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Multiple breaks in lending rate pass-through
A cross country study for the euro area

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Abstract

A new approach is proposed for searching multiple unknown breaks, possibly associated with EMU, in the short term business lending rate pass-through. Multiple breaks are detected in five out of nine countries of the euro area. The last break occurs much before the start of EMU for France, several months after that event for Austria, Italy and Germany. Long run pass-throughs decrease (except for France) sizably below one (except for the Netherlands); heterogeneity in the monetary transmission increases across countries. These results raise doubts on claims of a more effective monetary policy under EMU.

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 market rates, especially in a bank-centric economy such as the Economic and Monetary Union (EMU). A small but growing literature has investigated whether size and speed of the pass-through (PT) from market interest rates to retail ones in the euro zone increased in the wake of Stage Three of EMU, thus enhancing the effectiveness of the single monetary policy, and converged, thus making more uniform the transmission via the banking sector.

Angeloni and Ehrman (2003) provide evidence that since January 1999 lending and deposit rate PTs became on average higher, though no faster, in the four largest countries (the exception being Germany) and in the euro area as a whole. Doubts on the robustness of their findings are however raised by the mixed statistical significance of a structural break in coincidence with the introduction of the single currency (de Bondt *et al* 2005). In addition, criticism has been levelled at the assumption of an a priori break date, January 1999, when considering the *process* underlying a unique historical experiment like the EMU. An alternative empirical strategy is searching for an unknown break date, possibly related to expectations for or to the adjustment after the establishment of the unified monetary regime (Toolsema *et al* 2002, Sander-Kleimeier 2004a)[1].

Short term business loans - the first link in the transmission mechanism through banks - are singled out in most studies among the banking products because their PTs result faster in the post-break period; the findings are instead widely different for the long run PT (LPT). Short and long term rates for lending to business show the largest increase in the impact and peak PTs after January 1999 for the euro area[2]. This finding does not mean however uniformity across countries in the post-EMU period: the null of uniform impact PTs for short-term rates across countries[3] is soundly

[1] National studies concentrate instead on breaks after January 1999 (with aggregate series, Burgstaller 2003 for Austria, Coffinet 2005 for France, Bredin *et al* 2002 for Ireland, Gambacorta-Iannoni 2005 for Italy; with panel microdata, de Graeve *et al* 2004 for Belgium, Gambacorta 2004 for Italy, Lago-Gonzalez and Salas-Fumás 2005 for Spain) or after the (late) entry to EMU (Chionis-Leon 2005 for Greece).
[2] No break was however formally found when modelling an index of lending rates in an euro area monthly monetary model (Bruggeman-Donnay 2003).
[3] Actually 5 countries, due to data availability.

rejected both before and after January 1999 (Angeloni-Ehrman 2003); with national data, the null of no structural break is not rejected for Italy and Portugal (de Bondt *et al.* 2005). Within the single unknown break literature, a faster adjustment to LPT for short-term business loan rates is estimated after *the* structural break, with a date however often different from January 1999 (Sander-Kleimeier 2004a). The empirical findings are moreover somewhat puzzling: dates differ up to 4/5 years for Italy and Portugal, depending on the choice of the driving market rates; they are detected far after the launch of the euro for Germany; they are located approximately two years before the event for Austria, France and Spain, hinting at expectational effects (Sander-Kleimeier 2004a).

This paper argues that, in order to make sense of the variety of these results on short term lending rates PTs, a good starting point is to acknowledge that there is no ground, theoretical or empirical, to assume a single, known or unknown, structural break. Indeed, during the longest period usually considered in the literature - 1993 to late 2002[4] - the PT relation could have been affected by at least three events, somehow linked to EMU: the turbulence in the exchange rate markets in early 1995[5] for some Southern countries - Italy, Portugal, Spain -; the expectations induced by the preparation of Stage III of EMU and the ensuing convergence in market interest rates; the adjustment to the working of a unified monetary regime.

A theoretical econometric framework to deal with multiple unknown breaks in the case on integrated, I(1), regressors, as interest rates most often are, is however as yet missing (Perron 2005). This paper proposes therefore, as a plausible investigation methodology, to adopt the refinement approach laid out for the case of multiple unknown breaks with stationary regressors in Bai (1997), using to test the null of no break the critical values of the supremum F (supF) statistics computed for the case of one unknown structural break in regressions with I(1) regressors (Andrews 1993, Hansen 1992).

[4] The starting year, 1993 in Sander-Kleimeier (2004a,b) or 1994 in de Bondt *et al* (2005), is justified in order to avoid the turbulence derived from the September 1992 crisis of the European Monetary System (EMS). The end-year depends on the date papers were completed.

[5] The US$ depreciated by about 10% in the first quarter of 1995, causing tensions in the exchange rates within the EMS, with an official depreciation for the Portuguese and the Spanish currencies in early March; in addition, financial markets were hit by the crisis for the Mexican debt.

The proposed methodology is implemented for all the nine founding EMU countries that contributed interest rate national series for short term business lending to the National Retail Interest Rates (NRIR) data base, for the largest available sample for the recent past (at most September 2003). The market rates are alternatively the overnight or an interbank one, in order to investigate on the issue of possible differences in the break date because of the choice of the driving rate, brought to the fore by Sander-Kleimeier (2004a). Focusing on the short term lending rate should help better match the (short) maturity of the underlying credit aggregates with the market interest rates relevant for an appropriate pricing[6].

This paper makes several contributions to the literature. Two to three breaks, instead of a single one, are indeed detected in five out of nine countries. Break dates are identical or quite close (up to four months), irrespective of the driving market rate. Comparing the last two break-free periods, LPTs *decrease* (except for France) well below one (except for the Netherlands); the adjustment to the equilibrium is however often *faster*. Cross country heterogeneity in the transmission through the short term business lending rates overall *increases*.

The paper is organized as follows. Section 2 surveys the related literature. Section 3 lays out a methodology for searching multiple unknown break-points in cointegrated relations and describes the data. Section 4 reports and discusses the empirical results. Section 5 concludes.

2. Related literature

The empirical literature on lending rate 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 a standard Klein-Monti model of a monopolistic bank - with risk neutrality, perfect information, no switching or adjustment costs, no joint production of loans and deposits -, easily extended in a symmetric Cournot equilibrium (Freixas-Rochet 1997). The lending rate is determined as a mark-up over the marginal (opportunity) cost, identified either with

[6] If credit aggregates with longer maturity were considered, the (average) market interest rate relevant for their pricing would depend on the mix of fixed and floating rate instruments included, which could vary widely through time and across countries. As a consequence, the analysis could spot a change in the PT, when in fact there is nothing but a different mix of instruments/interest rate fixation characteristics.

the money market rate directly influenced by the central bank or with the interbank rate with the same maturity of loans. Assuming a linear approximation, the marginal cost coefficient can be interpreted as the LPT, with a transfer unitary value to the retail rate of changes in the driving market rate in a competitive market (Lago-Gonzalez and Salas-Fumás 2005). Within this framework, studies differ mostly on how to date structural breaks and proxy the marginal cost, in order to match the maturity of the credit aggregate underlying the lending rate.

The estimates of impact PT (IPT) and LPT in cross country studies are usually obtained in a single equation setting, reparametrizing an Autoregressive Distributed Lags (ARDL) specification, originally suggested in Cottarelli-Kourelis (1994), as an Error Correction Mechanism (ECM), following the Granger representation theorem for cointegrated variables. Interest rates are in fact usually I(1) processes.

Let an equilibrium, or cointegrated, relation between I(1) interest rates:

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

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

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

Eq. (1) includes as a deterministic component only a constant, being the presence of a linear trend in interest rates theoretically inconsistent (Hamilton 1994).

