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# Asymmetric Adjustment of Commercial Bank Interest Rates in the Euro Area: An Empirical Investigation into Interest Rate Pass-Through

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Shooting at a moving target in the fog is no easy task.
*Dornbusch, Favero and Giavazzi (1998)*
The markets are asymmetric; we are not.
*Alan Greenspan*

## I. Introduction

Since January 1, 1999 the new European Central Bank (ECB) has to conduct a "one-size-fits-all" monetary policy based on her assessment of the average economic conditions of the member countries of the European Monetary Union (EMU). Next to the usual issues and controversies in monetary policy making this implies three new challenges: (1) determining the appropriate average monetary policy in case of diverging economic conditions in the euro area, (2) dealing with possible asymmetric effects of that monetary policy in different member countries, i.e. a divergent monetary transmission mechanism which (3) is most likely subject to dramatic changes (convergence?) as financial market integration and restructuring alongside EMU evolves. While the first challenge has always been at the heart of the controversies about a common currency, the second issue has only recently become an important topic in empirical research. While the latter development is to be welcomed, the third challenge should remind us that judgements about the workings of the monetary mechanism that are based on past data could be misleading in the

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context of a regime change. However, this does not mean that analysing past data is of no use. Our study, therefore, focuses on the latter two challenges by providing new evidence on the financial market side of the monetary mechanism in European Union (EU) countries. In particular, we are examining the pass-through of money market rates to commercial bank lending rates by allowing for the existence of and possible changes in a long-run equilibrium relationship between these two interest rates in the period preceding EMU. We extend on the existing literature by considering a number of different symmetric and asymmetric adjustment mechanisms within and across countries to gain a deeper understanding of the nature and diversity of the financial sector in EU countries and its implication for monetary policy making in the EMU.

In the early 1990, arguably following up on two publications by the Bank for International Settlements (BIS 1994, 1995), a number of studies have investigated asymmetric responses of output to monetary policy innovations across countries which may complicate the implementation of a single monetary policy in the euro area (e.g. Britton and Whitley 1997, Ramaswamy and Sloek 1997, Barran, Coudert and Mojon 1997, Dornbusch, Favero and Giavazzi 1998). While most studies argue the case of asymmetric effects across countries, it is not undisputed that the evidence provided so far is clearly in favour of this hypothesis. For example, Kieler and Saarenheimo (1998) argue that the "econometric evidence does not provide a coherent picture of such differences". They attribute the failure to provide clear econometric evidence to the issues of correctly identifying monetary policy actions[1] and their causal effect on the economy in the "current (or more precisely, historical) set-up". This way the authors provide evidence in favour of no statistically significant differences in monetary transmission for Germany, France and the United Kingdom. However, most empirical papers are based on an estimation period from the 1970s up to date, a time period over which one has to account for changing and differing exchange rate regimes, the 1992/93 EMS crises, and a number of exchange rate re-alignments that all have had an impact on the workings of monetary policy. While it is very clear that these differences will disappear with the adoption of a single currency[2], it is less clear to what extent there will be a convergence in the monetary transmission mechanism itself.

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1 *Kieler* and *Saarenheimo* (1998) use a one percentage point increase in the three month money market rate which is sustained over a period of four years, or, more precisely, they simulate a series of monetary shocks such that they produce a sustained one percentage point increase in the money market rate.

Since Franco Modigliani (1963) the monetary mechanism has been described to consist of two parts: the financial market reaction and the wage-price mechanism. While the studies reviewed above have been examining the impact of monetary policy on the real economy, our study concentrates on the financial market reaction. There are a number of good reasons to do so: First, given the high proportion of bank finance in Europe relative to the USA and UK as shown in Table 1, the "lending channel" is an important element in the monetary mechanism in Europe (e.g. Bernanke and Gertler 1995, Kashyap and Stein 1993). If loans and bonds are imperfect substitutes in the balance sheets of banks and firms, and firms cannot simply access the capital markets but have to rely on bank finance, the transmission of monetary policy impulses is necessarily linked to bank behaviour. Second, if the structure of the financial system matters as a "conveyer" of monetary policy, these structural differences can lead to asymmetries in European banking market reaction and thus monetary policy transmission. In Germany, for example, the close bank-firm relationship tends to weaken the money market rate – lending rate link, while in economies like the British the like is known to be much more direct. Third, while there is evidence that the wage-price process is different across Europe, the Lucas principle suggests that this very process may adopt to the European focus of the ECB's monetary policy (Dornbusch, Favero and Giavazzi 1998). Banking markets, however, may be more resistant to convergence. E.g. Cecchetti (1999) argues that "differences in financial structure are the proximate cause for these national asymmetries in the monetary policy transmission mechanism" and adds that "unless legal structures are harmonised across Europe, financial structures will remain diverse, and so will the monetary transmission mechanism". In a similar vein in a recent ECB working paper Mojon (2000) argues in favour of concentrating his analyses on the pass-through of interest rate innovations to retail banking rates:

> "National segmentation in the European retail banking industry may remain significant in spite of EMU, because retail banking involves heavy investments in brand names, in a network of branches and in relationships with customers (Gual 1999) as well as country-specific legal expertise (Cecchetti 1999). As a consequence, the pass-through from policy-controlled interest rates to bank retail interest rates and the effects of those rates on spending decisions may remain country specific. This potential source of asymmetry across countries is particularly relevant in the euro area where bank rates are a key determinant of the cost of capital and the yield on savings."

