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The Term Structure of the Spreads between Portuguese and German Interest Rates during Stage II of EMU*

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Abstract

The spread between interest rates denominated in different currencies represents the expectations on exchange rate changes, according to the uncovered interest rate parity condition. In the present research the short- and long-term spreads between Portuguese and German Treasury bonds interest rates are studied, using weekly data covering the period from 1993-08-02 to 1998-12-14, supplied by the Banco de Portugal. The interdependence of the two spreads is estimated using cointegration methods, and their dynamic adjustment to the long-term relation is determined using impulse response analysis. The main conclusions of this research are that there was a structural break in the long-term relation between the two spreads in mid 1994, and that that relation was afterwards dominated by the consistent convergence of the Portuguese interest rates to European levels.

Keywords: term structure, interest rate parity, cointegration, structural break

JEL Classification: E43

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I. Introduction

The theoretical support for the present research are the theories on the term structure of interest rates, and the uncovered interest rate parity condition. The domestic term structure of interest rates is explained by expectations for future interest rates and by risk premiums. According to the interest rate parity condition, the spread between two interest rates, denominated in different currencies, is equal to the expectations for changes in the exchange rate. The period studied in the present research covers the Stage II of European Monetary Union. During that period, in accordance with the nominal convergence criteria, the Portuguese long-term interest rate had to converge to the lowest levels of European interest rates, here represented by the German long-term interest rate. The short-term interest rate has also eventually followed a similar process of convergence. However, the short-term interest rate must also be an instrument of monetary policy, used to keep the escudo exchange rate within the fluctuation margins, defined by the EMU rules. How these circumstances affected the processes of the short-term and long-term interest rates spread is the main subject research conducted in this paper, using cointegration and VAR methodology. The second part of this article describes the term structure and interest rate parity, which are the theoretical foundations of the empirical research. In the third part, cointegration and VAR analysis is applied to the data, where the short-term interest rates are represented by the one-week Portuguese and German money market interest rates, and the long-term interest rates are represented by the three-year Treasury bond interest rates operating in the two countries.

II. The term structure of interest rates and the uncovered interest rate parity

Using continuous time compounding discount functions, the prices of a zero coupon bond with a maturity of n and n-1on dates t and t+1, respectively, can be represented as follows:

$$P_{t,n} = \exp\left(-nR_{t,n}\right) \quad (1)$$

and

$$P_{t+1,n-1} = \exp\left[-\left(n-1\right)R_{t+1,n-1}\right] \quad (2).$$

The return obtained from holding the bond between t and t+1 is:

$$h\left(n\right)_{t,t+1} = \ln\left(\frac{P_{t+1,n-1}}{P_{t,n}}\right) \quad (3),$$

or, alternatively:

$$h\left(n\right)_{t,t+1} = nR_{t,n} - \left(n-1\right)R_{t+1,n-1} \quad (4).$$

According to the hypothesis of risk premiums in the term structure, the return from holding that zero coupon bond between date t and date t+1 is usually superior to the risk-free (one period) interest rate, $R_{t,1}$, their difference being the *return premium*[1] :

$$h(n)_{t,t+1} - R_{t,1} = \phi(n)_t \quad (5).$$

[1] This measure of risk premium is used both in the traditional approach to the term structure analyis, by Fama (1984) and McCulloch (1987), and in the dynamic approach, based on the Vasicek (1977) and Cox, Ingersoll and Ross (1985) models.

When $h(n)_{t,t+1}$ is replaced in (5), according to its definition given by (4), the following relation between interest rates and return premium results:

$$R_{t+1,n-1} - R_{t,n} = \frac{1}{n-1}\left(R_{t,n} - R_{t,1}\right) - \frac{1}{n-1}\phi\left(n\right)_t \qquad (6).$$

Finally, applying the law of iterated expectations to (6), from t to t+n-1, the following relation between the long-term interest rate and the short-term interest rates expected between t and t+n-1 is obtained:

$$R_{t,n} = \frac{1}{n}\sum_{i=1}^{n} E_t\left(R_{t+i-1,1}\right) + \theta\left(n\right)_t \qquad (7),$$

where $\theta(n)_t$ is a risk premium defined as:

$$\theta\left(n\right)_t = \frac{1}{n}\sum_{i=1}^{n} E_t\left(\phi\left(n-i+1\right)\right)_{t+i-1} \qquad (8).$$