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

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

where Δ is the first difference operator[7].

The key parameters, from an economic point of view, are γ (IPT), β (LPT), and θ, that is the speed the error is corrected; the latter parameter (also known as loading factor) should result statistically significant if cointegration holds.

The empirical choice of the (weakly) exogenous driving market rate[8] motivates the recently proposed distinction between a "monetary policy approach" (*MPA*), with the overnight rate taken as a proxy for the monetary policy rate, 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). 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* (2005), should match the range of maturities of the credit aggregate underlying the short term lending rate.

Recent literature on short term lending rate PT in the euro area provides quite different results as to the date of *the* structural break, possibly coincident with the advent of Stage Three of EMU, as well as to the changes in the LPTs and the speed of adjustment. Angeloni and Ehrmann (2003) argue that a single bank reserves market and the reduction in market interest rates volatility, due to the operating procedures of the European Central bank (ECB), have already produced larger and faster bank rate PTs. They report, having identified informally, via rolling-window regressions, January 1999 as a break-point, that both impact and peak 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. The impact and peak PTs for short and long term business lending rates in the euro area show the largest increase among the entire set of bank rates considered (from 0.35 to 0.53, from 0.81 to 1.11,

[7] In a bivariate relation, with at most one cointegration relation, the EG procedure is preferable to the Johansen one, being more robust to misspecification and reduced sample size (Maddala-Kim 1998).
[8] The weak exogeneiy of market rates to the lending rate is explicitly or implicitly assumed in the literature, since bank rates are not expected to affect market rate developments.

respectively). De Bondt (2005), on the contrary, estimates that LPTs for all euro area retail bank rates, except the mortgage rate, are *lower* in the post-EMU sample compared to the extended period (January 1996-June 2001). In particular, β shrinks from 1.53 to 0.88 for the short term lending rate to firms, having rejected with a Chow test the null of no structural break in January 1999.

A possible explanation for these conflicting findings may be related to the computation of a so-called euro-area data set, built upon national unharmonized bank rates, chosen by each country as the most representative for a given category. However, country-specific and cross country studies yield even more heterogeneous outcomes - on break dates and PTs - in the case of short term lending rates to business. In the following, we concentrate on those studies with a sample period including 2002 data[9].

Hofmann (2003), who *assumes* a unitary LPT and as a driver the 3-months interbank rate, obtains quite different results across the largest four countries countries. First, the break date in January 1999 is not statistically significant for Spain; second, whereas θ increases substantially over the EMU sub-sample everywhere, though always to a puzzling low value for Germany, γ, instead, increases in France and Italy and falls in Germany and Spain (Table 1).

De Bondt *et al.* (2005) *do not* detect a structural break in January 1999 in Italy and Portugal. The test is carried out within 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 yields, under the assumption that the latter provides a signal on the persistence of changes of the policy rates. Overall, in the last period the long-term market rate becomes statistically insignificant and LPTs decrease well below one (except for the Netherlands). The γs rise in Austria, France, Netherlands and Portugal and fall in Italy and Spain. The estimates for Germany are always very imprecise.

Sander and Kleimeier (2004a,b) estimate Eq. (1) with alternative driving market rates, under *MPA* and *CoFA*, and empirically determine whether a *single* structural break has occurred; once

[9] For earlier cross country studies see Donnay-Degryse (2001) and Heinemann-Schüler (2003).

detected it, they provide an EG estimate of Eq. (2) before and after the break date[10]. Their findings vary a lot across countries. Breaks as early as July 1994 and February 1995 under *MPA* and as late as July and October 1999 under *CoFA* are detected for Italy and Portugal; dates differ by one year (August 1997 and 1998), depending on the driving rate, for the Netherlands. Under both approaches break dates are located much before the introduction of the single currency for France (June 1997), Austria (August 1997) and Spain (September/November 1996) as well as much later for Germany (July 2000/February 2001). The βs show, as in de Bondt *et al* (2005), a pattern of reduction (Germany, Italy, Spain) and increase (France), with a mixed evidence, depending on the driving market rate, for Portugal. Overall, the post-break LPTs fall sizably below one; IPTs always slightly increase or at least stay constant[11]. The market for short term business lending has become more homogeneous, in spite of the findings of increases in PTs under *MPA* and reductions under *CoFA* (Sander-Kleimeier 2004a).

[TABLE 1 APPROXIMATIVELY HERE]

3. Econometric methodology and data

3.1 Econometric methodology

The simple point raised in this paper is that the introduction of the single currency is a *process*, announced well before its formal implementation and likely to imply a protracted adjustment in banks' pricing policy. The assumption of a *single* structural break in the PT relation in coincidence with the formal launch of the euro is hardly motivated on economic grounds; a *single unknown* break, though a better starting point, is still an unduly restrictive assumption, because forward looking behaviour on the one hand and late adjustments after EMU on the other hand cannot be ruled out. Our preferred maintained hypothesis is therefore of *multiple unknown* structural

[10] More precisely, when cointegration holds, first-stage EG estimates are obtained; at the second stage, five variations of threshold autoregressive models that allow for asymmetries are used. When cointegration does not hold, standard ARDL specifications are estimated. The published output allows to recover only the point estimates for IPT and LPT.

[11] Two studies confirm the findings on LPT and IPT for France and Germany. Under the assumption of a break in January 1999, a slight *decrease* in the LPT (well below one) but a quicker IPT are found for France (Coffinet 2005), while a reduction in both parameters occurs for Germany, with a sample extending only to May 2001 (de Bondt 2005).

breaks. However, in order not to fall in a data-mining trap we surmise that, in the period 1993-2002 or later, the break dates are at most three. The first one could be motivated by the financial turbulence in the exchange rate markets in early 1995; the second one could be justified because of the expectations set into motion by the announced advent of the euro, once the number of countries entering the EMU was agreed (approximately late 1996 - first half of 1997); the third one could be located after the introduction of a single monetary regime, as national banking systems tried to adapt to it.

The econometric literature does not provide however a framework to implement the proposed approach within the single equation specification in the case on I(1) regressors, as interest rates almost invariably turn out to be. As Perron (2005, 10) flatly states, "No results [for estimation and inferences about break dates] are yet available for multiple structural changes in regressions involving integrated regressors".

In order to circumvent this obstacle adopting a plausible framework for the empirical investigation we rely on some key papers. Hansen (1992), building on Andrews (1993), provides aymptotical critical values for supremum F (supF) statistics - i.e. the largest of the standard rolling Chow test statistics computed under the hypothesis of a break occurring in each subsequent date through the mid-70% sample period - in order to detect a *single unknown* structural break in the case of cointegrated I(1) regressors[12]. Bai (1997), in a linear model with stationary regressors, proposes an efficient procedure to estimate *multiple unknown* break dates. He suggests, in order to detect them, using basically the supF approach, to run first a nested procedure starting from the entire sample and subsequently in the subsamples. Let t_1, t_2 and t_3 be the break dates found accordingly in the full sample $t_0 - T$. In order to get efficiency he then advocates a *refinement* of this procedure, that is searching a break date in the samples $t_0 - t_2$, $t_1 - t_3$ and $t_2 - T$. In both stages the procedure assumes a maximum number of unknown breaks for the sample $t_0 - T$ (Bai-Perron 1998).

[12] The asymptotic distributions of these test statistics are non-standard because, when the break date is unknown, it is a nuisance parameter that appears only under the alternative hypothesis of structural break (Andrews 1993).