[2] The studies by *Dornbusch, Favero* and *Giavazzi* (1998) and *Kieler* and *Saarenheimo* (1998), respectively, explicitly account for intra-EMU exchange rate effects.

*Table 1:* **The Relative Importance of Bank Finance in Europe**

Panel A: Country Specific Characteristics in 1996

| Country | Market Capitalization as a Percentage of GDP | Corporate Debt as a Percentage of GDP | Bank Loans as a Percentage of all Forms of Finance |
|---|---|---|---|
| *EU Member Countries* | | | |
| Austria | 15 | 46 | 65 |
| Belgium | 45 | 60 | 49 |
| Denmark | 41 | 105 | 25 |
| Finland | 50 | 34 | 39 |
| France | 38 | 49 | 49 |
| Germany | 29 | 58 | 55 |
| Greece | 20 | 3 | 48 |
| Ireland | 18 | 13 | 80 |
| Italy | 21 | 37 | 50 |
| Netherlands | 96 | 48 | 53 |
| Portugal | 23 | 19 | 62 |
| Spain | 42 | 11 | 58 |
| Sweden | 99 | 73 | 32 |
| United Kingdom | 150 | 45 | 37 |
| *Other Countries* | | | |
| Japan | 67 | 39 | 59 |
| USA | 111 | 64 | 21 |

Panel B: Euro-Area Characteristics in June 1999[a]

| | Euro Area | USA | Japan |
|---|---|---|---|
| Bank Loans | 100.4 | 48.4 | 107.0 |
| Outstanding domestic debt securities | 88.8 | 164.6 | 126.5 |
| – *issues by corporates* | *3.3* | *29.0* | *14.6* |
| – *issued by financial institutions* | *31.0* | *45.4* | *18.8* |
| – *issued by the public sector* | *54.5* | *90.2* | *93.1* |
| Stock Market Capitalization | 71.1 | 163.3 | 137.7 |

Source: Cecchetti (1999) for Panel A, ECB Monthly Bulletin, January 2000, for Panel B. [a] All data are in percent of GDP and are given for June 1999 except for stock market capitalization which are for October 1999.

Until recently, the literature has often neglected the explicit empirical analysis of the pass-through of monetary policy in the financial sector. Exceptions are e.g. Cottarelli and Kourelis (1994) who focus on the impact of money market rate and policy rate changes on the lending rate, Cottarelli, Ferri, and Generale (1995), BIS (1995) and International Monetary Fund (1996), Sander and Kleimeier (2000), and Toolsema et al. (2000, 2001). Dornbusch et al. (1998) review the pre-1998 literature with respect to the financial market reaction in potential EMU member countries and find that the characteristics of the financial system "go some way towards explaining the observed asymmetries in the transmission mechanism". Our study extends on this literature in four important aspects:

1. The traditional pass-through model that utilizes an autoregressive distributed lag specification is extended by an error correction mechanism, which drives the rates back toward their long-run equilibrium relationship.
2. This long-run equilibrium relationship is analysed in a cointegration approach that tests and allows for structural breaks in order to examine the impact of changing conditions on financial market performance so far.
3. Recent research has shifted toward analysing asymmetric adjustment in interest rates (see Tong 1983, Scholnick 1996 and 1999, Balke and Fomby 1997, Enders and Granger 1998; Baum and Karasulu 1998, Ender and Siklos 2000). We therefore test for cointegration in the presence of asymmetric adjustment of interest rates.
4. After these three steps, we select for all EU countries the best-specified error correction pass-through model and obtain impact multiplier, long-run multiplier, and speed of adjustment coefficients that incorporate the relevant symmetric or asymmetric autoregressive decay.

While our findings largely confirm the results of earlier pass-through studies, such as the lack of convergence in the financial part of the monetary transmission mechanism, we provide these results within a more refined empirical analysis, which allows us to also identify the nature of the adjustment process itself, which again is found to be heterogeneous across European countries.

## II. Data

In order to analyse central bank policy rates, money market rates, and commercial bank lending rates, monthly interest rates have been collected from the CD-ROM version of the IMF's International Financial Statistics (IFS) for all EU member countries from January 1985 to December 1998. As lending rates the rates listed in line 60p of the IFS have been used, the central bank discount rates listed in line 60 have been used as policy rates, and money market rates as listed in line 60b have been used. Exceptions to this sampling procedure were the following: Due to changes in central bank policy, rates from line 60a were used for France as of July 1989 and for the Netherlands as of January 1994. For Luxembourg, the national rates were used for lending rates but not for policy rates. Due to the monetary union between Belgium and Luxembourg, Belgian money market rates are the appropriate policy rates to be used for Luxembourg. If a series was not available on the IFS, the series has been obtained from Datastream. This applies to Austrian, Danish, and Swedish lending rates where Datastream's commercial bank prime lending rates have been used and to French money market rates, where one-month money market rates have been available. British central bank policy rates and Greek money market rates were missing on both, the IFS and Datastream and have thus not be included in our analysis. One should note that in particular lending rates are often heterogeneous across countries. Only recently, also the ECB has started to publish retail lending rates on a regularly base for EMU member countries, however, these data are also coming with the warning that these data are not fully harmonized. In interpreting any estimation results as evidence for heterogeneity across countries one should therefore bear in mind these limitations in the database. A detailed description of the data can be found in table A-1 in the appendix.