From (7) it is possible to define the spread between the long-term interest rate, $R_{t,n}$ and the short-term interest rate, $R_{t,}1$, as a function of the changes in expected short-term interest rates and the risk premium:

$$R_{t,n} - R_{t,1} = \sum_{i=1}^{n-1}\left(1 - \frac{i}{n}\right)\left[E_t\left(R_{t+i,1}\right) - E_t\left(R_{t+i-1,1}\right)\right] + \theta\left(n\right)_t \qquad (9),$$

Forecasts of future interest rates are given by using equations of type (9), relative to a succession of adjacent interest rates, provided that the risk premiums $\theta(n)_t$ are constant, or at least stable, as Evans and Lewis (1994) and Hardouvelis (1994) propose.

According to the uncovered interest rate parity hypothesis, the difference between two interest rates, with respect to debts maturing at the same time, but denominated in different currencies, represents the market's expectations for exchange rate changes. By defining the long-term spread and the short-term spread between the interest rates denominated in A and B currencies with

$$S_{t,n} = R(A)_{t,n} - R(B)_{t,n} \quad (10),$$

and

$$S_{t,1} = R(A)_{t,1} - R(B)_{t,1} \quad (11)$$

respectively, these differences can, according to the hypothesis, be taken as proxies of the expected long-term and short-term rates of change in exchange rate. Taking the link between interest rate differentials and expected changes in the exchange rate, Lund (1999) used the term structure to find the probabilities of the capability of entering the European Monetary Union, for several EU countries. In the present study, the cointegration analysis between two spreads, is used to provide evidence for the consistency or disparity between the different expectations regarding the capability of the escudo joining the EMU, observed during the last decade.

III. Cointegration analysis

The cointegration method used in this article was proposed by Johansen (1988,1991). It uses a Maximum Likelihood Procedure, to estimate a *p*th-order nonstationary VAR model of the following type:

$$\Delta y_t = \Gamma_1 \Delta y_{t-1} + ... + \Gamma_k \Delta y_{t-k} + \Pi y_{t-1} + \mu + \varepsilon_t \quad (12).$$

where y_t is a vector of p cointegrated variables, $\Delta y_{t-1} \ldots \Delta y_{t-k}$ are the vectors of their changes, μ is a constant, , $\Gamma_1, \ldots, \Gamma_K$, are parameter vectors, Π is the product of a vector α of the reversion intensities of the variables to their long run relation, and a matrix β of the coefficients of cointegration, also called the *error correction model*, which is the long run relation between the variables, and ε_t is an error term satisfying the condition: $\varepsilon_t \sim N(0,\sigma_\varepsilon)$. The rank of the matrix Π, which can not be higher than the number p of variables included in the model, is the number of cointegrating vectors and it is determined by the statistics λtrace and λmax, proposed by Johansen (1988,1995).

III.1 Tests on the non-stationarity of the variables

The most significant aspect of cointegration methodology is to make possible the distinction between the long run path of each variable, which is determined by its cointegration with the others, and its short run changes. This distinction only makes sense if the variables under analysis are non-stationary, which has been estimated by applying the following Dickey-Fuller tests to the each of the two variables in question. These were, respectively, the spread between Portuguese and German one-week money market rates (defined as the short-term spread) $S_{t,1w}= R(P)_{t,1w}-R(G)_{t,1w}$, and the spread between Portuguese and German 3 year Treasury bond rates (defined as the long-term spread), $S_{t,3Y}= R(P)_{t,3Y}-R(G)_{t,3Y}$. These tests consist of estimating the ADF-Z statistic in the following three types of regressions, for any variable Y:

$$\Delta Y_t = \gamma_a Y_{t-1} + \sum_{i=2}^{p} \beta_i \Delta Y_{t-i} + \varepsilon_t \qquad (13),$$

$$\Delta Y_t = a_0 + \gamma Y_{t-1} + \sum_{i=2}^{p} \beta_i \Delta Y_{t-i} + \varepsilon_t \qquad (14),$$

and

$$\Delta Y_t = a_0 + a_1 t + \gamma Y_{t-1} + \sum_{i=2}^{p} \beta_i \Delta Y_{t-i} + \varepsilon_t \qquad (15).$$

The lagged changes for the dependent variable in the regressions were included to ensure that error terms are white noise. The number of lagged changes in each regression was determined by the Akaike-Schwartz criterion. In equation (14), the Z statistic was calculated and the joint hypothesis of a unit root and no constant, $\alpha_0=\gamma=0$ (statistic ϕ), was tested. In the same equation, a test was performed to determine the significance of the constant term. In equation (15), the joint hypothesis of a unit root, no constant and no trend, $\alpha_0=\alpha_1=\gamma=0$ (statistic ϕ), was tested together with ADF Z statistic. The results of these tests are presented in Table I.