Building on this literature the econometric methodology we propose and implement is the following. We first check, using the Augmented Dickey Fuller (ADF) statistics, that the retail and the chosen market rates are I(1) processes over the entire sample, January 1993 to the latest month included in the NRIR database. Second, we check for (at most three) structural breaks for equation (1) in every 70% central part of a sample,[13], applying the refinement procedure of Bai (1997), but adopting the Hansen (1992) 1% critical value[14]. Third, provided that the I(1) property holds for the lending and the market rates within each break-free period, we implement the two-stage EG procedure, estimating the same Eq. (1) used in the endogenous date search exercise and, if the null of cointegration is not rejected, Eq. (2) as well. Note that testing in a break-free period mitigates the well known problems of low power of tests for cointegration in the presence of breaks (Maddala-Kim 1998). In order to enhance comparison across countries the short-run dynamics for Eq. (2) is the same, leaving out lagged first-differenced regressors[15]. When, within a break-free sample, regressors are not I(1) and/or the null of cointegration for (1) is rejected by both Cointegrating Durbin Watson (CRDW) and ADF test statistics, IPT and LPT are estimated via a general-to-specific procedure starting from an ARDL (2,2) specification.

3.2 Data

The short term lending rate to firms is the series coded "N4" for each of the nine contributing countries to the unharmonized NRIR database[16]. The sample period runs from January

[13] This implies to neglect findings of statistically significant supF statistics in the 15% (so called trimming parameter) tails of a sample, in order to restrict each break date to be asymptotically distinct and bounded from the boundaries of the sample (Perron 2005). In addition, having to use tests for structural change, the minimum length of any sub-sample is set at a 15% of the entire sample, following a popular choice in the literature (Andrews 1993, Perron 2005). Empirically, with at least 10 years data, this implies a minimum sub-sample of 18 months.

[14] The critical value for the supF statistics at the 1% significance level is 16.2 (Hansen 1992, Table 1).

[15] We checked that including further lags did not change the estimates for the key parameters, with improvements in some misspecification test statistics in very few cases (results available upon request).

[16] http://www.ecb.int/stats/money/interest/html/retail.en.html. In the case of Belgium, Italy and Portugal the rates are two, coded as N4.1 and N4.2 (in this paper r_1 and r_2).

1993 to at most September 2003[17], depending on the country. The underlying aggregates rates refer to new businesses, except for Italy (outstanding stocks with a maturity up to 18 months)[18].

Among the driving market rates, overnight (*O/N*, for the *MPA*), and 1 to 12 months interbank rates (*INTi*, *i* being the monthly maturity, for the *CoFA*) are drawn from the national central banks' websites or, when not available, either from Datastream (Portugal) or from the IMF's IFS (Ireland). The chosen interbank rate is the most correlated with the bank rate (in first differences), following de Bondt (2002)[19].

Over the full sample period both bank and market rates fall dramatically since early 1995, inverting this trend in the first two years of euro and subsequently reaching low historical levels. This pattern, obviously identical across countries in the post EMU period, is qualitatively similar in previous years, though with a steeper reduction, due to initial higher levels, for Italy, Portugal and Spain. The same trio and to a lesser extent France are the countries most affected by the turbulence in financial markets in early 1995 (Figure 1).

The pattern of spreads (short term lending rate net of market rate) is quite varied within and across countries. Overall, a visual inspection suggests the usual stickiness for retail rates, a feature that does not seem mitigated in the post-EMU period. The benchmark of a fast adjustment to a unitary LPT is a constant spread, with no correlation with market rate levels, as it is for instance the case of US by mid-1990s (Sellon 2002). Only Netherlands, Portugal and Spain seem close to this pattern in the post-EMU period. All other countries show, within the last break-free period detected under *CoFA* (see par. 4.2 and Table 2), trending spreads, with end-sample levels sometimes higher

[17] 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 market rates in the reporting period hinder econometric exercises focused on long run parameters (see also Baele *et al*. 2004, Sørensen-Werner 2006).

[18] This feature should not represent much of an inconsistence, as the correlation, both in levels and first differences, with the average rate on overdrafts - not included in the NRIR database - is almost one (Di Lorenzo-Marotta 2005).

[19] For Spain, the estimation sample starts in June, instead of January, 1993, in order to have I(1) regressors (results are available at request). For Belgium, the sample starts at December 1995 under *MPA*, due to the availability of data for the overnight rate. We were unable to find the overnight rate series for Ireland.

than at the beginning (e.g. Germany[20]) and high negative correlations with market rates. This first inspection of data would then suggest an a priori case *against* a higher LPT in the post EMU period, a surprising result if in the meantime monetary uncertainty has fallen.

INSERT FIGURE 1 APPROXIMATIVELY HERE

4 Empirical results

4.1 Break dates

The proposed approach for searching multiple unknown breaks yields a single date for four countries; two or three ones are detected instead in the other five ones (Table 2)[21]. The dates are in general similar or differ at most up to four months, irrespective of the driving market rate. The main exception is Spain, where the break date according to *MPA* - March 1997 - is 15 months earlier than the one under *CoFA.* A multiple breaks approach shows therefore how pursuing both *MPA* and *CoFA* in order to detect changes in LPTs, as suggested by Sander-Kleimeier (2004a; Table 1), yields a low payoff, at least for the short term business lending rate, as it should be expected given the very close correlation among overnight and interbank rates.

The puzzle of very different dates for Italy and Portugal, depending on the driving rate, can be explained because each approach picks, in single break framework, just one out of two or three breaks. This is indeed a general result, in spite of the longer sample used in this paper: one date coincides or is very close to the ones detected in Sander-Kleimeier (2004a), because of *the* highest supF, confirming the findings of the encompassing exercise for a Italy and Portugal in Di Lorenzo-Marotta (2005).

An interesting feature suggested by the search exercise is that an expectational rationale for structural breaks in LPTs before the start of EMU, once the process had reached a defined aspect - say late 1996/first half 1997 - , could be suggested only for France. Further breaks, detected some months after the effective launch of the euro, for Austria, Italy and, even more, for Germany (with

[20] For similar evidence see Weth (2002, Figure 1).
[21] Detailed results on the supF statistics in the refinement procedure are available upon request.

the latest break in July 2000), hint at protracted adjustments. For Portugal no expectational effects can be inferred, being the breaks in early 1995 likely caused by the international financial turbulence at that time.

Finally, two cases deserve a closer scrutiny. Spain is the only country where a single break date is detected considerably later with respect to Sander-Kleimeier (2004a), especially under *COFA* (June 1998 instead of November 1996). This result, that raises doubts on the claim that the country would have experienced early the impact of the run up to EMU, is explained by the choice, as a driving market rate, of the three months interbank rate, instead of the one month one in this study. If the first were adopted, two breaks would be detected (May 1994 and March 1997), but the choice would be in contrast with the criterion of the highest correlation with the retail rate[22].

In the case of the lending rate to the 10% top-rated borrowers for Italy - hence hardly representative of the bulk of the market - , besides a very early break, as in Sander-Kleimeier (2004a), an additional one is detected in December 1996. This break could be rationalized as possibly induced by expectations for the advent of the euro, on the grounds that Italy entered again the EMS in November 1996, fulfilling a requirement for the inclusion in the EMU founding group, four years after the exit because of the 1992 crisis.

[TABLE 2 APPROXIMATELY HERE]

4.2 Pass-through estimates

Given the focus of this paper on structural changes in PTs possibly linked to EMU only the estimates for the break-free periods after 1995 are reported. The methodology laid out in Section 3 is implemented with the only exception for the penultimate period of Germany: owing to the poor

[22] The correlation coefficient for variables in levels is 0.99 for both interbank rates in Sander-Kleimeier (2004b, Table B1). The correlation coefficients for first differences are instead 0.84 for the 1 month one and 0.79 for the 3 months one.

statistical quality of the retail rate series[23], nor a cointegration relation or an ARDL(2,2) specification satisfying minimal statistical properties could be estimated[24].

