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A less effective monetary transmission in the wake of EMU? Evidence from lending rates pass-through

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Abstract

An approach to search for structural breaks in lending rates pass-through in the wake of EMU is proposed and implemented for Italy and Portugal. Breakpoints cluster in the second semester 1999 and the equilibrium pass-through on short term lending is, in contrast with earlier research, sizeably *lower* in the post-break period; in the Italian case, the adjustment to the equilibrium is however *faster*. The recently proposed distinction between monetary policy and cost-of-funds approaches does not yield different break-points. These results challenge the widely held view that EMU has in its wake enhanced the effectiveness of monetary transmission. A strengthened relationship lending could at least partly explain the reduced equilibrium pass-through in the Italian case.

JEL Codes: E43; E52; E58; F36

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1. Introduction*

The transmission of monetary policy hinges on how bank rates react to changes in the money market rates, especially in a bank-based economy. Several studies investigate whether size and speed of the pass-through (PT) of monetary policy impulses to retail bank interest rates in the euro zone have increased in the wake of Stage Three of EMU, thus enhancing the effectiveness of the single monetary policy, and converged, thus rendering more uniform the transmission via the banking sector.

Angeloni and Ehrman (2003) argue that since January 1999 PTs have become on average larger and faster across the euro area and in most of the largest countries, thus strengthening the transmission of monetary policy. Some papers however challenge this influential view, first of all because break-points should not be exogenously assumed. A second debated issue, within the PT empirical literature, is the often arbitrary choice of the driving market rate, that should instead match the maturity of the loans the lending rate refers to. Thirdly, no structural break has been found when modelling an index of lending rates in an euro area monthly monetary model (Bruggeman-Donnay 2003).

The variety of results in the literature motivates this paper, that proposes a new approach to reassess the evidence of structural breaks in the equilibrium PT for short term lending rates in two case studies: Italy and, though less thoroughly examined, Portugal. These are the two EMU countries where, according to recent studies, break-point dates are up to 4/5 years apart, when different driving market rates are chosen (Sander-Kleimeier 2004b, Table 1). The main contribution of this paper is to show that in both countries recent break-points cluster instead in the second semester 1999, thus excluding expectations effects of the monetary unification, with sizably *reduced* equilibrium PTs; in the case of Italy, the adjustment to the equilibrium is however *faster*.

The paper is organized as follows. Section 2 surveys the recent literature on retail interest rates PT in the EMU, with special reference to Italy and Portugal. Section 3 proposes a new methodology to search for break-points in cointegrated relations and implements it for Italy. Section 4 replicates more concisely the exercise for Portugal. Section 5 offers a tentative interpretation of the findings for Italy. Section 6 concludes.

2. The pass-through in the wake of Stage Three of EMU

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The empirical literature on bank rates PT in the EMU shares the same theoretical framework but often produces conflicting results, owing to different approaches in the econometric investigation. The reference setting is the standard Klein-Monti model of a monopolistic bank - with risk neutrality, perfect information, no switching costs and no joint production of loans and deposits -, easily extended in an oligopolistic structure of the banking industry (Freixas-Rochet 1997). The lending rate is determined as a mark-up over the marginal (opportunity) cost, identified with the money rate directly influenced by the central bank. In empirical applications, assuming a linear approximation, the marginal cost coefficient is interpreted as the equilibrium PT, with a reference unitary value in a competitive market and lower values (i.e. higher variable mark-up) in the monopolistic case. Within this framework, studies differ mostly on how to proxy the marginal cost, in order to match the maturity of the credit aggregate underlying the lending rate.

The estimates of the impact and equilibrium PT parameters are usually obtained reparametrizing an Autoregressive Distributed Lags (ADL) specification, as originally proposed in Cottarelli-Kourelis (1994), as an Error Correction Mechanism (ECM), following the Granger representation theorem for cointegrated variables.

Let a long run equilibrium or cointegrated relation between interest rates integrated of order one, or I(1):

$$r_t = \alpha + \beta\, rm_t + \varepsilon_t \qquad \varepsilon_t \sim NID(0, \sigma_\varepsilon^2) \tag{1}$$

with I(0) OLS residuals, *ecm*, at the first stage of the Engle-Granger (1987) two-step estimation procedure (EG), where:

- r = bank rate;
- rm = driving market interest rate;
- ecm = stationary deviation ("error" in the ECM acronym) of the bank rate from its long run equilibrium value, assumed to be a linear transformation of rm.

In the EG second stage the short term dynamics parameters are estimated from:

$$\Delta r_t = \theta ecm_{t-1} + \gamma_0 \Delta rm_t + \sum_{i=1}^{k} \gamma_i \Delta rm_{t-i} + \sum_{j=0}^{k'} \lambda_j \Delta r_{t-j} + u_t \qquad u_t \sim NID(0, \sigma_u^2) \tag{2}$$

where Δ is the first difference operator.

The key parameters, from an economic point of view, are γ_0 and β, that is the impact and equilibrium PTs, and θ, that is the speed the error is corrected.

The empirical choice of the (weakly) exogenous driving market rate motivates the recently proposed distinction between a "monetary policy approach" (MPA), with the overnight rate taken as a proxy for the monetary policy stance, and an industrial organization inspired "cost-of-funds approach" (CoFA), with the market interest rate better proxying the marginal cost of loaned funds (Sander-Kleimeier 2004a, b). The difference between the two approaches depends on how the monetary stance is thought to influence the very short end of the yield curve, possibly in relation with agents' expectations. The choice of a specific market rate or, alternatively, of a combination of several ones, to proxy the "true" marginal cost, as in de Bondt *et al* (2003), depends on the range of maturities of the credit aggregate underlying the lending rate. Similar choices can in fact distort comparisons across countries, if lending practices have quite different repricing schedules. This is one likely cause of the heterogeneity in the results of the literature on short term lending rates PT in the EMU, the issue this paper focuses on.

The heterogeneity refers to the date of structural breaks, possibly associated with the advent of Stage Three of EMU, as well as to the size of equilibrium PTs and the speed of adjustment. Angeloni and Ehrmann (2003) argue that a single bank reserves market and the reduction in the market interest rates volatility due to the operating procedures of the ECB have already produced larger and faster bank rates PTs. They report, having identified informally, via rolling-window regressions, January 1999 as a break-point, that both the impact and the maximum size of PTs for a set of lending and deposit rates have on average sizably increased in the period 1999-2002, compared to 1990-1998, in four of the largest EMU countries, Germany being the exception, and in the euro area as a whole. Business loans in the euro area show between the pre-1999 and the successive period the largest increase in the impact and peak coefficients, the latter reaching 1.1. de Bondt (2005), on the contrary, finds that for all euro area retail bank rates, except the mortgage rate, the equilibrium PT in the sample period after the introduction of euro (January 1999-June 2001) is lower than in the extended period (January 1996-June 2001). In particular, using a one-step ECM model, β shrinks from 1.53 to 0.88 for the short term lending rate to firms. A Chow test rejects at the 5% significance level the null of no structural break at January 1999.

These conflicting findings could be, *inter alia*, motivated by the unharmonised national data for bank rates underlying euro-area data. The empirical literature exploring national data series, however, provides even more heterogeneous outcomes - on break-points and PTs – as shown in the case of short term lending rates, the ones that better match for maturity the driving market rates (short-term lending rate to firms, r^{ST} , for Italy and rate on commercial bills, r^{CB} , for Portugal; on data definition see Section 3 and Appendix 1; Tables 1-2).

de Bondt *et al.* (2003) *do not* detect a structural break in January 1999 in Italy, in an empirical framework with the distinguishing feature of a driving market rate proxied by a combination, with estimated weights, of the 3-months interbank and of the 10-years Government bond rates, under the assumption that the latter provides a signal on the persistence of changes of the short term rates. This finding notwithstanding, the paper provides estimates for the entire April 1994-December 2002 period and for the subsample beginning January 1999. They show a small *reduction* in the impact PT and a larger one in the equilibrium PT of the interbank rate (from 0.93 to 0.76), whereas the equilibrium parameter with respect to the 10-years rate reverses sign, from positive (0.12) to negative (-0.15). The estimation results are similar and even sharper for Portugal, with equilibrium PT falling from (sizably) above to less than one.

Sander and Kleimeier (2004a) estimate Eq. (1) with alternative driving market rates - overnight (rm^{ON}; MPA) and one-month interbank (rm^{INT}; CoFA) - and empirically determine whether, between January 1993 and October 2002, a single structural break is occurred; once detected it, they provide an EG estimate of Eq. (2) before and after the break-point. Their findings for Italy are that *a*) according to the MPA, the break-point is February 1995 and the equilibrium PT reduces slightly below unity, with a larger decrease of the impact PT, whereas *b*) for the CoFA the break-point is July 1999, and a slight increase in the impact PT is associated with a sizable reduction for the equilibrium one (from 1 to 0.7). Considering also r^{TOP}, namely the minimum lending rate for top rated firms, a break-point occurs in February-95 only with the MPA; the equilibrium PT remains however pretty unchanged, near unity. The findings are more striking across the two approaches for Portugal[1]. The break-points are more than 5 years apart - July-94 and October-99 – and the equilibrium PT varies very considerably, rising with the MPA (from 0.26 to 1.52) and falling with the CoFA (from 1.24 to 0.65) for r^{CB}. The same pattern in dating break-points and estimates of equilibrium PTs occurs when considering the short term lending rate to firms, r^{STF}.

[TABLES 1 AND 2 APPROXIMATIVELY HERE]

3 – The econometric investigation in the Italian case

3.1 - Methodology

To endogeneously search for possibly EMU-related structural breaks and, once detected, successively estimate the impact and equilibium PTs, as well as the speed of correction of the "error", the methodology proposed in this paper is as follows.

1) Endogeneous search for break-points in the long run model (Eq. 1), adopting the supremum F (supF) testing procedure, where the date is associated with the largest of the standard

rolling Chow F-statistics computed under the hypothesis of a break occurring in each subsequent period through the mid-70% sample period (Andrews 1993), for the case of I(1) regressors (Hansen 1992). The innovation we propose, following Bai (1997), to this procedure, implemented by Kleimeier and Sander (2004a,b)[2] and Toolsema *et al* (2002), is that, with several local maxima, statistically significant[3], the algorithm should be repeated, starting from the earliest break-point, to pick up the latest one, which it is more interesting from an EMU perspective.

