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Structural breaks in the interest rate pass-through and the euro
A cross-country study in the euro area and the UK

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Abstract

We search for multiple unknown structural breaks in the short term business lending rate pass-through in euro countries, possibly associated with the introduction of the single currency. One break is detected in five EMU countries, two are found in other four, and in the UK as well. The last break occurs much before the event for France, several quarters later for Austria, Germany, Italy and Portugal, and the UK, hinting at best at a loose link with the inception of EMU. Long run pass-throughs decrease (except for France), becoming even more incomplete (except for the Netherlands and the UK); though the adjustment to equilibrium has become faster, cross-country heterogeneity in the euro area has barely changed. An incomplete lending rate pass-through, even in the long run, for the least sticky bank rate and the persistence of cross-country heterogeneity make tougher for the ECB to realize an effective area-wide monetary policy.

JEL Classification: E43; E52; E58; F36

Keywords: Interest rates; Monetary policy; European and Monetary Union (EMU); Cointegration analysis; Structural breaks

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1. Introduction*

The transmission of monetary policy hinges on how policy rate changes, via changes in market interest rates, affect bank rates, that are likely to influence aggregate demand at least to some extent. A monetary policy impulse, obeying the Taylor principle - that a central bank should raise its interest rate instrument more than one-to-one with increases in inflation (Woodford 2003, 91) - can however fail to be stabilizing if the pass-through (PT) to retail rates is incomplete even in the long run. Is this the case even after the introduction of the euro in a bank-based financial system such as the European Monetary Union (EMU)?

The issue has been investigated in the literature considering whether size and speed of PT have *increased* in the wake of EMU, thus enhancing the effectiveness of the single monetary policy, and *converged*, thus making more uniform the transmission via the banking sector across countries. 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 cast by the conflicting tests on a structural break in coincidence with the introduction of the euro (de Bondt *et al* 2005). In addition, criticism has been levelled at the assumption of January 1999 as a break date. An alternative empirical strategy is to search for a single unknown break date and to estimate PTs in the two periods (Toolsema *et al* 2002, Sander-Kleimeier (SK) 2004a, b).

The simple point raised in this paper is that a natural starting point to assess whether the making of EMU has modified bank interest rate PTs is to acknowledge that there is no theoretical nor empirical ground to assume a single structural break. The historical innovation of the euro is in fact the outcome of a *process*, announced well before its formal implementation and unlikely to imply an instantaneous adjustment in banks' pricing policy. During the longest sample usually considered in the literature - 1993 to late 2002[1] - PTs could have been altered because of at least three factors closely intertwined with the conduct of monetary policy: the turbulence in the exchange rate markets in early 1995 for Italy, Portugal and Spain[2]; the expectations induced by the final steps in the transition to EMU and the ensuing convergence in market interest rates; the

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[1] The choice of the starting year, 1993 in SK (2004a, b) or 1994 in de Bondt *et al* (2005), is meant to avoid the turbulence derived from the September 1992 crisis of the European Monetary System (EMS).

[2] 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 Mexican debt crisis.

adjustment of national banking sectors to the working of a unified monetary regime. *The* latest break-date is obviously the most interesting one in an investigation of the effects, if ever, of EMU on the transmission of monetary impulses to bank rates.

We explore the implications of this view focusing on short term business loans (BL) - the first link in the transmission mechanism through banks - because their interest rate PT turns out to be the largest one among the banking products in the literature. This finding is however, perhaps, the only area of agreement. The null of equality for impact PTs (IPT) of short term BL rates across countries is soundly rejected both before and after January 1999[3] (Angeloni-Ehrman 2003); the null of no structural break is not rejected instead only for Italy and Portugal in de Bondt *et al* (2005). The date of the endogenously detected break is approximately two years before the inception of EMU for Austria, France and Spain, a finding that is taken as a hint at expectational effects in these countries; it is instead far after that event for Germany. A puzzling result is that break-dates differ up to 4/5 years for Italy and Portugal, depending on the choice of the driving market rate (SK 2004a). Adopting the empirical strategy of searching for an additional break-date in the latest sub-sample, instead of stopping at the first one as in Toolsema *et al* 2002 and SK (2004b), Di Lorenzo and Marotta (2006) show however that break-dates are very similar for the two countries, regardless of the driving market rate. The literature provides also different findings for LPT changes through time.

To make sense of the variety of the results on dating breaks and on estimating size and speed of PTs this paper adopts the empirical strategy of searching for multiple breaks in nine founding EMU countries, and in a control non-EMU country like the UK as well, using the longest available common sample. The robustness of the findings is checked investigating three issues. First, how different are LPT estimates splitting the sample at January 1999, possibly because the short track of EMU makes the data not very informative to detect breaks? Second, does the expected competition enhancement of the single monetary policy translates in predictable changes in IPTs, when the econometric specification allows for asymmetric responses to market rates increase or decrease? Third, are the findings on breaks robust to a refinement approach, originally laid out for the case of multiple unknown breaks with stationary regressors in Bai (1997), when extended to the case of regressors integrated of order one, or I(1), as interest rates most often turn out to be? The tentative nature of the last exercise must be stressed because a formal framework to search for multiple unknown breaks with I(1) regressors is as yet missing (Perron 2006).

This paper makes several contributions to the literature. Two breaks are indeed detected in four countries, as well as in the United Kingdom, a single one in five and none only in the case of a

[3] Actually 5 countries, due to data availability.

retail rate in Belgium; the starting date of the latest break-free period varies across countries from mid-1977 to early 2001. Comparing the last two break-free periods, LPTs *decrease* (except for France) well below one (except for the Netherlands and the UK); the adjustment to equilibrium is generally *faster*; the monetary transmission across countries has not become more uniform. The results of the main exercise on break-dates and LPTs survive the robustness checks. It is worth noticing that there is evidence of an asymmetric IPT in a few EMU countries and in the UK as well, hinting at an enhanced competition in loan markets through time.

The paper is organized as follows. Section 2 describes the data set and provides an overview of lending spread patterns across countries; Section 3 surveys the related literature. Section 4 lays out the empirical strategy to search for multiple unknown break-dates in cointegrated relations, whereas the empirical results are shown and discussed in Section 5. Finally, Section 6 summarizes and concludes.

2. Data description

A visual inspection of the data is useful to set the stage for the empirical investigation. The short term BL rate is the series coded "N4" for each of the nine contributing countries to the unharmonized National Retail Interest Rates (NRIR) database at the European Central Bank (ECB)[4]. The sample starts, at the earliest, at January 1993 (see fn 1) and ends, at most, at September 2003[5]. The rate is computed as an average for new businesses, except for Italy (outstanding stocks with a maturity up to 18 months)[6]. For comparison with a non-EMU, but a member of the European Union, country we consider also the 4 UK clearing banks' base rate, downloaded from the Bank of England website. The chosen interbank market rate, that should match the (short) maturity of the underlying credit aggregates for an appropriate pricing[7], is the most correlated (in first differences) with the retail rate among those with a maturity of 1, 3, 6 or 12 months, following de Bondt (2002)[8].

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

[5] As of January 2003 the ECB collects a new set of harmonized bank rates statistics (denoted with the MIR acronym), 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 hinders econometric exercises focused on long run parameters (see also Baele *et al*. 2004, Sørensen-Werner 2006, ECB 2006).

[6] This feature should not represent much of an inconsistency, because the correlation, both in levels and in first differences, with the average rate on overdrafts - not included in the NRIR database - is almost one (Di Lorenzo-Marotta 2006).

[7] 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 through time and/or across countries, when in fact there is nothing but a different mix of instruments with different interest rate fixation characteristics.

[8] Results available upon request

An inspection of the lending spread (short term BL rate net of the interbank rate) for the nine EMU countries and the UK, against the backdrop of a dramatic fall of market rates since early 1995, in particular for Italy, Portugal and Spain, with an inversion in the first two years after the introduction of the euro and a subsequent - mid 2001 - further decline to low historical levels, yields several interesting features (Fig. 1).

Lending spreads - approximately constant in the benchmark case of a complete LPT - and interbank rate changes should be uncorrelated if the adjustment to LPT is fast. The effective patterns for the two series are indeed quite varied through time and across countries. Only the spreads for France, the Netherlands, Portugal and Spain come close, in recent years, to the benchmark case, as it happens for the UK and for the US by mid-1990s (Sellon 2002). The other EMU countries show instead upwards trending spreads, with end-sample levels sometimes higher than at the beginning (e.g. Belgium, Germany[9], Ireland).

The visual inspection of the data, corroborated by a formal test of the null of stationarity for the lending spreads, would then suggest for the euro area an a priori case against a complete LPT during the entire sample and, perhaps more unexpectedly, even *against* a definite tendency towards it in the EMU period as well[10]. This would be a rather puzzling result, if confirmed by the empirical investigation, when at the same time monetary policy has become more predictable and, as a consequence, the competition in banking markets has supposedly increased.

