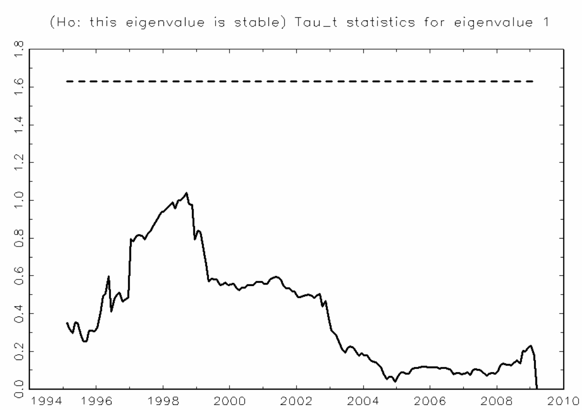
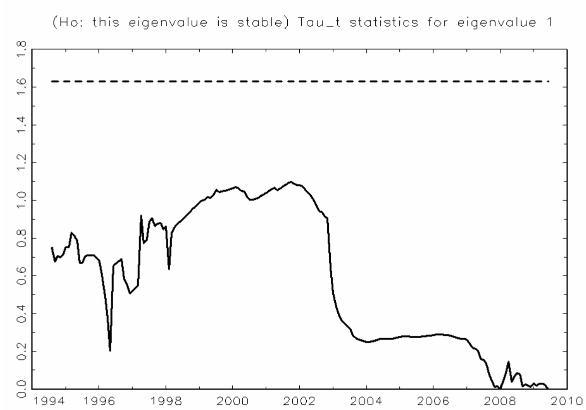