2) Check that lending and driving market rates are I(1) through Augmented Dickey-Fuller (ADF) tests in the break-free periods.

3) Check that in the same periods the OLS estimation at the first stage of Eq. (1) generates I(0) residuals, thus rejecting the null of no cointegration. This should help mitigate the well known problems of low power of tests for cointegration in the presence of breaks (Maddala-Kim 1998).

4) If the no cointegration hypothesis is rejected, estimation of an ECM specification (Eq. 2) following the EG procedure. The optimal number of lags, allowing for a maximum of three, is determined according to the minimum Akaike Information criterion. The first-stage β estimate is superconsistent, but biased in small samples, and the bias is inversely proportional to the fit (Banerjee *et al.*, 1986). For robustness, an alternative one-step, general-to-specific (Hendry 1995), procedure is therefore also used, to jointly estimate short term and equilibrium PTs, in an ECM specification combining Eqs. (1) and (2) through suitable non-linear restrictions:

$$\Delta r_t = \theta(r_{t-1} - \alpha - \beta\, rm_{t-1}) + \gamma_0 \Delta rm_t + \sum_{i=1}^{k} \gamma_i \Delta rm_{t-i} + \sum_{j=1}^{k'} \lambda_j \Delta r_{t-j} + v_t \qquad v_t \sim NID(0, \sigma_v^2) \quad (3)$$

This procedure, justified when the explanatory variable is (weakly) exogenous for β, as it is safe to assume for a driving market rate set in the highly integrated European money market, has two interesting features. First, an alternative check of cointegration can be implemented through the Ericsson-MacKinnon (2002; EM) test, adjusted for the degrees of freedom, specifically designed for this estimation procedure; second, it allows to assess the precision in estimating β, the t-ratio at the EG first stage being not interpretable as usual.

5) The PT can be potentially asymmetric for positive or negative changes in the driving rate: as a consequence, a robustness check of the results obtained for the basic symmetric specification is carried out introducing separate regressors, according to their sign, for the short

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[1] Owing to the availability of the data used, the starting date is January 1993 for MPA and October 1994 for CoFA, with the consequences on the results detailed in Section 4.
[2] Mid-80% sample period in their studies.

term dynamics. An extension of the exercise to the equilibrium parameters was deemed problematic given the limited length of the sample period.

3.2 The data

The short term lending rates analyzed are, as in the literature surveyed, r^{ST} for the larger credit aggregate and r^{TOP}, both drawn from the ECB *National Retail Interest Rates* (NRIR) database (see Appendix 1). This paper adds the overdraft rate r^{OD}, because it may provide further insights to the PT empirical analysis because it improves on r^{ST} on two accounts: a) it refers to contracts with homogeneous characteristics for (low) maturity and interest rate fixation; b) the underlying stocks outstanding are not the result of a stratification of past 18 months market conditions, being rather more similar to new businesses, as it happens for the NRIR lending rates of all other EMU countries. In spite of the high correlation even of the differenced series of r^{ST} and r^{OD} (see Table A1), the econometric investigation for the two rates produces indeed quite different results.

The sample period goes from January 1993[4] to February 2004, the last month the series are available in the NRIR database (Figure 1).

[FIGURE 1 APPROXIMATIVELY HERE]

3.3 – Results

The results are reported following the sequence of steps of the proposed methodology.

1) The rolling Chow F-test statistics, after a local maximum at the beginning of 1995 for all bank rates, follow differentiated paths afterwards[5]. The statistics settle around a plateau between 1998 and 1999, and the results are invariant with the choice of the driving rate, for both r^{ST} and r^{OD}. The path is much smoother for r^{TOP}, signalling no structural change in the relation with rm^{INT} after the first break, whereas with rm^{ON} a statistically significant local maximum is detected at end-1996 (Figures 2a-b).

Repeating the algorithm for the period after the last break the supF procedure indicates for each bank rate only one statistically significant absolute maximum (Figures 3a-b):

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[3] The critical asymptotic values of the supF with I(1) regressors are 16.2, 12.4 and 10.6, at the 1%, 5% and 10% significance levels, respectively (Hansen 1992, Table 1).

[4] Data for many EMU countries are available only since early 1990s. Following the literature, the sample period starts after the 1992 EMS crisis.

- r^{ST} and r^{OD} (CoFA): June 1999;
- r^{ST} (MPA): July 1999;
- r^{OD} (MPA): November 1999;
- r^{TOP} (MPA): October 1997.

[FIGURES 2 E 3 APPROXIMATIVELY HERE]

It is worth remarking that the date of the latest structural break for r^{ST} is, according to the MPA, more than 4 years later than in Sander-Kleimeier (2004a,b), while it almost coincides with the date suggested by the CoFA..

As an encompassing check we investigated whether these earlier results could be explained following our proposed approach. Using the same sample period – from January 1993 to October 2002 - and data set[6] the supF procedure considering only the absolute maxima detects in fact almost identical break-points (April 1995 with MPA and June 1999 with CoFA for r^{ST} and March 1995 for r^{TOP} with MAP[7]; Figure A1). The interesting feature is, however, that repeating the supF procedure after these dates yields findings very similar to our own: the late break-points are June 1999 for r^{ST} with *both* approaches and October 1997 for r^{TOP} with MAP (Figure A2).

2) Checking the order of integration of each interest rate series in the pre- and post-break periods shows that the null of integration of order one cannot be rejected, at high confidence levels, for the market interest rates; the same does not hold for the lending rates (Table A2). More precisely, using first the ADF test for the series in levels and, if satisfied, also in differences, the null is rejected in the pre-break period for each bank rate in level form, though the statistic is near the critical values; in the case of r^{OD} the confidence level reaches the 1% *p*-value. The test for the differenced series is instead always satisfied, at least at the 5% significance level. Due to the low power of these asymptotic tests for relatively small samples the EG procedure is anyway implemented; it is stopped at the first stage when the no cointegration null fails to be rejected.

3) In the pre- and post-break periods Eq. (1) is estimated at the first stage associating each lending rate with either rm^{ON} or rm^{INT}; only one relation is estimated for the couple r^{TOP} and rm^{INT}. Standard Cointegrating Regressions Durbin-Watson (CRDW) and ADF tests are computed, under the null of non-stationary OLS residuals (Table 3).

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[5] The first local maxima occur in January and March 1995 with rm^{INT} and rm^{ON} as driving rate, respectively. To search for the successive break-point the sample starts at April 1995, in order to exclude, as in earlier studies, the March outlier, with approximately 200 basis points increase in market rates.

[6] Overnight and 1 month interbank rates are drawn from Datastream.

[7] For unspecified reasons the estimation sample for r^{TOP} with CoFA starts at July 1994 in Sander-Kleimeier (1994b), the month with an absolute maximum for supF according our encompassing exercise.

These tests, with a well known low power in small samples, provide divergent outcomes and the statistics are generally close to critical values. The null is however almost always rejected at the 10% significance level at least by one test. The exception is r^{OD} in the pre-break period, only when the driver is rm^{INT}. This result is interesting, because the earlier finding of the r^{OD} series failing the unit root test, even at the 1% significance level, would have led to expect an outcome of no cointegration with *any* I(1) market rate.

4) The EG procedure with Eq. (2) for the three lending rates produces overall smaller and less precise estimates of the speed of error correction, θ, compared with the single-equation alternative. This is reassuring, because a persistent disequilibrium, if the cointegration relation is accepted, is implausible from an economic point of view. Symmetrically, a cautionary note is raised when, though the long run relationship between r^{OD} and rm^{ON} passes the cointegration test in the pre-break period, the estimated θ in the second step turns out to be not significantly different from zero, and the same happens for r^{ST} in the pre-break period.

The EM test, computed as a t-ratio of the OLS estimate θ for Eq. (3), is useful because it focuses on the economic meaning of the parameter: considering again the r^{OD} case, only with rm^{INT} as a driving rate the test is passed and a plausibile θ is obtained (one fifth of the error is corrected in a month). The EM test turns out to be more severe, as it can be gathered from the several instances of failures with rm^{ON} as driving rate. These results suggest that rm^{INT} is empirically preferable to rm^{ON}, despite of their high correlation, at least partly because of the lower variability of the former.

Let us consider now the more interesting parameter for a PT study. The results of the EG procedure can be summarized as follows.

[TABLE 3 APPROXIMATIVELY HERE]

r^{ST} and r^{OD} . As expected, owing to the high correlation between the market rates and even more among the bank rates (Table A1), the results are similar with either driving rate. The estimated β in the post-break period is sizably below unity and shrinks by at least one fourth compared to the previous period (from 0.95 to 0.7 for r^{ST}); in the EMU period, the β for r^{ST} is lower by about one sixth compared to r^{OD}, thus signalling how "special" are overdraft loans..

The main findings with the one-step procedure do not differ greatly, though they provide some interesting integrations. β is significantly different from unity in the post break period: the largest estimate, at a 95% confidence level, is less than 0.8 for r^{ST} and even 0.7 for r^{OD}. The estimates are systematically lower than those obtained with the EG procedure: for r^{ST}, the difference is of the order of magnitude of the 95% confidence interval, a clue to a small sample bias (Banerjee *et al* 1986). β is instead always above or equal to unity in the pre-break period.

The estimates for the impact PT, γ_0, a typical index of rate stickiness, are almost identical between the two periods for r^{ST}, whatever the estimation procedure or the driving market rate. The results for r^{OD} are comparable to the long run ones: for instance, with rm^{INT}, cointegration tests passed in both periods, γ_0 shrinks by almost a half (0.11 instead of 0.19) and is about one third of the corresponding parameter for r^{ST}.

To sum up on the adjustment speed, θ doubles between the early and the late period for r^{ST} and increases by about a half for r^{OD} [8]. Considering only, to save space, the results with rm^{INT}, 90% of the adjustment toward the equilibrium PT is realized within three months (66% in the pre-break period) and completed within a semester (within a year before); Table 4. The evidence is therefore of a *faster* PT for the main lending rates.

r^{TOP}. The findings in the case of the lending rate for the top rated firms are similar across approaches for the equilibrium PT: close to unity for the extended period April 1995 - February 2004 in the CoFA; shrinking from unity to approximately four fifths in the MPA. The one-step estimates of θ are again higher in the post break-period, though smaller in comparison with the other bank rates; γ_0 is slightly higher than for r^{ST}.