INSERT FIGURE 1 APPROXIMATIVELY HERE

3. Related literature

Recent literature on short term BL rate PTs provides a wide range of results as to the date of a *single* structural break, possibly coincident with the start of EMU, as well as to the changes in LPTs and the adjustment speed to equilibrium. Angeloni and Ehrmann (2003) argue that a single bank reserves market and the reduction in money market interest rates volatility, due to the ECB operating procedures, have already produced larger and faster bank rate PTs. First, they identify via rolling-window regressions January 1999 as a break-point. Second, they find 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. In particular, impact and peak PTs for short and long

[9] For similar evidence see Weth (2002, Figure 1).

[10] The Kwiatkowski-Phillips-Schmidt-Shin (KPSS) (level) test statistic, adjusted for sample size (Sephton 1995), rejects always the null of stationarity of lending spreads at least at the 5% significance level for the common period starting April 1995, except for one of the two rates of Belgium and Italy and for Spain; the rejection rate is only slightly lower after January 1999 (Table A.1, in the Appendix).

term BL rates computed for the euro area show the largest increases (from 0.35 to 0.53, from 0.81 to 1.11, respectively)[11]. De Bondt (2005), on the contrary, finds that LPTs for all euro area bank rates, except the mortgage one, are *lower* in the EMU period - with a Chow test rejecting the null of no break at January 1999 - compared to the extended one (January 1996-June 2001). In particular, LPT for the short term BL rate shrinks from 1.53 to 0.88.

Considering this bank rate, cross-country and national studies disagree even more, mostly because of the choice of the driving market rate and of how to deal with EMU-related breaks[12].

Hofmann (2003), who *assumes* a unitary LPT and as a driver the 3-months interbank rate, finds that the break at January 1999 is not statistically significant for Spain and that whereas the adjustment to equilibrium becomes faster after the introduction of euro, though remaining puzzlingly slow for Germany, IPTs increase in France and Italy and fall in Germany and Spain (Table 1).

De Bondt *et al* (2005) adopt as a driver a combination, with estimated weights, of the 3-months interbank rate and of the 10-years Government bond yield, under the assumption that the second one provides a signal on the persistence of changes of the policy rate. They also assume January 1999 as a break date and, even if a Chow test *does not* rejects the null for Italy and Portugal, they run estimates for all countries over an extended sample and over the EMU one. Their findings are that in the last period the long term market rate becomes statistically insignificant, LPTs *decrease* well below one (except for the Netherlands), IPTs rise in Austria, France, the Netherlands and Portugal and fall in Italy and Spain. The estimates for Germany are always very imprecise.

SK (2004a,b) endogenously search for a *single* break in PT equations with alternative driving market rates. They propose in fact a 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 a market rate to better proxy the marginal cost of loaned funds. The findings are rather heterogeneous 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[13]; 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 euro for France (June 1997), Austria (August 1997) and Spain

[11] No break was however detected in the equation for an index of lending rates in an euro area monthly monetary model (Bruggeman-Donnay 2003).

[12] We survey studies with up to 2002 data; earlier cross-country studies are Donnay-Degryse (2001) and Heinemann-Schüler (2003).

[13] Di Lorenzo and Marotta (2006) show that a second break date, much nearer to the start of EMU, can be found in the last period and is very similar under *MPA* and *COFA,* as it should be expected given the very close correlation among overnight and interbank rates.

(September/November 1996) as well as much later for Germany (July 2000/February 2001). LPTs show opposite patterns over time (on average, from 0.91 to 0.72 under *CoFA*, from 0.71 to 0.87 under *MPA*); IPTs increase at most slightly.

In two national studies, under the assumption of January 1999 as a break-date, 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, though with a sample extending only to May 2001 (de Bondt 2005). Gambacorta and Iannotti (2005, Table 4) find for Italy a unitary LPT but a rather low speed of adjustment (0.19), in an Asymmetric Vector Error Correction Model including also the current accounts and the three months interbank rate, estimated during the period 1993:09 2002:12. They impose however in the LPT relation a "convergence" dummy variable for the constant term, taking the value 1 between 1995:03-1998:09 and zero elsewhere, instead of modelling a change in the slope coefficient.

[TABLE 1 APPROXIMATIVELY HERE]

4. Econometrics

The assumption of a *single known* structural break in the interest rate LPT in coincidence with the introduction 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 protracted adjustments on the other hand cannot be ruled out. The maintained hypothesis in this paper is therefore of *multiple unknown* breaks. The econometric literature does not provide however as yet a suitable framework for a search in the case of I(1) regressors, as interest rates almost invariably turn out to be (Perron 2006, 287).

To circumvent this obstacle, we adopt an econometric approach aimed at providing robust answers to the key questions: do LPTs and the speed of adjustment towards equilibrium have changed, and how much? To this end, besides including the longest available sample after the introduction of the euro for the short term BL rates self-selected as the most representative by each EMU country, we follow and extend Di Lorenzo-Marotta (2006), that generalizes the approach of Toolsema *et al* (2002) and SK (2004b).

The reference setting, as in the literature surveyed, 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 (Klein 1971, Monti 1972). The lending rate is determined as a mark-up over the marginal (opportunity) cost, proxyied by a market rate, matching the maturity of loans. Assuming a linear approximation, in a competitive market the marginal cost coefficient can be interpreted as the LPT, with a complete transfer of changes of the driving market rate to the retail one (Lago-Gonzalez and Salas-Fumás 2005). For estimation purposes, the Autoregressive Distributed Lags (ARDL) specification in Cottarelli-Kourelis (1994) is reparametrized as an Error Correction Mechanism (ECM), following the Granger representation theorem for cointegrated variables[14].

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

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

with I(0) OLS residuals, *ecm*, at the first stage of the Engle-Granger (1987) two-step estimation procedure (EG) [15], 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.

Eq, (1) includes only a constant, the presence of a linear trend in interest rates being theoretically inconsistent (Hamilton 1994, 501). Short term dynamics parameters are obtained in the EG second step dropping sequentially insignificant regressors according to the general-to-specific approach (Hendry 1995):

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

where Δ is the first difference operator.

The key parameters are γ (i.e. IPT), β (i.e. LPT), and θ (i.e. the adjustment speed to equilibrium), also known as loading factor, that should result statistically significant if cointegration holds. Within this basic framework the empirical investigation in this paper proceeds as follows.

First, we search for a single unknown break-date in the long run model (Eq. 1), adopting the supremum F (supF) testing procedure. This means that the date is associated with the largest (and statistically significant) rolling standard Chow F-statistics computed under the null of a break occurring in each subsequent period through the mid-70% sample period (Andrews 1993)[16]. When the algorithm yields several local maxima, it is rerun, starting from the earliest break-point, to detect the successive one, and so on. We consider only an interbank driving rate, owing to the expected

[14] 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.

[15] 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 to reduced sample size (Maddala-Kim 1998).

[16] The asymptotic distribution is non-standard because, when the break-date is unknown, it is a nuisance parameter that appears only under the alternative hypothesis of structural break. For critical values see Table 2.

better match with the short term BL maturity and because of the Di Lorenzo-Marotta (2006) findings on the similar dating of breaks when considering also the alternative of an overnight rate.

Second, we check that in the last two break-free periods the EG first step estimation 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). If cointegration holds, we proceed with the EG second step for (Eq. 2). In order to enhance comparison across countries, and owing to sample size constraints, the same short-run dynamics is imposed, leaving out lagged first-differenced regressors[17]. The first-step estimate of β is superconsistent, but biased in small samples, and the bias is inversely proportional to the fit (Banerjee *et al* 1986); asymptotic cointegration tests can have low power in relatively small samples. For that reason we adopt also an alternative one-step procedure, jointly estimating an ECM specification that combines Eqs. (1) and (2). This is important on two grounds: a) we obtain confidence intervals for all three key parameters whereas, as it is well known, the t-statistic in the EG first-step procedure cannot be interpreted in the standard way, owing to the super-consistency of the estimates; b) we are able to use an alternative cointegration test, adjusted for the degrees of freedom (Ericsson-MacKinnon 2002; EM).

Finally, we check the robustness of the main results exploring three issues: January 1999 as a break-date; asymmetric IPTs; break detection mimicking a refinement search procedure.

5 Results

5.1 Break dates

To implement the proposed approach the first step is choosing the market rate and testing the order of integration for the regressors. The one month interbank rate turns out to be the most correlated (in first differences) with the bank rate; only for Belgium the chosen interbank rates are the 6- and 3-months for the retail rates r_1 and r_2, respectively[18]. Augmented Dickey-Fuller (ADF) tests show that most interest rates are I(1) over the available samples (Table A.2, in the Appendix).

