

**EUROPEAN UNIVERSITY INSTITUTE**  
DEPARTMENT OF ECONOMICS

EUI Working Paper ECO No. 99/8

Interest Rate Spreads between Italy and Germany  
1995–1997

Marcello D'AMATO and Barbara PISTORESI

**BADIA FIESOLANA, SAN DOMENICO (FI)**

All rights reserved.

No part of this paper may be reproduced in any form  
without permission of the author.

©1999 M. D'Amato and B. Pistoiesi

Printed in Italy in March 1999

European University Institute

Badia Fiesolana

I-50016 San Domenico (FI)

Italy

# Interest Rate Spreads between Italy and Germany: 1995-1997\*

Marcello D' Amato, *University of Salerno*

Barbara Pistoresi, *EUI and University of Modena*

## Abstract

In this paper we study the determinants of the long term yield spread between Italian and German government bonds using daily observations for a period 1 January 1995- 28 October 1997. We split total spread into two main factors: an exchange rate factor, that we approximate by a differential on swap contracts (same maturity) and a default risk factor, that we consider as a residual. Using cointegration analysis we test if the interest rates parity condition holds in the period considered and also study the dynamic adjustment of total spread and its components using impulse response analysis. The main result is that an uncovered parity condition cannot be rejected in the sample only if the relationship is augmented by the German short term interest rate. Impulse response analysis shows that this latter variable permanently affects the default risk. The main conclusion is that the reduction of the total spread in the period studied was due both to credibility gains and to favorable dynamics in the German interest rate.

---

\* We are indebted to Mario Forni for the usual helpful comments. We are also grateful to Nunzio Cappuccio, Michael Ehrmann, Massimiliano Marcellino and an anonymous referee for useful suggestions. We thank Andrea Landi and Cosimo Musiello (Prometeia) for providing us with the data. Barbara Pistoresi thanks the European University Institute for hospitality and financial support. *E-mail address:* pistores@datacomm.iue.it, damato@bridge.diima.unisa.it.

# 1 Introduction

In this paper we study the determinants of the total interest rate spread on government bonds (BTP, BUND) between Italy and Germany within the period 1995-1997. The analysis is performed by decomposing the total spread within its components, as suggested by the uncovered interest rate parity condition: the exchange rate component and the default risk component.

An estimation of the default risk component is important, in this framework, since, after the establishment of EMU, the exchange rate component of the spread will disappear and any interest rate premium will depend on the (expected) default risk assigned by financial markets to a given country.

In the period considered the total interest rates spread between interest rates exhibited a steady reduction and the aim of this paper is to give account: a) of the role that in this reduction has been played by the exchange rate component with respect to the default risk component and b) the role that international factor (long run dynamics of the German interest rate) with respect to idiosyncratic factor (credibility) may have played in the dynamics of the default risk.

Both of these issues will be addressed empirically by estimating a long run relationship based on the interest rate parity condition. Point a) will be addressed by proxing the exchange rate component by interest rate differential on swap contracts (same maturity) denominated in different currencies and treating the default risk as a residual as, for example, in Favero et al. (1997). Point b) involves variables which cannot be easily proxied (particularly at the daily frequencies analysed in this paper) and hence will be addressed in a less direct approach. Both the international factor (world interest rate) and the national factor (credibility factor) can be taken as exogenously given with respect to the variables considered in the equilibrium relationship we use, and hence much of the answer to the question raised in point b) has to rely on the dynamic response of interest rate spread to these exogenous components.

The paper is organised as follows. Section 2 introduces the usual uncovered interest rate parity condition that decomposes total spread on government bond issued by Italy and Germany (BTP, BUND) into two factors: an exchange rate factor and a default risk factor. Section 3 shows the methodology and the results: (i) cointegration analysis is used to test the existence of a long run uncovered parity condition; (ii) impulse response analysis is used to study the dynamic adjustment of each variables to shocks. Finally, section 4 concludes.

## 2 The determinants of government bonds spread

We start from the following uncovered interest rate parity condition

$$(1 + I^{it})(1 - p) + p(1 - \alpha)(1 + I^{it}) = (1 + I^{ge})(E^e/E)$$

where  $I^{it}$  represents the annually compounded interest rate on Italian government BTP denominated in domestic currency,  $I^{ge}$  represents the annually compounded interest rate on German Government BUND denominated in D-Marks;  $p$  represents the probability of a default of Italian government occurring before the maturity expiration of the title,  $\alpha$  represents the fraction of the cost to the creditor due to default;  $E^e$  represents the expected exchange rate at time  $t$  about the time of maturity of the asset;  $E$  represents the current exchange rate (number of Italian liras for one D-Mark).

Some algebra shows that, by disregarding second order terms, this equilibrium condition can be rewritten, as follows:

$$I^{it} - I^{ge} \cong (E^e - E)/E + \alpha p \tag{1}$$

where  $I^{it} - I^{ge} = Sp$  can be defined as the total spread between Italian and German long term bonds,  $(E^e - E)/E = Er$  represents the expected exchange rate changes, while  $\alpha p = Dr$  represents the expected

cost due to default and can be defined as the default risk term. Hence (1) can be rewritten as follows

$$Sp = Er + Dr \tag{2}$$

## 2.1 Measures and definition of the variables

The main problem in testing such equilibrium condition and in decomposing  $Sp$  in its components is that these latter are not observable. In recent papers different routes have been proposed to escape this problem.