The main result is that β in the last break-free period *shrinks* everywhere, with the exception of France[25]; θ increases in most countries, but rather unevenly, and with the noticeable exception of Germany; γ on average does not change. As these findings hold broadly irrespective of the chosen driving market rate, in contrast to Sander-Kleimeier (2004a), detailed comments refer only to those under *CoFA*: the preference is motivated by a better maturity matching of an interbank rate with the retail one as well as by more precise estimates (Table 3).

LPT size. β in the post-EMU period falls everywhere - on average from 0.9 to 0.7 -, except for France[26]. The range of values, as suggested by the pattern of spreads, goes from about one for Netherlands to 0.25 (down from 0.78 until 1997[27]) for Germany, with a cluster around 0.7 for the other countries. When multiple break dates suggest expectational effects as well as ex-post (EMU) adjustments, β keeps reducing (Austria, Italy). Taking into account the relative size of Germany in the EMU area, these results hardly suggest that the single monetary policy has produced more uniformity in LPTs across countries (the unweighted coefficient of variation increases from 0.27 to 0.30; excluding Germany, instead, it falls from from 0.27 to 0.18; Table 4).

[23] The series, even with data in levels, fluctuates very little, possibly because, as explained in the Bundesbank web site, *the average rates are calculated as unweighted arithmetic means from the interest rates reported to be within the spread. The spread is ascertained by eliminating the reports in the top 5% and the bottom 5% of the interest rate range.* Germany is the country that shows the highest differences between the unharmonized (in the NRIR database) and harmonized (in the new database, MIR) series for short term enterprises loan rates (on average, approximately 450 basis points; Sørensen-Werner, 2006, Chart a3.B).

[24] Test statistics for misspecification (normality and serial correlation of the first order for residuals in the ECM or in the ARDL specifications) are significant at the 5% level in about one fourth of the cases; both tests reject the null in only 3 cases (2 refer to the Italian r_2 equations). Results are available upon request.

[25] The monthly series for France is almost a quarterly one. We prefer, given the focus on the break dates search for LPT, to stick to the original series, as also Coffinet (2005) does, instead of interpolating somehow arbitrarily, as in de Bondt *et al.* (2005).

[26] Only the EG point estimates are reported, as it is well known that the standard errors of the OLS estimates are not interpretable in the usual way; this is instead possible – and standard errors are reported - with the ARDL estimates.

[27] Comparable results, based on panel estimates for the period April 1993 – December 2000, can be found in Weth (2002, Table 4).

Speed of adjustment and IPT. θ, a parameter that can be estimated only in an ECM framework[28], increases in the last break-free period in most countries (on average, from 0.39 to 0.57), with a coefficient of variation rising from 0.20 to 0.45. Combining (the point estimates of) speed of adjustment to the and LPT, i.e. βθ, because there can be a trade off between the two parameters for the working of the market, two outcomes are worth remarking in the cross country comparison in the last break-free periods: *i*) the ranking is similar, except for Austria, France and Netherlands; *ii*) range and coefficient of variation are much larger in the last period. Changes of γ split about evenly across countries (up in Austria, France, Germany, down in Italy, Portugal, Spain), leading to averages and coefficients of variation almost identical overtime.

Overall, these results suggest therefore only a slightly faster transfer of market rates changes into the lending ones.

[TABLES 3 AND 4 APPROXIMATIVELY HERE]

4.3 Discussion

The bottom line of the econometric investigation is that in a cross country comparison banks' pricing policies underwent more than a single structural change, at different dates, during the period the process of preparation and implementation of EMU took place. These changes are not however the ones – reduced heterogeneity and tendency towards unitary LPT - expected with the implementation of an uncertainty-reducing single monetary policy. Some possible offsetting factors are, against the backdrop of a sluggish growth after the peak at mid-2000 in the EMU area and in some countries in particular, the consolidation of the banking industry, mostly within national borders, and the Basel 2 process towards the revision capital requirements[29].

[28] A high statistical significance, as obtained in this paper, is a further check that the ECM specification is data-consistent. Interestingly, the significance level is generally lower when the overnight, rather than an interbank, is the driving market rate.

[29] Domestic market structure features can have further interacted. For instance, in 2002 the EU Commission convicted seven large Austrian banks for having arranged an interest rate cartel (Burgstaller 2003).

The slowdown of the pace of growth led to slower lending to the corporate sector. The negative effects on the financial position of firms led to a deterioration of the asset quality of banks, as witnessed by the increase in loan-loss provisions and the adoption of stricter lending criteria (ECB 2004). In the run up towards Basel 2 these developments are likely to have led to higher risk premia embedded in the lending rate, thus reducing PTs (Figure 1)[30].

Domestic consolidation of the banking industry is likely to have increased the lenders' market power relative to SMEs. An instance can be gleaned by the divergent pattern of LPTs for r_2 - the minimum rate for the 10 percent top-rated borrowers - in comparison with r_1 - the lending rate to non-primary borrowers - in the Italian case (Table 3; Figure 1). This pattern is compatible with the working of a dual credit market. The best borrowers exploited their bargaining power, obtaining lending rates, r_2, close to money market ones; enhanced relationship lending with the bulk of customers[31] could have produced the expected intertemporal smoothing for the broad-based lending rate, r_1 (Berlin-Mester 1998).

The difficulties in disentangling the different factors suggest caution in linking the structural changes detected in PT relations to the advent of EMU. That said, the results on LPTs - a generalized reduction (except for France), well below one (except for the Netherlands) - that derive from an agnostic view on dating breaks strongly support the view of a dampening of the impulses of a single monetary policy through the short term business lending rate. This outcome is somewhat in agreement with the scepticism of de Bondt *et al.* (2005, 15) towards the view of an increase of PTs in the euro area since the start of EMU (Angeloni-Ehrman 2003) or in the last break-free period (Sander-Kleimeier 2004a)[32].

[30] The average lending margin for short and long term corporate lending increased, between May 98-May 99 and May 01-May 02, in four countries. Germany, in particular, had an increase of 36 basis points, and became the second most expensive lender after Ireland (Cabral *et al* 2002, Table 17).

[31] The developments for two indicators between June 1999 and September 2003, such as the number of multiple lending relationships, decreased by one sixth, and the share of the main bank's loans, increased by about seven percentage points, lend support to this view (Di Lorenzo-Marotta 2005).

[32] Gambacorta and Iannoni (2005, Table 4) find for Italy a unitary LPT but a rather low speed of adjustment (0.19) for the short-term lending rate, in an Asymmetric Vector Error Correction Model, including also the current accounts and the three months interbank rates, estimated during the period 1993.9 2002.12. However, they impose in the long run PT

The outcome of an incomplete LPT in the last break-free period is in agreement also with a panel study for the post-EMU period, where the new harmonized series, from January 2003 to June 2004, are reconstructed backwards to January 1999, using the NRIR series. The findings of Sørensen-Werner (2006, Tables A4, A10) on short-term lending rate to firms LPTs, for the same countries examined in this paper, are very similar to those for the last break-free period (average 0.78 vs 0.70 in this paper, standard deviation of 0.18 vs 0.21), except for Portugal, that like Netherlands has a unitary LPT. Remarkable differences emerge instead for the speed of adjustment: overall the estimated θs are lower (the average is 0.41 vs 0.57)[33].