For analysing the impact of monetary policy on lending rates, there is an issue of what proxy for monetary policy to choose. Table 2 presents some basic correlation among interest rates in European countries. While in the whole period from 1985 to 1998 we find a varied picture of correlation between money market and policy rates, the sub-period[3] from 1994 to 1998 shows in all cases, with the notable exception of Austria, a correlation close to one. This justifies in particular for the second subperiod to use money market rates as a proxy for the monetary policy

3 The later conducted structural break tests for the cointegration regression leads to the choice of this sub-period.

stance as European central banks increasingly tried to exercise influence over the money market rate. This is in line with Bernanke and Blinder (1992) who favour the US-federal funds rate as a proxy for the monetary policy stance. When investigating whether the money market or the policy rate is more closely related to the lending rate the evidence is mixed. Basing the judgement on correlation coefficients, in the first sub-period one would favour the policy rate, while in the second sub-period the evidence clearly speaks in favour of the money market rate, with the exception of Finland and (to a much lesser degree) Germany and Italy. It therefore appears that a "discount rate addiction" (Cottarelli and Kourelis, 1994), i.e. the announcement effect of a discount rate change that induces banks to change their lending rates, has lost in importance as banks increasingly seem to base their pricing decision on cost of funds considerations. The latter observation is also reflected in the fact that basically in all cases the correlation coefficients between money market and lending rates have increased, with the notable exception of Germany[4]. In conclusion, we will concentrate here on the relationship between money market and lending rates[5].

## III. Analysis of Symmetric and Asymmetric Adjustment of Lending Rates in Europe

### *1. The Pass-Through Approach*

Beginning with Cottarelli and Kourelis (1994), a growing literature is discussing the response of lending rates to monetary policy impulses as an important part of the monetary transmission process. These approaches typically model the transmission process in a dynamic model for the lending rate such as

$$(1) \qquad L_t = \beta_1 + \sum_{i=1}^{k^*} \beta_{L.i} L_{t-i} + \beta_2 M_t + \sum_{i=1}^{n^*} \beta_{M.i} M_{t-i} + \varepsilon_t$$

where $L_t$ and $M_t$ are lending and money market rates, respectively. $k*$ and $n*$ are defined as the model's optimal lag-length. The estimated coef-

4 *Mojon* (2000) reports similar results (with respect to short-term credits to firms) when comparing the sub-periods 1979–1988 and 1988–1998. In his sub-samples all correlation increase except for Germany and Italy. In our sample division it becomes clear that Mojon's result with respect to Italy was basically due to the low correlation in the 1985–1993 segment.

5 The only exception is Greece where the money market rate is not available and is substituted by the policy rate.

*Table 2*

**Correlation Between Interest Rates in Europe 1985 to 1998**

| Correlation / Country | Money Market – Policy Rate 1985–1993 | Money Market – Policy Rate 1994–1998 | Policy – Lending Rate 1985–1993 | Policy – Lending Rate 1994–1998 | Money Market – Lending Rate 1985–1993 | Money Market – Lending Rate 1994–1998 |
|---|---|---|---|---|---|---|
| Austria | 0.95 | 0.97 | –0.39 | 0.26 | –0.22 | 0.17 |
| Belgium | 0.38 | 0.96 | 0.53 | 0.96 | 0.94 | 0.98 |
| Denmark | 0.50 | 0.98 | N.A. | 0.81 | N.A. | 0.80 |
| Finland | 0.73 | 0.94 | 0.80 | 0.99 | 0.63 | 0.94 |
| France | 0.28 | 0.90 | 0.62 | 0.93 | 0.63 | 0.94 |
| Germany | 0.97 | 0.98 | 0.99 | 0.92 | 0.95 | 0.90 |
| Greece | N.A. | N.A. | 0.46 | 0.95 | N.A. | N.A. |
| Ireland | 0.94 | 0.84 | 0.97 | 0.87 | 0.89 | 0.97 |
| Italy | 0.86 | 0.95 | 0.76 | 0.98 | 0.86 | 0.97 |
| Luxemburg | 0.40 | 0.96 | 0.55 | 0.90 | 0.77 | 0.87 |
| Netherlands | 0.98 | 0.99 | 0.98 | 0.98 | 0.98 | 0.97 |
| Portugal | 0.85 | 0.89 | 0.68 | 0.88 | 0.60 | 0.94 |
| Spain | 0.94 | 1.00 | 0.92 | 0.99 | 0.93 | 0.99 |
| Sweden | 0.21 | 0.95 | 0.71 | 0.93 | 0.37 | 0.97 |
| UK | N.A. | N.A. | N.A. | N.A. | 0.98 | 0.85 |

Note: N.A. indicates that one or both rates needed to calculate correlations were not available.