Figure 1 *(continued)*

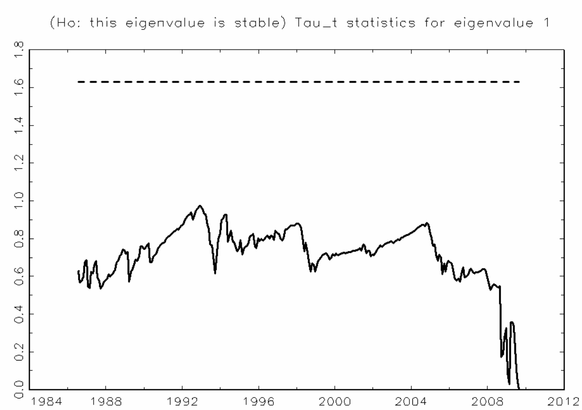
Greece: 5 year – 12 month VECM



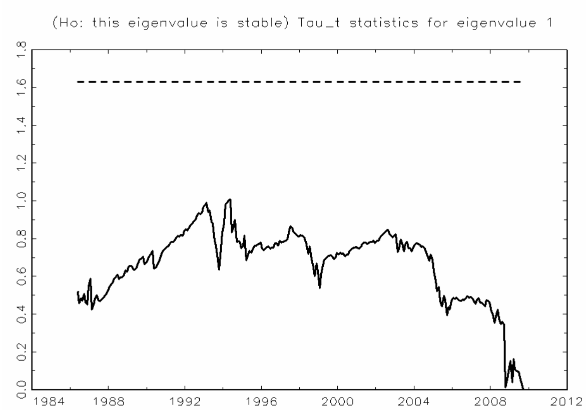
Greece: 3 year – 12 month VECM



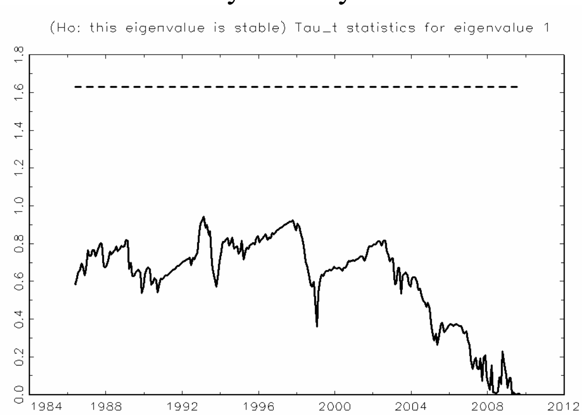
Ireland: 10 year – 5 year VECM



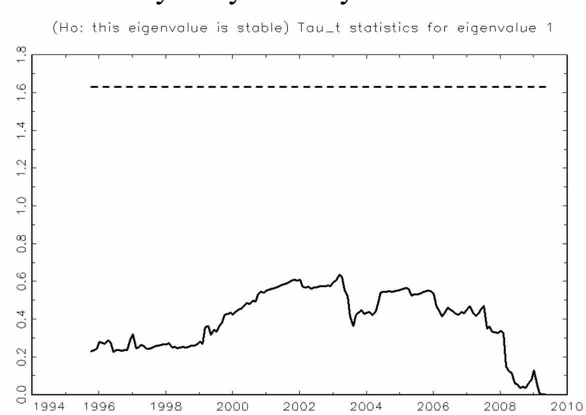
Ireland: 10 year – 2 year VECM



Ireland: 5 year – 2 year VECM



Italy: 10 year – 5 year VECM

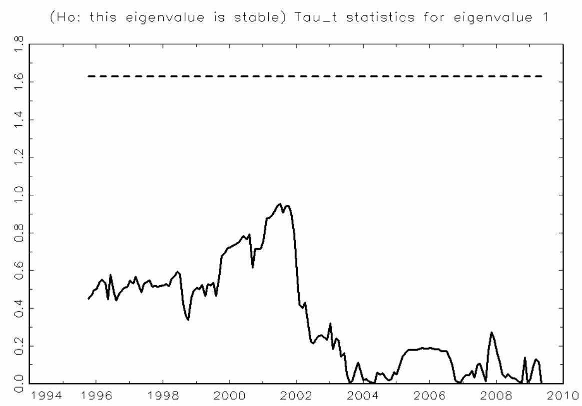


Italy: 10 year – 3 year VECM

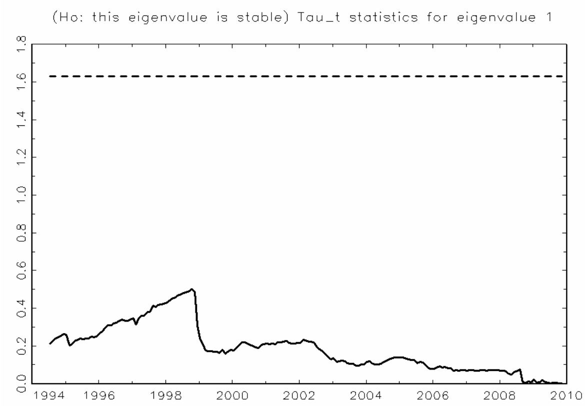
Italy: 10 year – 12 month VECM

Figure 1 *(continued)*

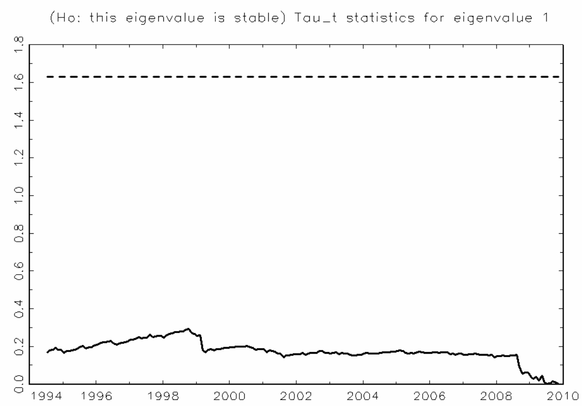
Italy: 5 year – 3 year VECM



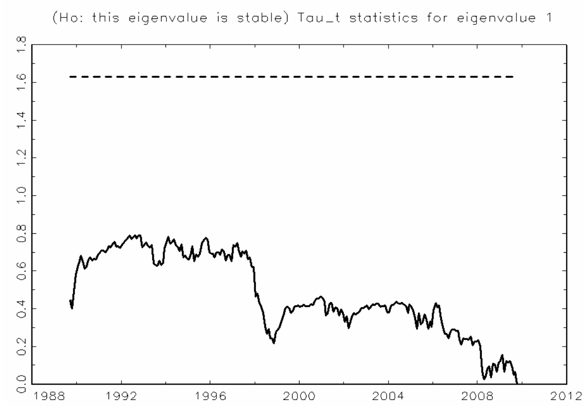
Italy: 5 year – 12 month VECM



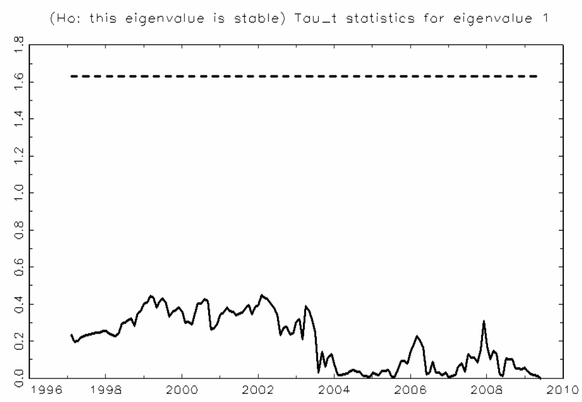
Italy: 3 year – 12 month VECM



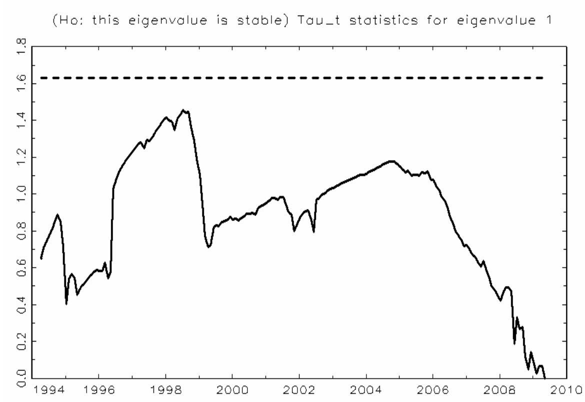
The Netherlands: 10 year – 3 month VECM



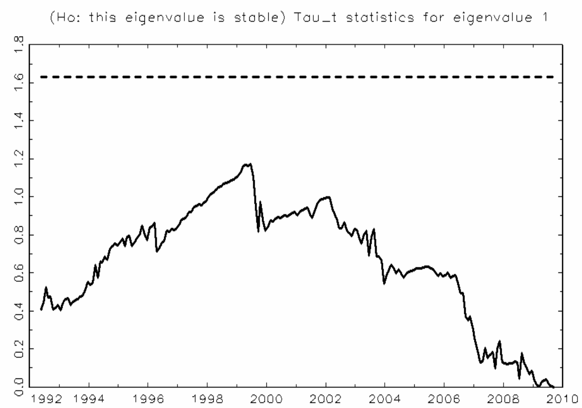
Portugal: 10 year – 5 year VECM



Portugal: 5 year – 2 year VECM



Spain: 10 year – 5 year VECM



Spain: 10 year – 3 year VECM

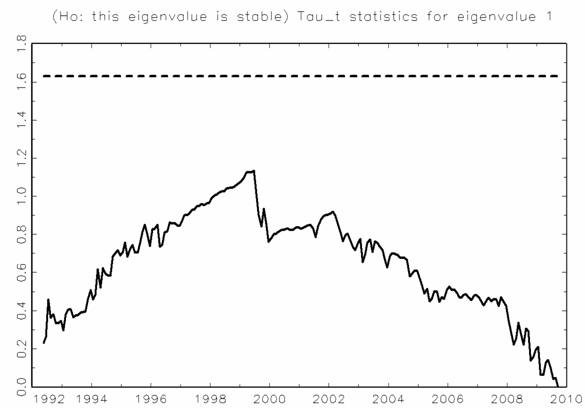
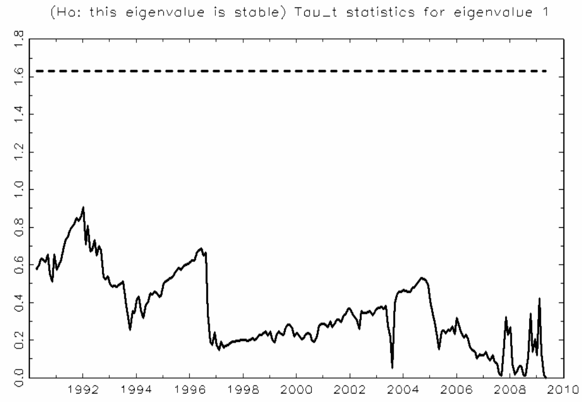
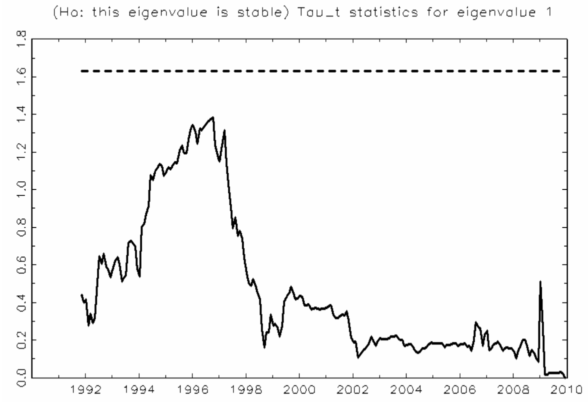


Figure 1 *(continued)*

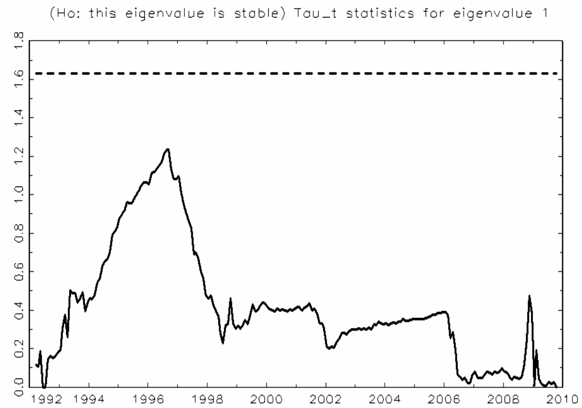
Spain: 5 year – 3 year VECM



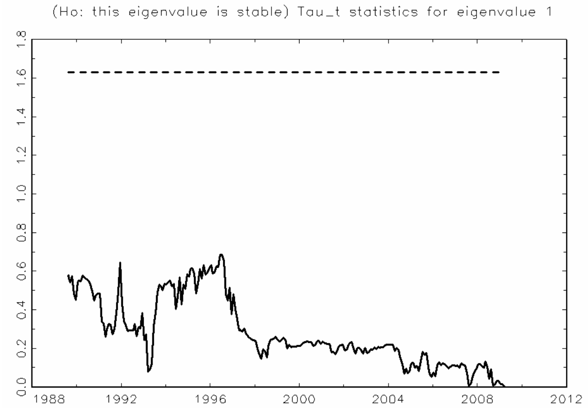
Spain: 5 year – 3 month VECM



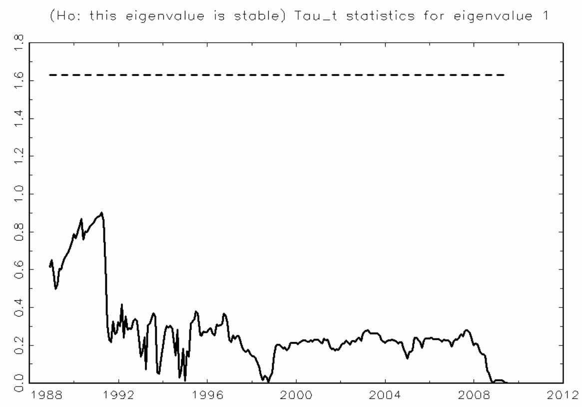
Spain: 3 year – 3 month VECM



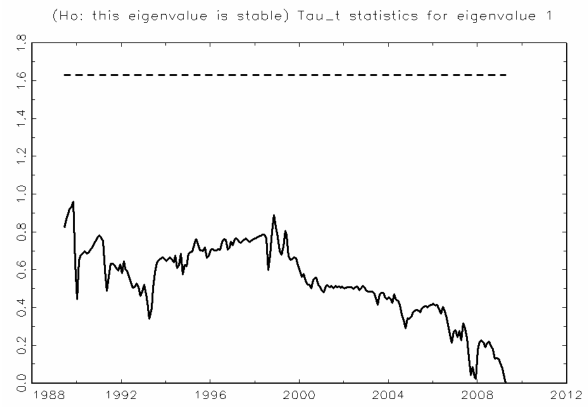
Sweden: 10 year – 3 month VECM



Sweden: 5 year – 3 month VECM



Sweden: 2 year – 3 month VECM



United Kingdom: 5 year – 3 month VECM

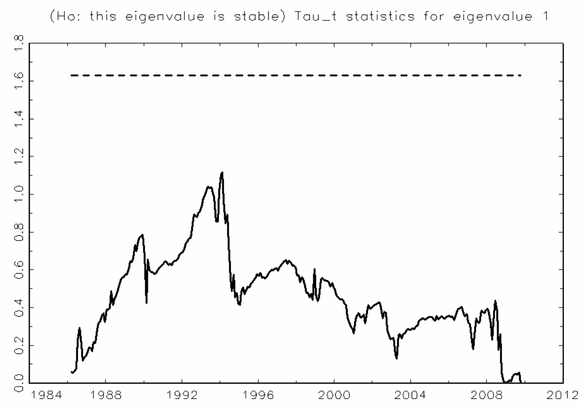


Figure 2

Term structure: deviations from means of actual spread $\left(S_{(n,m),t}\right)$ and theoretical spread $\left(S^{'}_{(n,m),t}\right)$