5. Conclusion

This paper makes several contributions to the empirical literature investigating the likely effects of EMU on the short term business lending rate PT. This retail rate is the natural benchmark to assess whether EMU has made the transmission mechanism more uniform, given that some earlier studies find that its PT in the post-EMU period is the largest and fastest, with a resulting reduction in cross country heterogeneity, among all lending and deposit rates.

An econometric methodology is proposed in order to endogenously search for *multiple unknown* structural breaks when retail and market rates are integrated I(1) regressors. Break dates are therefore left to be determined by the data, instead of being assumed to coincide with the introduction of the single currency or with a single unknown date. This approach allows therefore for both expectational effects and adjustments after the implementation of the new monetary regime. The methodology is implemented using the most extended sample available after the introduction of the euro in each of the nine EMU countries contributing to the NRIR database.

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relation a "convergence" dummy variable for the constant term, taking the value 1 between 1995.3-1998.9 and zero elsewhere, instead of modelling a change in the slope coefficient.

[33] The estimates are however hardly plausible for Germany (-0.05), Austria (-0.03) and Belgium (-0.17). In addition, for the last two countries they are not statistically different from zero even at the 10% significance level, raising the suspicion that cointegration could not hold.

The empirical investigation yields a single break for France, Ireland, Netherlands and Spain; two or three breaks are detected in the remaining five countries. The dates are in general similar or differ at most four months, irrespective of the driving market rate, be an overnight or an interbank one. The distinction between a monetary policy and a cost of funds approach, identified respectively with the chosen type of market rate in Sander-Kleimeier (2004a), does not seem therefore to yield significative payoffs, at least for this retail rate. Overall, because of an a priori better maturity matching and of more reliable statistical results, the second and more traditional approach turns out to be preferable.

A case for an early structural break much before the start of EMU, possibly based on an expectational rationale, can be empirically made only for France; further breaks are instead detected some quarters after the advent of euro for Austria, Italy and Germany. The empirical results support the view of a dampening of the impulses of a single monetary policy through the short term business lending rate. The shrinking in the last break-free period of the long run PT, with the exception of France, well below one, except for Netherlands, is only partially compensated by the speedier adjustment to the (lower) equilibrium value in some countries, but with Germany being a notable outlier.

These results are broadly in agreement with the scepticism of de Bondt *et al* (2005, 15) towards the view of an increase of PTs in the euro area since the start of EMU (Angeloni-Ehrman 2003) or in the last break-free period (Sander-Kleimeier 2004a); they match the bulk of the findings for the EMU period, with a different database, of Sørensen-Werner (2006). The transmission through the market for short-term business lending has become more heterogeneous, contrary to earlier studies conclusions (Sander-Kleimeier 2004a).

The overall picture contrasts with the economic intuition that a reduced volatility in money market rates, owing to the introduction of the euro, is bound to mitigate uncertainty and therefore to ease the transfer of monetary impulses to retail rates. These expected effects could have been offset by other contemporaneously evolving processes, such as the consolidation and concentration of the

banking industry, mostly within national borders, and the revision of Basel capital requirements, during a prolonged period of low output growth and of lenders' deteriorating creditworthiness.

Panel studies with microdata could help disentangle the effects of these different factors on lending rate PT, along the lines of Gambacorta (2004), de Graeve *et al* (2004), Lago-Gonzalez and Salas-Fumás (2005), provided they include a proper treatment of multiple unknown structural breaks.

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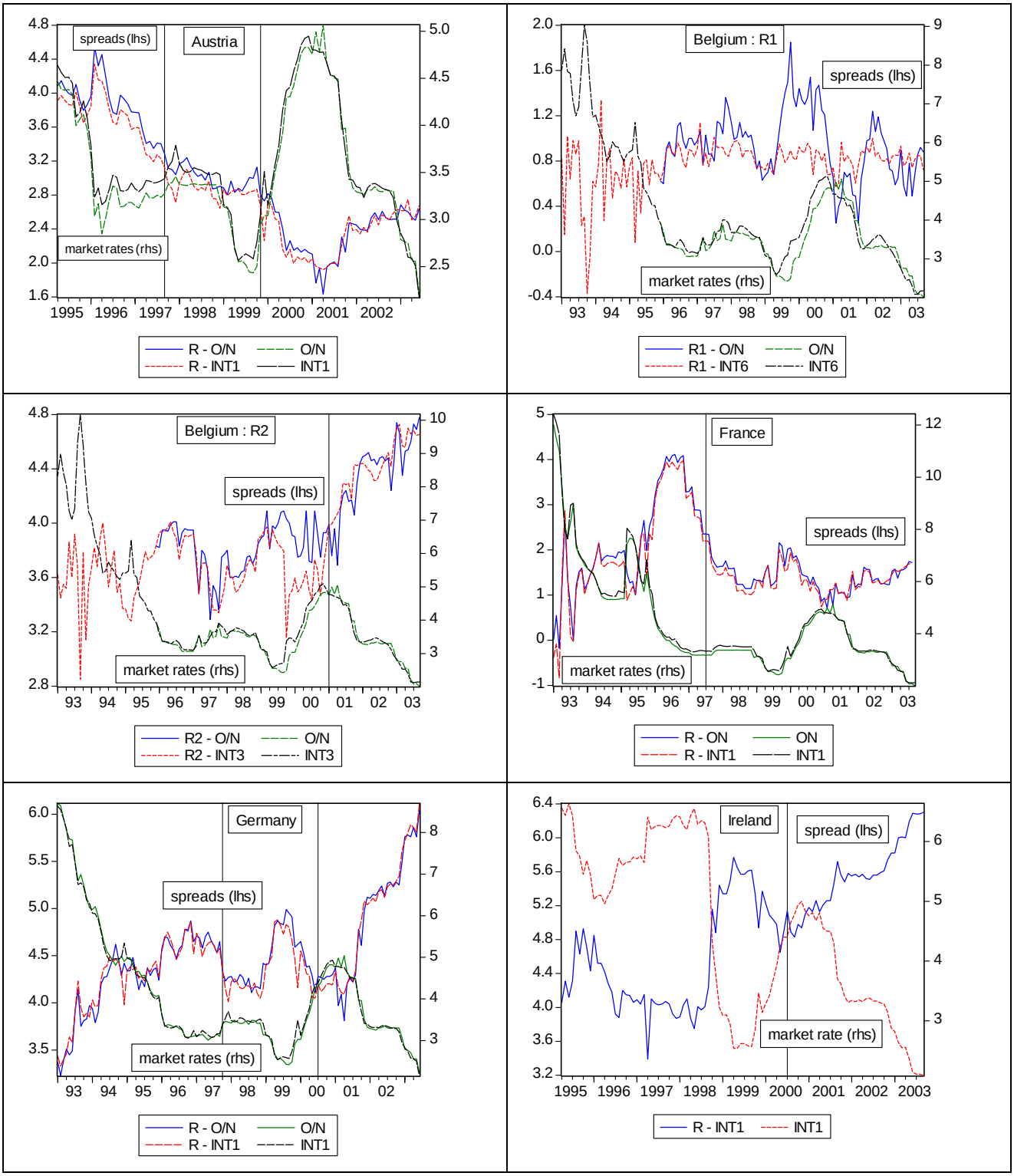
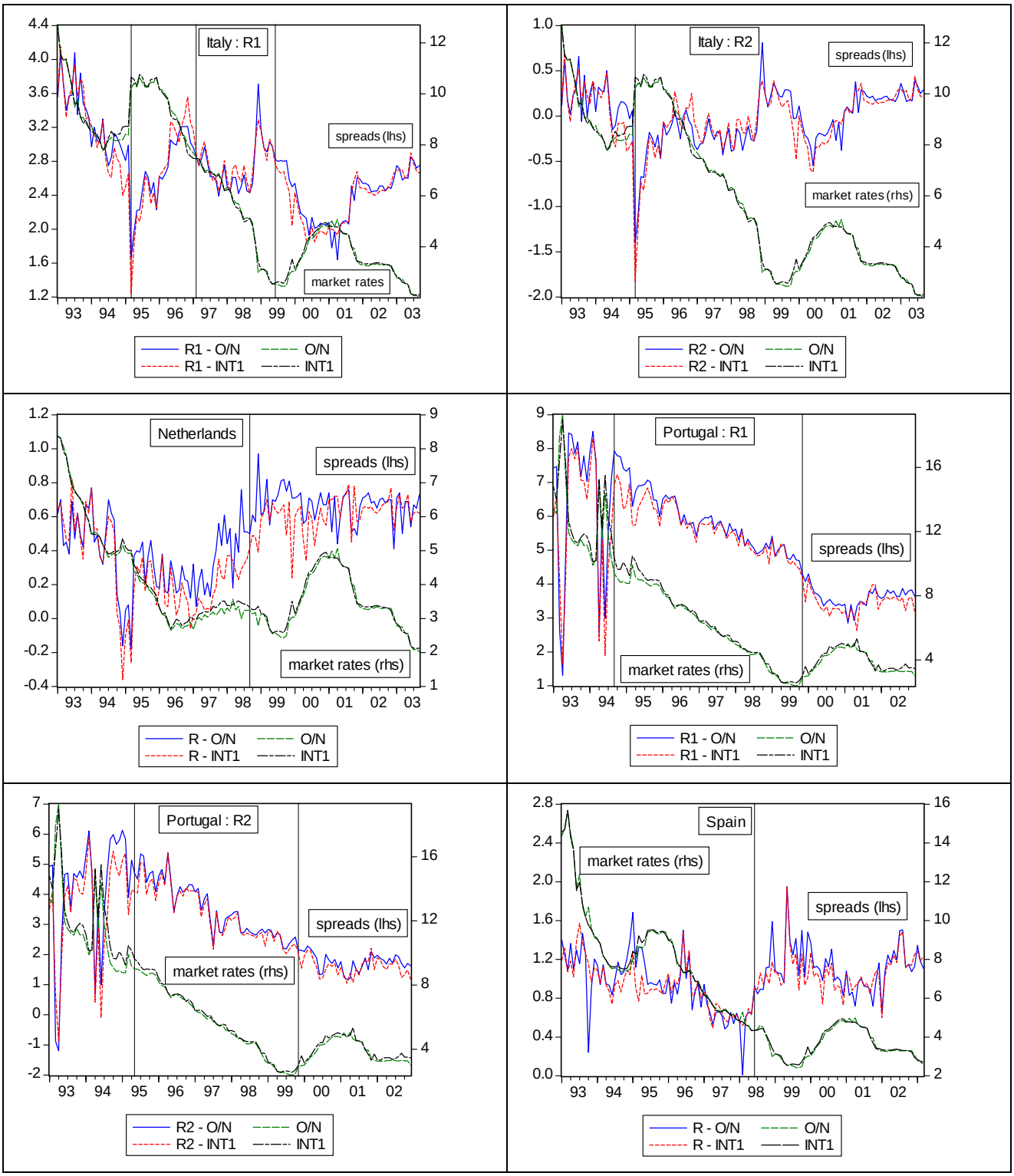
Figure 1 **Spreads for short term lending rates to business, market rates and break dates (percentage points)**