ficient $\beta_2$ is the impact multiplier. A value of less than 1 indicates sluggish adjustment of lending rates to money market rates, also known as lending rate stickiness. The long-term multiplier can be calculated from (1) as

$$\theta = \frac{\beta_2 + \sum_{i=1}^{n*} \beta_{M.i}}{1 - \sum_{i=1}^{k*} \beta_{L.i}} \tag{2}$$

In the long run equation (1) therefore has the form of

$$L_t = \theta_0 + \theta M_t + u_t \tag{3}$$

Cottarelli and Kourelis (1994) argue that this formulation is consistent with the monopolistic competition model relating the lending rate to the money market rate. If $\theta$ is equal to one, we speak of a full pass-through in the long run, whereas equation (1) models the partial adjustment process over time towards the long-run equilibrium in the case of lending rate stickiness, that is $\beta_2 < 1$. Since it is widely accepted that the time series for interest rates typically exhibit an I(1) property, pass-through models like equation (1) are regularly estimated in first differences to avoid spurious regression problems.

Pass-through studies make three major points: First, since impact multipliers are typically below one this is a sign for lending rate stickiness. Reasons given are (1) information asymmetries leading to adverse selection or adverse incentives (moral hazard) à la Stiglitz and Weiss (1981), (2) "menu costs" of changing prices, (3) lack of competition among banks (barriers to entry etc.), (4) lack of competition between banking finance and direct finance (like commercial papers), (5) implicit interest rate insurance through banks, which are interested in maintaining long-term relationship with their customers, and (6) longer maturities. The second point made is that considerable differences in pass-through coefficients prove the case for financial asymmetry across countries, which are held responsible for a differential impact of a single monetary policy in the future.

Focussing on the specific multiplier estimates reveals substantial differences between the point estimates of different pass-through studies. This could be the result of a methodological sensitivity to the choice of sample period and modelling. For example, while Toolsema and de Haan (2000) conduct CUSUM tests to analyse parameter instability and found evidence in favour of stability, Mojon (2000) separates periods of increasing and decreasing interest rates and finds differences in the multipliers. In particular, he shows that the pass-through is faster in periods of increasing rates than in phases of decreasing rates. Moreover, recent pass-through studies are increasingly using an error correction process as part of the pass-through specification (Sander and Kleimeier 2000, Mojon 2000, Toolsema, Sturm, and de Haan 2001). Table 3 summarizes the findings of some recent pass-through studies for six different European countries. For comparative reasons, we indicate in this table already some of the results of this study for the period from 1994 to 1998 that preceded the introduction of the single currency.

## *2. Connecting Short-Term Pass-Through and Long-Run Equilibrium*

Building on the reviewed literature, we propose to base pass-through measurement on a well specified error correction model that explicitly incorporates the long-run relationship between lending and money market rates provided the series are cointegrated as given by equation (4):

$$(4)\qquad \Delta L_t = \beta_1 + \sum_{i=1}^{k*} \beta_{L.i} \Delta L_{t-i} + \beta_2 \Delta M_t + \sum_{i=1}^{n*} \beta_{M.i} \Delta M_{t-i} + \beta_{ECT} ECT_{t-1} + \varepsilon_t$$

where the error correction term *(ECT)* contains the estimated residuals from the long-run equilibrium relationship defined by equation (3), provided such a relationship can be established by cointegration testing procedures.

This formulation has a number of advantages over the standard pass-through model of equation (1). First, the reformulation directly deals with the issue of non-stationary interest rates, a feature generally found in empirical studies and also confirmed here for our data. This is important because only in cases when no cointegration is present or when the underlying time series are stationary, the standard pass-through model is appropriate and ought to be estimated in first differences or levels, respectively. Second, next to the impact multiplier we can directly obtain the speed of adjustment towards the long-run equilibrium via the estimated coefficient of the *ECT* in equation (4). Third, in estimating the long-run multiplier we follow the two-step procedure suggested by Engle and Granger (1987), i.e. we obtain $\theta$ directly from the cointegrating re-

*Table 3*

**Pass-Through Coefficients**

| Country | Study | Impact Multiplier $\beta_2$ | 3-Months Interim Multiplier | Long-Run Multiplier $\theta$ |
|---|---|---|---|---|
| Belgium | present study | 0.77 | | 0.94 |
| | CK | 0.21 | 0.67 | 0.87 |
| | BF | 0.61 | | 1.27 |
| | TSH | 0.75 | 0.76 | 1.02 |
| | M | | 0.64; 0.47; 0.96[1] | |
| France | present study | 0.07 | | 0.51 |
| | CK | – | – | – |
| | BF | 0.43 | | 0.74 |
| | TSH | 0.08 | 0.53 | 0.62 |
| | M | | 0.81; 0.75; 0.86[2] | |
| Germany | present study | 0.19 | | 1.01 |
| | CK | 0.37 | 0.87 | 1.00 |
| | BF | 0.11 | | 1.05 |
| | TSH | 0.33 | 0.72 | 0.90 |
| | M | | 0.67; 0.95; 0.68[1] | |
| Italy | present study | 0.20 | | 0.87 |
| | CK | 0.12 | 0.60 | 0.83 |
| | BF | 0.26 | | 1.22 |
| | TSH | 0.18 | 0.61 | 0.62 |
| | M | | 0.54; 0.55; 0.55[1] | |
| Netherlands | present study | 0.12[4] | | 0.98 |
| | CK | 0.52 | 0.82 | 0.82 |
| | BF | 1.08 | | 1.08 |
| | TSH | 0.85 | 0.84 | 0.97 |
| | M | | 1.03; 0.91; 0.99[1] | |
| Spain | present study | 0.71 | | 1.07 |
| | CK | 0.36 | 0.78 | 0.94 |
| | BF | 0 | | 1.17 |
| | TSH | 0.90 | 1.03 | 1.14 |
| | M | | 0.51; 0.56; 0.65[3] | |