[TABLE 4 APPROXIMATIVELY HERE]

5) Taking into account, to save space, only the one-step procedure to investigate whether asymmetric effects can be detected for positive and negative changes of rm^{INT}, the estimation of separate short run dynamics coefficients does not produce a better fit (Table 5). We report the estimates only for the post-break period, that differently from the earlier one includes both negative and positive changes in the driving rates. Two features are worth mentioning: first, the null of symmetry cannot be rejected for r^{ST} and r^{OD} in the recent period; second, the null for r^{TOP} is instead rejected and impact PT is *higher* for *negative* changes in the driving rate.

[TABLE 5 APPROXIMATIVELY HERE]

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[8] The somehow compensating changes of β and θ in the post-break period could raise suspicions of identification difficulties owing to the relatively short sample period and the low variability of regressors. As a check we used a linear specification for the one-step procedure, leaving the data to determine the coefficient for the (lagged) driving rate. To this end we imposed β=1 in the ECM regressor and introduced as a further regressor the lagged once market rate, in levels. The results (available on request) provide support to the estimates presented in the paper.

Overall, in the Italian case, the econometric exercise points to a *lower equilibrium PT* associated with Stage Three of EMU for r^{ST} and even more for r^{OD}, with values sizably below unity, while no change is detectable for r^{TOP}.

The findings differ from some recent research, with the partial exception of CoFA results of Sander-Kleimeier (2004a, b). The one-step estimation procedure, beside the widely used EG one corroborates these results, because it overcomes the critique of small sample bias for superconsistent estimates. The distinction between MPA and CoFA, linked to the choice of how to proxy the driving market rate, does not provide, in contrast to Sander-Kleimeier (2004a, b), a distinctive contribution to date structural breaks in equilibrium PTs, provided a suitable search methodology is adopted. Overall, the interbank as the driving market rate proves empirically more reliable than the overnight one.

The proposed approach tested in the Italian case is corroborated when considering the Portuguese one (Section 4); a tentative interpretation for Italy, based on banking structure considerations, is put forward in Section 5.

4. The Portuguese case

Portugal is the other country, with Italy, that according to Sander-Kleimeier (2004b, Table 1) has, for r^{CB} and r^{STF} (Figure 4), break-points very far apart, up to 5 years, according to the MPA or the CoFA (Table 2). Moreover, in contrast to Italy, the cointegration hypothesis at the first stage of the Engle-Granger procedure is sometimes rejected[9]; the range and the direction of changes in the equilibrium PTs between pre- and post-break periods are hardly plausible.

This section investigates whether, on a sample that for both bank rates starts in January 1993 and ends at October 2002, those findings are confirmed adopting the proposed approach to search for structural breaks. The focus is on dating break-points and on estimating equilibrium PTs, because earlier studies do not provide benchmark estimates for γ_0 and θ. It is worth recalling that the results are not comparable with the Italian ones, owing to the different underlying credit aggregates (see Appendix 1).

[FIGURE 4 APPROXIMATIVELY HERE]

The supF testing procedure, using either driving market rate, detects a first local maximum at the beginning of 1995 (Figure 5). Replicating the procedure for the subsequent period, starting as

[9] In these cases the estimation procedure relies on ADL specifications.

usual from April 1995, to remove the outlier of March with its rates spike, indicates a maximum, *common* to both bank rates and for either approach, in October/November 1999 (Figure 6). The EG first stage estimates of Eq. (1) in the pre- and post-break periods reject the null of no cointegration (Table 6)[10]. This outcome reinforces the proposition, suggested by the Italian case, that the distinction between MPA and CoFA is not empirically useful when searching for structural breaks. It is worth noticing that our results are very similar to Sander-Kleimeier (2004a; Table 2) when considering the early break-point with rm^{ON}. The almost 5 years difference when considering rm^{INT} turns out to be explained by their choice to use a series, drawn from Datastream, available only from October 1994 (results available on request). Extending the series back to January 1993 using national data yields, in the supF exercise, the results shown in Figures 5-6.

A remarkable result in the Portuguese case is the almost halving of β across periods for both lending rates, and the very similar values of the parameter. An even larger reduction of the PT than in Italy, at least for r^{STF} , adds further evidence against the Angeloni-Ehrman (2003) view for the euro area, though Portugal is not specifically analyzed in that study.

[FIGURES 5 AND 6 APPROXIMATIVELY HERE]
[TABLE 6 APPROXIMATIVELY HERE]

5. A tentative interpretation for the Italian case

A lower equilibrium PT on short term lending in the wake of Stage Three of EMU implies a less effective monetary transmission of the ECB impulses to the Italian and the Portuguese credit markets and, plausibly, owing to the banking intermediation key role for bank dependent SMEs, to the real economy.

In the Italian case, an interpretation, within a static industrial organization scheme *à la* Klein- Monti, of a mark-up increase runs contrary to recent studies documenting that the banking sector has become more competitive during the 90s (Angelini-Cetorelli 2003). The declining, though still higher than the euro area average, differential between lending and borrowing rates through 2003 (Oecd 2005) supports this view. The static industrial organization scheme can be however hardly applied to lending rates determination, especially in a period when bank-customer relationships underwent dramatic changes.

The econometric investigation in this paper provides a suggestive piece of evidence: the divergent pattern of equilibrium PTs for the minimum rate r^{TOP} in comparison with the lending

[10] The ADF tests, always passed, for the I(1) feature of the bank and market rates series, in the pre- and post-break

rates to non-primary borrowers, r^{ST} and r^{OD}. This pattern is compatible with a credit market where top rated borrowers have kept exploiting their bargaining power with banks, with lending rates close to money market ones, whereas enhanced relationship lending with the bulk of customers has produced the expected intertemporal interest rates smoothing (Berlin-Mester 1998). Asymmetric effects for the short run PT lend additional support to this interpretation: while the parameter for r^{TOP} is in absolute terms lower when the market rate increases and viceversa, there is no such evidence for the other rates (Table 5).

Two classical indicators of credit market structure – the extent of multiple lending relationships and the share of the main bank's loans – point to a strengthened relationship lending, especially after the introduction of euro. According to the BIP database of Bank of Italy, the average number of banks per borrower between end-June 1999 and end-March 2004 decreased by one sixth for non financial enterprises with credit lines between 500,000 and one million euro and more than one fifth for larger amounts; the share of the main bank rose by about 5 to 9 percentage points. It is worth remarking that these indicators moved much more slowly between end-March 1998[11] and end-June 1999, when a break-point occurs in the equilibrium PT for r^{ST}: the former decreased by about 5 per cent, the latter did not change. Taking into account that the indicators did not appreciably differ from early 1994, it can be safely inferred that in a decade their changes overlap with the post-break period.

Additional factors, internal and external to the banking industry, support the hypothesis of intensified relationship lending:

i) the consolidation process helped reducing the average number of banks per borrower partly because of a mechanical effect - fewer banks -, partly because the increase of banks belonging to a group produces more uniformity in lending practices, thus mitigating arbitrage opportunities across lenders;

ii) the mutual convenience, for banks and firms, to reduce information asymmetries through more stable relationships may have been enhanced by the Basel Accord revision process, started almost in coincidence with the advent of euro, that emphasizes a better assessment of credit risk via external and internal ratings[12].

The intertemporal smoothing of lending rates is beneficial not only to borrowers, but also to banks. They can in fact stabilize their interest margin, which is affected by a low PT to market rates

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periods, are not reported, to save space (results available on request).

[11] The quarterly series on the BIP data base, though with a slight discontinuity in 2000, starts on March 1998.

[12] This is presumably more the case for SMEs, whose creditworthiness may be more sensitive to qualitative elements that need a deeper knowledge of their business and a closer monitoring or a stronger relationship with their management. On the contrary, larger firms, with presumably more reliable and market-disciplined financial managers, may have their ratings driven mostly by quantitative elements.

for about two fifths of their liabilities, namely demand and saving deposits. The equilibrium PT for demand deposits has considerably reduced in recent years, characterized by low money market rates and almost nil depositors' remuneration, down from 0.61 to 0.44 in the same break-free periods identified for the lending rates[13]. Such a pattern implies a greater risk of margin erosion on fund raising when market rates decline, as in most of the post-break period. A reduction of the PT on short term lending rates, as in fact happened according to the evidence provided in this paper, could have helped to mitigate this risk.

6. Conclusion

The paper makes several contributions to the empirical literature on lending rates PT.

Methodology. First, a strategy is suggested to endogenously search for structural breaks in cointegrated relations, when the researcher's interest is for the latest break-point. Second, the econometric investigation through the two-step Engle-Granger approach is supplemented with a one-step single-equation procedure, in order to better assess the precision of key-parameter estimates and to implement the Ericsson-MacKinnon test for cointegration adjusted for degrees of freedom.

Data. The determination of short term lending rates is related, following the literature, to two alternative driving market rates. In the Italian case the analysis is extended, aside from the sample period, to the overdrafts interest rate, because of an *ex ante* better maturity matching with the marginal cost proxy.

EMU. The econometric investigation shows that Stage Three of EMU has not implied in its wake a strengthened monetary transmission in the Italian case, at least considering the size of equilibrium PT for short term lending rates: the estimated parameter has shrunk to around 0.7 (0.6 for the overdraft component), down from almost unity, in the post-break period. No structural change is detected in the equilibrium, slightly less than one, PT for the lending rate to top rated borrowers. The equilibrium PT has fallen even more, by almost a half, in Portugal. These results are only partly offset, for the two main short term lending rates in the Italian case, by a faster adjustment, that is complete within a semester instead of a year.