The hypothesis of a structural break in each series was also tested. For this purpose, a research procedure, proposed by Perron (1997) was used to determine the moment of the structural break, in which two models of structural break are considered. Perron's first model, hereafter designed by Perron I, for testing for a structural break in any variable, *Y*, provides for a change in the intercept, the null hypothesis being the unit root, and it may be represented as follows:

$$Y_t = \mu + \theta DU_t + \beta t + \delta D\left(T_b\right)_t + \alpha Y_{t-1} + \sum_{i=1}^{k} c_i \Delta Y_{t-i} + \varepsilon_t \qquad (16),$$

where T_b+1 denotes the time at which the break occurs in the trend function, with $D(T_b)$ being a dummy with a value equal to 1 at Tb+1 and equal to zero at all the other dates. The dummy variable $DU_t(t>T_b)$, represents a change in the trend intercept, due to the innovation that begins at T_b.

The second Perron model, hereafter designed by Perron II, allows a change in both the intercept and the slope, at time T_b+1, and its representation is:

$$Y_t = \mu + \theta DU_t + \beta t + \gamma DT_t + \delta D\left(T_b\right)_t + \alpha Y_{t-1} + \sum_{i=1}^{k} c_i \Delta Y_{t-i} + \varepsilon_t \qquad (17),$$

where DT_t represents the change in the slope. The results of these tests are presented on Table II.

Table I –Tests on non-stationarity

	SP1W (s.t spread)			SP3Y (3y spread)		
Model	**ADF_Z**	**Lags**	**ϕ**	**ADF_Z**	**Lags**	**ϕ**
a) without constant and trend	-2.0566$^{(*)}$	13		-1.1702$^{(*)}$	3	
b) with constant	-8.1607$^{(*)}$	14	1.3673$^{(*)}$	-0.5799$^{(*)}$	3	0.8505$^{(*)}$
c) with constant and trend	-178.0878$^{(**)}$	14	9.0685$^{(**)}$	-5.6215$^{(*)}$	3	1.9588$^{(*)}$

$^{(*)}$ Less than the critical value at 5%
$^{(**)}$ Greater than the critical value at 5%

Table II - Tests on structural break on the interest rate spreads

a) Model: Perron I

(Student T within parenthesis)

	SP1W	SP3Y
T stat. of $\alpha = 1$	-4.33867 (*)	-0.46289(*)
μ	0.02600	0.0002
	(4.4440)	(0.20489)
θ	0.00596	-0.00208
	(1.9452)	(-3.6950)
β	-0.0001	0.000007
	(-4.05061)	(1.51538)
δ	0.00941	-0.00055
	(0.72653)	(0.28375)
α	0.55567	0.994
	(5.42594)	(79.062)

(*)Less than the critical value at 5%

b) Model: Perron II

(Student T within parenthesis)

	SP1W	SP3Y
T stat. of $\alpha = 1$	-4.77216(**)	-1.10129(*)
μ	-0.01551	-0.00123
	(-1.86731)	(-1.11646)
θ	0.04893	0.00231
	(5.10944)	(2.21202)
β	0.00130	0.00008
	(5.31932)	(3.0715)
γ	-0.00141	-0.00009
	(-5.55357)	(-3.22427)
δ	-0.0111	0.00035
	(-0.81566)	(0.179229)
α	0.51072	0.98054
	(4.98118)	(55.482)

(*)Less than the critical value at 5%
(**)Less than the critical value at 5% and greater than the critical value at 10%

According the results of the Dickey-Fuller tests, none of the spreads is a stationary variable. The statistic ϕ in the third of these models leads to the non-rejection of the existence of a trend as the explanation for the non-stationarity of the short-term spread. The tests on the structural break show, in the case of the short-term spread, that the hypothesis of a structural break, consisting in a change of both the intercept and the slope, can not be clearly excluded. The break date estimated in this test is 1994-06-27, which is just after the periods of exchange rate instability of 1992 and 1993.