Note: CK refers to Cottarelli and Kourelis (1994), BF refers to Borio and Fritz (1995), TSH refers to Toolsema, Sturm, and de Haan (2001), and M refers to Mojon (2000). [1] Data refer to period: 79–98; 79–88; 88–98, [2] Data refer to period: 84–98, 84–93; 88–98, [3] Data refer to period: 80–98; 80–88; 88–98. [4] Not significant, but a strong error correction mechanism is present, see Table 6.

gression[6]. Fourth, this error correction specification allows us to analyse a variety of asymmetric adjustment mechanisms, thus showing more openly the differences in the financial part of the monetary transmission mechanism. Finally, using models with asymmetries allows us to detect cointegration in cases where there are asymmetries and where other methods would thus fail to detect cointegration.

### a) Cointegration Testing[7]

Before cointegration analysis can proceed, it must *first* be ensured that all interest rate time series have unit roots. As Kleimeier and Sander (2000) have shown, time series of European money market and lending rates exhibit structural breaks in the early 1990s which reflect – among others – the structural changes in the European banking market. Such a structural break has two specific implications for cointegration analysis: First, if a structural break at an a priori unknown point in time is present in a time series, the unit root tests proposed by Engle and Granger (1987) have very little power. Better specified test statistics as proposed by Banerjee, Lumsdaine, and Stock (1992), which are consistent even in the presence of structural breaks, will therefore be employed. Based on these unit root test statistics[8] we are confident that cointegration analysis can proceed for all national series with the exception of Ireland, where lending rates appear to be I(0). Thus, special care should be taken when interpreting the results for Ireland. Consequently, we proceed with cointegration testing.

*Secondly*, we test whether or not the cointegration vector is characterised by a structural break and if so, when this break takes place. This is important since in the presence of a structural break, the standard cointegration tests such as those proposed by Engle and Granger have low power, i.e. the rejection frequency of the ADF test is clearly reduced

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[6] Following *Engle* and *Granger* (1987) and *Wickens* and *Breusch* (1988) one can obtain a *consistent* estimator for the long-run multiplier from the cointegrating regression. Wickens and Breusch, however, suggest alternative formulations of (1) also in the case of non-stationary cointegrated variables in order to estimate both, long-term and short-term multipliers in one step.

[7] The empirical analysis is conducted using the RATS software package. However, all non-standard statistics have been programmed by the authors, in particular the structural break tests, unit root tests, and asymmetric cointegration models.

[8] For a detailed description of the methodology used see *Kleimeier* and *Sander* (2000). The results for the unit root test statistics can be obtained from the authors upon request.

(e. g. Gregory et al., 1996). To test for a structural break in the cointegration relationship described in equation (3), a supremum F (supF) test is calculated. This test was first proposed by Quandt (1960) and has recently been the focus of various studies (e. g. Andrews 1993, Diebold and Chen 1996, Hansen 1992). This test can be seen as a rolling Chow test and is more flexible than the standard Chow test because it allows simultaneously to test for the significance and the timing of a structural break in the cointegration relationship[9]. Based on the critical values reported by Hansen (1992), Table 4 provides evidence for the presence of a structural break in the cointegration relationship in all countries with the exception of Denmark[10]. The breaks occur between January 1987 and July 1993, are similar to those detected by Kleimeier and Sander (2000). Cointegration analysis is particularly vulnerable to small sample sizes, however, as pointed out by Enders and Siklos (2000) we have no clear way of determining whether the gains from estimating over a longer sample period outweigh the losses resulting from applying the model over different underlying economic structures. Since the evidence points so strongly to the presence of structural breaks, we decide in favour of the small sample size. As it is the objective of this study to focus on the transmission mechanism as it was in place right at the introduction of the EMU, the subsequent analysis will only focus on the post-break period. Consequently, a sample period free of breaks and common to all countries will be selected which ranges from January 1994 until December 1998. In more recent empirical work (Kleimeier and Sander, 2002) we also experimented with data covering the first years of the EMU. Doing so and using ECB retail lending rates, we found that the single currency had a strong impact on the structure of the EU banking markets. We therefore decided not to extend the sample beyond 1998 but rather focus in this study on convergence in the pre-EMU phase. Since our results are by and large in line with the evidence provided in the literature (see Table 3) one can conclude that the gains from estimating over a break-free period may outweigh the limitations of the brief estimation horizon.

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9 In particular, a series of standard Chow tests are conducted for a series of different break points b, which move through the mid-70% of the sample. SupF equals the largest Chow F-statistic and is compared to critical values as reported by *Hansen* (1992). The sequence of F-statistics can give an indication about the timing of the break.

10 For Denmark, structural breaks can only be investigated in the time period from February 1994 to December 1998. Thus, the insignificant supF statistic for Denmark indicates that the common period used in subsequent analyses is free of structural breaks.