Source: authors' calculations. Vertical lines at break dates detected when using a cost of funds approach (see Table 2).

Table 1 Review of the literature on the pass-through to business short term loan interest rates

Study	Market rate	Break date	Sample	Short run pass-through (γ)	Long run pass-through (β)	Adjustment speed (θ)
			Austria			
Sander-Kleimeier (2004b) EG procedure or, absent cointegration, ARDL estimation	Overnight	August 1997	95.04-97.08	0.03	1.02	
			97.09-02.10	0.24	0.52	
	Government 10 years bond		95.04-97.08	0.05	1.19	
			97.09-02.10	0.26	0.56	
de Bondt *et al.* (2005) Single-step ECM estimation	3 months interbank / Government 10 years bond	January 1999, *a priori*	94.04-02.12	0.24***/-0.02	0.38***/0.65***	-0.12***
			99.01-02.12	0.38***/-0.01	0.62***	-0.37***
			Belgium : r_1			
Sander-Kleimeier (2004b)	Overnight	April 1995	93.01-95.04	0.41	0.43	
			95.05-02.10	-0.01	0.80	
	6 months interbank	March 1995	93.01-95.03	0.20	0.44	
			95.04-02.10	0.39	0.84	
de Bondt *et al.* (2005)	3 months interbank / Government 10 years bond	January 1999, *a priori* (Chow test *p*-value = 0.10)	94.04-02.12	0.75***/0.31***	0.59***/0.21*	-0.23
			99.01-02.12	0.96***/0.38***	0.81***/0.28**	-0.52**
			Belgium : r_2			
Sander-Kleimeier (2004b)	Overnight	January 1994	93.01-95.04	n.a.	n.a.	
			95.05-02.10	0.27	0.84	
	3 months interbank	December 1993	93.01-93.12	n.a.	n.a.	
			94.01-02.10	0.29	0.85	
			France			
Hofmann (2003); Single-step ECM estimation	3 months interbank	January 1999, *a priori*	95.01-02.11	-0.11	1 *a priori*	-0.11***
			99.01-02.11	0.62***		-0.42***
Sander-Kleimeier (2004b)	Overnight	June 1997	93.01-97.06	0.06	0.56	
			97.07-02.10	0.21	0.72	
	6 months interbank		93.01-97.06	0.27	0.54	
			97.07-02.10	0.32	0.77	
de Bondt *et al.* (2005)	3 months interbank / Government 10 years bond	January 1999, *a priori*	94.04-02.12	0.35 / -0.09	0.86 / 0.37*	-0.30***
			99.01-02.12	0.90 / -0.36	0.78***	-0.77
Coffinet (2005) Single-step ECM estimation	3 months interbank	January 1999, *a priori*	86.01-98.12	0.08	0.79***	-0.17
			99.01-03.09	0.48***	0.77***	-0.13
			Germany			
Hofmann (2003)	3 months interbank	January 1999, *a priori*	95.01-02.11	0.28***	1 *a priori*	-0.06***
			99.01-02.11	0.23***		-0.08***