A single break-date is detected for Belgium (r_2), France, Ireland, Netherlands, Spain; two are found for Austria, Germany, Italy (r_1), Portugal and none for Belgium (r_1) [19]; Table 2 and Figure A.1 in the Appendix. These findings suggest first that, in contrast with the gist of SK (2004a), an expectational rationale for structural breaks in LPTs before the start of EMU, once the process had

[17] 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).
[18] Results available upon request.
[19] We checked that the dates are indeed the same or differ at most up to four months, irrespective of the driving market rate, interbank or overnight. An exception is Spain, where the break date according to *MPA* - March 1997 - is 15 months earlier than under *CoFA* (results available upon request).

reached a defined aspect - say late 1996/first half 1997 - , could fit only the French experience. No such effects can be inferred for Portugal, being the breaks in late 1994-early 1995 likely caused by the international financial turbulence at that time. Second, breaks detected some months after the launch of the euro for Austria, Italy and, even more, for Germany (with the latest break in March 2001), hint at protracted adjustments of these national banking industries. Third, a note of caution in associating structural changes to EMU is suggested by the break dates - January 1996 and November 2000 - detected in the UK: the second date is in particular barely distinguishable from one that could have been motivated by a process of adjustment to a new monetary environment in an euro country.

A case deserves a closer scrutiny. Spain is the only country where a single break date is detected considerably later with respect to SK (2004a) under *COFA* (June 1998 instead of November 1996). This result, that casts doubts on the claim that the country would have experienced early the impact of the run-up to the EMU, can be explained by the choice of the three months interbank rate in that study, in contrast with the advocated criterion of the highest correlation with the retail rate[20].

[TABLE 2 APPROXIMATELY HERE]

5.2 Pass-through

Owing to the focus on structural changes, possibly linked to the introduction of the euro, we report the results only for the last two break-free periods (Table 3). Overall, most estimates are highly statistically significant and pass at least one of the cointegration tests[21].

The main results are:

i) β shrinks everywhere in the last period, with the exception of France[22], well below one. These findings hold irrespective of the estimation method, except for a few cases (Netherlands and, only for the absolute size but not for the change through time, for Germany and Italy (r_1); Table 3);

[20] The correlation coefficient for variables in levels is 0.99 for 1 month and 3 months interbank rates, but are 0.84 and 0.79, respectively, for the first-differences (SK 2004b, Table B1).

[21] The exceptions are France and Portugal (r_2) in one period, but the loading factor θ is statistically significant, and most especially Germany. The likely cause in the last case is a data problem. The retail rate series fluctuates very little, possibly because, as explained in the Bundesbank web site, the average rates are computed as unweighted arithmetic means from the interest rates reported by banks, after eliminating those in the top 5% and the bottom 5% of the interest rate range. Germany is also the country that shows the highest differences between the unharmonized (in the NRIR database) and the harmonized series for short term enterprises loan rates (on average, approximately 450 basis points; Sørensen-Werner, 2006, Chart a3.B).

[22] 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).

ii) θ increases in most countries, except for Ireland;, also γ increases on average, but with sizably differences across countries. The speed of adjustment has therefore become faster and the mean lag much shorter.

A detailed examination follows, concentrating on the one-step ECM estimates.

LPT. β falls on average from 0.9 to 0.7. The range of estimates, as hinted by the pattern of lending spreads, cluster around 0.7 for the most countries, with an outlier of 0.21 (down from 0.97[23]) for Germany; the unitary value is outside the upper end of the 5% confidence interval everywhere, except for the Netherlands and Portugal (r_2). These results hardly suggest that in the period overlapping (at least partially) with the introduction of the euro LPTs have come closer to being complete. They also provide mixed evidence on the claim that cross-country LPTs have become more uniform: the unweighted coefficient of variation increases from 0.26 to 0.31 (though from 0.27 to 0.16 excluding Germany; Table 4).

Speed of adjustment and IPT. θ increases in most countries (on average, from 0.32 to 0.54), with a slightly larger coefficient of variation (from 0.39 to 0.46). It could be argued that from a policy point of view a lower LPT could be acceptable if the adjustment to it were faster. The averaged indicator $\beta\theta$ indeed increases (from 0.30 to 0.43), but so does also the coefficient of variation across countries (from 0.46 to 0.68). Changes of γ follow the same pattern (average from 0.43 to 0.51, coefficient of variation from 0.50 to 0.75). Combining γ and θ, the mean lag halves on average (from 2.5 to 1.2 months; from 1.8 to 1 excluding Germany), though some countries show a higher value in the last period (Belgium, Ireland, Spain).

The estimation results for the UK are useful to put in perspective these results for the EMU countries. Whereas β slightly increases to one, with a complete LPT, just like the Netherlands (and the US as well; Sellon 2002), the increased mean lag, 0.4 instead of 0.24 months, is as small as one third of the corresponding average value for the euro countries.

[TABLES 3 AND 4 APPROXIMATIVELY HERE]

5.3. Robustness

January 1999 as break-date. A first exercise to evaluate the robustness of the results so far examined is to compare the estimates for the last break-free period with those obtained in a sample starting at January 1999 (Table 5). Unsurprisingly, the main differences can be spotted for the countries with breaks that do not cluster around this date. βs are higher, considering point estimates

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

that differ from those obtained in the main exercise by more than one standard deviation, for Belgium (r_2) and Germany, though with hardly believable low θs, Ireland and Italy, These findings could explain, in the vein of as an informal encompassing exercise, claims in the literature about larger LPTs after the introduction of the euro. Even within this set of estimates, however, LPT is complete only for two countries (Belgium and the Neherlands).

Asymmetric IPT. A stylized fact in the PT literature is the asymmetric pattern of lending rates when market rates rise or decline. We check the symmetric specification for the short term dynamics in the main exercise adding a regressor picking only positive changes in the interbank rate (has zero values otherwise). The expected sign of this slope-dummy regressor (γ^+) is negative if, owing to the competition in loan markets, banks are more reluctant to transfer to borrowers increases in market rates than lowering them when the opposite happens.

We report, to save space, the one-step ECM estimates with γ^+ statistically significant at least at the 5% confidence level (Table 6). Two noteworthy results are: a) the temporal pattern of positively signed γ^+s in the last but one period for France and Germany and of negatively signed ones in the last period for the Netherlands and the UK hint at a process of increasing competition, at least in some countries; b) even when there are no recent structural breaks, as in the case of Italy's r_2 - the minimum rate for the 10 percent top-rated borrowers -, the negatively signed γ^+ can be rationalized because of the competitive features characterizing that market niche (Di Lorenzo-Marotta 2006). The Irish case is rather puzzling, and deserves further research.

These results show that an asymmetric *IPT* can be detected only in few countries/periods, thus backing overall the symmetric specification in the main exercise. The result for a non-EMU country like the UK suggests that it is a hard task to disentangle the effects of the inception of single monetary policy from those associated with the making of a single financial market in the European Union and of the ensuing liberalization and consolidation processes in the banking industry.

Refinement in the search for multiple breaks. The efficient procedure to detect *multiple unknown* break dates in a linear model with stationary regressors proposed by Bai (1997) relies basically on the supF approach. In the first stage, as in this paper, when the algorithm yields several statistically significant local maxima, it is rerun, starting from the earliest break-point, to detect the successive one, and so on. 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, the *refinement* means further searching a break-date in the samples $t_0 – t_2$, $t_1 – t_3$ and $t_2 – T$. In this second stage the intermediate t_i could change and the search stops when the dates become stable[24]. The procedure assumes a maximum number of unknown breaks over the

[24]Actually, implementing the procedure is a bit messy, because it is not obvious the sequence to refine further when an intermediate t_i changes. Suppose, refining over the interval $t_1 – t_3$, that an intermediate break point is found, different

entire sample; the intervals between dates must also be sufficiently large in order to apply asymptotic theory (Bai-Perron 1998). This is why we surmise that break dates are at most three in a sample starting at January 1993. 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 adoption of a single currency area, once the number of the founding countries was agreed (approximately late 1996 - first half of 1997); the third one could be located after the inception of EMU, as national banking systems adjusted to it. Mimicking the refinement procedure yields differences with respect to previous findings only for Italy (r_1), where a third break date – February 1997 – is detected, and Germany, where the last break is anticipated to July 2000 (results available on request). In the first case, the estimates of the key parameters for the last period but one remain pretty the same (β = 0.93, θ = -0.35, γ = 0.56); in the second case, the poor quality of the data hinders an informed assessment. Our final evaluation is therefore that the main exercise findings are robust to a refinement-like procedure.

[TABLES 5 AND 6 APPROXIMATIVELY HERE]

5.4 Discussion

The bottom line of the empirical investigation is that all nine EMU countries underwent one or two structural changes in banks' pricing policies, at different dates, in the period the process of preparation and implementation of EMU took place. These changes, though resulting in faster lending rate PTs, did not produce the expected, owing to a more predictable monetary policy, larger and more uniform LPTs. Some candidate offsetting factors were, against the backdrop of a sluggish growth after the peak at mid-2000 in the EMU area and in some large countries in particular, the consolidation of the banking industry, mostly within national borders, and the Basel 2 process towards the revision of capital requirements[25].