*Table 4*

**Structural Break Test in the National Cointegration Vectors of Lending and Money Market Rates from 1985 to 1998**

| Country | supF | Break point |
|---|---|---|
| Austria | 328.5 | October 1990 |
| Belgium | 15.9 | December 1990 |
| Denmark | 2.6 | none |
| Finland | 63.6 | March 1990 |
| France | 153.3 | April 1992 |
| Germany | 369.0 | July 1992 |
| Greece | 140.3 | May 1993 |
| Ireland | 76.8 | August 1992 |
| Italy | 20.1 | January 1987 |
| Luxembourg | 56.3 | July 1993 |
| Netherlands | 32.3 | June 1988 |
| Portugal | 56.5 | February 1992 |
| Spain | 45.0 | May 1987 |
| Sweden | 533.0 | September 1992 |
| United Kingdom | 15.9 | September 1992 |

Note: For Greece, policy rates have been used instead of money market rates.

*Thirdly*, we will now proceed with cointegration testing for the period 1994 to 1998. In doing so we will consider six different specifications. Next to the familiar Engle and Granger (1987) symmetric cointegration approach, we estimate five alternative specifications for asymmetric and threshold adjustment towards long-run equilibrium. As Enders and Granger (1998) argue, if adjustment is approximately symmetric the Dickey-Fuller test is more powerful than any other test for cointegration in models allowing for asymmetry. However, with several plausible types of asymmetry the power of the test statistics for the alternative models is superior to the corresponding Dickey-Fuller testing procedure. Therefore we proceed as follows: First, all six models are estimated. If asymmetric cointegration can be established by means of standard tests for coefficient significance and coefficient equality, the model that exhibits the best fit to the data will be selected. If not, the Engle and Granger specification will be given preference.

Starting with the well known Engle and Granger (1987) *symmetric* cointegration model, the usual Dickey-Fuller (DF) and augmented Dickey-Fuller (ADF) tests have to be conducted. The DF tests are based on the estimated residuals of the cointegration regression

$$\Delta u_t = -\delta_0 u_{t-1} + \varepsilon_t \tag{5}$$

where the $t$-statistic for the estimated coefficient $\delta_0$ provides an indication regarding the cointegration of the two series. The ADF test is obtained from the regression

$$\Delta u_t = -\delta_0 u_{t-1} + \sum_{i=1}^{c^*} \delta_i \Delta u_{t-i} + \varepsilon_t \tag{6}$$

where the optimal lag length $c^*$ is found based on the minimum Akaike information criterion (AIC) criteria for lags up to 12. According to Kremers et al. (1992) a residual-based test imposes a common factor restriction which lowers the power of the DF test procedure. In order to avoid rejecting the symmetric cointegration model too often, we therefore perform an additional t-test on $\beta_{ECT}$ of equation (4) as in indicator for cointegration.

However, if interest rates are not adjusting symmetrically to deviations from the long-run equilibrium, the Dickey-Fuller tests may fail to detect cointegration. Fortunately, a growing number of different specifications of asymmetric adjustment models are available. We consider five different specifications for *asymmetric* adjustment of interest rates.

The *first* model we consider is the threshold autoregressive model ($TAR^0$) developed by Tong (1983). The model makes a distinction whether the explained interest rate (lending rate in our case) is above or below its equilibrium level. Thus, the $TAR^0$ allows for asymmetric adjustment depending on the sign of equilibrium-deviation. For example, if the money market rate decreases without an immediate adjustment in the lending rate, we obtain a positive realization of the error term $u_t$. When in this case the autoregressive decay is faster than in the case of money market rate increases, then the lending rate adjustment is faster downward than upward. An appropriate test procedure is to set a Heaviside indicator $I_t$ for different states of $u_{t-1}$.

$$I_t = \begin{cases} 1 & if \quad u_{t-1} \geq 0 \\ 0 & if \quad u_{t-1} < 0 \end{cases} \tag{7}$$

Using this definition, we test for cointegration by estimating equation (8), which represents an modification of the ADF test. The null of no cointegration is rejected if the estimated F-statistic for $H_0 : \rho_1 = \rho_2 = 0$ is significant based on critical values provided by Enders and Siklos (2000).

$$\Delta u_t = I_t \rho_1 u_{t-1} + (1 - I_t) \rho_2 u_{t-1} + \sum_{i=1}^{m^*} \rho_{2+i} \Delta u_{t-i} + \varepsilon_t \tag{8}$$

with the optimal lag length $m*$ determined via the minimum AIC criteria for models with up to 12 lags. When cointegration is established, an F-test for equality of $\rho_1$ and $\rho_2$ indicates the presence of asymmetry.