Sander-Kleimeier (2004b)	Overnight	July 2000	93.01-00.07	0.16	081	
			00.08-02.10	0.30	0.44	
	1 month interbank	February 2001	93.01-01-02	0.23	0.78	
			01.03-02.10	0.26	0.25	
de Bondt *et al.* (2005)	3 months interbank / Government 10 years bond	January 1999, *a priori*	94.04-02.12	0.18***/ -0.02	0.36	-0.02
			99.01-02.12	0.08/0.01	- / 0.73	-0.02
de Bondt (2005)	1 month interbank	January 1999, *a priori*	96.01-01.05	0.12	1.05	-.13**
			99.01-01.05	0.02	0.89	-.23**
Ireland						
Sander-Kleimeier (2004b)	Overnight	November 1995	93.01-95.11	0.40	0.65	
			95.12-02.10	0.26	0.53	
	3 months interbank	December 1993	93.01-93.12	n.a.	n.a.	
			94.01-02.10	0.43	0.57	
de Bondt *et al.* (2005)	3 months interbank / Government 10 years bond	January 1999, *a priori*	94.04-02.12	0.43***/-0.14**	0.55***	-0.09
			99.01-02.12	0.21**	0.87***	-0.19***
Italy: r_1						
Hofmann (2003)	3 months interbank	January 1999, *a priori*	95.01-02.11	0.17***	1 *a priori*	-0.18***
			99.01-02.11	0.25***		-0.23***
Sander-Kleimeier (2004b)	Overnight	February 1995	93.01-95.02	0.31	1.09	
			97.03-02.10	0.16	0.96	
	1 month interbank	July 1999	93.01-99.07	0.27	1.02	
			99.08-02.10	0.31	0.68	
de Bondt *et al.* (2005)	3 months interbank / Government 10 years bond	NO (Chow test *p*-value = 0.20)	94.04-02.12	0.19***/-0.01	0.93*** / 0.12*	-0.15***
			99.01-02.12	0.16***/ -0.07	0.76*** / - 0.15***	-0.60***
Italy: r_2						
Sander-Kleimeier (2004b)	Overnight	February 1995	93.01-95.02	0.43	0.94	
			95.03-02.10	0.21	0.92	
	1 month interbank	June 1994	94.07-02.10	0.31	0.95	
Netherlands						
Sander-Kleimeier (2004b)	Overnight	August 1997	93.01-97.08	0.44	1.08	
			97.09-02.10	0.40	0.99	
	1 month interbank	August 1998	93.01-98.08	0.19	1.06	
			98.09-02.10	1.01	1.00	
de Bondt *et al.* (2005)	3 months interbank / Government 10 years bond	January 1999, *a priori*	94.04-02.12	0.57***/-0.02	1.15***/-0.31***	-0.31***
			99.01-02.12	0.44***/-0.01	1.05***	-0.77***
Portugal: r_1						
Sander-Kleimeier (2004b)	Overnight	July 1994	93.01-94.07	-	0.26	
			94.08-02.10	0.22	1.52[a]	
	1 month interbank	October 1999	94.10-99.10	0.25	1.24	
			99.11-02.10	0.23	0.65	
de Bondt *et al.* (2005)	3 months interbank / Government 10 years bond	NO (Chow test *p*-value = 0.57 in January 1999)	94.04-02.12	0.36***/-0.37***	1.24***	-0.25***
			99.01-02.12	0.64***/-0.28	0.93***	-0.27**
Portugal: r_2						

Sander-Kleimeier (2004b)	Overnight	February 1995	93.01-95.02	0.15	0.33	
			95.03-02.10	0.50	1.51	
	1 month interbank	November 1999	94.10-99.11	0.61	1.33	
			99.12-02.10	0.78	0.77	
Spain						
Hofmann (2003)	3 months interbank	NO (Chow test *p*-value = 0.19 in January 1999)	95.01-02.11	0.64***	1 *a priori*	-0.52***
			99.01-02.11	0.52***		-0.65***
Sander-Kleimeier (2004b)	Overnight	September 1996	93.01-96.09	0.24	0.85	
			96.10-02.10	0.39	0.78	
	3 months interbank	November 1996	93.01-96.11	0.64	0.97	
			96.12-02.10	0.64	0.79	
de Bondt *et al*. (2005)	3 months interbank / Government 10 years bond	January 1999, *a priori*	94.04-02.12	0.76***/0.03	0.96***	-0.41***
			99.01-02.12	0.58***/0.08	0.87***	-0.73***

Sources: Hofmann (2003), Table 1; Sander-Kleimeier (2004b), Tables B3-B4; de Bondt *et al*. (2005), Table 4; de Bondt (2005), Table A1; Coffinet (2005), Tableau A2. For Belgium, Italy and Portugal, the NRIR database includes two rates, coded N4.1 and N4.2. Unless specified the literature deals only with the first of the two rates. [a] Computed as the long run coefficient in an ARDL specification. ***, **, *: statistically significant at the 1, 5 and 10 per cent level.

Table 2 Break dates for short-term business lending rates long run pass-throughs

Country	Sample	Overnight	1 month interbank rate[a]
		Break date	
Austria	1995.04-2003.06	*October 1997*	*September 1997*
		March 2000	Novembre 1999
Belgium			
- r_1	1995.12/ 1993.01-2003.09	July 1999	No break
		December 2000	
- r_2	1995.12/ 1993.01-2003.09	May 2000	January 2001
France	1993.01-2003.08	*June 1997*	*June 1997*
Germany	1993.01-2003.06	April 1995	October 1997
		October 1997	
		July 2000	July 2000
Ireland	1995.04-2003.09	n.a.	July 2000
Italy			
- r_1	1993.01-2004.02	*March 1995*	March 1995
		February 1997	February 1997
		June 1999	*June 1999*
- r_2		March 1995	March 1995
Netherlands	1993.01-2003.09	June 1998	*September 1998*
Portugal			
- r_1	1993.01-2002.12	*August 1994*	September 1994
		November 1999	*November 1999*
- r_2		*March 1995*	May 1995
		November 1999	November 1999
Spain	1993.06-2003.03	March 1997	June 1998

Source: authors' calculations. In italics, break dates common with Sander-Kleimeier (2004a). [a]6- and 3-months interbank rate for r_1 and r_2 for Belgium, respectively.

Table 3 Short run business lending rate pass-throughs across EMU countries

(Engle-Granger two-step procedure unless otherwise stated; standard errors in brackets[1])

Market rate	Sample Period	α	β	θ	γ_0	Cointegration tests: CRDW[2], ADF[3]
			Austria			
1 month interbank	95:04-97:09		1 (0.28)		0.08 (0.04)	ADL(1,0)
	97:10-99:11	2.67	1.06	-0.39 (0.16)	0.62 (0.13)	CDRW = 1.07*** ADF = -3.63**
	99:12-03:06	3.50	0.69	-0.38 (0.15)	0.55 (0.09)	CDRW = 0.94*** ADF = -3.67***
Overnight	95:04-97:10		1 (0.32)		0.06 (0.04)	ADL(1,0)
	97:11-00:03		1 (0.11)		~0.53 (0.13)	ADL(1,1)
	00:04-03:06	3.45	~0.70	~-0.24 (0.13)	~0.46 (0.09)	CDRW = 0.84*** ADF = -2.08**
			Belgium: r_1			
6 months interbank	93:01-03:09	1.09	0.93	-0.80 (0.18)	0.86 (0.11)	CDRW = 1.68*** ADF = -6.54***
Overnight	95:12-99:07	1.03	0.98	-0.21 (0.12)	0.61 (0.13)	CDRW = 0.68*** ADF = -3.10***
	99:08-00:12	1.99	0.82	-0.83 (0.36)	0.74 (0.30)	CDRW = 1.55*** ADF = -3.09***
	01:01-03:09	1.17	0.87	-0.21 (0.12)	0.56 (0.17)	CDRW = 0.65*** ADF = -2.48**
			Belgium: r_2			
3 months interbank	93:01-01:01	3.87	0.95	-0.47 (0.12)	0.79 (0.09)	CDRW = 1.01*** ADF = -5.60***
	01:02-03:09	5.20	0.77	-0.67 (0.20)	0.71 (0.14)	CDRW = 1.21*** ADF = -3.88***
Overnight	95:12-00:05	4.84	0.69	-0.29 (0.12)	0.52 (0.12)	CDRW = 0.63*** ADF = -2.51**
	00:06-03:09	5.43	0.69	-0.55 (0.19)	0.48 (0.12)	CDRW = 0.99*** ADF = -4.46***
			France			
1 month interbank	93:01-97:06	4.82	0.53	-0.29 (0.15)	0.27 (0.12)	CDRW = 0.53*** ADF = -2.17**
	97:07-03:08	2.14	0.78	-0.36 (0.09)	0.73 (0.14)	CDRW = 0.61*** ADF = -4.65***
Overnight	93:01-97:06	4.82	0.55	-0.25 (0.14)	0.36 (0.13)	CDRW = 0.47** ADF = -2.19**
	97:07-03:08	2.31	0.74	-0.30 (0.09)	0.63 (0.14)	CDRW = 0.53*** ADF = -4.31***
			Germany			
1 month interbank	93:01-97:10	5.34	0.78	-0.36 (0.10)	0.46 (0.06)	CDRW = .95*** ADF = -3.98***
	97:11-00:07				0.15 (0.05)	first differences regression
	00:08-03:06	7.72	0.25	-0.28 (0.12)	0.26 (0.08)	CDRW = 0.47** ADF = -2.62**
Overnight	95:05-97:10	5.70	0.67	-0.62 (0.18)	0.30 (0.10)	CDRW = 1.30*** ADF = -3.30***
	97:11-00:07				0.14 (0.07)	first differences regression
	00:08-03:06	7.71	0.25	-0.30 (0.11)	0.16 (0.05)	CDRW = 0.61*** ADF = -2.70***
			Ireland			
1 month interbank	95:04-00:07	6.64	0.58	-0.44 (0.11)	0.29 (0.10)	CDRW = 0.71*** ADF = -12.16***
	00:08-03:09	7.07	0.58	-0.29 (0.10)	0.47 (0.07)	CDRW = 0.39** ADF = -5.77***
			Italy : r_1			
1 month interbank	95:04-97:02	6.03	0.65	-0.37 (0.13)	0.34 (0.11)	CDRW = 0.46** ADF = -1.93*