One route is to consider different measures for  $Er$  (free of any default risk) treating  $Dr$  as a residual. The most common measure for  $Er$  is given by interest rates differential on swap contracts (same maturity) denominated in different currencies:  $Er = SI^{it} - SI^{ge}$  (for example Favero et al., 1997, Tosato, 1996, Seghelini, 1996). The swap differential between ITL and DEM is usually assumed to embody the currency risk. Other candidates free from default risk are interest rates on long term bonds issued by a supra-national organisations (such as World Bank or the European Investment Bank). A different strategy would require using an independent measure of default risk treating  $Er$  as a residual. In this paper we use interest rates on swap contracts as a proxy for the exchange rate risk<sup>1</sup> for references about this issue see the discussion in Favero et. al. (1997).

---

<sup>1</sup>The excess of the yield on government bonds over the rate on swap contracts is used as a rough measure of the default risk by the Bank of Italy, see Banca d'Italia, *Bollettino Economico*, no.24, 1995). Favero et al. (1997) criticizes this approach for not considering the different taxation treatment between returns on bonds from different countries. In this paper we use the less sophisticated measure for the default risk as in Banca d'Italia, since this should not lead to substantial distortion. The main reason can be argued as follows: by looking at Fig.3, top panel, in Favero et al. (1997), the importance of the tax factor is continuously decreasing in Italy from January 1992 to December 1995 and approaches a constant level. Henceforth, the tax factor has remained constant in the period under analysis and may only affect the constant term in the cointegrating relationship. Tests on the cointegrating parameters show the negligible role of the restricted constant in the cointegrating vector. The results are available on request.

Data are daily observations for a period 1 January 1995, 28 October 1997 of the following variables:

*btp10y* : interest rate on 10 years Italian Government bonds (BTP denominated in Liras)

*bund10y* : interest rate on 10 years German Government bonds (BUND denominated in D-Marks)

*lit10y* : interest rate on swap contract (EuroITL)

*dm10y* : interest rate on swap contract (EuroDEM)

*lit3m* : interest rate on 3-months Italian bond (EuroITL)

*dm3m* : interest rate on 3-months German bonds (EuroDEM)

$Spr10y = btp10y - bund10y$

$Swpr10y = lit10y - dm10y$

### 3 Methodology and results

A natural framework for testing the interest rate parity condition as an equilibrium condition due to the degree of integration of the variable is given by Maximum Likelihood Procedure developed by Johansen (1991) to investigate cointegration properties of the data. The cointegration testing procedure appeals to the fact that deviations from equilibrium condition(s) for two or more variables, which are nonstationary (we return on this definition later), when taken by themselves, should be stationary. The intuition is that economic forces should avoid persistent long run deviation(s) from equilibrium condition(s), although significant short run deviation(s) - that is transitory but that could be long lasting - may be observed. An implication is that, while individual interest rates may wander extensively, certain groups of such series should not diverge from one to another in the long run. The economic counterpart of this statistical definition, is that interest rates exhibit considerable volatility and persistence to a national and international shocks representing changes in domestic and international policies and finance conditions,

however the parity condition in international financial markets should ensure that the individual rates do not wander arbitrarily far.

### 3.1 Cointegration analysis

Now we illustrate the basic econometric tools used. Suppose that a  $(n \times 1)$  vector  $X_t$  can be represented as a nonstationary  $p$ th-order vector autoregression:

$$X_t = \mu + B_1 X_{t-1} + \dots + B_p X_{t-p} + v_t \quad (3)$$

where  $v_t \sim (0, \Sigma_v)$ ,  $\det \Sigma_v \neq 0$ . Any VAR of this form can be equivalently written as

$$X_t = \mu + \Phi_1 \Delta X_{t-1} + \dots + \Phi_{p-1} \Delta X_{t-p+1} + \Gamma X_{t-1} + v_t \quad (4)$$

where

$$\begin{aligned} \Phi_s &= -(B_{s+1} + \dots + B_p) \text{ for } s = 1, 2, \dots, p-1 \\ \Gamma &= B_1 + \dots + B_p \end{aligned}$$

Subtracting  $X_{t-1}$  from both sides of (4) produces the following VECM (Vector Error Correction Model)

$$\Delta X_{t-1} = \mu + \Phi_1 \Delta X_{t-1} + \dots + \Phi_{p-1} \Delta X_{t-p+1} + \Pi X_{t-1} + v_t \quad (5)$$

where

$$\Pi = \Gamma - I = -(I - B_1 - \dots - B_p)$$



The rank  $r$  of the  $\Pi$  matrix can at most be equal to the number of variables included in the model, that is  $n$ . If  $r = n$ , the vector  $X$  is integrated of order zero, i.e.e stationary. If  $r < n$ , it determines the number of cointegrating vectors (representing  $r$  long run relationships among  $n$  variables) in the VAR model explaining  $X$ . In this case the  $\Pi$  matrix can be decomposed as  $\Pi = \alpha\beta'$  where  $\beta'$  is a  $(r \times n)$  matrix of cointegrating vectors and  $\alpha$  is a  $(n \times r)$  matrix of the adjustment coefficients to the long run relationships. If the rank is equal to zero the variables are not cointegrated and a VAR in levels as in (3) does not exist.

For example, if the parity condition (2) holds, we expect to find one cointegrating relationship between  $Sp$  (defined by  $btp10y - bund10y$ ) and  $Er$  (defined by  $lit10y - dm10y$ ) with the particular form  $\beta = (1, -1)$  or equivalently one cointegrating relationship between the four variables:  $btp10y, bund10y, lit10y, dm10y$  with the particular form  $\beta = (1, -1, -1, 1)$ .