Interestingly, the latest break-points cluster around the second semester 1999 in both countries, irrespective of the driving market rate, in sharp contrast with structural changes occurring as early as 1994 and 1995, a result that previous studies rationalized on the grounds of expected reduced variability of money market rates and of inflation convergence induced by the monetary

[13] EG first stage estimates; the one-step estimation results (available on request) are similar.

unification (Sander-Kleimeier 2004b). The claim that different driving market rates, following a monetary policy or a cost of funds approach, may yield different break-points is disputed: once a suitable methodology of search is adopted, the dates are shown to cluster very closely.

The overall results of this paper challenge the view, recently put forward by several authors, that EMU has in its wake enhanced the effectiveness of monetary transmission via the banking sector and made it more homogeneous across countries, because of rising and converging equilibrium PTs. The findings of this study uncover an as yet incomplete path to an even monetary transmission of ECB's impulses to national credit markets, should the PT increase in countries other than Italy and Portugal after the introduction of euro be confirmed.

Two research themes worth pursuing in future, in the light of these results are, first, to extend the proposed methodology to endogenously search for break-points to other EMU countries, and, second, to assess on panel data the suggested interpretation, for the Italian case, of a link between the PT reduction on short lending rates and the strengthening of relationship lending in coincidence with (and possibly owing to) the monetary unification.

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Appendix 1 – The data

The national, unharmonized, retail bank rates (*National Retail Interest Rates*, NRIR) collected by the ECB refer to types of loans and deposits representatives of a country banking industry and are grouped into 6 macro categories of lending rates and 5 of deposit rates[14].

Italy. The short term lending rates analysed in this paper for Italy[15] are:

- r^{ST} : average rate, weighted by stocks, on short term (maturity up to 18 months) loans, with lending to enterprises accounting for about a half (NRIR acronym: N4_1). The aggregate accounts for about half of total loans;
- r^{OD} : average rate, weighted by stocks, on overdrafts[16]. The aggregate amounts to slightly less than a half of short term loans (Figure A3);
- r^{TOP}: minimum rate, computed as a weighted average by stocks, on short term term loans to firms (NRIR acronym : N4_2)[17].

It is worth noticing that Italy is the only country, in the NRIR database, with interest rates referring to outstanding stocks instead of new businesses.

The sample period goes from January 1993[18] to February 2004, the last month the series are published. The discontinuity is due to the introduction of new harmonized interest rate series[19]. For instance, in the overlapping period (January 2003 – February 2004) of the new and old overdraft interest rate series[20] the levels are quite different, because of a new sample of reporting banks and of new methods to collect rates (end-of-month instead of an average of ten-days data).

Portugal. The short term lending rates analysed in this paper for Portugal are r^{CB} and r^{STF} , average rates for commercial bills (NRIR acronym: N4_1) and loans (NRIR acronym: N4_2) to private non-financial enterprises firms with 91 to 180 days maturity, respectively.

[14] http://www.ecb.int/stats/money/interest/html/retail.en.html.
[15] Two other series, N2 and N5, average rates, weighted by new medium-long term (maturity beyond 18 months) businesses to firms and households, respectively, are not analyzed owing to the absence of a reliable market rate that could satisfactorily proxy their marginal cost throughout the entire time span. The financial characteristics of the aggregates (average maturity and interest rate fixation) have indeed changed during the period, spanning across a wide range of maturities and comprising fixed- and variable rate contracts, with time-varying proportions.
[16] Source: Banca Informativa Pubblica (BIP) of the Bank of Italy.
[17] The data refer to the first decile of short term loans to firms, ordered by increasing interest rates.
[18] Data for many EMU countries are available only since early 1990s. Following the literature, the sample period starts after the 1992 EMS crisis.
[19] As of January 2003 the ECB collects a new set of harmonized bank rates statistics, that relate to aggregates with common features across the EMU countries, such as, for instance, the initial horizon of rate determination, an aspect that provides a synthetic representation of the contract maturity and of the rate fixation. Though bound to be the ideal data base for PT empirical analysis across countries, the as yet short sample and the low variability of the money market rates in the reporting period hinder econometric exercises focused on long run parameters (see also Baele *et al.* 2004).
[20] In the period January 2003-February 2004 the levels of the interest rates in the new harmonized series were on average higher by 43 basis points.

The sample period goes from January 1993[21] to October 2002.

For both countries, the driving money market rates are the one-month interbank (rm^{INT}) and the overnight (rm^{ON}) rates, chosen, as in Sander-Kleimeier (2004a), because they are the most correlated with the bank rates (see Figure 1 and Table A1 for Italy). The data for Italy were downloaded from the Bank of Italy website. For Portugal, the overnight rate series was downloaded from the Bank of Portugal website; for the one-month interbank rate series, owing to a missing value at April 2001, the Datastream series was used as in Sander-Kleimeier (2004a). The series starts however at October 1994; it was extended back to January 1993 inserting the data downloaded from the Bank of Portugal website[22].

[21] Data for many EMU countries are available only since early 1990s. Following the literature, the sample period starts after the 1992 EMS crisis.
[22] The NRIR series ends at December 2002, with another missing value at November 2002.

Table 1 **Italy: Review of the literature on short term loan interest rates pass-through**

Study	Sample period	Market rate	Break-point	Estimation procedure	Estimation sample	Short term pass-through (γ_0)	Equilibrium pass-through (β)	Adjustment speed (θ)
Short term lending rate (r^{ST}) [a]								
de Bondt *et al.* (2003, Table 4)	1994.04 2002.12	• 3 months interbank; • Government 10 years yield.	NO (Chow test *p*-value with break at January-99 = 0.20)	ECM: single-equation	1994.04 2002.12	• 0.19 for 3 months interbank; • 0 for 10 yrs yield	• 0.93 for 3 months interbank; • 0.12 for 10 yrs yield	-0.15
					1999.01 2002.12	• 0.16 3 months interbank; • -0.07 for 10 yrs yield	• 0.76 for 3 months interbank; • - 0.15 for 10 yrs yield	-0.60
Sander-Kleimeier (2004a, Table B6)	1993.01 2002.10	Overnight (Monetary policy approach)	YES: February-95	ECM: 2-step EG	1993.01 1995.02	0.31	1.09	n.a.
					1997.03 2002.10	0.16	0.96	
Sander-Kleimeier (2004a, Table C6)	1993.01 2002.10	One-month interbank (Cost of funds approach)	YES: July-99		1993.01 1999.07	0.27	1.02	
					1999.08 2002.10	0.31	0.68	
Minimum short term lending rate to firms (r^{TOP})[a]								
Sander-Kleimeier (2004a, Table B6)	1993.01 2002.10	Overnight (Monetary policy approach)	YES: February-95	ECM: 2-step EG	1993.01 1995.02	0.43	0.94	n.a.
					1995.03 2002.10	0.21	0.92	
Sander-Kleimeier (2004a, Table C6)	1994.07 2002.10	One-month interbank (Cost of funds approach)	NO		1994.07 2002.10	0.31	0.95	

[a] For data description see Appendix 1.

Table 2 Portugal: Review of the literature on short term loan interest rates equilibrium pass-through

Study	**Sample period**	**Market rate**	**Break-point**	**Estimation procedure**	**Estimation sample**	**Equilibrium pass-through (β)**
			Commercial bill rate (r^{CB})[a]			
de Bondt *et al.* (2003, Table 4)	1994.04 2002.12	3-months interbank[b]	NO (Chow test *p*-value with break at January-99 = 0.57)	ECM: single-equation	1994.04 2002.12	1.24
					1999.01 2002.12	0.93
Sander-Kleimeier (2004a, Table B4)	1993.01 2002.10	Overnight (Monetary policy approach)	YES: July-94	ECM: EG first stage	1993.01 1994.07	0.26
				No cointegration; momentum threshold ADL	1994.08 2002.10	1.52[c]
Sander-Kleimeier (2004a, Table C4)	1994.10 2002.10	1-month interbank (Cost of funds approach)	YES: October-99	ECM: EG first stage	1994.10 1999.10	1.24
					1999.11 2002.10	0.65
			Short term lending rate to firms (r^{STF})[b]			
Sander-Kleimeier (2004a, Table B4)	1993.01 2002.10	Overnight (Monetary policy approach)	YES: February-95	ECM: EG first stage	1993.01 1995.02	0.33
					1995.03 2002.10	1.51
Sander-Kleimeier (2004a, Table C4)	1994.10 2002.10	1-month interbank (Cost of funds approach)	YES: November-99	No cointegration; ADL	1994.10 1999.11	1.33[c]
				ECM: EG first stage	1999.12 2002.10	0.77

[a] For data description see Appendix 1. [b] 10-years Government rate not statistically significant in equilibrium PT estimation. [c] Computed as the long run coefficient in an autoregressive distributed lags (ADL) specification.

Table 3 **Italy: Alternative ECM estimation procedures**

(other short-run dynamics coefficients omitted; std error in brackets[1])