The sluggish growth led to slower lending to the corporate sector. The negative effects on the financial position of firms produced 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

from t_2, implying a modification of the original $t_0 - t_2$ and $t_2 - T$ periods. It is up to the researcher to choose over which of the two sample refine first.

[25] 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).

run up towards Basel 2 these developments are likely to have led to higher risk premia embedded in the lending rates, thus reducing interest rate PTs at least in some countries (Figure 1)[26].

Domestic consolidation of the banking industry is likely to have increased lenders' market power relative to SMEs. A piece of evidence is suggested in the Italian case 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 (Table 3; Figure 1). This fact fits in fact the working of a dual credit market. The best borrowers exploited their bargaining power, paying interest rates, r_2, close to money market ones (see also Table 6); enhanced relationship lending with the bulk of customers[27] could have instead produced the expected intertemporal smoothing for the broad-based lending rate, r_1 (Berlin-Mester 1998).

The difficulties in disentangling the different factors on aggregate time series, as well as the breaks in the UK case, suggest caution in linking the structural changes detected in interest rate LPTs to the inception, expected and effective, of EMU. Panel studies exploiting the richness of microdata could help, along the lines of Gambacorta (2004), de Graeve *et al* (2004), Lago-Gonzalez and Salas-Fumás (2005), provided they were integrated with a proper treatment of the multiple unknown structural breaks.

That said, the results this paper offers on LPTs - a generalized reduction (except for France), well below one (except for the Netherlands) - that derive from endogenously dating breaks strongly support the view of a dampening of the impulses of a single monetary policy in the long run via the short term BL rate. This finding reinforces the scepticism of de Bondt *et al* (2005, 15) about the view of an increase of PTs in the euro area since the introduction of the euro (Angeloni-Ehrman 2003) or in the last break-free period (SK 2004a, 474).

An incomplete LPT is also in agreement with a panel study in a cointegrated framework, where the new harmonized bank rate series, from January 2003 to June 2004, are reconstructed backwards to January 1999, using the NRIR data set. The findings on short term BL rate LPTs, are very similar to those for the last break-free period (average of 0.78 vs 0.71 in this paper, standard deviation of 0.18 vs 0.22), except for Portugal, that has a complete LPT like the Netherlands

[26] 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).

[27] 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 some support to this view (Di Lorenzo-Marotta 2006).

Sørensen-Werner (2006, Tables A4, A10). Remarkable differences emerge instead for the speed of adjustment: overall, the estimated θs are lower (the average is 0.41 vs 0.54 in this paper)[28].

6. Conclusion

This paper makes several contributions to the empirical literature investigating the structural break(s), possibly associated to the introduction of the euro, in the pass-though of monetary policy impulses, via changes in market rates, to bank interest-rates. The short term business lending rate is the natural choice to assess whether the monetary transmission has become more effective and uniform across countries, because in previous studies its pass-through is the largest and fastest in the EMU period among bank rates.

Instead of assuming a *single* break - either dated January 1999 or endogenously detected - we searched for *multiple unknown* breaks, allowing for expectational effects or adjustments after the implementation of the new monetary regime. The data set includes the longest available national interest rate series after the introduction of the single currency - up to September 2003 - for nine euro countries and for the UK, a non-euro member of the European Union, taken as a control country.

The empirical investigation yields a single break-date for Belgium (r_2), France, Ireland, Netherlands, Spain and none for Belgium (r_1); two are found for Austria, Germany, Italy (r_1), Portugal. A case for a break much before the inception of EMU, based on an expectational rationale, can be made only for France; further breaks are instead detected several quarters after January 1999 for Austria, Germany, Italy and Portugal. The findings of two breaks also in the UK cast doubts on linking structural changes in banks' pricing policies exclusively or predominantly to the introduction of the euro.

A comparison of the estimates for the last two break-free periods points to a dampening of the impulses of a single monetary policy via the short term business lending rate in the euro area. While the long run interest rate pass-through shrinks, with the exception of France, well below the unitary value found for the Netherlands, the adjustment to equilibrium has become generally faster, with a mean lag that has halved to slightly more than one month, still almost three times the corresponding value for the UK. The empirical evidence is mixed on whether the transmission of monetary policy has become more uniform across countries, contrary to earlier studies conclusions. An incomplete pass-through, even in the long run for the least sticky bank rate, and the persistence

[28] The estimates look 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, casting doubts that cointegration could not hold.

of cross-country heterogeneity make tougher for the ECB to realize an effective area-wide monetary policy.

The overall picture contrasts with the economic intuition that a reduced volatility in money market rates, owing to a single monetary policy, 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 developing processes in the sample period, 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 in the euro area.

Panel studies with microdata could help disentangling the effects of these different factors on lending rate pass-throughs, provided they include a proper treatment of multiple unknown structural breaks. Another interesting issue for future research is to investigate the implications of an incomplete bank interest rate pass-through in the euro area on the use of standard Taylor rules in assessing the monetary policy of the ECB in comparison with the central banks of countries, like the US and the UK, where the transfer of policy rate changes to bank rates is complete, at least in the case of business lending.

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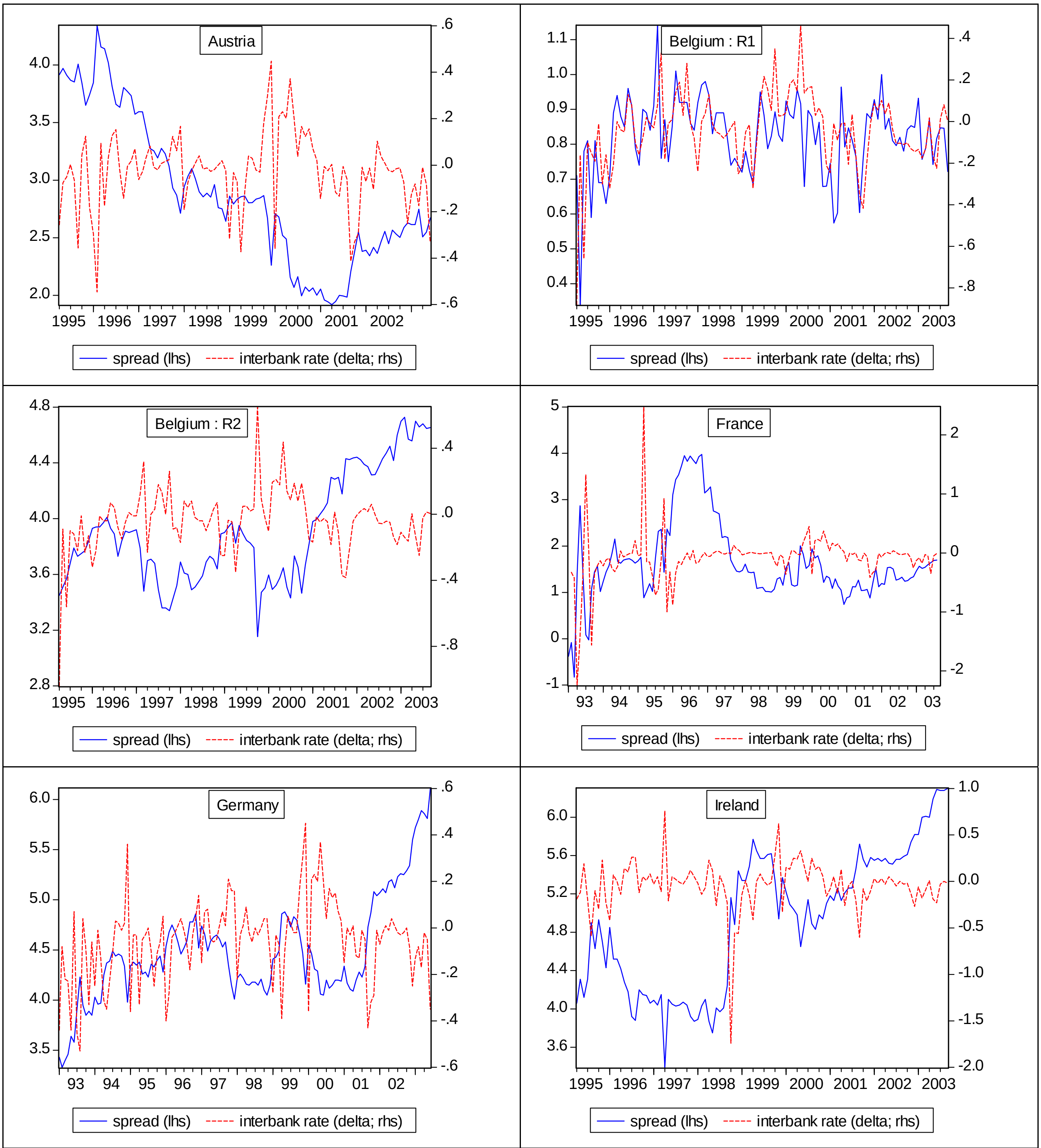
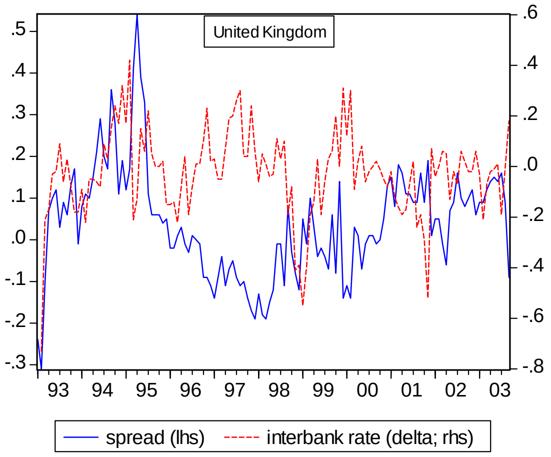
Figure 1 **Short term business lending spread and interbank rate changes**



ctd.