### 3.2 Introduction to the identification and IRF analysis

In the econometric exercise we present in the next section we also perform impulse response analysis (IRF), that is the study of the dynamic adjustment of each variable to shocks. To this aim we implement identifying restrictions on model (4), that can be considered as the reduced form of a general dynamic structural model, to recover structural disturbances from reduced form disturbances.

The autoregressive representation of (4) is given by:

$$B(L)X_t = v_t \tag{6}$$

By inverting  $B(L)$  we get the (*reduced form*) moving average representation:

$$\Delta X_t = E(L)v_t \tag{7}$$

For the representation (6) and (7) the following conditions hold

$$v_t \sim (0, \Sigma_v), \det \Sigma_v \neq 0$$

$$E(0) = I_n$$

$$B(0) = I_n$$

$$E(L) = I + E_1L + E_2L^2 + \dots$$

$$B(L) = I - B_1L - B_2L^2 - B_3L^3 - \dots - B_pL^p$$

Now we assume that  $\Delta X_t$  has the following *structural* moving average representation

$$\Delta X_t = C(L)\varepsilon_t$$

where  $\varepsilon_t \sim (0, \Sigma_\varepsilon)$ . The innovations  $v_t$  of the reduced form are assumed to be linear combinations of the structural disturbances  $\varepsilon_t$ , i.e.  $v_t = S\varepsilon_t$  for some  $(n \times n)$  full rank matrix  $S$ .

Hence, the following relation holds

$$S\Sigma_\varepsilon S' = \Sigma_v$$

Since  $\Sigma_v$  can be estimated from the reduced form, the problem of the identification relates to the conditions under which the structural parameters in  $S\Sigma_\varepsilon S'$  can be recovered from  $\Sigma_v$ .

The structural model, i.e.e the coefficients of  $C(L)$  will be identified introducing enough restrictions to determine  $S$  uniquely. The orthonormality of the variance-covariance matrix  $\Sigma_\varepsilon = I$  provides  $n(n + 1)/2$  non linear restrictions on the elements of  $S$ . To just-identify the model, we need  $n(n - 1)/2$  additional restrictions. Since  $C(L) = E(L)S$ , we can choose different identification strategies, for example: (i) pre-specify three of the parameters of the contemporaneous matrix  $S$  (short run restrictions)<sup>2</sup>, (ii) pre-specify particular long run relationships between the variables to constraint the matrix of long run multipliers  $C(1)$  given that also the relation  $C(1) = E(1)S$  holds (long run restrictions), (iii) pre-specify a mix of long run and short run restrictions, (iv) choose the standard Choleski identification that imply the contemporaneous effects in the model are triangular, while lagged dynamics are unrestricted<sup>3</sup>. We discuss the restrictions of our model in the next section. Suitably transformed, the estimates of  $C(L)$  based on the identifying restrictions allow us to express all the  $n$  variables as the sum of  $p$  distributed lags of the structural shocks,  $\varepsilon_t$ .

### 3.3 Estimation of the interest rates parity condition

The necessary condition to perform Johansen cointegration procedure concerning model (5) is that all the variables in the VAR contain a unit root, that is they are I(1).<sup>4</sup>

We test the null hypothesis of the presence of a unit root by using Dickey - Fuller regressions, augmented when necessary with a number of

---

<sup>2</sup>Premultiplying both sides of (6) by  $S^{-1}$  we obtain the structural VAR representation  $A(L)X_t = \varepsilon_t$ , where the matrix of the contemporaneous effects is  $A(0) = S^{-1}$ .

<sup>3</sup>For a summary on the identification in VAR models, see Canova (1995a) and Canova (1995b).

<sup>4</sup>Shea (1992) reviews the debate on the nonstationary behavior of interest rates. The debate has proponents of linear time series model with unit root (for example Campbell and Shiller, 1987) and proponents of linear mean variance stationary model, I(0) (for example, Fama and Bliss (1987)). Goodwin and Grennes (1994) also review the limitation of conventional regression tests of interest rate equalisation in the case of non stationary series.

lags in order to prevent autocorrelated errors, and non parametric measure of persistence at zero frequency due to Cochrane (1988)<sup>5</sup>. Table 1 presents the results. The estimates of persistence at zero frequency and the ADF tests show that all the variables: *btp10y*, *bund10y*, *lit10y*, *dm10y*, *lit3m*, *dm3m* contain a unit root. The VECM (5) is a common representation used for variables that are trending. It implies that  $X$  has a linear trend but this can be eliminated by cointegrating relationships in  $\beta$ . In other words,  $\beta$  that eliminates stochastic non stationarity also eliminates deterministic non stationarity (this is the usual case of *deterministic cointegration*)<sup>6</sup>. Figure 1 shows that the variables of interest are trending in the relevant period and this is compatible with a VECM representation with an unrestricted constant.

---

<sup>5</sup>The persistence measures how important the random walk component is to the behaviour of a series. It describes how much a shock changes the forecast of a variable in the long run. If this change is zero, the innovations are viewed only to have transitory effects. When a series has a random walk component, that is it is I(1), innovations are expected to persist into the future and the persistence measures must be significantly different to zero. If the series is a pure random walk the persistence is equal to one. However, a persistence equal to one not necessarily implies that a series is a pure random walk, while the reverse occur. A persistence equal to one implies that a series is a pure random walk only for the class of models ARMA(1,1), see, for example, Pistoresi (1997). A persistence different from zero only implies that the series contains a random walk component.