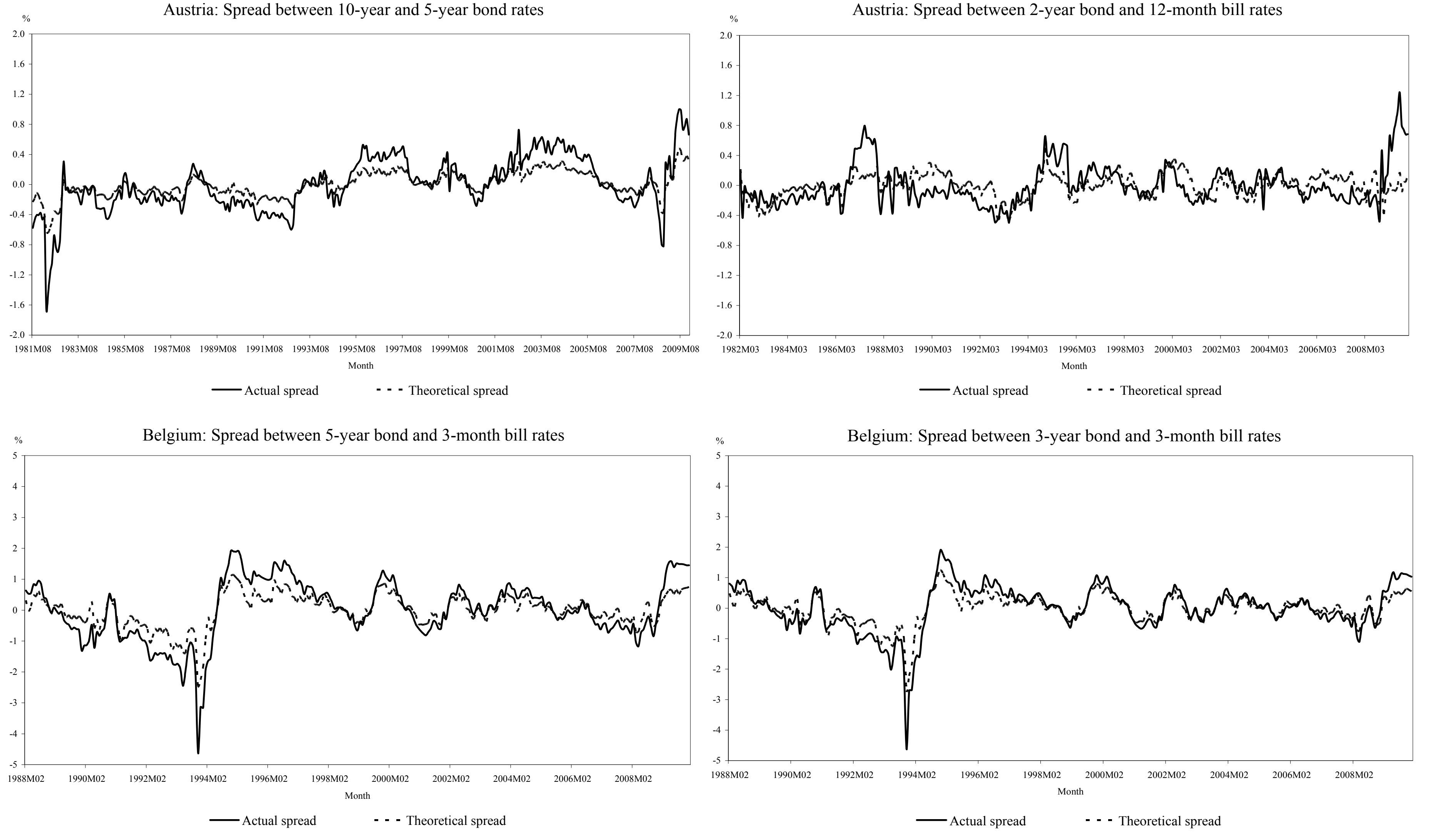
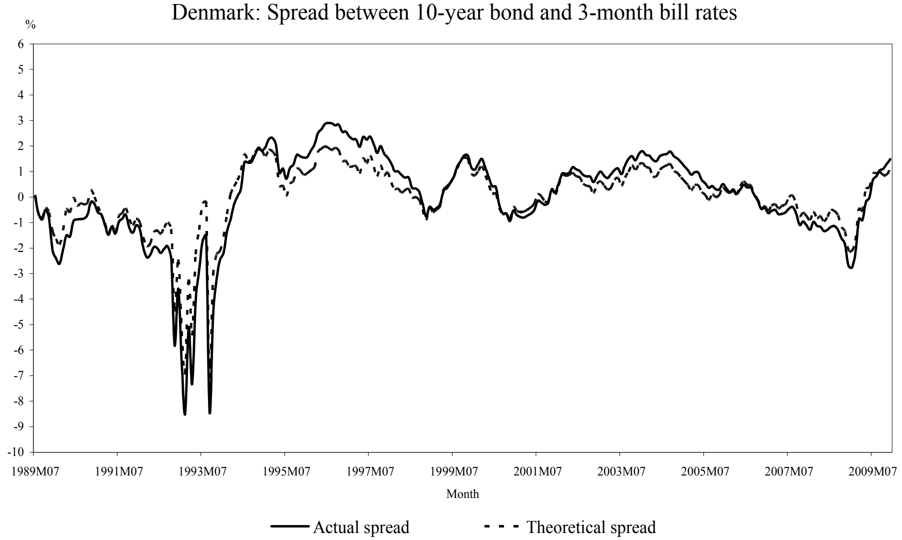
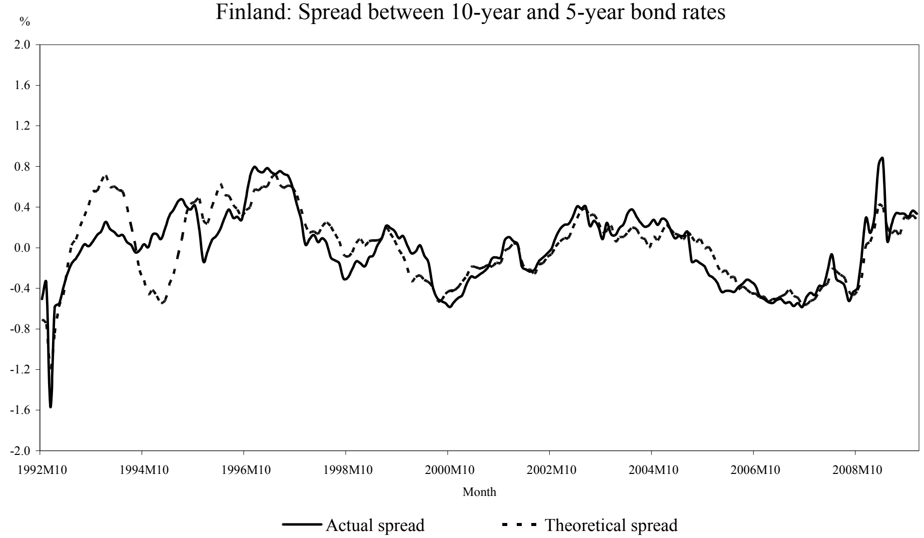