Market rate (estimation procedure)	**Sample Period**	α	β	θ	γ_0	**Adj Rsq**	**DW**	**N[2]**	**LM[3]**	**Cointegration tests: CRDW[4], ADF[5] and EM[6]**
r^{ST}: post-break										
1 month interbank (EG 2-step)	99.06-04.02	3.39	.70	-.32 (.10)	.21 (.05)	.75	2.06	.27	.52	CDRW = .52*** ADF = -2.94**
1 month interbank (one-step ECM)		3.21 (.60)	.75 (.02)	-.46 (.04)	.27 (.05)	.77	1.70	.68	.20	EM = -11.05***
Overnight (EG 2-step)	99.07-04.02	3.45	.70	-.22 (.08)	.26 (.04)	.75	2.01	.97	.96	CDRW = .85*** ADF = -2.42
Overnight (one-step ECM)		3.30 (.12)	.73 (.04)	-.22 (.08)	.30 (.05)	.73	1.74	.98	.23	EM = -2.59
r^{ST}: pre-break										
1 month interbank (EG 2-step)	95.04-99.05	3.14	.95	[-.03] (.04)	.24 (.05)	.73	2.14	.57	.30	CDRW = .33* ADF = -2.84*
1 month interbank (one-step ECM)		1.82 (.25)	1.07 (.02)	-.22 (.04)	.21 (.06)	.83	1.88	.07	.80	EM = -5.34***
Overnight (EG 2-step)	95.04-99.06	3.06	.95	[-.04] (.04)	.25 (.04)	.73	2.14	.65	.18	CDRW = .44** ADF = -2.98**
Overnight (one-step ECM)		2.05 (.41)	1.03 (.04)	-.11 (.05)	.23 (.04)	.84	2.18	.55	.50	EM = -2.35
r^{OD}: post-break										
1 month interbank (EG 2-step)	99.06-04.02	4.65	.61	-.34 (.08)	[.10] (.06)	.65	1.90	.01	.45	CDRW = .37* ADF = -2.34
1 month interbank (one-step ECM)		4.52 (.09)	.64 (.03)	-.36 (.07)	.11 (.07)	.67	1.87	.15	.30	EM = -5.01 ***
Overnight (EG 2-step)	99.11-04.02	4.71	.60	-.34 (.09)	.19 (.05)	.68	1.57	.20	.03	CDRW = .53*** ADF = -2.74*
Overnight (one-step ECM)		4.62 (.09)	.62 (.03)	-.36 (.07)	.22 (.05)	.68	2.00	.85	.65	EM = -5.00***
r^{OD}: pre-break										
1 month interbank (EG 2-step)	95.04-99.05	3.99	.90	[-.04] (.04)	.22 (.05)	.64	2.19	.54	.14	CDRW = .31 ADF = -2.83
1 month interbank (one-step ECM)		2.76 (0.23)	1.02 (0.02)	-.25 (.04)	.19 (.06)	.76	1.68	.51	.44	EM = -7.16***
Overnight (EG 2-step)	95.04-99.10	3.91	.91	[-.06] (.04)	.25 (.05)	.68	2.16	.77	.13	CDRW = .39** ADF = -3.04**
Overnight (one-step ECM)		3.61 (.46)	.90 (.06)	-.08 (.05)	.24 (.04)	.78	2.07	.92	.76	EM = -1.71
r^{TOP}										
1 month interbank (EG 2-step)	95.04-04.02	.31	.93	-.14 (.04)	.29 (.04)	.87	2.22	.09	.08	CDRW = .42** ADF = -4.74***
1 month interbank (one-step ECM)		.23 (.08)	.93 (.01)	-.15 (.04)	.29 (.04)	.89	2.27	.27	.04	EM = -3.39**
r^{TOP}: post-break										
Overnight (EG 2-step)	97.10-04.02	.66	.84	-.16 (.08)	.27 (.03)	.91	1.96	.04	.46	CDRW = .84*** ADF = -4.41***
Overnight (one-step ECM)		.70 (.07)	.81 (.02)	-.24 (.07)	.26 (.03)	.91	1.83	.06	.42	-4.06**
r^{TOP}: pre-break										
Overnight (EG 2-step)	95.04-97.09	.35	.93	-.25 (.09)	.33 (.08)	.80	2.31	.48	.23	CDRW = .56*** ADF = -4.24***
Overnight (one-step ECM)		-.36 (.30)	1.00 (.03)	-.34 (.06)	.31 (.08)	.88	2.08	.39	.70	-5.32***

[1] Heteroskedasticity consistent whenever the White test is below the 5% significance level. [2] *p*-values for the Jarque-Bera test under the null of normality of residuals. [3] *p*-value for the Breusch-Godfrey test under the null of no first order correlation of residuals. [4] Critical values, computed for samples of 100 observations, under the null of I(1) first stage residuals, at the 1% (***), 5% (**) and 10% (*) significance: 0.51, 0.38, 0.32. [5] Asymptotic critical values under the null of I(1) first stage residuals at the 1% (***), 5% (**) and 10% (*) significance; ADF with constant and up to 2 lags (MacKinnon 1991). [6] Critical values, adjusted for the degrees-of-freedom, in the single-equation ECM procedure (Ericsson-MacKinnon 2002). [.] : coefficients not significantly different from zero at the 10% significance level.

Table 4 Italy: Pass-through at 3, 6 and 12 months and speed of adjustment (%) to equilibrium (computed out of pre- and post-break single-equation estimates with one month interbank rate[1])

Bank rate	Estimation period	3 months	6 months	12 months	β
r^{ST}	Post-break	0.64 (90%)	0.74 (98%)	0.75 (100%)	0.75
	Pre-break	0.71 (66%)	0.90 (84%)	1.03 (96%)	1.07
r^{OD}	Post-break	0.59 (92%)	0.65 (101%)	0.64 (100%)	0.64
	Pre-break	0.67 (66%)	0.87 (86%)	0.99 (98%)	1.02

[1] See Table 3, including other short-term dynamics coefficients omitted there.

Table 5 Italy: Asymmetries in short term pass-throughs of one-month interbank rate (estimation procedure: single-equation ECM; std error in brackets[1])

Bankrates	Sample	α	β	θ	γ_0+	γ_0-	γ_1+	γ_1-	Adj Rsq	DW	N[1]	LM[1]	Symmetry test[2]
r^{ST}	99.06-04.02	3.21 (.06)	.75 (.02)	-.45 (.05)	.31 (.11)	.24 (.04)			.76	1.73	.79	.24	.61
r^{OD}	99.06-04.02	4.54 (.08)	.64 (.03)	-.37 (.07)	.04 (.17)	.16 (.07)			.66	1.89	.23	.35	.54
r^{TOP}	95.04-04.02	.29 (.06)	.94 (.01)	-.20 (.04)	.18 (.09)	.32 (.04)	.05 (.03)	.20 (.04)	.90	2.20	.69	.14	.00

[1] See Table 3, including other short-term dynamics coefficients omitted there. [2] *p*-value for a Wald test under the null of equality between the sum of $\gamma+$'s and of $\gamma-$'s.

Table 6 Portugal: Equilibrium lending rates pass-throughs (Engle-Granger procedure first step)

Bankrates	Market rates	Sample	α	β	CRDW[1]	ADF[1]
post-break						
r^{CB}	1 month intb	99:11 – 02:10	4.87	.66	1.48***	-5.05***
	Overnight	99:12 – 02:10	4.99	.64	1.48***	-4.54***
r^{STF}	1 month intb	99:12 – 02:10	2.57	.78	1.38***	-4.17**
	Overnight	99:12 – 02:10	2.83	.72	1.30***	-3.91**
pre-break						
r^{CB}	1 month intb	95:04 – 99:10	4.17	1.24	0.97***	-5.25***
	Overnight	95:04 – 99:11	3.97	1.30	1.13***	-4.41***
r^{STF}	1 month intb	95:06 –99:11	1.31	1.36	1.38***	-5.54***
	Overnight	95:04 – 99:11	1.29	1.39	1.43***	-5.49***

[1] See Table 3.

Figure 1 **Italy: Short term lending and market rates**

(percentage points)

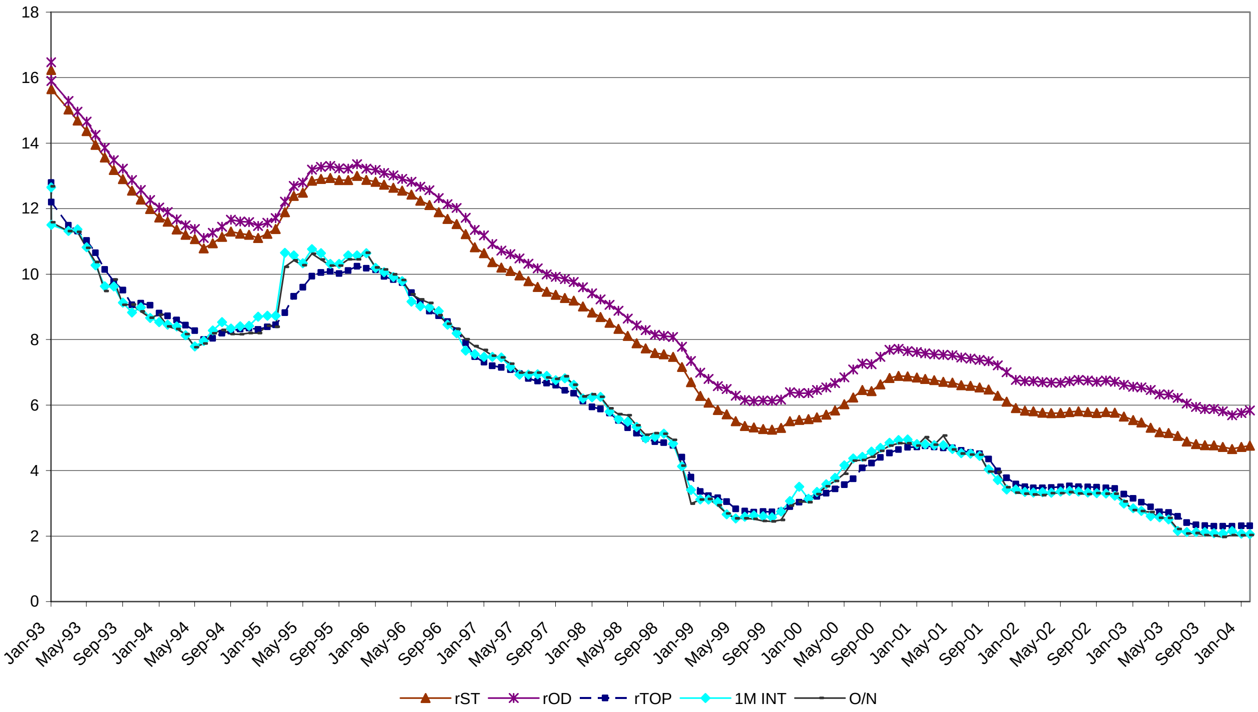
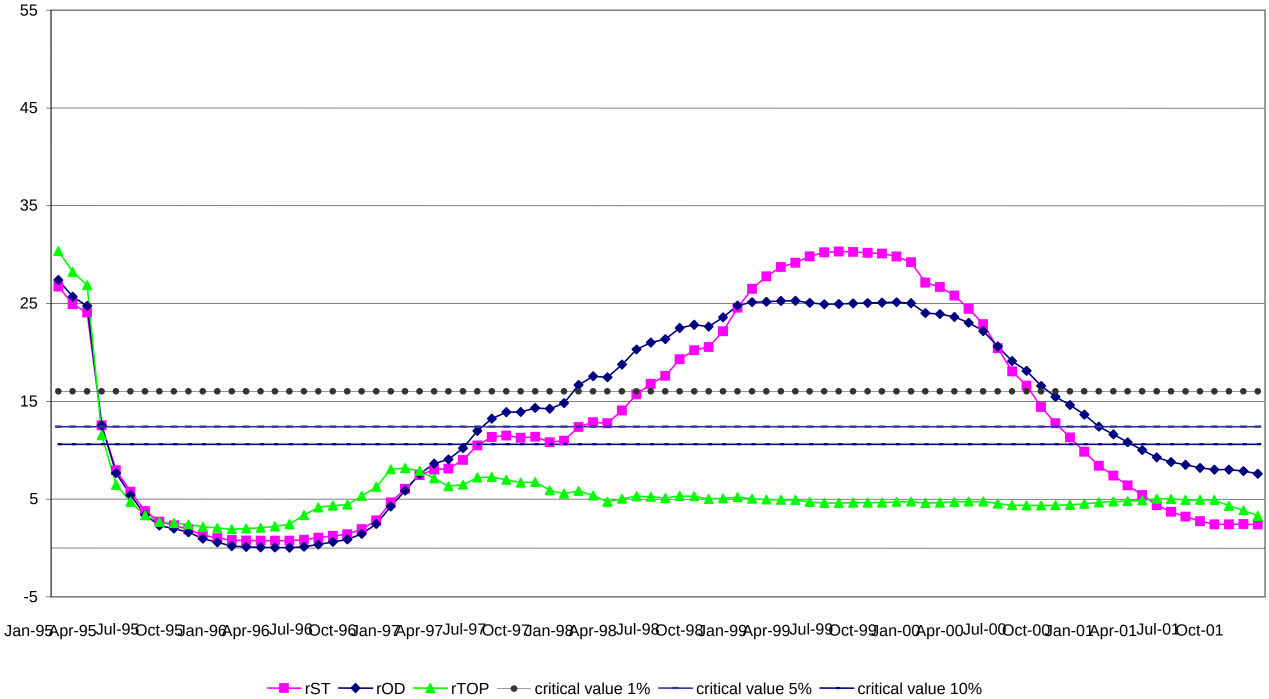