<sup>6</sup>See Campbell and Perron (1991) for a discussion on deterministic and stochastic cointegration.

**Table 1** Spectral density function and unit root tests

Variables	spectrum $f(0)$ ,(s.e.)	ADF (lags)
<i>btp10y</i>	1.14 (0.36)	-0.15 (1)
<i>bund10y</i>	0.79 (0.25)	-2.07 (1)
<i>lit10y</i>	0.79 (0.25)	-0.09 (0)
<i>dm10y</i>	0.50 (0.16)	- 2.00 (1)
<i>lit3m</i>	0.74 (0.23)	- 0.55 (0)
<i>dm3m</i>	0.50 (0.16)	-2.62 (2)
<i>spr10y</i>	1.19 (0.37)	0.08 (1)
<i>swspr10y</i>	0.98 (0.31)	0.21 (1)

*Notes:* Spectral density function of the first differences of the variables at zero frequency  $f(0)$  is the spectrum at zero corresponding to a period equal to infinity, i.e. a trend. Bartlett window estimates, window size equal to 52. Asymptotic standard errors in parentheses, (s.e.). ADF: unit root tests from augmented Dickey-Fuller regression with a constant included (levels of the variables).Lag length chosen by LM tests. Critical values: 5% = -2.86.

**Model 1.** Now we start testing for cointegration among 4 variables entering the pure interest parity condition (2) that is  $X = (btp10y, bund10y, lit10y, dm10y)$ . The VAR has  $p = 4$  that is the number of lags supported by traditional lag truncation criteria.<sup>7</sup> Standard tests of the hypothesis of reduced rank of the  $\Pi$  matrix enable us to accept one cointegrating vector at the 95% quantiles. (Table 2).

**Table 2** Cointegration tests on  $X = (btp10y, bund10y, lit10y, dm10y)$ 

$H_0$	Trace	$T - nm$	95%
$r = 0$	48.7	47.57	47.2
$r \leq 1$	19.82	19.36	29.7

*Notes:* Critical values in Ostervald - Lenum (1992, table 1). Critical values adjusted taking into account the degrees of freedom in Reimers (1992).

<sup>7</sup>The choice of the VAR was performed by using information criteria tests and diagnostic tests for autocorrelation. Results are available on requests.

Having found one cointegrating vector we next test the hypothesis that it has the form required by (2). The restricted cointegrating vector  $\hat{\beta} = (1, -1, -1, 1)$  yields the following statistics:  $\chi^2(3) = 23.61$  ( $pvalue = 0.00$ ) which enable us to reject the hypothesis.

In other words the rejection of the hypothesis above means that data cannot allow a representation of the form given by the pure interest parity condition,  $Sp = Er + Dr$ . The result is that total spread cannot be decomposed in the two unobservable components by using the swap spread and treating the default risk term as an I(0) residual. In other words the usual term  $Dr = Sp - Er$  turns out to be an I(1) non stationary variable whose dynamics, as we will see, is driven not only by country specific factor, but also by permanent international shocks.

**Model 2.** After rejecting model 1 we propose a more general model for estimating the interest rates parity condition which considers the role that the 3-months Eurospread, i.e. the spread between Italian and German short term rates ( $lit3m, dm3m$ ) and/or the level of the European interest rate (proxied by the German short term interest rate) may have in explaining and forecasting the dynamics of either the Exchange rate factor or the Default risk factor (See also Seghelini, 1996 and Tosato, 1996). These variables, (a) may give account of different position of Italy and Germany in the business cycle, (b) may contain relevant information of the policymaker about the credibility of future monetary stance, (c) may affect directly the total spread via the default risk, as in the case when an increase (decrease) in the level of short term rates has negative (positive) effects on public expenditure for interest repayment. The Eurospread, i.e.  $lit3m - dm3m$ , represents the elements of the total spread that can be ascribed to points (a) and (b) not totally captured in the data on swap contracts. Moreover  $dm3m$  represents the international component (point c).<sup>8</sup>

---

<sup>8</sup>This latter component could have been proxied by the US short term interest rate instead of the German one. However, different studies support the view that there is a strong degree of integration between US and German financial markets, in the form

The candidate variables are Italian and German three months interest rates. To test this possibility we run cointegration tests on  $X = (btp10y, bund10y, lit10y, dm10y, lit3m, dm3m)$ , also in this case  $p = 4$ . Cointegration tests are reported in Table 3 below.

**Table 3** Cointegration tests on  
 $X = (btp10y, bund10y, lit10y, dm10y, lit3m, dm3m)$

$H_0$	Trace	$T - nm$	95%
$r = 0$	101.1	97.6	98.33
$r \leq 1$	70.74	68.28	71.80

Notes: see notes in Table 2

The resulting cointegration relationship is given by the following vector:  $\beta = (1, 1.97, -0.66, -3.00, -0.12, -0.25)$ , and the corresponding vector of adjusting coefficients is given by:  $\alpha = (-0.046, -0.006, -0.015, 0.010, -0.006, -0.015)$ . The following hypothesis are implemented both to restrict cointegrating coefficients, as suggested by the interest rate parity condition, and to test for weak exogeneity.