Figure 2 *(continued)*

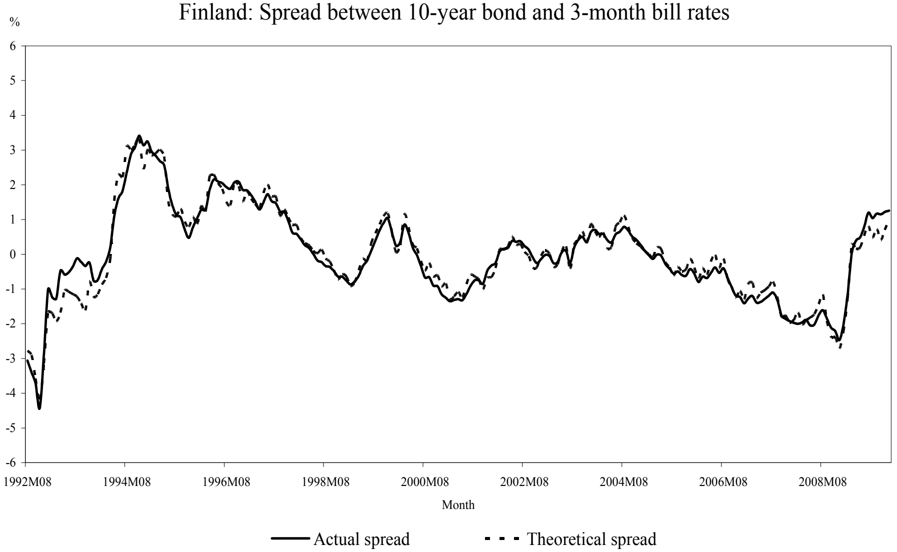
Denmark: Spread between 10-year bond and 3-month bill rates



Finland: Spread between 10-year and 5-year bond rates



Finland: Spread between 10-year bond and 3-month bill rates



Finland: Spread between 5-year bond and 3-month bill rates

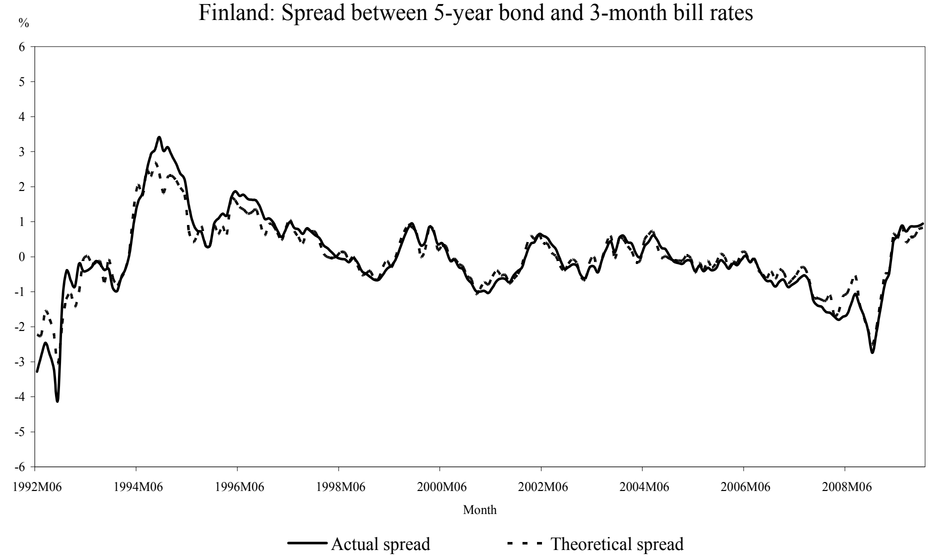
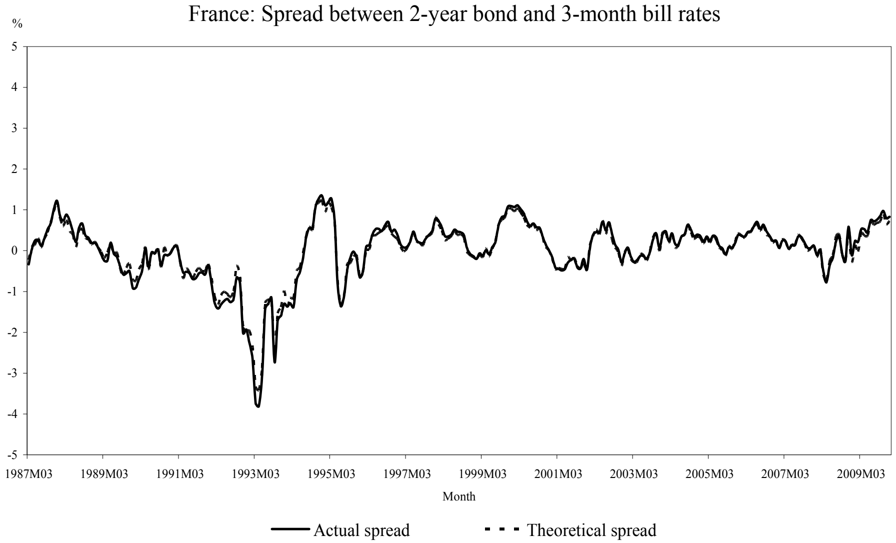
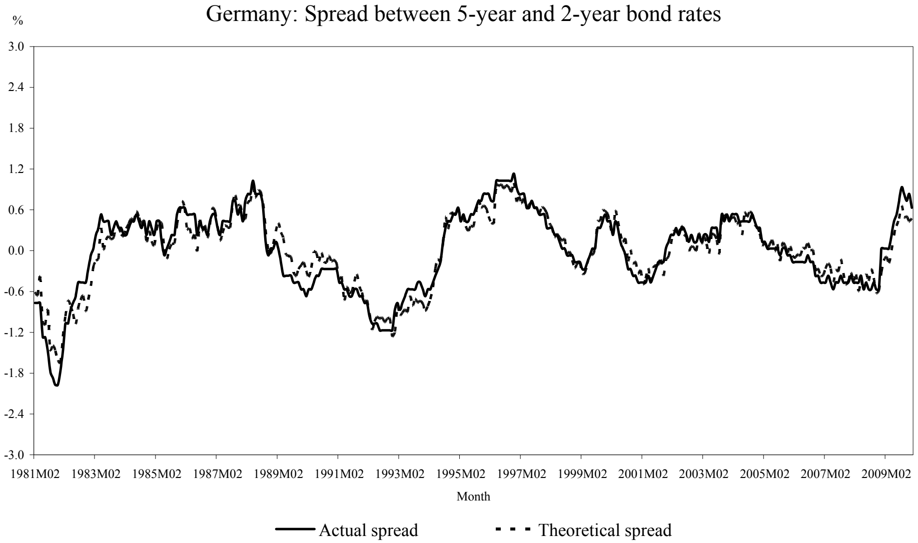
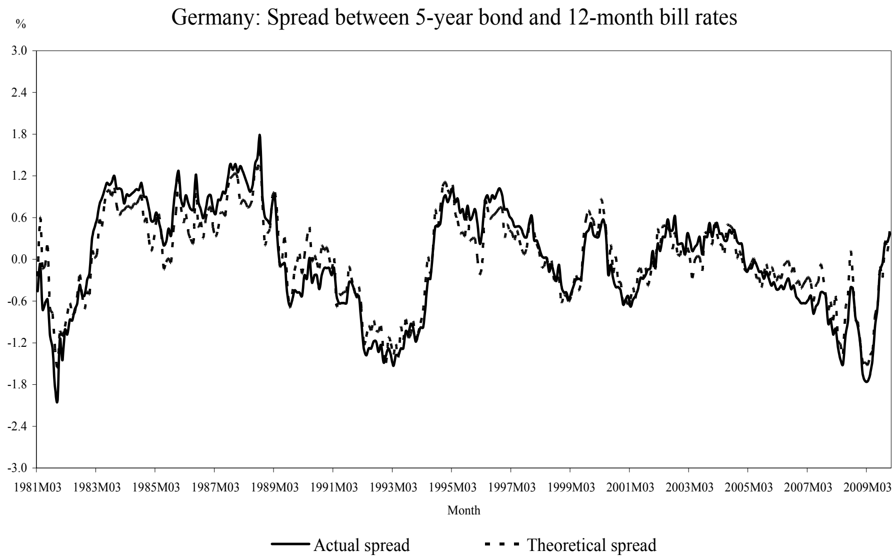