The *second* model (TAR*) is a modification of the $\text{TAR}^0$ in the sense that the threshold that was formerly implicitly set at zero is now allowed to deviate from that value. The rationale behind such a non-zero threshold is that one or both variables may only adjust to a disequilibrium once it exceeds a certain minimum deviation in one direction. For example, the lending rate will adjust fast only when out of an equilibrium situation the money market rate drops in a way that the deviation from equilibrium exceeds an optimal threshold of, say, 0.5 percentage points. For lower deviations or increases in the money market rate, adjustment takes place at a significantly slower pace. Now the Heaviside indicator in conjunction with equation (8)[11] is defined as

$$I_t = \begin{cases} 1 & if \quad u_{t-1} \geq a_0^* \\ 0 & if \quad u_{t-1} < a_0^* \end{cases} \tag{9}$$

Following Chan's (1993), the optimal threshold $a_0*$ is found by searching over the mid-80% of the distribution of $u_t$ and selecting the model for which the residual sum of squares is minimized. Cointegration and asymmetry testing proceeds with the above described F-tests.

The *third* variation is a Band-TAR model (B-TAR*), which defines the Heaviside indicator as

$$I_1 = \begin{cases} I_1 = 1 & if \quad u_{t-1} \geq a_0^* \quad and \quad 0 \quad otherwise \\ I_2 = 1 & if \quad |u_{t-1}| < a_0^* \quad and \quad 0 \quad otherwise \\ I_3 = 1 & if \quad u_{t-1} < -a_0^* \quad and \quad 0 \quad otherwise \end{cases} \tag{10}$$

11 For both, the TAR* and the following B-TAR* model, the optimal lag length $m^*$ of the $\text{TAR}^0$ specification is used.

while equation (8) has to be modified to

(11) $$\Delta u_t = \sum_{j=1}^{3} \rho_j I_j u_{t-1} + \sum_{i=1}^{m^*} \rho_{3+i} \Delta u_{t-i} + \varepsilon_t$$

Procedures for optimal lag length $m^*$ and optimal threshold $a_0^*$ are corresponding to those of the TAR* and the F-tests for cointegration and asymmetry are applied to all three coefficients $\rho_j$. Such a model has often been applied in particular to model interest rate cointegration where infrequent and discrete adjustments in the rates occur (e.g. Balke and Fomby 1997, Baum and Karasulu 1998). For example, if deviations from equilibrium are small and will therefore not lead to an adjustment of the dependent interest rate, one may find no cointegration within a narrow band bordered by $a_0^*$ and $-a_0^*$ while outside this band cointegration and thus an error correction mechanism may be present. In the context of our study, such behaviour could be related to the "menu cost" argument of lending rate stickiness such that banks only adjust lending rates when deviations are sufficiently large. However, if it happens that inside the band cointegration is found but not outside, this could indicate that banks implicitly insure their customers against excessive deviations from equilibrium by smoothing the response of the lending rate.

Finally, our *fourth* and *fifth* models are so-called momentum threshold autoregressive (M-TAR) models. Whereas in the TAR models the autoregressive decay always depends on the degree of deviation from equilibrium, one could also imagine situations where the adjustment speed depends on how fast the rates move away from or towards equilibrium. Enders and Granger (1998) therefore propose an M-TAR model where the Heaviside indicator depends as follows on the change in error correction term, $\Delta u_t$

(12) $$I_t = \begin{cases} 1 & if \quad \Delta u_{t-1} \geq a_0 \\ 0 & if \quad \Delta u_{t-1} < a_0 \end{cases}$$

Similar to the TAR$^0$ and TAR* specifications, the threshold in the M-TAR can either be set at zero leading to the M-TAR$^0$ specification or be optimised at $a_0^*$ leading to the M-TAR* specification[12]. Cointegration and asymmetry testing proceeds based on equation (8) above. The M-TAR models have successfully been applied to the term structure of interest

12 When optimising the threshold in the M-TAR* model, the optimal lag length $m^*$ from the M-TAR$^0$ is used.

rates by Enders and Granger (1998) and Enders and Siklos (2000). According to the latter authors, M-TAR adjustment can be especially useful when decision makers (in our case banks) are viewed as attempting to smooth out large changes in a series. In sum, the analysis proceeds in seven steps:

(1) Estimation of the cointegration relationship to obtain the estimated residuals $u_t$.

(2) Estimation of all five TAR-type models.

(3) Cointegration test for each TAR-type model.

(4) Asymmetry test for equality of coefficients in case of cointegration for each TAR-type model.

(5) Repetition of steps (2) to (4) with varying lag-length to optimise AIC for each TAR-type.

(6) Selection of optimal TAR-type model based on the minimum AIC across all model specifications.

(7) Symmetric cointegration testing if step (6) can not establish asymmetric cointegration based on the optimal TAR-type model.

The results of the cointegration analysis are summarized in Table 5[13]. Looking first at the Engle-Granger model of symmetric cointegration, the Dickey-Fuller test statistics reveal that out of 15 countries there are five countries where no cointegration could be found and in three countries only marginal evidence for cointegration could be detected. When looking additionally at the t-tests for $\beta_{ECT}$ these statistics indicate symmetric cointegration for two more countries and in two of the three marginal cases cointegration can be confirmed. However, in ten cases we find evidence for asymmetric adjustment. In such cases TAR-type models are better suited to detect cointegration. In particular, there is stronger evidence in favour of cointegration in the cases of Austria, Denmark, Finland, France, Italy and Spain where formerly there was conflicting, marginal, or no evidence for cointegration. In total, for these ten countries we selected an asymmetric cointegration model. Out of these five cases the M-TAR* model was most appropriate whereas the B-TAR* was selected three times and the TAR* twice. It is worth noticing that all asymmetric models selected are indicating that there exist non-zero

[13] Results for the Engle-Granger cointegration regression are given in Table A-2. DW, DF and ADF tests are available from the authors upon request as are the results regarding the estimated coefficients and test statistics of equations (4), (8) and (11).