$H(1)$	$\hat{\beta} = (1, -1, -1, 1, 0, 0)$	$\chi^2(5) = 13.14(pvalue = 0.02)$
$H(2)$	$\hat{\beta} = (1, -1, -1, 1, 0, \beta_6)$	$\chi^2(4) = 5.95(pvalue = 0.20)$
$H(3)$	$\hat{\beta} = (1, -1, -1, 1, 0, \beta_6)$ $\hat{\alpha} = (\alpha_1, \alpha_2, 0, 0, 0, 0)$	$\chi^2(8) = 12.52, (pvalue = 0.12)$ $\beta_6 = -0.63, \alpha_1 = -0.034, \alpha_2 = 0.011$
$H(4)$	$\hat{\beta} = (1, -1, -1, 1, 0, \beta_6)$ $\hat{\alpha} = (\alpha_1, \alpha_2 = -\alpha_1, 0, 0, 0, 0)$	$\chi^2(9) = 16.04, (pvalue = 0.066)$ $\beta_6 = -0.69, \alpha_1 = -\alpha_2 = 0.016$

of a statistical relationship given by a cointegrating vector (1,-1) between long term and short term rates in the two countries. See, for example, Goodwin and Grennes (1994) and Katsimbris and Miller (1993). This would imply that the role played by either rate in affecting the long term component of the default risk for the long term interest rate in Italy will be very similar. Moreover, the analysis is performed on a sample period close by to the establishment of the EMU, which makes the German rate particularly relevant.

H(1) rejects the hypothesis that both the three-months interest rates can be ignored in the long run relationship between  $Sp$  and  $Swspr10$ . H(2) does not reject the exclusion of  $lit3m$  from the cointegrating relationship, and hence we can reject the hypothesis that the 3 months Eurospread adds relevant information and only the level of the German short term interest rate must be considered. H(3) imposes a further restriction on the weak exogeneity structure across the variables considered; finally, H(4), restricts the adjustment coefficient to be the same in magnitude for  $btp10y$  and  $bund10y$ . H(3) and H(4) suggest that the only variable that adjusts to long run disequilibrium is  $btp10y - bund10y$ , i.e.  $Sp$ , in other words  $Swspr10$  and  $dm3m$  turn out to be weakly exogenous in the system<sup>9</sup>. The three-months Italian interests rate does not enter in the long run relationship and moreover does not enter into the short run dynamics of the equations for  $Sp10y$ ,  $Swspr10y$  and  $dm3m$ , in other words, it does not Granger cause any other variable (the Wald test  $\chi^2(12) = 18.33$  ( $p - value = 0.11$ ) enables us to accept these short run restrictions). For this reason, it has been excluded from the final model for the determinants of the total spread fluctuations. Let us now move to the final model.

**Model 3.** Having found a representation for the data which allows us not to reject an augmented interest rate parity condition as an equilibrium relationship among the following variables:  $X = (Spr10y, Swspr10y, dm3m)$  where  $Spr10y = btp10y - bund10y$ , and  $Swspr10y = lit10y - dm10y$ .

Table 4 shows the same long run properties of the data as described by H(4) in model 2 but estimated within a more parsimonious VAR. Trace statistics for testing the hypothesis of reduced rank enables us to accept one cointegrating vector at the 99% quantile. At the 95% quantile the hypothesis of full rank cannot be rejected and hence the variables in the

---

<sup>9</sup>In the sample period at hand, Italian lira was pegged to the DM. Provided that the pegging is credible we expect the interest rate parity condition to determine the long run level of interest rate in the Italian money market. The final implication, given the exogeneity of the interest rate on Bund, is that the endogenous variable within system the must be the total spread.



system would be considered as I(0) variables. However this cannot be the case, because this would contradict the results contained in Table 1, Table 2 and Table 3. In particular, Table 1 shows that the variables in the  $X = (Spr10y, Swspr10y, dm3m)$  vector are I(1). Table 2 also shows that there exists at most one cointegration vector among  $X = (btp10y, bund10y, lit10y, dm10y)$  henceforth, by adding another I(1) variable,  $dm3m$ , we can safely expect at most two cointegration vectors. Finally, cointegration tests on  $X = (btp10y, bund10y, lit10y, dm10y, lit3m, dm3m)$  show that there exists at most one cointegration vector (Table 3). By dropping,  $lit3m$  and by imposing H2 cannot increase the cointegration space. This evidence suggests us to rely on the results based on the 99% quantile.

In this case the (cointegration) residual represents only the country specific, stationary, component of the default risk. The default risk defined as  $Dr = Sp - Er$  also contains an international component whose dynamics, as we will see, is driven by the shock to  $Dm3m$ .

**Table 4** Cointegration tests on  
 $X = (Spr10y, Swspr10y, dm3m)$

$H_0$	<i>Trace</i>	$T - nm$	95%	99%
$r = 0$	40.23	39.53	29.7	35.65
$r \leq 1$	19	18.66	15.4	20.04
$r \leq 2$	6.38	6.27	3.8	6.65

*Notes:* see notes in Table 2

$H(5)$	$\hat{\beta} = (1, -1, \beta_3)$ $\hat{\alpha} = (\alpha_1, 0, 0)$	$\chi^2(3) = 4.62, (pvalue = 0.20)$ $\beta_3 = -0.61, \alpha_1 = -0.047$
--------	---	---

The weak exogeneity result in our data stated by  $H(5)$ , that is a negative significant  $\alpha_1 = -0.047$ , provides evidence that in the presence of deviations from the long run equilibrium the spread,  $Spr10y$ , adjusts to restore equilibrium while exchange factor and German short term interest rate do not. In particular,  $Sp10y$  adjusts to two *permanent* shocks (we have two common trends in the system) represented respectively by

shocks to *Swspr10y* and *dm3m*. We also expect to find that shocks to *Sp10y* have only *transitory* effects<sup>10</sup>

Having established the long run equilibrium properties of the system, we move to the analysis of its short run dynamics. Impulse response analysis (IRF) describes the dynamic response of the system to shocks of interest. However, this requires the identification procedure described in section 3.2, to allow the analysis of the system to a particular shock independently from other shocks. We have a system in three variables, hence the orthonormality of the variance-covariance matrix  $\Sigma_\varepsilon = I$  provides 6 non linear restrictions on the elements of  $S$ . To just-identify the model, we need 3 additional restrictions that come from choosing  $S$  to be the Choleski factor. This identification scheme implies that while the contemporaneous effects in the VAR are triangular, lagged dynamics are completely unrestricted leaving the data freely describe the dynamic interdependence we are interested in examining here. We are confident with the Choleski identification scheme because the innovations of our reduced form system are nearly uncorrelated<sup>11</sup>.