Italy : R1
4.4
4.0
3.6
3.2
2.8
2.4
2.0
1.6
1.2
1.6
1.2
0.8
0.4
0.0
-0.4
-0.8
93 94 95 96 97 98 99 00 01 02 03
spread (lhs) interbank rate (delta; rhs)
Italy : R2
1.0
0.5
0.0
-0.5
-1.0
-1.5
-2.0
1.6
1.2
0.8
0.4
0.0
-0.4
-0.8
93 94 95 96 97 98 99 00 01 02 03
spread (lhs) interbank rate (delta; rhs)
Netherlands
1.0
0.8
0.6
0.4
0.2
0.0
-0.2
-0.4
.8
.6
.4
.2
.0
-.2
-.4
-.6
93 94 95 96 97 98 99 00 01 02 03
spread (lhs) interbank rate (delta; rhs)
Portugal : R1
9
8
7
6
5
4
3
2
1
5
4
3
2
1
0
-1
-2
-3
93 94 95 96 97 98 99 00 01 02
spread (lhs) interbank rate (delta; rhs)
Portugal : R2
6
5
4
3
2
1
0
-1
5
4
3
2
1
0
-1
-2
-3
93 94 95 96 97 98 99 00 01 02
spread (lhs) interbank rate (delta; rhs)
Spain
2.0
1.6
1.2
0.8
0.4
0.5
0.0
-0.5
-1.0
-1.5
-2.0
93 94 95 96 97 98 99 00 01 02
spread (lhs) interbank rate (delta; rhs)

ctd.



Source: ECB and National Central Banks' websites. 1-month interbank rate, except for Belgium (6- and 3-months interbank rate for r_1 and r_2, respectively).

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

Study Estimation procedure	Market rate	Break date	Sample	Short run pass-through (γ)	Long run pass-through (β)	Adjustment speed (θ)
			Austria			
Sander-Kleimeier Engle-Granger two-step procedure	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*. one-step ECM	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	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*.	3 months interbank / Government 10 years bond	January 1999, *a priori*	94:04-02:12	0.75***/0.31***	0.59***/0.21*	-0.23
		(Chow test *p*-value = 0.10)	99:01-02:12	0.96***/0.38***	0.81***/0.28**	-0.52**
			Belgium (r_2)			
Sander-Kleimeier	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 one-step ECM	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	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*.	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 one-step ECM	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	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***
						ctd.

Study Estimation procedure	Market rate	Break date	Sample	Impact pass-through (γ)	Long run pass-through (β)	Adjustment speed (θ)
Sander-Kleimeier	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.*	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	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	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.*	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	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	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.*	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***
Di Lorenzo-Marotta one-step ECM	Overnight	June 1999 (last break)	95:04-99:06	0.25***	1.03***	-0.11**
			99:07-04:02	0.30***	0.73***	-0.22***
	1 month interbank	May 1999 (last break)	95:04-99:05	0.21***	1.07***	-0.22***
			99:06-04:02	0.27***	0.75***	-0.46***
Italy (r_2)						
Sander-Kleimeier	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	
Di Lorenzo-Marotta	Overnight	September 1997	95:04-97:09	0.31***	1.00***	-0.34***
			97:10-04:02	0.26***	0.81***	-0.24**
	1 month interbank	NO	95:04-04:02	0.29***	0.93***	-0.15***
Netherlands						
Sander-Kleimeier	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.*	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***

Study Estimation procedure	**Market rate**	**Break date**	**Sample**	**Impact pass-through (γ)**	**Long run pass-through (β)**	**Adjustment speed (θ)**
			Portugal (r_1)			
Sander-Kleimeier	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*.	3 months interbank / Government 10 years bond	NO (Chow test *p*-value = 0.57 at 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**
Di Lorenzo-Marotta	Overnight	November 1999	95:04-99:11		1.30	
			99:12-02:10		0.64	
	1 month interbank	October 1999	95:04-99:10		1.24	
			99:11-02:10		0.66	
			Portugal (r_2)			
Sander-Kleimeier	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	
Di Lorenzo-Marotta	Overnight	November 1999	95:04-99:11		1.39	
			99:12-02:10		0.72	
	1 month interbank	November 1999	95:04-99:11		1.36	
			99:12-02:10		0.78	
			Spain			
Hofmann	3 months interbank	NO (Chow test *p*-value = 0.19 at January 1999)	95:01-02:11	0.64***	1 *a priori*	-0.52***
			99:01-02:11	0.52***		-0.65***
Sander-Kleimeier	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*.	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: Coffinet (2005), Tableau A2 ; de Bondt (2005), Table A1; de Bondt *et al*. (2005), Table 4; Di Lorenzo-Marotta (2006) Tables 3, 6; Hofmann (2003), Table 1; Sander-Kleimeier (2004b), Tables B3-B4. [a]Long run coefficient in an Autoregressive Distributed Lags specification, owing to no cointegration. ***, **, *: statistically significant at the 1, 5 and 10 per cent level, when available.

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

Country	Full sample	1 month interbank rate[a]	
		Break date	**supF**[b]
Austria	1995.04-2003.06	*September 1997*	*256.54*
		November 1999	127.52
Belgium *(r_1)*	1993.01-2003.09	No break	14.84
Belgium *(r_2)*		January 2001	168.74
France	1993.01-2003.08	*June 1997*	*173.20*
Germany	1993.01-2003.06	October 1997	27.76
		March 2001	218.32
Ireland	1995.04-2003.09	July 2000	41.71
Italy *(r_1)*	1993.01-2004.02	March 1995	24.06
		June 1999	*60.30*
Italy *(r_2)*		August 1994	37.27
Netherlands	1993.01-2003.09	*September 1998*	*93.11*
Portugal *(r_1)*	1993.01-2002.12	September 1994	77.49
		November 1999	*296.04*
Portugal *(r_2)*		May 1995	124.89
		November 1999	115.67
Spain	1993.06-2003.03	June 1998	48.31
United Kingdom	1993.01-2003.09	January 1996	25.74
		November 2000	16.87

In italics, same break-dates as in Sander-Kleimeier (2004a, Table 1). [a]6- and 3-months interbank rate or Belgium (r_1 and r_2, respectively). [b]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). See also Figure A.1, in the Appendix.

Table 3 **Short term business lending rate pass-throughs**

(Heteroskedasticity consistent standard errors in brackets)