Figure 2 Italy: Rolling Chow tests

(sample period: 1993.01-2004.02)

a) Market rate: one-month interbank



b) Market rate: overnight

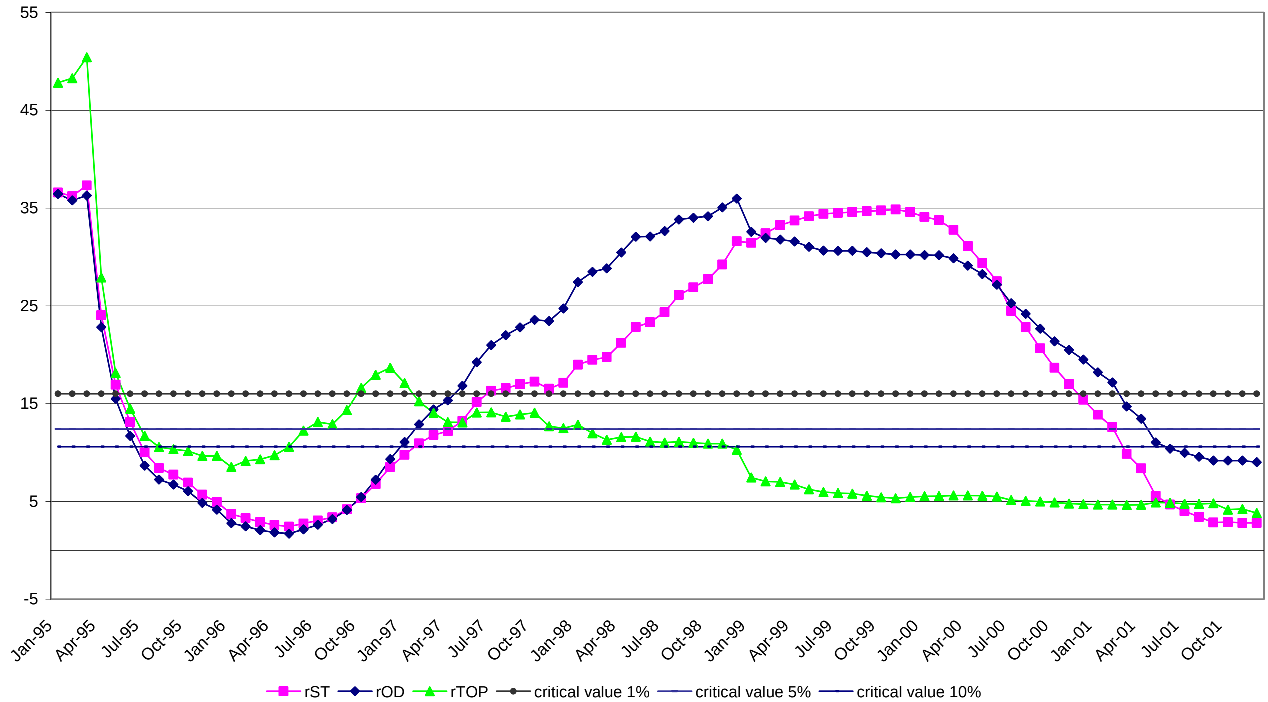
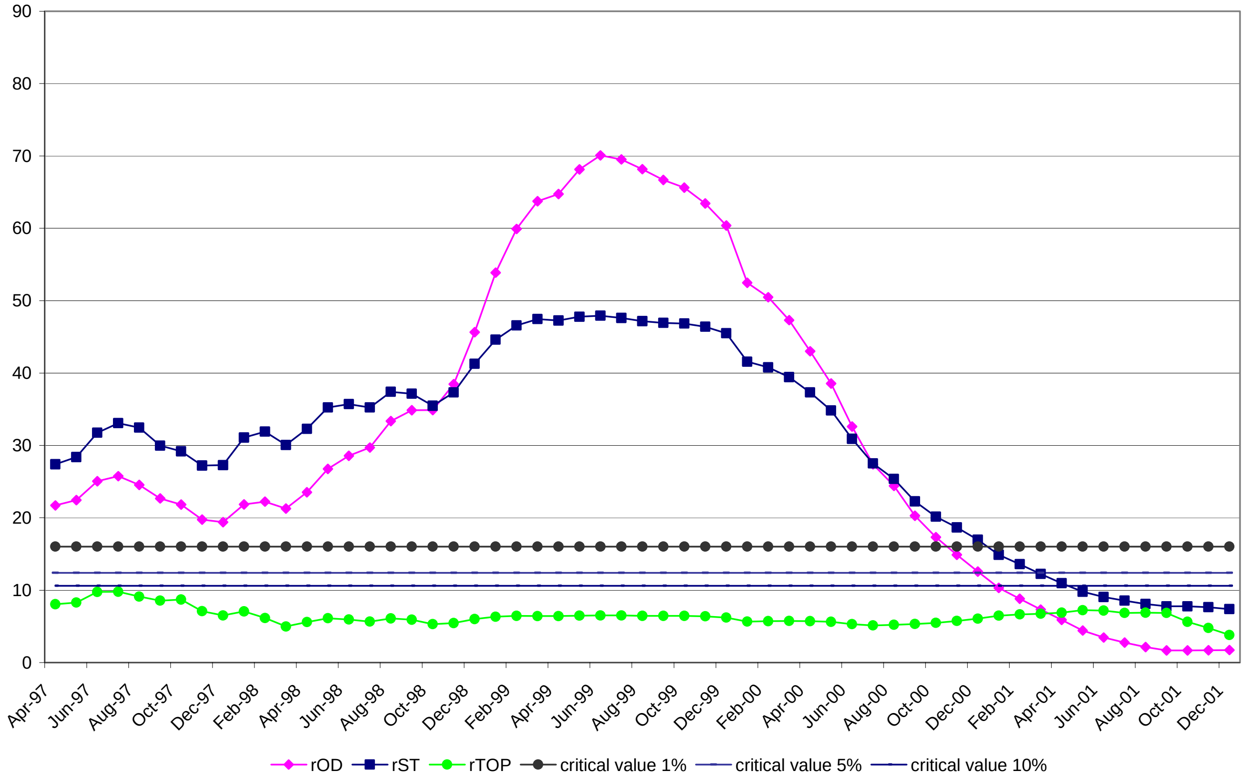


Figura 3 **Italy: Rolling Chow tests**

(period after first break-point: 1995.04-2004.02)

a) Market rate: one-month interbank



b) Market rate: overnight

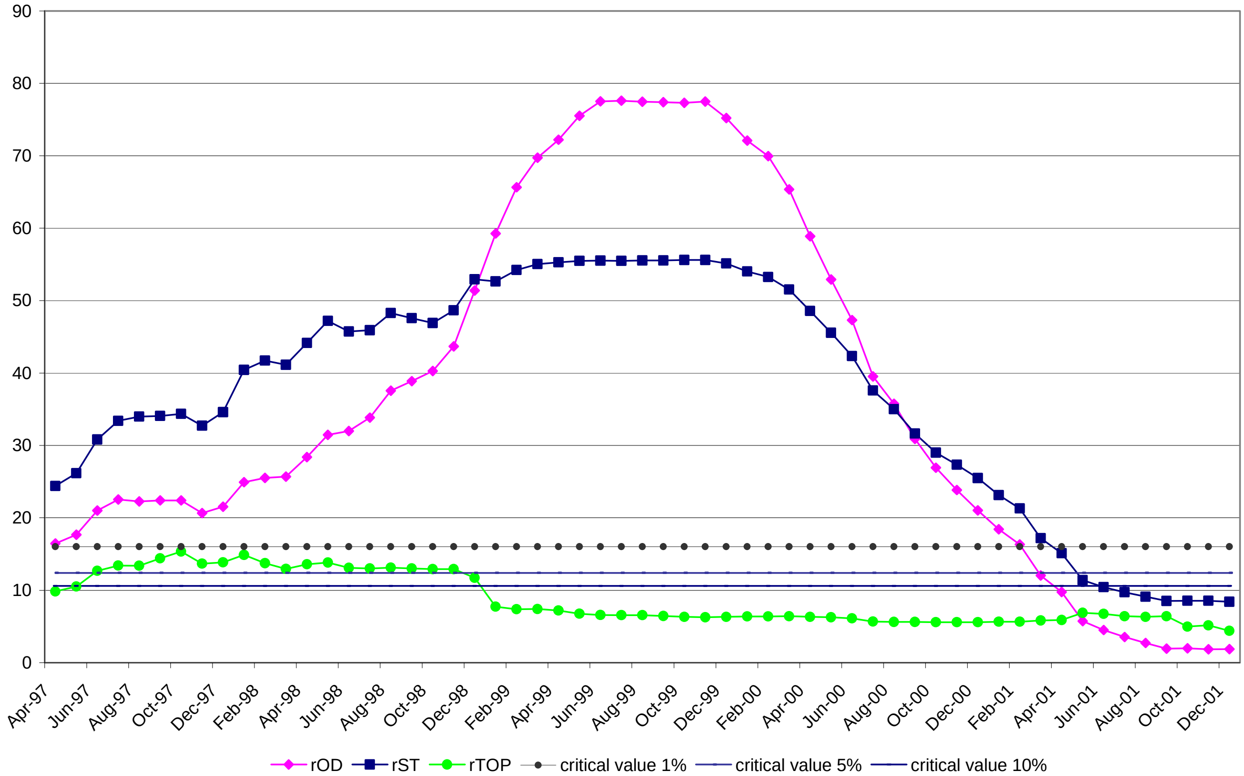


Figure 4 **Portugal: Short term lending and market rates**

(percentage points)