Figure 2 shows the estimates of  $C(L)$  transformed to express the variables as the sum of distributed lags of the structural disturbances. In particular, it summarises the responses to one shock respectively to *dm3m* (first column), *Swspr10y* (second column) and *Sp10y* (third column).

The transitory shock to *Sp10y* has a null effect on the other two

---

<sup>10</sup>At the end of November 1996, Italy rejoined the ERM following a four year absence. Our results are robust with respect to this change of regime. In particular, the cointegrating vector and the weak exogeneity structure, as restricted in *H5*, cannot be rejected by the data when estimated on the subsamples covering periods: Jan95-Ott96 ( $Pvalue(\chi_4) = 0.5$ ), Jan96-Nov96 ( $Pvalue(\chi_4) = 0.5$ ), Nov96-Ott97 ( $Pvalue(\chi_4) = 0.3$ ). Detailed results are available on request.

<sup>11</sup>This approach is used for example by Canova and De Nicolò' (1997): " If the innovations in the reduced form system are uncorrelated, all the identification schemes which impose restrictions on the contemporaneous impact of shocks will produce the same the results. In the less extreme case where innovations are of the reduced form system are *nearly* uncorrelated results will be qualitative robust to alternative identifications schemes which impose restrictions on the covariance matrix of the shocks". See also Canova (1995b) for a discussion of this point.

variables. On the other hand  $Sp10y$  is permanently affected by shocks to  $Swspr10y$  and  $dm3m$  which is consistent with the view that interest rate spread is the endogenous variable which adjusts to long run equilibrium, as required in an exchange rate pegging regime as the one working within the sample considered.

$dm3m$  is unchanged both in the short and long run when  $Sp10y$  and  $Swspr10y$  are shocked implying that the variable is not Granger caused by the other variables within the system as already tested in section 3.3. This is consistent with the interpretation of this shock as the exogenous level of the European interest rates. A shock to the  $dm3m$  (that is the international shock) has a permanent effect on the other two variables.

The dynamic responses of  $Sp10y$  and  $Swspr10y$  to a shock to  $dm3m$  have a similar shape. However, these responses are not identical. The difference between the two responses suggests a quite interesting interpretation and allows us to address the point raised before on the nature of the source of the reduction of the spread (idiosyncratic versus international).

If the dynamic responses were completely identical, we should have inferred that any shock to the international interest rate affected the total spread only through the exchange rate factor. In other words, since the difference  $Sp10y - Swspr10y = DR$  we would get the result that the shocks to the international interest rate would not affect the default risk on Italian bonds rates.

We get the opposite result: the effects of the international shocks are not totally captured by changes in the expectations embodied in the changes of the exchange rate factor. Instead, in our data the difference in the dynamic shape of the interest rate spread and interest rate spread on swap contracts suggests that a shock in the European level of the short term interest rate modifies the Italian default risk factor that responds to this shock as plotted in Figure 3. As a conclusion, the default risk dynamics in Italy during the period under analysis is driven not only by a country specific component (credibility gain), but also by an international component.

Finally, it is worth noticing that the long run value of the shock to  $Swspr10y$  is immediately reflected one to one in the long run value of  $Sp10y$ .

## 4 Conclusions

This paper studies the determinants of the total interest rate spread on government bonds between Italy and Germany on the period 1995-1997. The usual uncovered interest rate parity condition does not hold in the period considered. This equilibrium condition between 10-years BTP - BUND spread, spread on swap contracts ITL-DEM reflecting the currency risk (also defined as exchange rate factor) and default risk (also defined country risk, mainly due to the high Italian debt) must be augmented by German short term interest rate.

The impulse response analysis performed suggests that the level of the short term German interest rate permanently affects both the exchange rate factor and the default risk factor. The implication of this result is that the reduction of the spread in the period studied was due both to credibility gains and to favorable dynamics in the German interest rate which reduced both the exchange rate term and the default risk term.

## REFERENCES

Campbell, J.Y., and Shiller, R.J. (1987), Cointegration and Tests of Present Value Models, *Journal of Political Economy*, **95**, 1063-1088.

— and Perron, P. (1991), Pitfalls and Opportunities: What Macroeconomists Should Know about Unit Root, O.J. Blanchard & Fisher, eds. *NBER Macroeconomic Annuals*, Cambridge, MA:MIT Press, 141-201.

Canova F., (1995a), VAR models, in *Macroeconometrics*, edited by Hoover K.D., Kluwer Academic Publishers.

— (1995b), VAR Specification, Estimation, Testing and Forecasting, in Pesaran H., and M. Wickens, *Handbook of Applied Econometrics*, 31-65.

—- De Nicolo' G. (1997), Stock returns, term structure, inflation and real activity: an international perspective, *Materiali di Discussione* n. 192, Dipartimento Economia Politica, Universita' di Modena.