Estimation procedure	Sample Period	α	Long run pass-through (β)	Adjustment speed (θ)	Impact pass-through (γ)	Cointegration and misspecification tests: ADF[1], τ_c[2], JB[3], BG[4], EM[5]
			Austria			
Engle-Granger two-step procedure (A)	97:10-99:11	2.70	1.05	-0.36 (0.19)	0.61 (0.16)	ADF = -2.97** τ_c = -2.57
	99:12-03:06	3.50	0.69	-0.38 (0.15)	0.55 (0.09)	ADF = -3.67*** τ_c = -2.56
One-step ECM (B)	97:10-99:11	2.67 (0.31)	1.04 (0.10)	-0.36 (0.18)	0.59 (0.16)	JB = 0.13; BG = 2.24 EM = -2.03
	99:12-03:06	3.34 (0.21)	0.72 (0.05)	-0.35 (0.14)	0.48 (0.08)	JB = 1.38; BG = 2.97* EM = -2.48
			Belgium (r_1)			
A	93:01-03:09	1.09	0.93	-0.80 (0.18)	0.86 (0.11)	ADF = -6.54*** τ_c = -5.73***
B		1.12 (0.10)	0.92 (0.03)	-0.79 (0.18)	0.85 (0.10)	JB = 133.52***, BG = 1.56 EM = -4.42***
			Belgium (r_2)			
A	93:01-01:01	3.87	0.95	-0.47 (0.12)	0.79 (0.09)	ADF = -5.60*** τ_c = -3.34**
	01:02-03:09	5.20	0.77	-0.67 (0.20)	0.71 (0.14)	ADF = -3.88*** τ_c = -3.34**
B	93:01-01:01	3.89 (0.16)	0.95 (0.04)	-0.47 (0.12)	0.79 (0.10)	JB = 78.43***, BG = 9.09*** EM = -3.96***
	01:02-03:09	5.16 (0.09)	0.78 (0.03)	-0.62 (0.20)	0.70 (0.17)	JB = 0.13, BG = 0.67 EM = -3.07*
			France			
A	93:01-97:06	4.82	0.53	-0.29 (0.15)	0.27 (0.12)	ADF = -2.17 τ_c = -2.61
	97:07-03:08	2.14	0.78	-0.36 (0.09)	0.73 (0.14)	ADF = -4.65*** τ_c = -4.37***
B	93:01-97:06	4.81 (0.63)	0.49 (0.11)	-0.29 (0.15)	0.23 (0.16)	JB = 7.73***, BG = 0.47 EM = -1.95
	97:07-03:08	2.29 (0.37)	0.73 (0.10)	-0.31 (0.12)	0.68 (0.18)	JB = 30.50***, BG = 2.06 EM = -2.55
			Germany			
A	97:11-01:03	5.21	0.75	-0.10 (0.06)	0.33 (0.07)	ADF = -2.11 τ_c = -1.36
	01:04-03:06	7.89	0.20	-0.27 (0.13)	0.19 (0.07)	ADF = -1.95 τ_c = -1.51
B	97:11-01:03	4.62 (0.98)	0.97 (0.32)	-0.09 (0.07)	0.29 (0.06)	JB = 0.08, BG = 5.38*** EM = -1.33
	01:04-03:06	7.82 (0.17)	0.21 (0.04)	-0.27 (0.13)	0.17 (0.08)	JB = 1.30, BG = 1.26 EM = -2.06
			Ireland			
A	95:04-00:07	6.64	0.58	-0.44 (0.11)	0.29 (0.10)	ADF = -12.16*** τ_c = -2.50
	00:08-03:09	7.07	0.58	-0.29 (0.10)	0.47 (0.07)	ADF = -5.77*** τ_c = -2.64
B	95:04-00:07	6.53 (0.11)	0.60 (0.02)	-0.44 (0.11)	0.29 (0.10)	JB = 11.74***, BG = 2.68* EM = -3.85**
	00:08-03:09	6.93 (0.12)	0.61 (0.04)	-0.31 (0.12)	0.41 (0.09)	JB = 3.02, BG = 1.26 EM = -3.19*
			Italy (r_1)			
A	95:04-99:06	3.11	0.95	-0.24 (0.04)	0.47 (0.05)	ADF = -3.62*** τ_c = -2.48
	99:07-03:09	3.49	0.68	-0.52 (0.07)	0.25 (0.05)	ADF =-2.69* τ_c = -2.19
B	95:04-99:06	2.00 (0.18)	1.05 (0.02)	-0.27 (0.04)	0.23 (0.05)	JB = 1.18, BG = 2.60* EM = -7.46***
	99:07-03:09	3.29 (0.06)	0.73 (0.02)	-0.50 (0.05)	0.26 (0.05)	JB = 0.93, BG = 0.60 EM = -9.45***

ctd

Estimation procedure	**Sample Period**	α	**Long run pass-through (β)**	**Adjustment speed (θ)**	**Impact pass-through (γ)**	**Cointegration and misspecification tests: ADF[1] , τ_c [2], JB[3],BG[4],EM[5]**
			Italy (r_2)			
Engle-Granger two-step procedure (A)	94:08-03:09	0.33	0.93	-0.41 (0.03)	0.32 (0.06)	ADF =-4.47*** τ_c = -3.23*
One-step ECM (B)	94:08-03:09	0.16 (0.04)	0.94 (0.01)	-0.41 (0.03)	0.30 (0.05)	JB =11.05 ***, BG = 39.51*** EM = -16.27***
			Netherlands			
A	93:01-98:09	-0.14	1.09	-0.31 (0.09)	0.67 (0.10)	ADF = -3.81*** τ_c = -3.20*
	98:10-03:09	0.63	1.00	-0.95 (0.14)	0.80 (0.13)	ADF = -6.99*** τ_c = -4.40***
B	93:01-98:09	0.28 (0.24)	0.96 (0.07)	-0.26 (0.08)	0.40 (0.13)	JB = 2.00, BG = 1.43 EM = -3.38**
	98:10-03:09	0.59 (0.06)	1.01 (0.02)	-0.91 (0.14)	0.89 (0.13)	JB = 9.50***, BG = 2.95*** EM = -6.70***
			Portugal (r_1)			
A	94:10-99:11	4.06	1.25	-0.35 (0.09)	0.34 (0.11)	ADF = -5.64*** τ_c = -5.00***
	99:12-02:12	4.84	0.65	-0.55 (0.11)	0.16 (0.08)	ADF = -4.20** τ_c = -3.26*
B	94:10-99:11	3.62 (0.23)	1.24 (0.04)	-0.26 (0.08)	0.11 (0.12)	JB = 8.02**, BG = 8.34*** EM = -3.34**
	99:12-02:12	4.58 (0.31)	0.71 (0.07)	-0.54 (0.10)	0.18 (0.08)	JB = 6.45**, BG = 0.56 EM = -5.22***
			Portugal (r_2)			
A	95:06-99:11	1.26	1.36	-0.71 (0.11)	0.44 (0.26)	ADF = -2.76** τ_c = -4.95***
	99:12-02:12	2.53	0.77	-0.50 (0.17)	0.49 (0.14)	ADF = -2.33 τ_c = -2.27
B	95:06-99:11	1.11 (0.18)	1.35 (0.03)	-0.64 (0.11)	0.00 (0.28)	JB =15.86***, BG = 0.97 EM = -5.64***
	99:12-02:12	2.41 (0.52)	0.79 (0.13)	-0.44 (0.18)	0.47 (0.14)	JB = 0.76, BG = 0.01 EM = -2.41
			Spain			
A	93:01-98:06	0.42	1.06	-0.49 (0.12)	0.97 (0.04)	ADF = -4.68*** τ_c = -3.38**
	98:07-03:03	1.59	0.86	-0.91 (0.14)	0.80 (0.11)	ADF = -6.78*** τ_c = -5.08***
B	93:01-98:06	0.42 (0.13)	1.06 (0.02)	-0.49 (0.12)	0.96 (0.05)	JB =12.87***, BG = 1.95 EM = -4.24***
	98:07-03:03	1.57 (0.16)	0.86 (0.04)	-0.90 (0.14)	0.80 (0.11)	JB = 76.96***, BG = 0.05 EM = -6.37***
			United Kingdom			
A	96:02-00:11	0.26	0.95	-0.78 (0.13)	0.81 (0.05)	ADF = -5.63*** τ_c = -3.41**
	00:12-03:09	0.01	1.02	-0.67 (0.19)	0.82 (0.06)	ADF = -3.49*** τ_c = -2.75
B	96:02-00:11	0.22 (0.10)	0.96 (0.02)	-0.78 (0.13)	0.81 (0.05)	JB =4.38, BG = 1.26 EM = -6.11***
	00:12-03:09	0.03 (0.07)	1.01 (0.02)	-0.68 (0.16)	0.74 (0.08)	JB =2.63, BG = 1.62 EM = -4.13***

[1] Critical values under the null of I(1) EG first-step residuals for an ADF test statistic with no constant, computed for a sample size = 500 (Phillips-Ouliaris 1990). [2]Asymptotic critical values under the null of I(1) EG first-step residuals for a t-test statistic with constant and 1 lag (MacKinnon 1996). [3] Jarque-Bera test under the null of normality of residuals. [4] Breusch-Godfrey test under the null of no up to the second order correlation of residuals [5] Asymptotic critical values, adjusted for degrees-of-freedom, under the null of no cointegration (Ericsson-MacKinnon 2002). ***, **, *: statistically significant at the 1, 5 and 10 per cent level. Market rate: one-month interbank rate, except for Belgium (6 and 3 months interbank for r_1 and r_2, respectively)

Table 4 Key parameters for short term business lending rate pass-through
(absolute values; within brackets, statistics excluding Germany)