Figure 2 *(continued)*







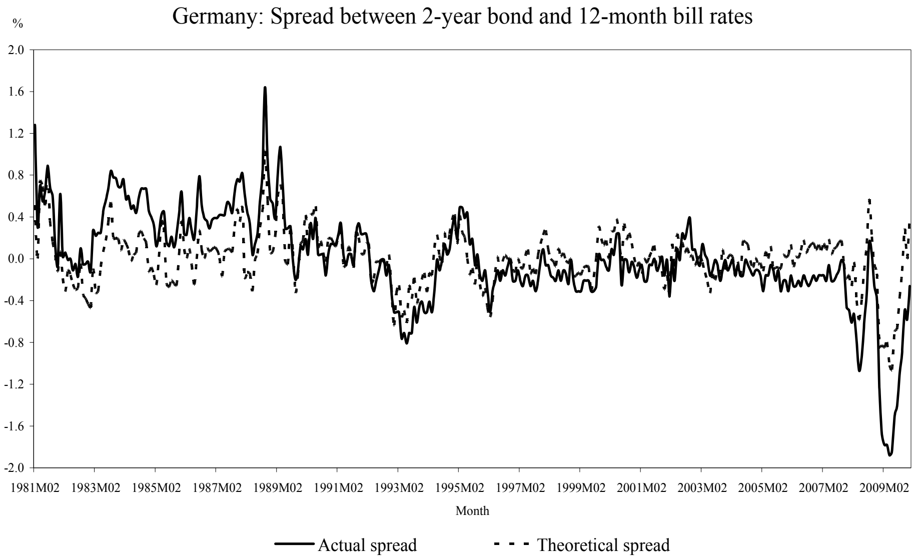
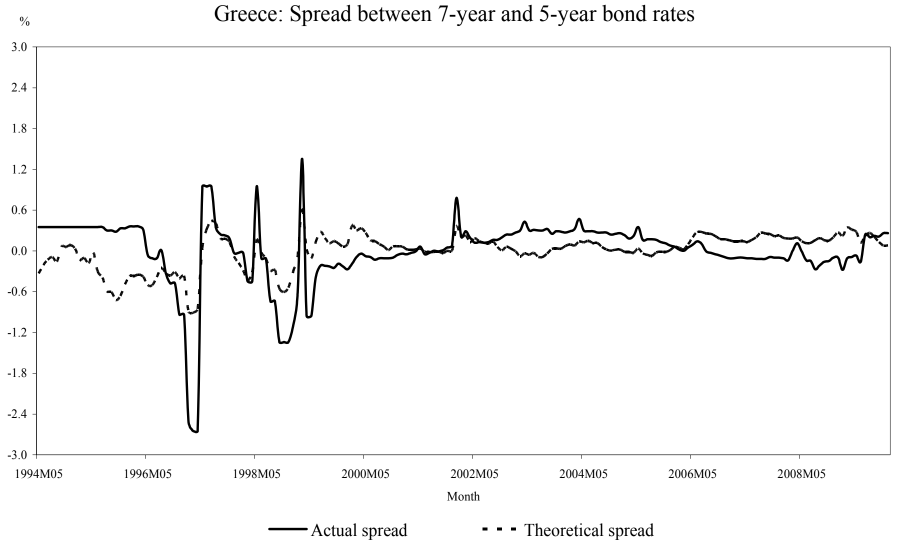
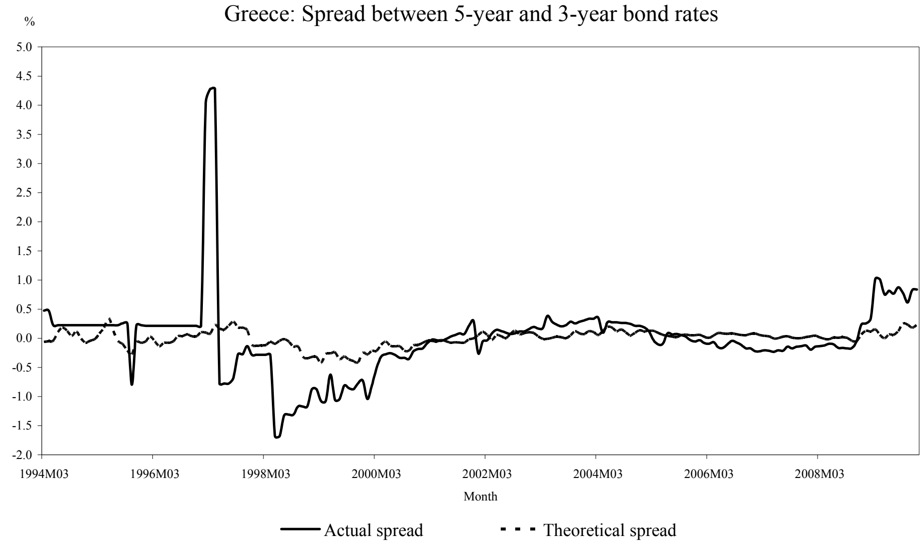


Figure 2 *(continued)*





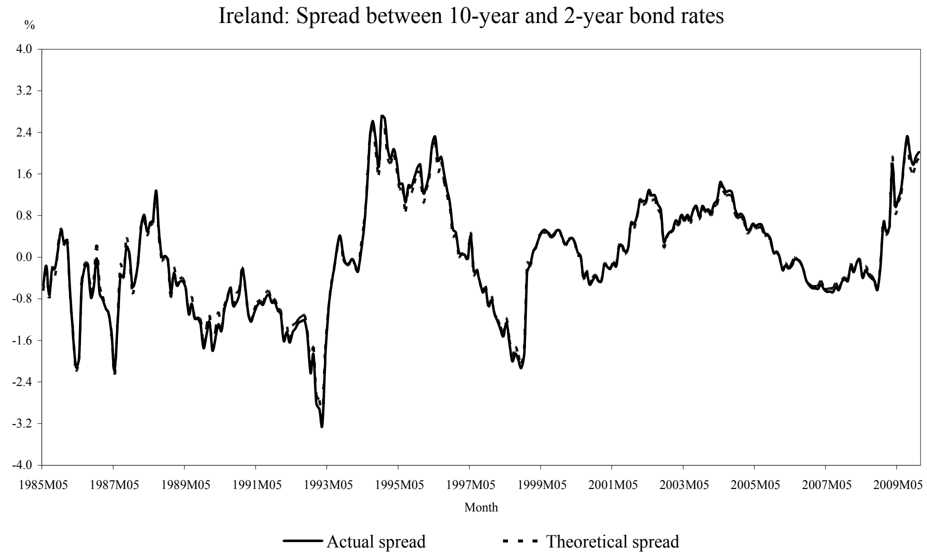
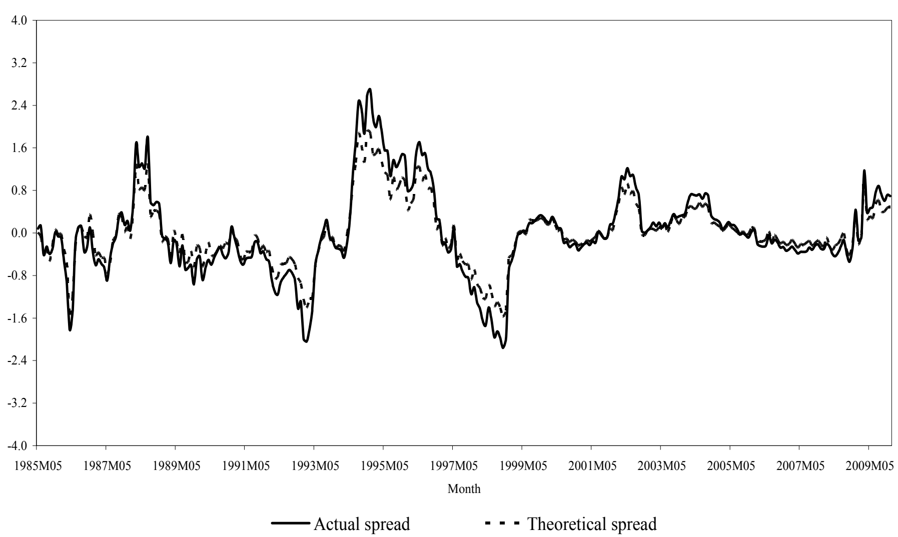
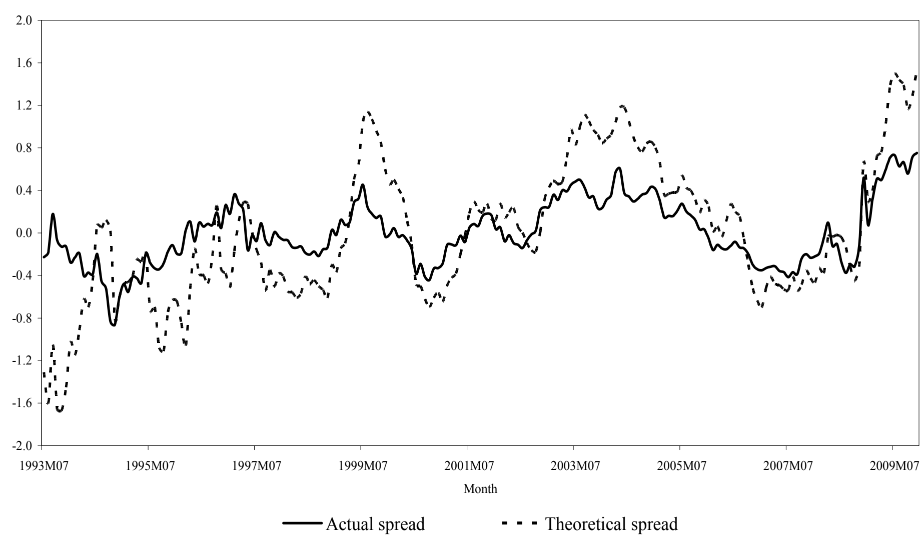