Caporale G.M. and Pittis N., (1997), Domestic and External Factors in Interest Rate Determination, *Applied Financial Economics*, **7**, 465-471.

Cochrane, J. (1988), How Big is the Random Walk in GNP?, *Journal of Political Economy*, **109**, 241-265.

Doornick, J.A., and Hendry, D. (1994), PcFiml 8.0 An Interactive Econometric Modelling of Dynamic System, International Thomson Publishing.

Fama, E. F., and Bliss, R.P. (1987), The Information in Long Maturity Forward Rates, *American Economic Review*, **77**, 680-692.

Favero, C., Giavazzi, F. and Spaventa, L., (1997), High Yields: The Spread on German Interest Rates, *Economic Journal*, **107**, 956-985.

Goodwin, B. K. and Grennes, T.J. (1994), Real Interest Rate Equalization and the Integration of International Financial Markets, *Journal of International Money and Finance*, **13**, 107-124.

Johansen, S. (1991), Estimation and Hypothesis Testing of Cointegrating Vectors in Gaussian Vector Autoregressive Models, *Econometrica*, **59**, 1551-1580.

Katsimbris G.M. and Miller S. M., (1993), Interest Rate Linkages within European Monetary System: Further Analysis, *Money, Credit and Banking* **25**, p.771-779.

Osterwald-Lenum, M. (1992), A Note with Fractiles of the Asymptotic Distribution of the Maximum Likelihood Cointegration Rank Test Statistics: Four Cases, *Oxford Bulletin of Economics and Statistics*, 54: 461-72

Pistoresi, B. (1997), Radici Unitarie e Persistenza: L'Analisi Univariata delle Fluttuazioni Economiche, *Rivista di Politica Economica*, **LXXXVII**, 111-152.

Reimers, H. E. (1992), Comparisons of Tests for Multivariate Cointegration, *Statistical Papers*, 33, 335-339.

Seghelini, M. (1996), The 10 year ITL/DEM swap spread equilibrium: the estimate of a cointegrating vector, in *Simulation and Forecast of the Italian Yield Spread*, Background Papers, September 1996, Deutsche Bank Research, Milan.

Shea, G. (1992), Benchmarking the Expectations Hypothesis of the Interest Rate Term Structure: An Analysis of Cointegration Vectors, *Journal of Business & Economic Statistics*, **10**, 3, 347-366.

Tosato, C. (1996), Measuring EMU expectations embodied in the 10 year BTP-BUND spread, in *Simulation and Forecast of the Italian Yield Spread*, Background Papers, September 1996, Deutsche Bank Research, Milan.

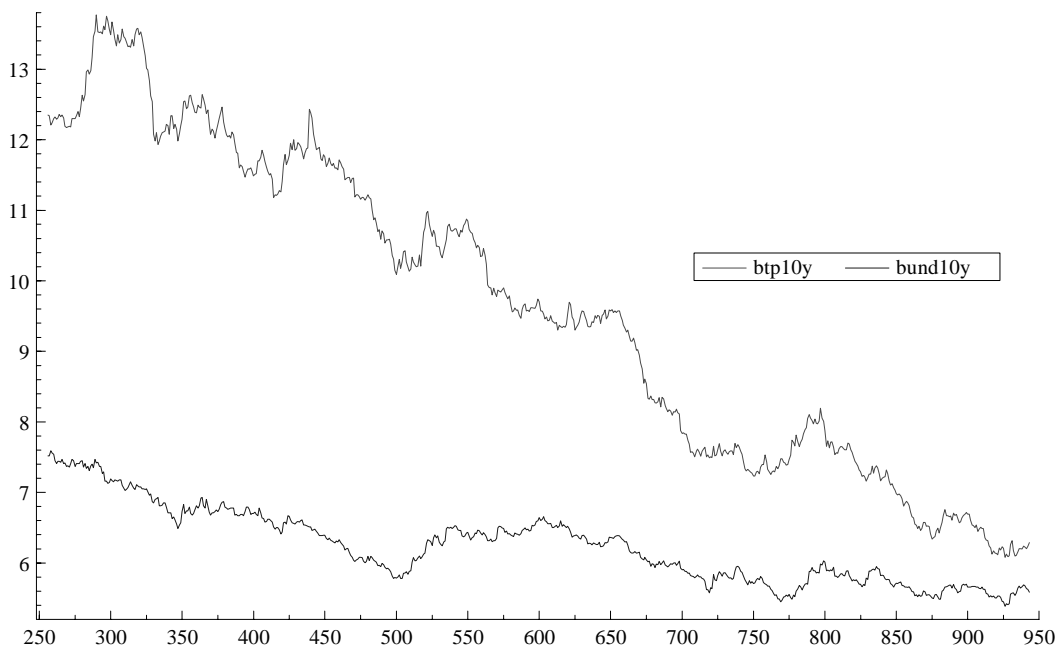


Figure 1: Long term interest rates



Figure 2: Short term interest rates



Response to One S.D. Innovations  $\pm 2$  S.E.