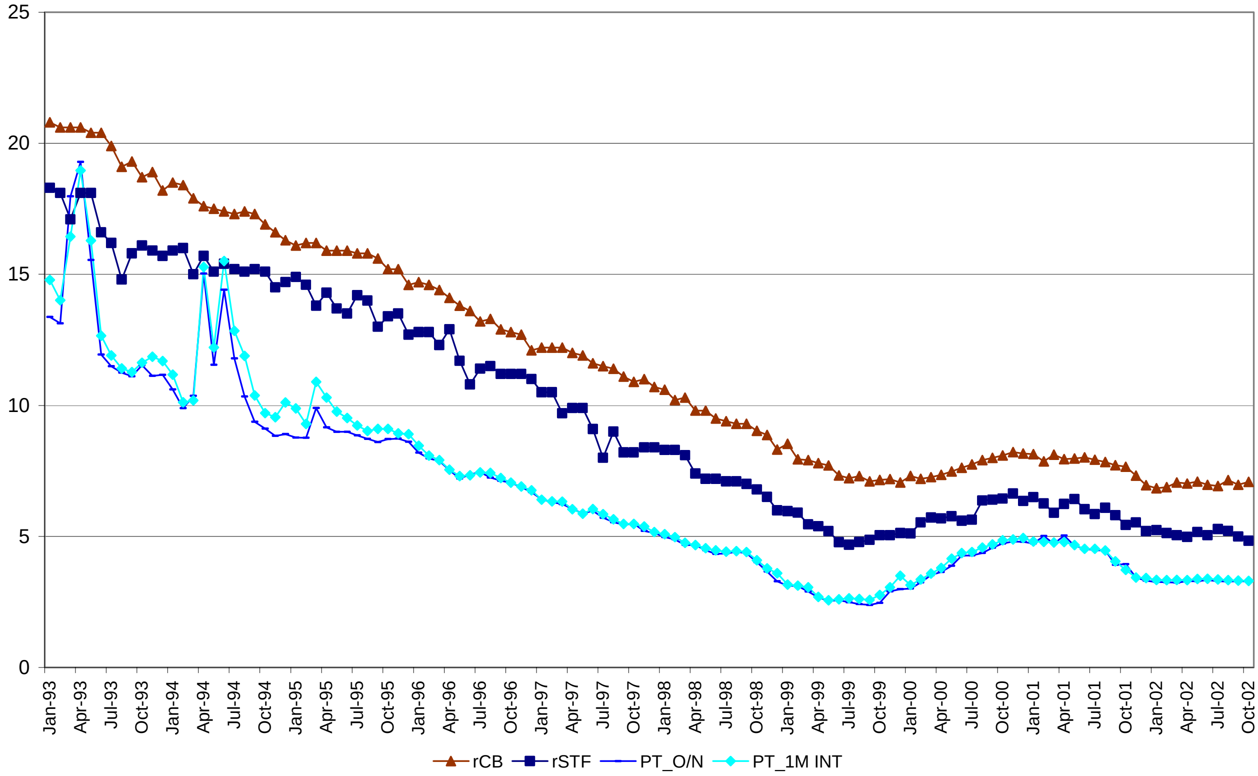


Figure 5 Portugal: Rolling Chow tests with one-month interbank and overnight rates

(sample period: 1993.01-2002.10)

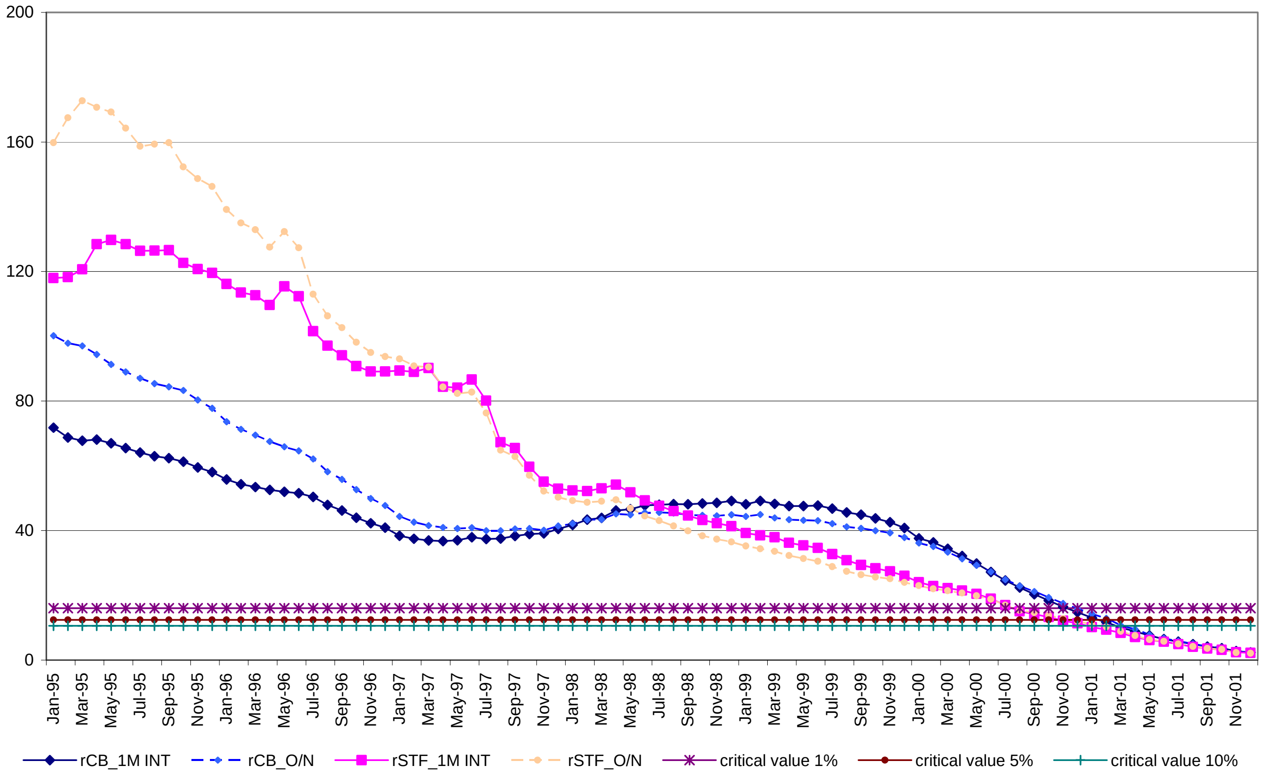
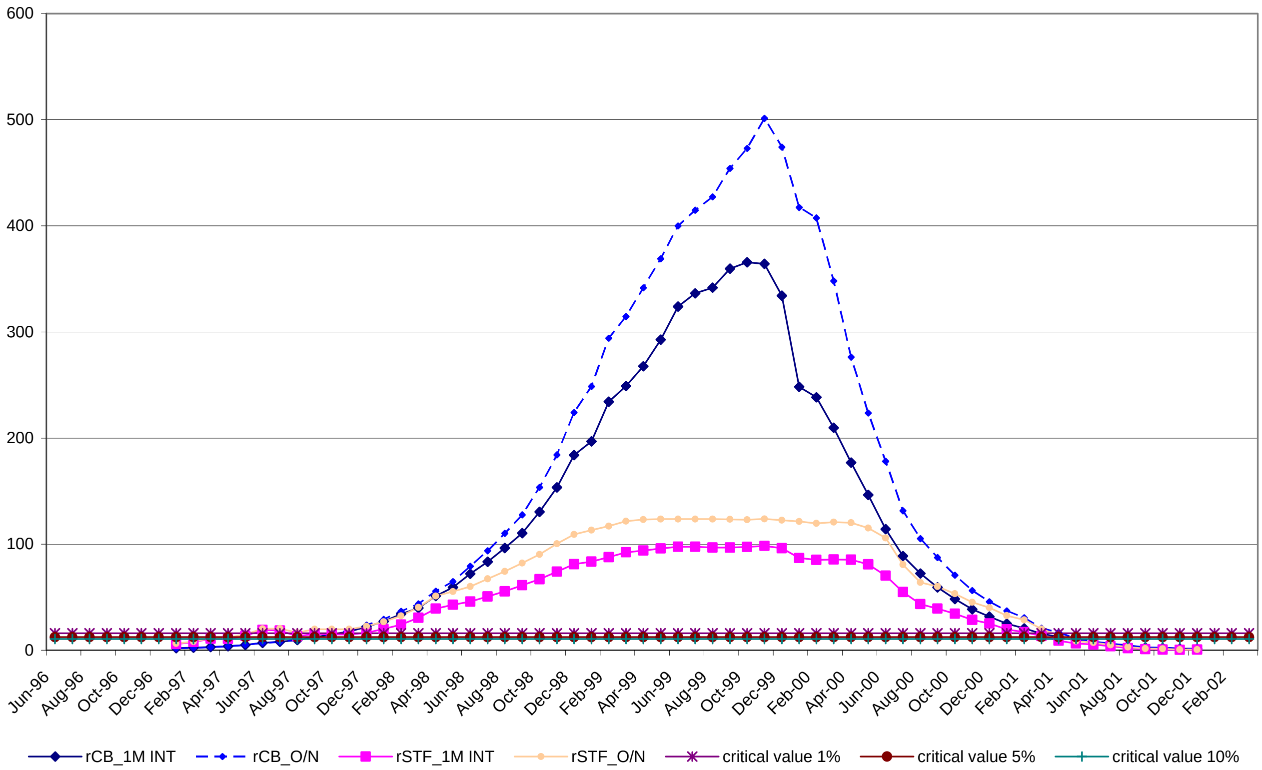


Figure 6 Portugal: Rolling Chow tests with one-month interbank and overnight rates

(sample period: from first breakpoint to 2002.10)



[1] See Figure 5 on first break-points.