Figure 2 *(continued)*

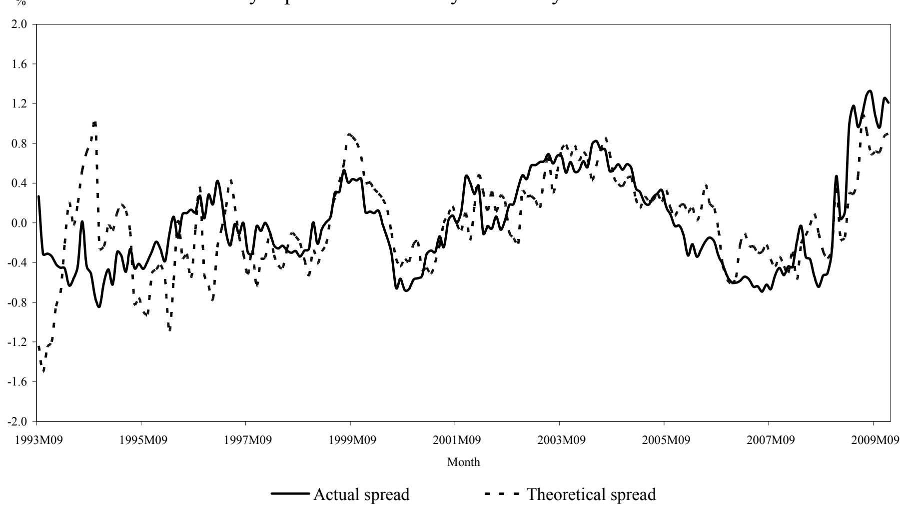
Ireland: Spread between 5-year and 2-year bond rates



Italy: Spread between 10-year and 5-year bond rates



Italy: Spread between 10-year and 3-year bond rates



Italy: Spread between 10-year bond and 12-month bill rates

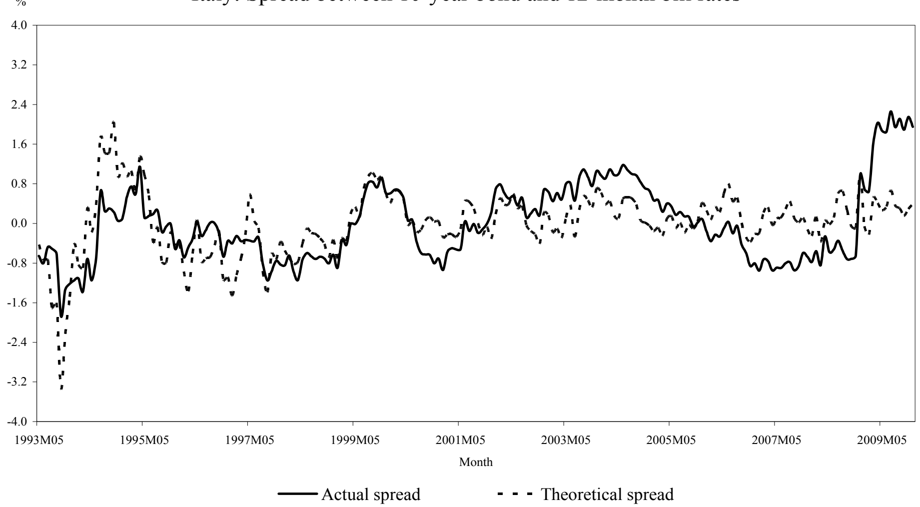
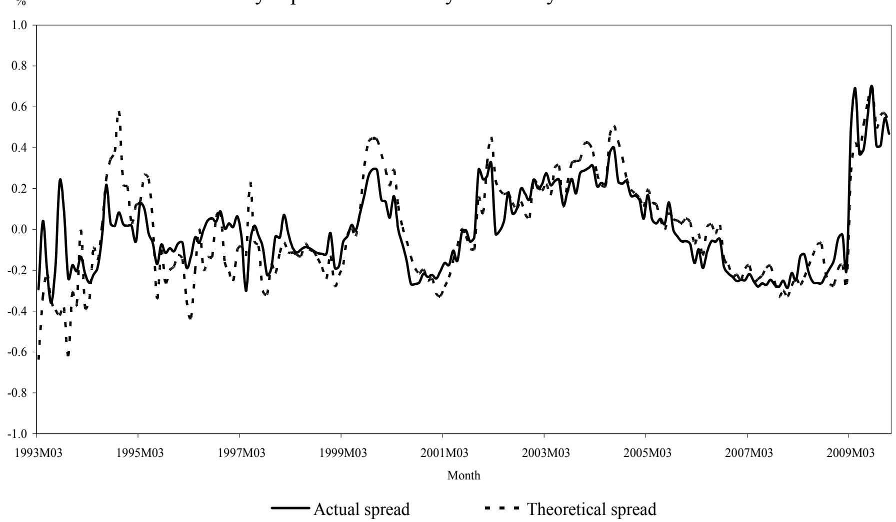
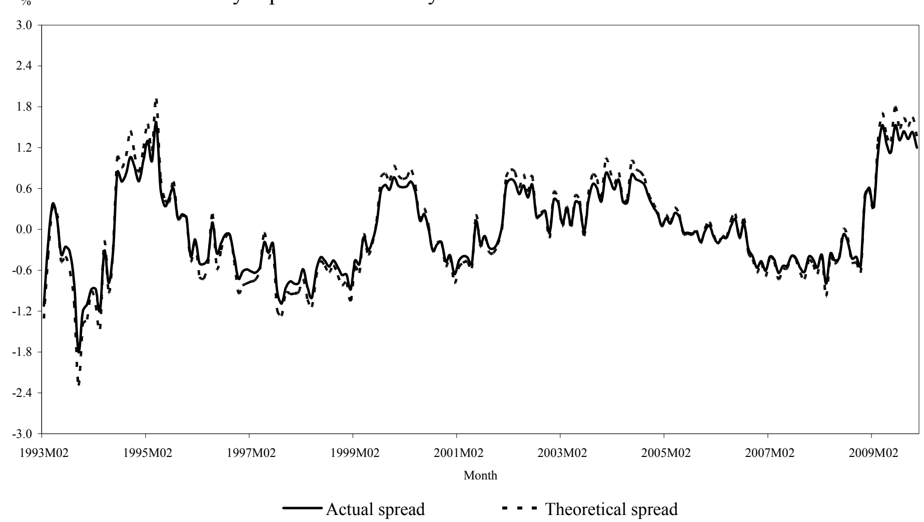