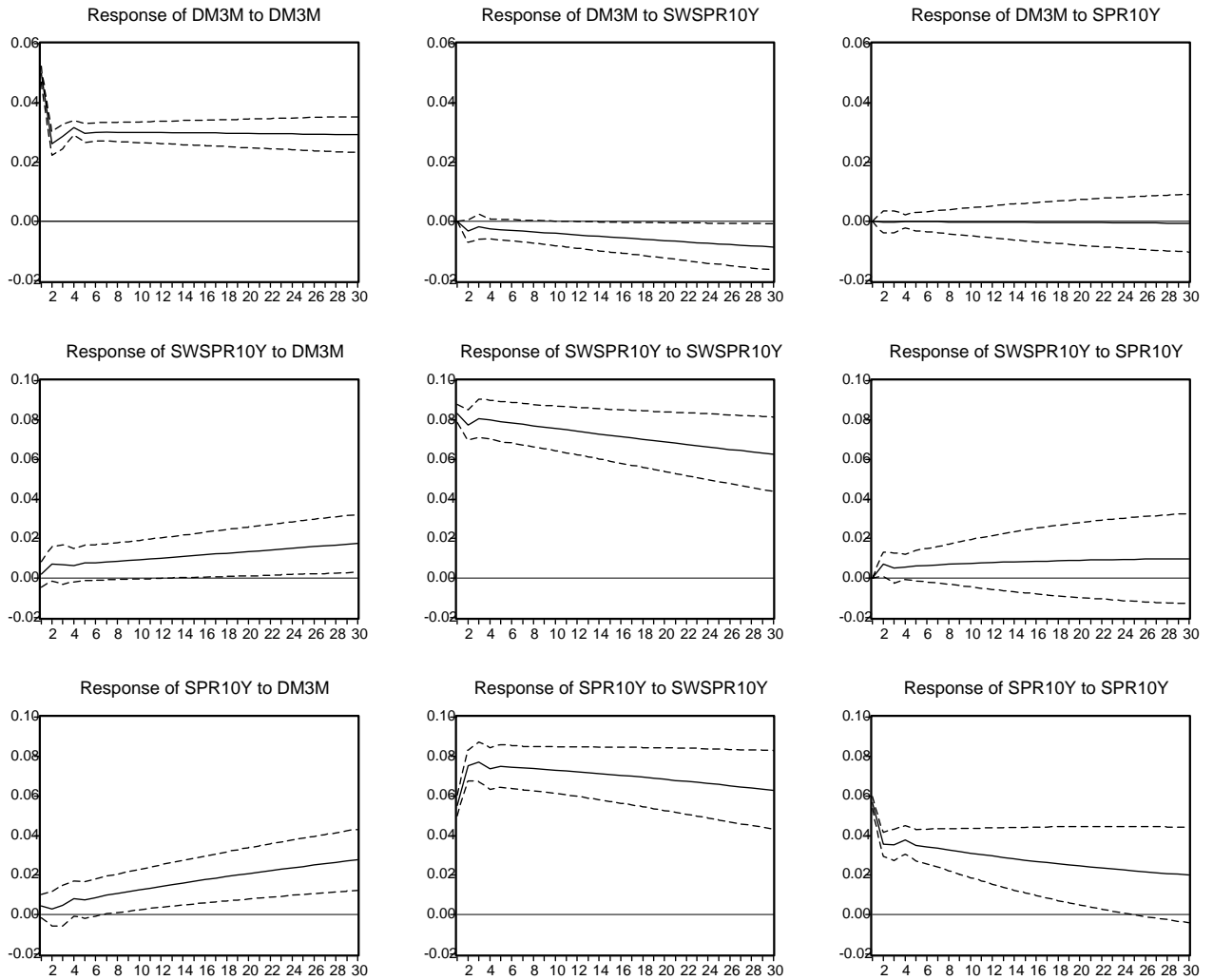


Figure 3: Impulse Response Functions

Response to One S.D. Innovations  $\pm 2$  S.E.

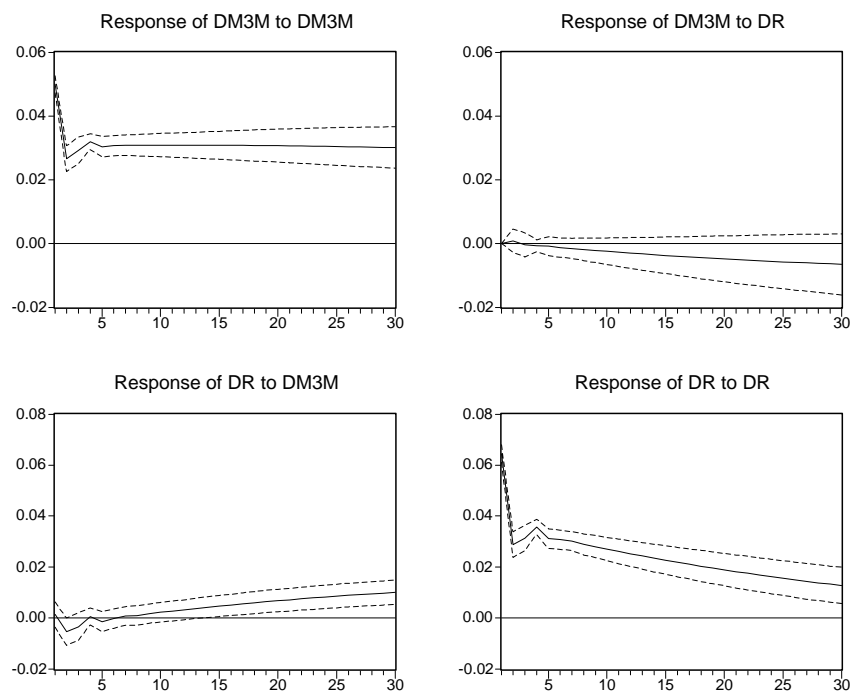


Figure 4: Impulse Response Functions