Appendix 2 – Tables and Figures

Table A1 Italy: Correlations between (first differenced) lending (r^{OD}, r^{ST} and r^{TOP}) and market (overnight, 1-, 3-, 6-, 12-months interbank) rates

	rm^{ON}	rm^{INT}	rm^{INT3}	rm^{INT6}	rm^{INT12}	r^{OD}	r^{ST}
rm^{INT}	0.88						
rm^{INT3}	0.77	0.94					
rm^{INT6}	0.68	0.87	0.96				
rm^{INT12}	0.57	0.75	0.87	0.95			
r^{OD}	0.60	0.58	0.54	0.48	0.40		
r^{ST}	0.61	0.61	0.56	0.51	0.42	0.99	
r^{TOP}	0.59	0.62	0.58	0.53	0.45	0.87	0.90

Table A2 Italy: Unit root tests for lending and market rates

Break-free periods	Lending and market rates	ADF[1]	
		Levels	Differences
	rm^{INT}		
95.04-99.05	r^{ST}; r^{OD}	-2.03	-8.72***
99.06-04.02		-0.81	-3.68***
95.04-04.02	r^{TOP}	-2.31	-9.16***
	rm^{ON}		
95.04-99.06	r^{ST}	-2.49	-7.72***
99.07-04.02		-0.54	-3.07**
95.04-97.09	r^{TOP}	-2.49	-7.58***
97.10-04.02		-1.97	-4.80***
95.04-99.10	r^{OD}	-2.72	-7.89***
99.11-04.02		-0.54	-2.92**
	r^{ST}		
95.04-99.05	rm^{INT}	-3.96**	-3.42**
99.06-04.02		-1.19	-2.99**
95.04-99.06	rm^{ON}	-4.10**	-3.48**
99.07-04.02		-1.18	-2.92**
	r^{OD}		
95.04-99.05	rm^{INT}	-4.45***	-3.42**
99.06-04.02		-1.37	-3.50**
95.04-99.10	rm^{ON}	-3.91**	-3.45**
99.11-04.02		-1.18	-3.10**
	r^{TOP}		
95.04-04.02	rm^{INT}	-1.81	-4.13***
95.04-97.09	rm^{ON}	-3.36*	-7.25***
97.10-04.02		-2.50	-3.39**

[1] With constant and trend in the pre-break period for the level series; with constant only, otherwise.
*,**,***: rejection of the unit root null hypothesis at 10%, 5%, 1% (asymptotic) significance levels.

Figure A1 **Replicating Sander&Kleimeier for Italy: Rolling Chow tests**

(sample period: 1993.01-2002.10)

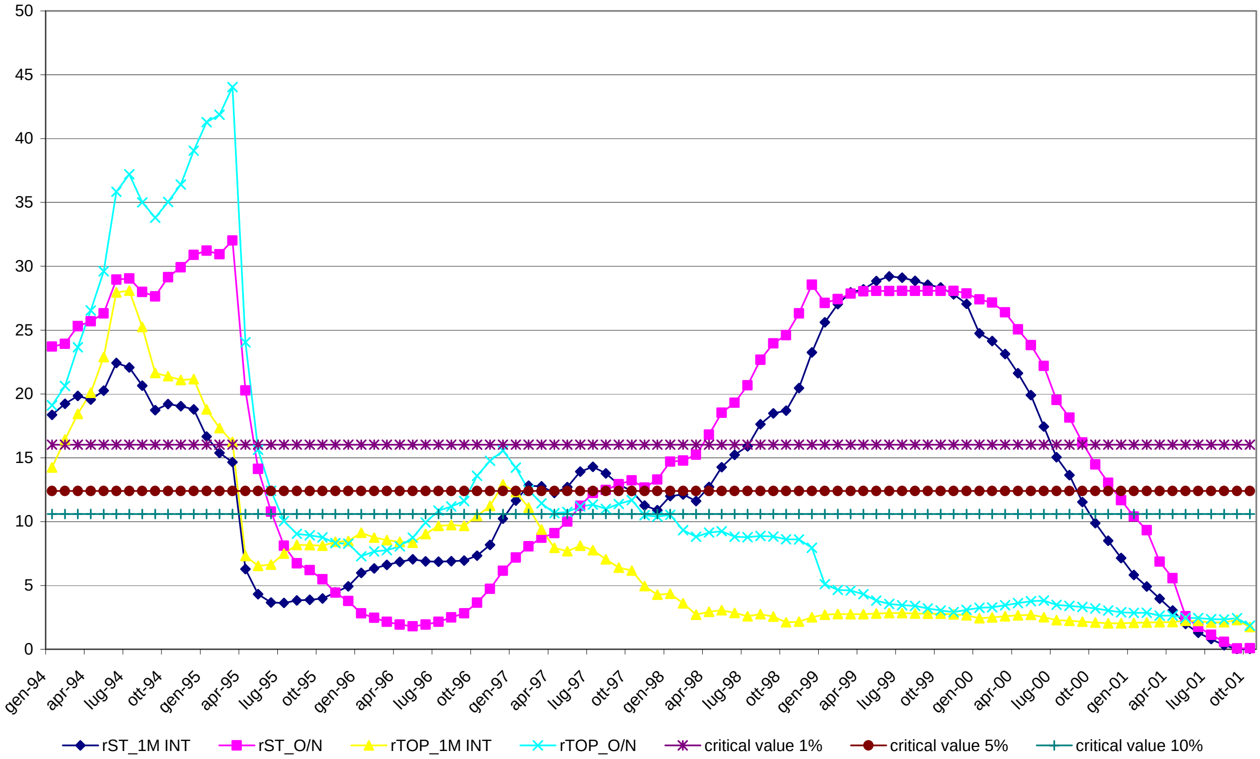


Figure A2 **Italy: Rolling Chow tests**

Sample period: from 1995.04 to 2002.10

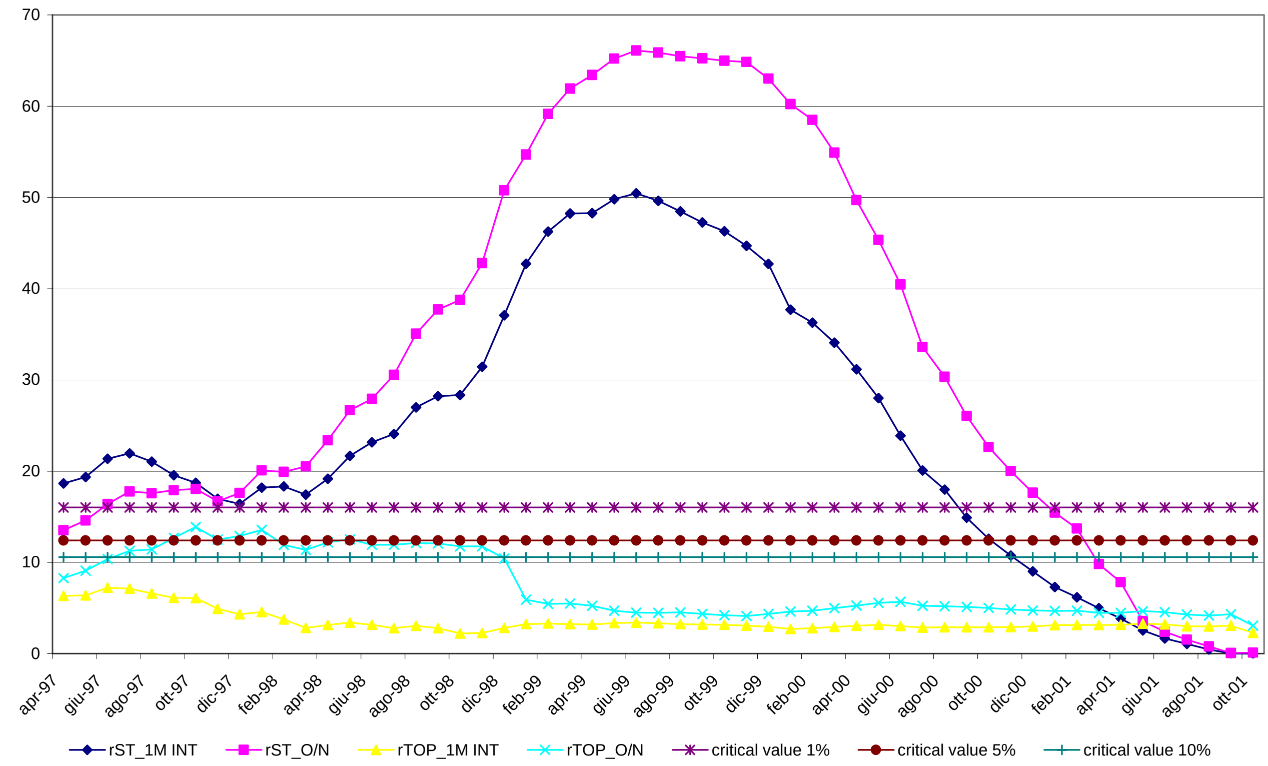


Figure A3 **Italy: Bank Overdrafts as a percentage of Short Term Loans**