III.2. The cointegration models

Three cointegration models have been tested in this paper. The first model includes constants in the cointegration space. In this model, the ith transposed vector in the β matrix is represented by $\{\beta_{i,0}\ \ \beta_{i,3y}\ \beta_{i,,1w}\}$, where $\beta_{i,0}$ is the constant. The ratio-likelihood criterion proposed by Sims (1980) has been used to determine number of lags included in the VAR, which are 20. Since the values calculated for the λtrace statistics, shown on Table III, are below the value corresponding to the critical level of 5%, no cointegration between the two spreads can be observed with this type of model. In the second model estimated, no constant was included in the cointegration space. In this model, the ith transposed vector in the β matrix can thus be represented by $\{\beta_{i,3y}\ \beta_{i,,1w}\}$. The λtrace statistics calculated for this model, also presented on Table III, indicate that the hypotheses of one cointegration vector between the two spreads cannot be rejected using this model. Autocorrelation of residuals can be rejected from the model, according to the results of the Ljung-Box and Lagrange-Multiplier tests shown in Table IV. However, heteroskedasticity can not be rejected in one of the equations in the VAR, as the results show (Table IV). Finally, the hypothesis of a structural break in the long run relation of the second cointegration model was tested as a possible cause of the heteroskedasticity observed. The results of the tests on the first and second Perron models are presented in Table V. The test on the first model shows that a structural break in the long-term relation between the two spreads occurred on 7 July 1994. During the spring of that year a very significant increase in money market interest rates occurred. It was only after the middle of that year that short-term interest rates were again capable of following a consistent path of convergence to the European levels. This could explain the structural break that was detected in that period. In order to correct the structural break in the error correction model, a dummy variable has been included in the cointegration space, with a value of zero for the break date, 1994:07:04, and the preceding dates, and a value of one for the dates after the break date. The values calculated for the λtrace statistics of that model are given in Table III.

The values obtained for the λtrace statistics in the third cointegration model confirm the hypothesis of one cointegration vector linking the two spreads, with a dummy variable separating the dates before and after the structural break. The results of residual analysis, shown on Table IV, show no evidence of residual autocorrelation or heteroskedasticity. The number of lags in the VAR, which is 20, as in the previous models, has been determined by ratio likelihood criteria. The values obtained for the coefficients of the β vector, $\{\beta_{3y} \quad \beta_{1w} \quad \beta_{DU}\}$, normalized by the first coefficent, are: {1 –1,643 0.019}, which means that the long-term relation between these variables can be represented by:

$$SP3Y_t = 1.643SP1W_t - 0.019DU_t$$

The Granger causality relation between the two variables, or the hypothesis of strong exogeneity, has also been tested, using an F test. Table VI gives the results of this test, and they show that, while changes in the long-term spread are caused by changes in the short-term spread, the inverse is not so clearly evident. Another hypothesis that has been tested in the third cointegration model is whether any of the spreads is weakly exogenous. The values of the α coefficients (speeds of adjustment of the variables to their long-run relation), their T statistics, and the chi-squared statistics of the test on weak exogeneity, are also listed in Table VI. The results of these tests clearly show that the hypothesis of weak exogeneity for the long-term spread can not be rejected.

Table III– Test on the cointegration rank

a) model with constant in the cointegration space

λtrace	H0: r	p-r	λtrace 95%
12.58	0	2	19.993
1.95	1	1	9.133

b) model without constant in the cointegration space

λtrace	H0: r	p-r	λtrace 95%
10.96	0	2	10.35
1.44	1	1	2.98

c) model without constant and with a dummy variable in the cointegration space

λtrace	H0: r	p-r	λtrace 95%
28.34	0	2	24.08
9.95	1	1	12.21
2.84	2	1	4.14

Table IV–Residual analysis of the cointegration models

a) model with constant in the cointegration space

Tests on residual autocorrelation	Tests on residual heteroskedasticity (ARCH)
L-B(65), $\chi^2(180)$ = 169.858, p-val = 0.69	$\chi^2(20)$ = 18.310 Sig. Level = 0.566995
LM(1), χ^2 (4) = 2.944, p-val = 0.57	$\chi^2(20)$ = 189.317 Sig. Level = 0.0000
LM(4), χ^2 (4) = 7.409, p-val = 0.12	