Figure 2 *(continued)*

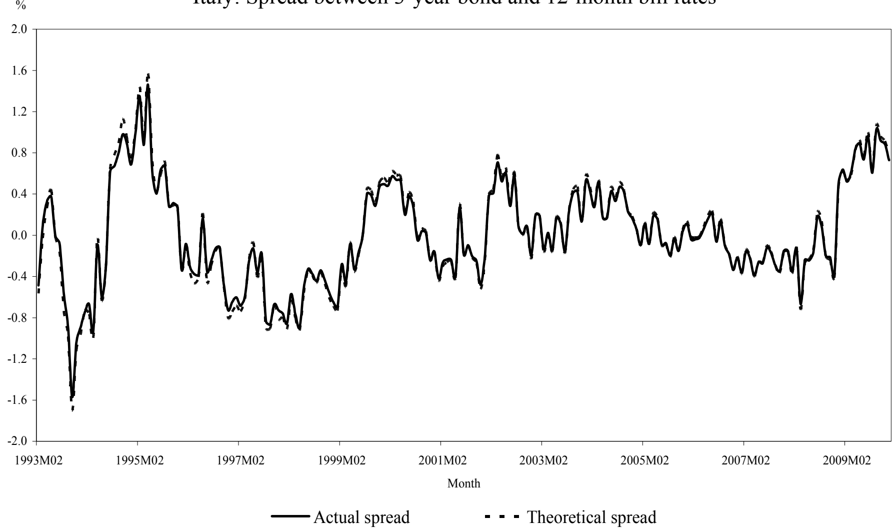
Italy: Spread between 5-year and 3-year bond rates



Italy: Spread between 5-year bond and 12-month bill rates



Italy: Spread between 3-year bond and 12-month bill rates



The Netherlands: Spread between 10-year bond and 3-month bill rates

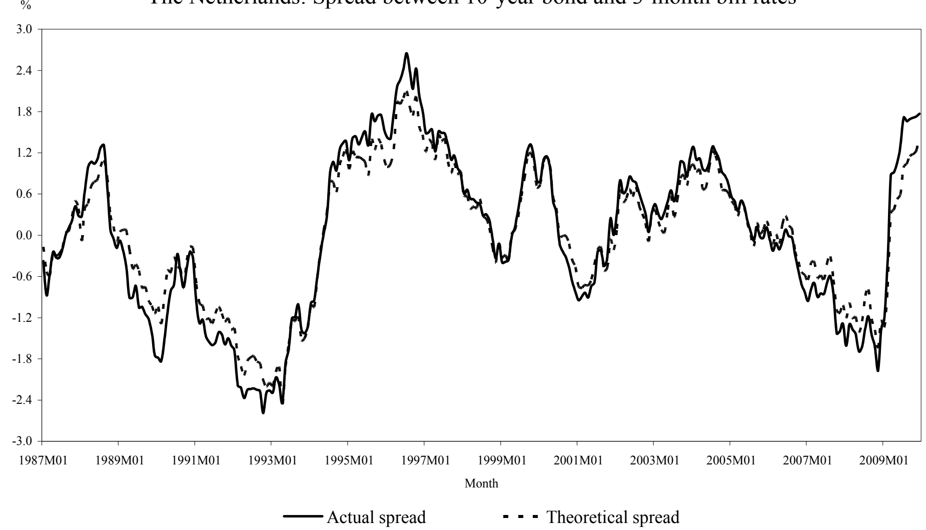
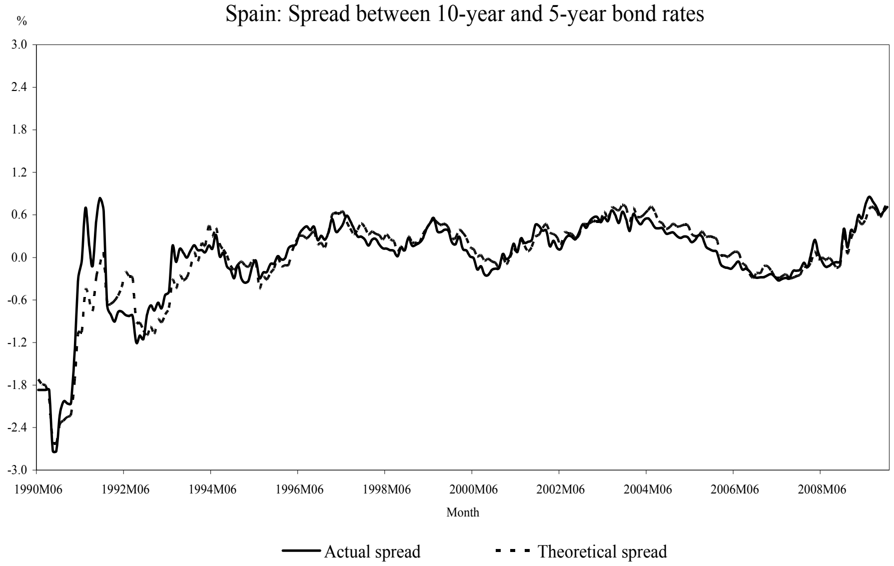
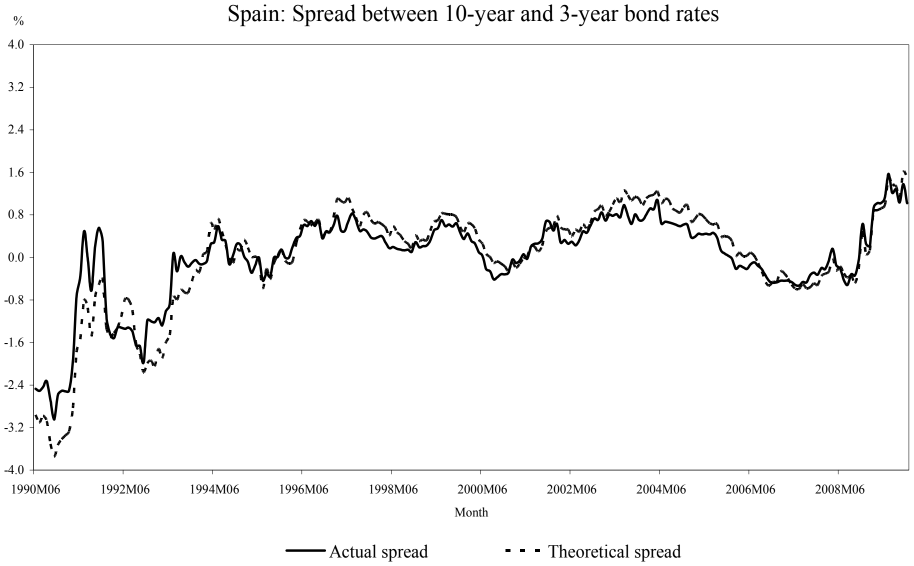
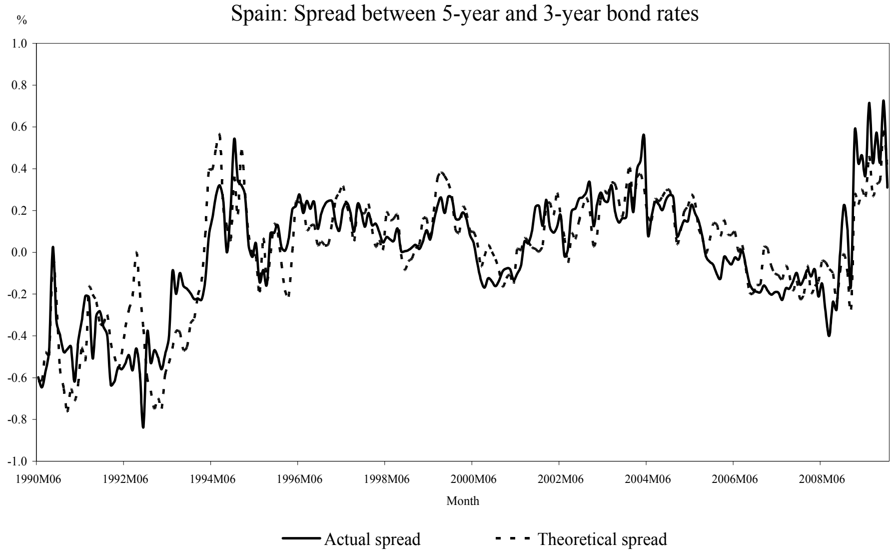