Country: last break date	Pre-break					Post-break				
	β	θ	βθ	γ	Mean lag (months)	β	θ	βθ	γ	Mean lag (months)
Austria: 1999:11	1.04	0.35	0.36	0.59	1.17	0.72	0.48	0.35	0.48	1.08
Belgium (r_2): 2001:01	0.95	0.47	0.45	0.79	0.45	0.78	0.62	0.48	0.70	0.48
France: 1997:06	0.49	0.29	0.14	0.23	2.66	0.73	0.31	0.23	0.68	1.03
Germany: 2001:03	0.97	0.09	0.09	0.29	7.89	0.21	0.27	0.06	0.17	3.07
Ireland : 2000:06	0.60	0.44	0.26	0.29	1.61	0.61	0.31	0.19	0.41	1.90
Italy (r_1): 1999:06	1.05	0.27	0.28	0.23	2.85	0.73	0.50	0.36	0.26	1.48
Netherlands: 1998:09	0.96	0.26	0.25	0.40	2.31	1.01	0.95	0.96	0.89	0.12
Portugal (r_1): 1999:11	1.24	0.26	0.32	0.11	3.42	0.71	0.54	0.38	0.18	1.52
Spain: 1998:06	1.06	0.49	0.52	0.96	0.08	0.86	0.90	0.77	0.80	0.22
Average (EMU)	**0.93 (0.92)**	**0.32 (0.35)**	**0.30 (0.32)**	**0.43 (0.45)**	**2.49 (1.82)**	**0.71 (0.77)**	**0.54 (0.58)**	**0.43 (0.45)**	**0.51 (0.55)**	**1.21 (0.98)**
standard deviation (EMU)	**0.24 (0.25)**	**0.13 (0.10)**	**0.14 (0.12)**	**0.29 (0.30)**	**2.31 (1.19)**	**0.22 (0.12)**	**0.25 (0.24)**	**0.29 (0.27)**	**0.27 (0.26)**	**0.93 (0.65)**
coefficient of variation (EMU)	**0.25 (0.27)**	**0.39 (0.28)**	**0.46 (0.37)**	**0.66 (0.67)**	**0.93 (0.66)**	**0.31 (0.16)**	**0.46 (0.42)**	**0.68 (0.58)**	**0.54 (0.47)**	**0.76 (0.67)**
UK: 2000:11	0.96	0.78	0.75	0.81	0.24	1.01	0.68	0.69	0.74	0.38

Source: Table 3, one-step ECM estimates. Mean lag = $(1-\gamma)/\theta$.

Table 5 Short term business lending rate pass-through in the EMU period

(One-step ECM; heteroskedasticity consistent standard errors in brackets)

Country sample period	α	β	θ	γ	Misspecification and cointegration tests: JB, BG, EM
Austria 1999:01-2003:06	3.49 (0.21)	~0.69 (0.06)	~-0.23 (0.10)	~0.48 (0.08)	JB = 0.53; BG = 0.60 EM = -2.31
Belgium (r_1) 1999:01-2003:09	0.92 (0.05)	0.97 (0.01)	-0.99 (0.18)	**1.27** **(0.07)**	JB = 1.69, BG = 1.12 EM = -5.91***
Belgium (r_2) 1999:01-2003:09	3.60 (0.99)	**1.15** **(0.27)**	**-0.10** **(0.06)**	0.60 (0.23)	JB = 13.96***, BG = 3.95** EM = -1.73
France 1999:01-2003:08	2.23 (0.22)	0.75 (0.06)	**-0.50** **(0.20)**	0.80 (0.22)	JB = 12.37***, BG = 0.80 EM = -2.55
Germany 1999:01-2003:06	5.96 (0.90)	**0.76** **(0.28)**	**-0.05** **(0.02)**	0.26 (0.04)	JB = 0.71, BG = 1.98 EM = -2.62
Ireland 1999:01-2003:09	6.36 (0.20)	**0.74** **(0.06)**	-0.22 (0.06)	0.30 (0.07)	JB = 1.08, BG = 1.19 EM = -3.19*
Italy (r_1) 1999:01-2003:09	3.24 (0.09)	0.75 (0.02)	**-0.39** **(0.03)**	0.27 (0.05)	JB = 2.49, BG = 0.30 EM = -11.25***
Italy (r_2) 1999:01-2003:09	0.31 (0.04)	**0.91** **(0.01)**	-0.44 (0.03)	0.25 (0.04)	JB =42.98***, BG = 3.24** EM = -9.14***
Netherlands 1999:01-2003:09	0.61 (0.05)	1.00 (0.01)	-1.02 (0.14)	0.81 (0.13)	JB = 12.01***, BG = 2.45* EM = -7.03***
Portugal (r_1) 1999:01-2002:12	5.05 (0.74)	**0.61** **(0.18)**	**-0.19** **(0.09)**	-0.00 (0.10)	JB = 1.79, BG = 1.29 EM = -2.17
Portugal (r_2) 1999:01-2002:12	2.91 (0.53)	0.68 (0.13)	-0.27 (0.10)	0.32 (0.13)	JB = 0.66, BG = 0.45 EM = -2.59
Spain 1999:01-2003:09	1.57 (0.16)	0.86 (0.04)	-0.89 (0.14)	0.79 (0.13)	JB = 57.88***, BG = 0.02 EM = -6.16***

In bold, estimates that differ more than one standard deviation from the ones in the last break-free period in Table 3. Test statistics: see Table 3.

Table 6 Asymmetric short term business lending rate pass-through

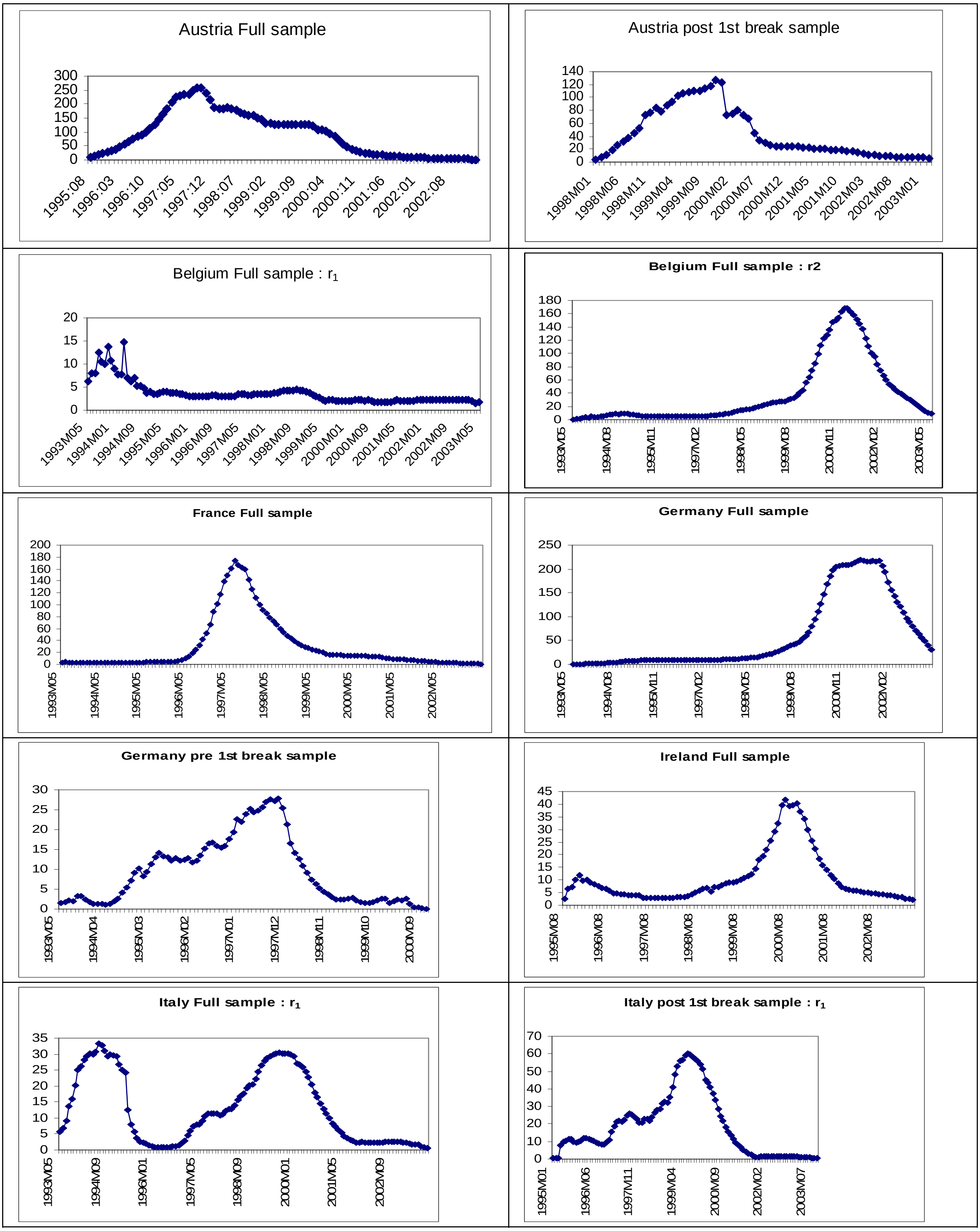
(Country/break-free periods with at least 5% statistically significant γ^+ estimates; one-step ECM; heteroskedasticity consistent standard errors in brackets)