b)model without constant in the cointegration space

Tests on residual autocorrelation	Tests on residual heteroskedasticity (ARCH)
L-B(65), $\chi^2(180)$ = 169.339, p-val = 0.70	$\chi^2(20)$ = 20.002 Sig. Level = 0.45780461
LM(1), χ^2 (4) = 2.800, p-val = 0.59	$\chi^2(20)$ = 188.146 Sig. Level = 0.0000
LM(4), χ^2 (4) = 7.199, p-val = 0.13	

c)model without constant and with a dummy variable in the cointegration space

Tests on residual autocorrelation	Tests on residual heteroskedasticity (ARCH)
L-B(65), χ^2 (411) = 435.087, p-val = 0.20	$\chi^2(20)$ = 29.977 Sig. Level = 0.07
LM(1), χ^2 (9) = 4.916, p-val = 0.84	$\chi^2(20)$ = 12.636 Sig. Level = 0.892
LM(4), χ^2 (9) = 7.342, p-val = 0.60	$\chi^2(20)$ = 10.154 Sig. Level = 0.965

Table V– Tests on the structural break in the cointegration relation in the model without constant

(Student T within parenthesis)

Model: Perron I		**Model: Perron II**	
T stat. of $\alpha = 1$	-5.5797[(**)]	T stat. of $\alpha = 1$	-4.615[(*)]
μ	-0.00467	μ	0.1201
	(-3.52411)		(2.99211)
θ	0.01043	θ	-0.00683
	(5.43061)		(-1.67393)
β	-0.00004	β	-0.0006
	(-4.4395)		(-4.67824)
δ	-0.0599	γ	0.00057
	(-7.1117)		(4.49797)
α	0.73723	δ	0.02361
	(15.655)		(2.50843)
		α	0.74419
			(13.427)

[(*)]Less than the critical value at 5%
[(**)] Greater than the critical value at 5% (**break date: 1994:07:04**)

Table VI– Tests of causality and of week exogeneity in the model without constant and with a dummy variable in the cointegration space

a) Causality tests

Dependent Variable: ΔSP1W			Dependent Variable: Δ SP3Y		
	F-Statistic	Sig. Level		F-Statistic	Sig. Level
ΔSP3Y*+ECM**	0.841	0.667	ΔSP1W*+ECM**	4.887	0.000
ΔSP1W*	6.824	0.000	ΔSP3Y*	1.903	0.014

* Lags 1 to 20
** Error Correction Model (Lag 1)

b)Speed adjustment estimators and week exogeneity tests

The alpha coefficients and statistics			Tests on week exogeneity	
	α	T-Statist.		
SP1W	-0.071	-2.270	SP1W	$\chi^2(1) = 3.24$ p-value = 0.07
SP3Y	0.005	0.855	SP3Y	$\chi^2(1) = 0.62$ p-value = 0.43

The independence of the money market spread relative to the long-term spread can be explained by the fact that the short-term spread was mainly governed by monetary policy actions which, during a significant part of the period under study, were determined by the exchange rate target for the escudo. The fact that the causality is more evident from the short-term to long-term spread, agrees with the expectation theory on the term structure of interest rates, which explains the level of the long-term rates by means of the expectations for future short-term rates.

III.3. The impulse response function analysis

The analysis of impulse response functions provides information on how a shock to one of the variables included in the cointegration space has an effect on the other. This type of analysis implies restricting the consideration that the contemporaneous effects on the VAR are triangular, which can be achieved using Choleski decomposition. The responses of each spread to a shock to itself or to the other, are represented on the graphs of Figure 1. The responses of each variable are plotted on the central lines of the graphs. The upper and lower lines are the response plus (less) twice the standard error. The impulse analysis graphs show that the dynamics of the responses of the long term spread to shocks to the short-term spread are more intense than in the inverse case. This can be explained by the fact that the causality relation between the two spreads has essentially one direction.

Conclusion

In the present research cointegration analysis has been used to test the relations beetwen the short-term and long-term spreads of Portuguese interest rates relative to German interest rates, during the Second Stage of European Monetary Union. The cointegration model obtained shows that the long run relation between the two spreads indicates a structural break at beginning of 1994. That date is after the end of a period of instability in the exchange markets, observed during 1992 and 1993, and it is the beginning of a consistent process of convergence of Portuguese interest rates to European levels, according to the parameters of the cointegration relation. According to the results of these tests, the causality relation between the two spreads is mostly directed from the short-term to the long term spread. The main relation in the causality shows that the capability of the long-term Portuguese interest rates to reach the European levels, has been strongly dependent on the capability of the monetary policy to ensure a stable external value for the escudo.

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