Figure 2 *(continued)*







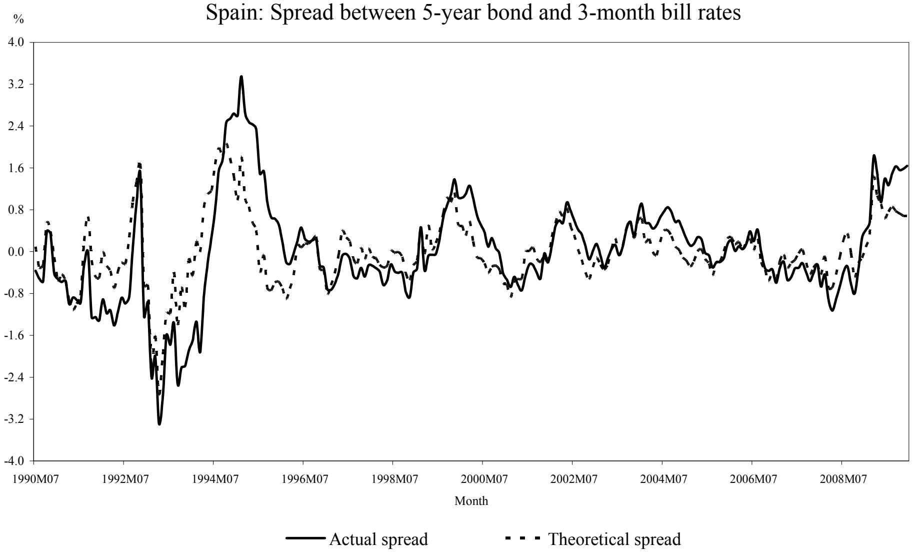
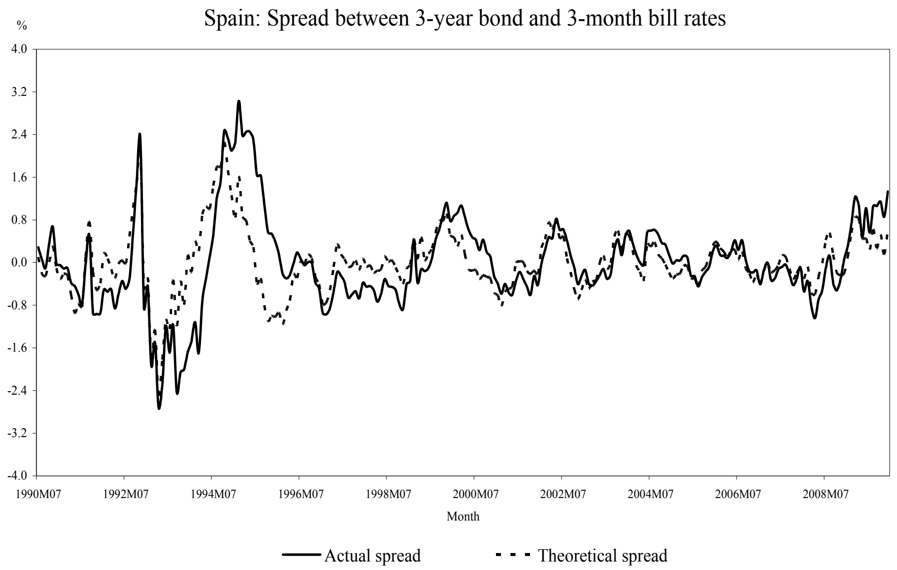
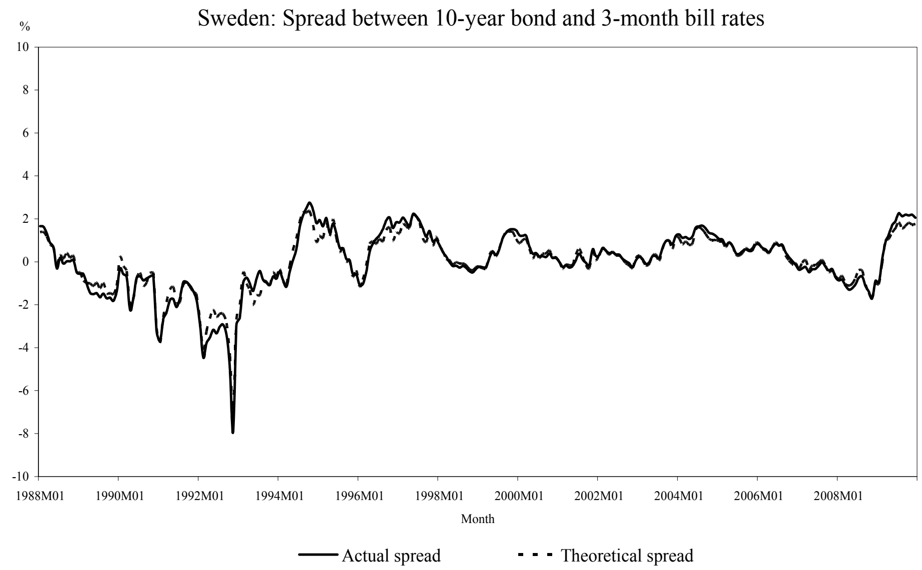
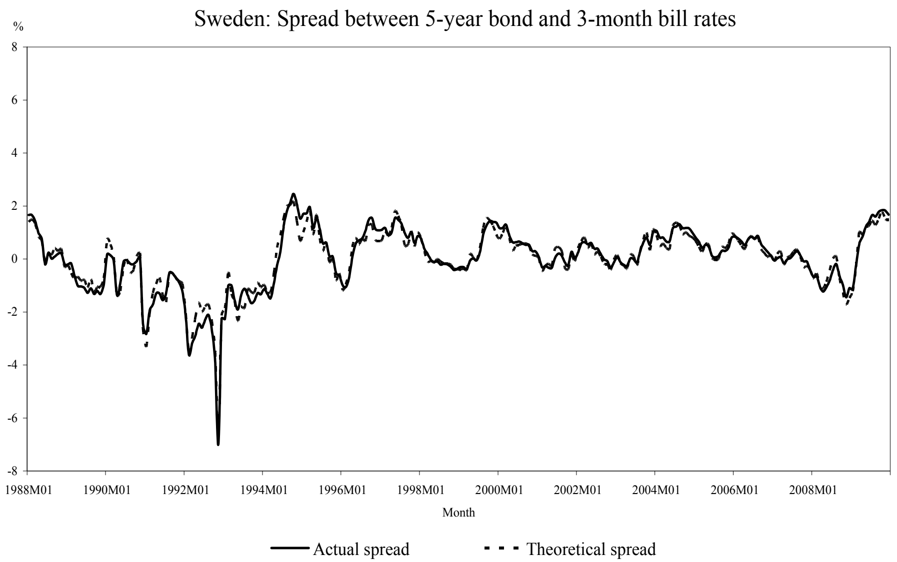


Figure 2 *(continued)*







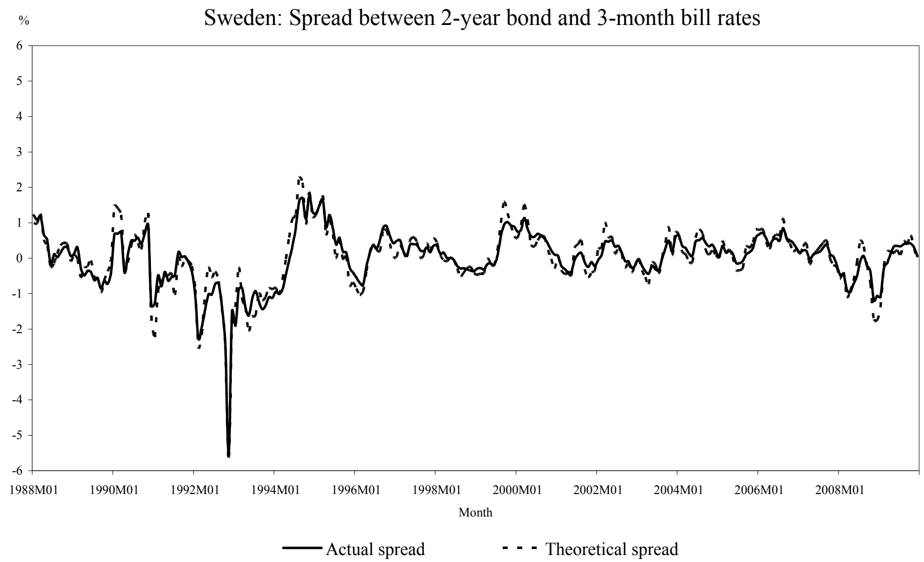


Figure 2 *(continued)*