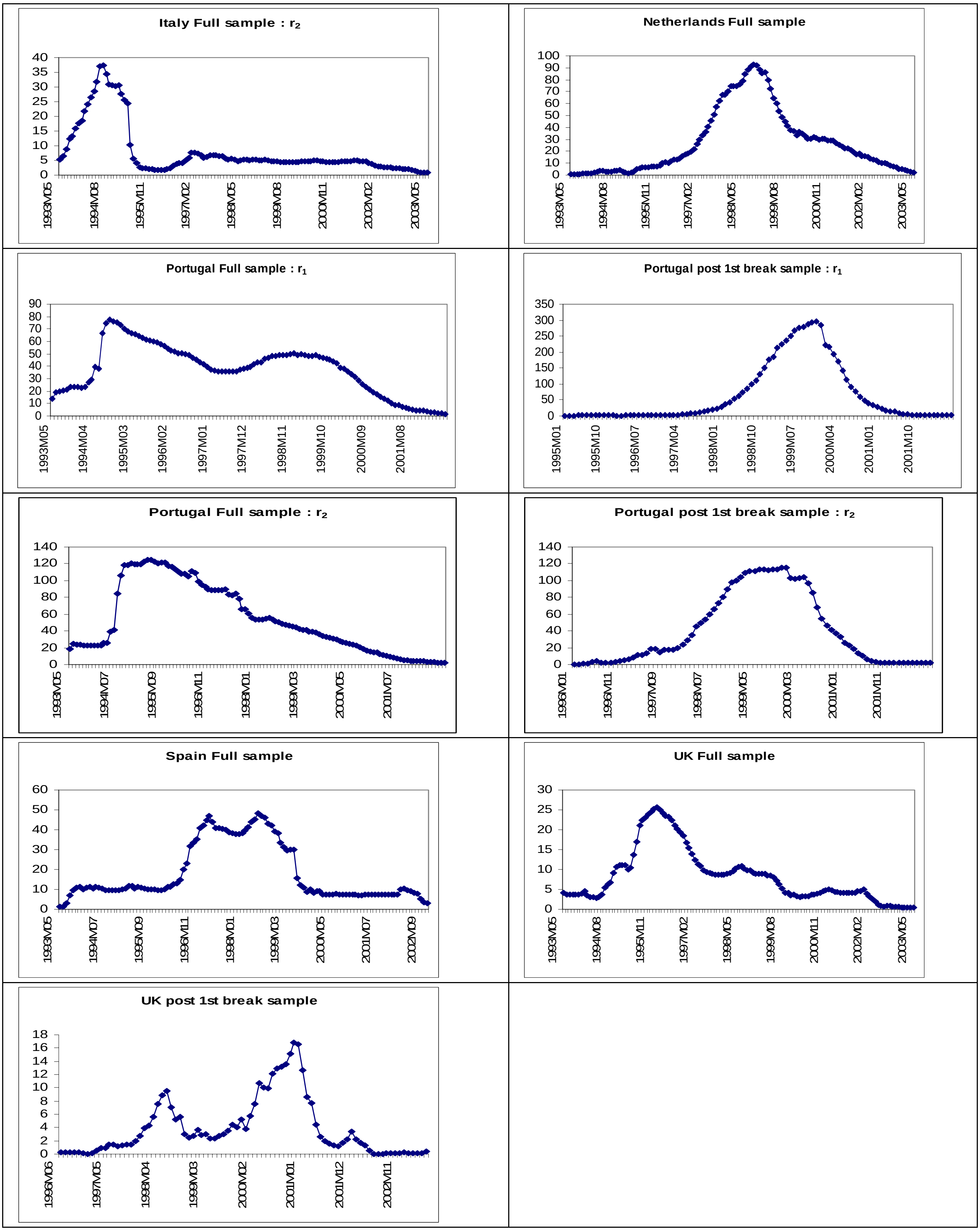
Country sample period	α	β	θ	γ	γ^{+a}	Misspecification and cointegration tests: JB, BG, EM
France 1993:01-1997:06	5.27 (0.61)	0.34 (0.12)	-0.29 (0.14)	-0.16 (0.14)	0.72 (0.20)	JB = 13.73***, BG = 0.04 EM = -2.06
Germany 1997:11-2001:03	3.77 (1.66)	1.10 (0.45)	-0.07 (0.06)	0.00 (0.06)	0.51 (0.14)	JB = 1.19, BG = 1.89 EM = -1.20
Ireland 1995:04-2000:07	6.63 (0.12)	0.60 (0.02)	-0.45 (0.12)	0.38 (0.09)	-0.33 (0.15)	JB = 9.40***, BG = 4.07** EM = -3.80**
Italy (r_2) 1994:08-2003:09	0.19 (0.03)	0.95 (0.01)	-0.41 (0.03)	0.44 (0.05)	-0.24 (0.06)	JB =6.01**, BG = 23.53*** EM = -16.08***
Netherlands 1998:10-2003:09	0.63 (0.04)	1.01 (0.01)	-1.00 (0.12)	1.19 (0.12)	-0.85 (0.24)	JB = 0.57, BG = 0.03 EM = -8.02***
United Kingdom 2000:12-2003:09	0.10 (0.06)	1.00 (0.02)	-0.63 (0.13)	0.90 (0.07)	-1.08 (0.18)	JB =1.48, BG = 1.16 EM = -4.76***

[a]The slope-dummy regressor includes interbank rate positive changes, zero values otherwise. Test statistics: see Table 3.

Appendix: Figure and Tables

Figure A.1 **SupF statistics[a]**




Italy Full sample : r_2
Netherlands Full sample
Portugal Full sample : r_1
Portugal post 1st break sample : r_1
Portugal Full sample : r_2
Portugal post 1st break sample : r_2
Spain Full sample
UK Full sample
UK post 1st break sample

[a]See Table 2.

Table A.1 **KPSS tests for lending rate spreads**
(short term business lending rate net of 1-month interbank rate[a])

Country	Extended sample	Test statistic	EMU sample	Test statistic
Austria	1995:04-2003:06	2.07***	1999:01-2003:06	0.36
Belgium (r_1)	1995:04-2003:09	0.13	1999:01-2003:09	0.07
Belgium (r_2)	1995:04:2003:09	1.26***	1999:01:2003:09	1.26***
France	1995:04-2003:08	0.90***	1999:01-2003:08	0.21
Germany	1995:04-2003:06	0.92***	1999:01-2003:06	0.92***
Ireland	1995:04-2003:09	1.59***	1999:01-2003:09	0.85***
Italy (r_1)	1995:04-2003:09	0.42	1999:01-2003:09	0.37*
Italy (r_2)	1995:04-2003:09	0.92***	1999:01-2003:09	0.63**
Netherlands	1995:04-2003:09	1.75***	1999:01-2003:09	0.44*
Portugal (r_1)	1995:04-2002:12	2.25***	1999:01-2002:12	0.71**
Portugal (r_2)	1995:04-2002:12	2.23***	1999:01-2002:12	0.74**
Spain	1995:04-2003:03	0.66*	1999:01-2003:03	0.24
United Kingdom	1995:04-2003:09	0.20	1999:01-2003:03	0.33

Critical values, adjusted for sample size, for the null of level stationarity are drawn from Sephton (1995, Table 2). ***, **, *: statistically significant at the 1, 5 and 10 per cent level. [a]Except for Belgium (see Table 2).

Table A.2 Unit root tests for short term business lending and market interest rates

Interest rates	**Augmented Dickey Fuller[a]**	
	Level	**First Differences**
Austria 1995:04-2003:06		
r	-1.52	-3.18***
1 month interbank	-1.51	-6.65***
Belgium 1993:01-2003:09		
r_1	-2.39	-10.29***
6 months interbank	-2.02	-7.06***
Belgium 1993:01-2003:09		
r_2	-2.60*	-9.68***
3 months interbank	-2.16	-7.41***
France 1993:01-2003:08		
r	-2.85*	-4.37***
1 month interbank	-4.51***	-8.55***
Germany 1993:01-2003:08		
r	-3.62***	-4.03***
1 month interbank	-3.97***	-3.40***
Ireland 1995:04-2003:09		
r	-1.85	-7.32***
1 month interbank	-1.27	-8.16***
Italy 1993:01-2003:09		
r_1	-1.29	-4.12***
r_2	-1.25	-4.35***
1 month interbank	-1.14	-9.03***
Netherlands 1993:01-2003:09		
r	-4.27***	-3.97***
1 month interbank	-3.67***	-7.02***
Portugal 1993:01-2002:12		
r_1	-3.01**	-3.06***
r_2	-2.08	-4.39***
1 month interbank	-1.95	-2.19**
Spain 1993:01-2003:03		
r	-3.50***	-3.71***
1 month interbank	-3.13**	-4.06***
United Kingdom 1993:01-2003:09		
r	-1.37	-4.32***
1 month interbank	-1.54	-4.57***

[a]ADF tests under the null of unit root with constant (level) and no constant (first differences); lags selected with the Schwartz Information Criterion. ***, **, *: statistically significant at the 1, 5 and 10 per cent level.