	97:03-99:06	3.14	0.93	-0.35 (0.10)	0.56 (0.04)	CDRW = 0.79*** ADF = -3.62***
	99:07-03:09	3.49	0.68	-0.52 (0.07)	0.25 (0.05)	CDRW = 0.55*** ADF =-2.69***
Overnight	95:04-97:02	5.31	0.73	-0.37 (0.13)	0.48 (0.11)	CDRW = 0.45** ADF = -2.21**
	97:03-99:06	3.26	0.90	-0.30 (0.15)	0.47 (0.07)	CDRW = 0.78*** ADF = -2.58**
	99:07-03:09	3.55	0.67	-0.51 (0.09)	0.34 (0.05)	CDRW = 0.95*** ADF =-2.36**
Italy : r_2						
1 month interbank	95:04-03:09	0.30	0.94	-0.40 (0.04)	0.42 (0.04)	CDRW = 0.42** ADF =-4.71***
Overnight	95:04-03:09	0.35	0.93	-0.41 (0.04)	0.45 (0.04)	CDRW = 0.63*** ADF = -5.24***
Netherlands						
1 month interbank	93:01-98:09	-0.14	1.09	-0.31 (0.09)	0.67 (0.10)	CDRW =0.75*** ADF = -3.81***
	98:10-03:09	0.63	1.00	-0.95 (0.14)	0.80 (0.13)	CDRW =1.78*** ADF = -6.99***
Overnight	93:01-98:06	0.13	1.05	-0.35 (0.10)	0.58 (0.10)	CDRW =0.96*** ADF = -4.20***
	98:07-03:09	0.74	0.98	-0.95 (0.14)	1.01 (0.10)	CDRW =1.81*** ADF = -7.40***
Portugal : r_1						
1 month interbank	94:10-99:11	4.06	1.25	-0.35 (0.09)	0.34 (0.11)	CDRW =1.28*** ADF = -5.64***
	99:12-02:12	4.84	0.65	-0.55 (0.12)	0.15 (0.08)	CDRW = 1.40*** ADF = -4.20**
Overnight	94:09-99:11	3.80	1.34	-0.27 (0.08)	0.47 (0.12)	CDRW = 0.87*** ADF = -4.61***
	99:12-02:12	4.94	0.66	-0.59 (0.16)	0.31 (0.12)	CDRW = 1.48*** ADF = -4.60***
Portugal : r_2						
1 month interbank	95:06-99:11	1.26	1.36	-0.71 (0.12)	0.44 (0.23)	CDRW = 1.53*** ADF = -2.75***
	99:12-02:12	2.53	0.77	-0.50 (0.17)	0.49 (0.14)	CDRW = 1.05*** ADF = -2.33**
Overnight	95:04-99:11	1.29	1.39	-0.70 (0.13)	0.75 (0.26)	CDRW = 1.43*** ADF = -5.49***
	99:12-02:12	2.71	0.75	-0.59 (0.18)	0.53 (0.19)	CDRW = 1.22*** ADF = -3.89***
Spain						
1 month interbank	93:06-98:06	0.21	1.09	-0.53 (0.12)	1.03 (0.06)	CDRW = 1.04*** ADF = -4.44***
	98:07-03:03	1.59	0.86	-0.91 (0.14)	0.80 (0.12)	CDRW = 1.81*** ADF = -6.78***
Overnight	93:06-97:03	0.83	1.02	-0.61 (0.08)	0.90 (0.09)	CDRW = 1.38*** ADF = -4.83***
	97:04-03:03	2.01	0.74	-0.67 (0.12)	0.62 (0.11)	CDRW = 1.41*** ADF = -6.16***

[1] Heteroskedasticity consistent whenever the White test is below the 5% significance level. [2]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. [3] Asymptotic critical values under the null of I(1) first stage residuals at the 1% (***), 5% (**) and 10% (*) significance; ADF with no constant (MacKinnon 1996).

Table 4 **Key parameters of lending rate pass-through across countries (absolute values; within brackets, statistics excluding Germany)**

Country	Penultimate break-free period				Last break-free period			
	β	**θ**	**βθ**	**γ**	**β**	**θ**	**βθ**	**γ**
Austria	1.06	0.39	0.41	0.62	0.69	0.38	0.26	0.55
Belgium r_2	0.95	0.47	0.45	0.79	0.77	0.67	0.52	0.71
France	0.53	0.29	0.16	0.27	0.78	0.36	0.28	0.73
Germany	0.78	0.36	0.28	0.46	0.36	0.28	0.07	0.26
Ireland	0.58	0.44	0.26	0.29	0.58	0.29	0.17	0.47
Italy r_1	0.93	0.35	0.33	0.56	0.68	0.52	0.35	0.52
Netherlands	1.09	0.31	0.34	0.67	1	0.95	0.95	0.8
Portugal r_1	1.25	0.35	0.44	0.34	0.65	0.55	0.36	0.15
Spain	1.09	0.53	0.58	1.03	0.86	0.91	0.78	0.8
average	**0.90 (0.93)**	**0.39 (0.39)**	**0.35 (0.37)**	**0.55 (0.57)**	**0.70 (0.75)**	**0.57 (0.58)**	**0.43 (0.46)**	**0.56 (0.59)**
standard deviation	**0.24 (0.25)**	**(0.08) (0.08)**	**0.13 (0.13)**	**0.25 (0.27)**	**0.21 (0.13)**	**0.25 (0.25)**	**0.29 (0.27)**	**0.23 (0.22)**
coefficient of variation	**0.27 (0.27)**	**0.20 (0.21)**	**0.36 (0.35)**	**0.46 (0.46)**	**0.30 (0.18)**	**0.45 (0.43)**	**0.66 (0.60)**	**0.42 (0.37)**

Source: Table 3. For Germany, the period before the penultimate one.