*Table 5*

**Model Selection**

| Country | AIC values at optimal lag length | | | | | Cointegration based on best TAR | Symmetric cointegration | | Selected pass-through model |
|---|---|---|---|---|---|---|---|---|---|
| | $TAR^0$ | TAR* | B-TAR* | $M\text{-}TAR^0$ | M-TAR* | | DW & DF tests | $t(\beta_{ECT})$ | |
| Austria | 60.96 | **51.13** | 53.14 | 61.37 | 55.34 | (yes), asymmetric | no | no | TAR* |
| Belgium | –30.52 | –32.50 | –35.32 | –30.32 | **–39.20** | yes, asymmetric | (yes) | yes | M-TAR* |
| Denmark | 55.91 | 38.87 | 40.82 | 51.24 | **38.11** | yes, asymmetric | yes | no | M-TAR* |
| Finland | –17.74 | –34.99 | **–37.40** | –18.14 | –20.58 | yes, asymmetric | no | yes | B-TAR* |
| France | 44.84 | **36.58** | 37.36 | 44.58 | 38.29 | yes, asymmetric | yes | no | TAR* |
| Germany | 6.92 | 2.20 | **–1.03** | 4.62 | –0.12 | no | no | yes | SYM |
| Greece | 96.90 | 95.66 | 97.37 | 95.80 | **94.40** | no | no | no | STD |
| Ireland | 12.40 | 10.71 | **8.55** | 12.16 | 10.92 | yes, asymmetric | yes | yes | B-TAR* |
| Italy | 65.10 | 62.83 | 64.58 | 61.43 | **58.48** | yes, asymmetric | (yes) | no | M-TAR* |
| Luxembourg | 30.88 | 30.41 | 30.33 | 30.88 | **23.84** | (yes), asymmetric | (yes) | yes | M-TAR* |
| Netherlands | 33.76 | 30.35 | **22.53** | 32.03 | 26.77 | no | yes | yes | SYM |
| Portugal | 118.02 | **115.39** | 116.60 | 124.08 | 117.94 | no | yes | yes | SYM |
| Spain | 57.22 | 56.10 | **52.93** | 57.46 | 56.93 | (yes), asymmetric | yes | no | B-TAR* |
| Sweden | 55.55 | 50.79 | **48.53** | 58.21 | 51.98 | no | no | no | STD |
| United Kingdom | 24.43 | 23.96 | 22.34 | 20.44 | **15.19** | yes, asymmetric | yes | yes | M-TAR* |

Note: The "best" TAR model is highlighted in bold. A "yes" in parentheses indicates marginal evidence in favour of cointegration.

thresholds, a finding that is in line with the "menu cost" argument. In the cases of Germany, the Netherlands, and Portugal no asymmetric cointegration was found but the adjustment process is basically symmetric in which case the Engle-Granger cointegration model applies. Finally, in the case of Greece and Sweden no cointegration and thus no error correction mechanism could be found.

### b) The Pass-Through of Money Market Rate Innovations in Europe

Based on the cointegration testing we select the appropriate model for analysing the pass-through of interest rates as given in the last column of Table 5 and – for the sake of convenience – again in the second column of Table 6. In the case where no cointegration was found, we use the standard pass-through model (STD). This can be done by estimating the error correction model of equation (4) with $\beta_{ECT}$ set to zero. For this as well as for all other specifications of equation (4), we have chosen an optimal lag length $k*$ and $n*$ for lending and money market rates, respectively, by applying the minimum AIC criteria for all models with up to 12 lags in either interest rate. Consequently, in the STD model the impact multiplier is given by the estimated coefficient $\beta_2$ and the long-run multiplier $\theta$ is calculated according to equation (2).

When cointegration was found, the long-run multiplier $\theta$ is directly obtained from the cointegrating regression reported in Table A-2 of the appendix while again the impact multiplier is $\beta_2$ obtained from the appropriate specification of equation (4). The error correction mechanism itself depends on the optimal model selected in Table 5. In the case of the symmetric cointegration model (SYM), the *ECT* is equal to the estimated residuals of the cointegrating regression. $\beta_{ECT}$ is therefore estimating the speed of a symmetric adjustment process towards a long-run equilibrium. In the models with asymmetric adjustment, $\beta_{ECT}$ and the *ECT*s are 2-dimensional or, in the case of the B-TAR*, 3-dimensional vectors which give the speed of adjustment depending on the definition of the *ECT*s of equations (7), (9), (10), or (12), respectively. Furthermore, where appropriate, the value of the optimal threshold $a_0^*$ is given.

Looking at the results of the pass-through analysis, we first can confirm the findings of earlier pass-through studies that within Europe the stickiness of the lending rate as measured by the impact multiplier varies considerably. It ranges from zero (i.e. an insignificant impact multiplier) in Austria, Denmark, Luxembourg, and the Netherlands up to

*Table 6*

**Pass-Through of Money Market Rate Innovations onto Lending Rates in Europe 1994 to 1998**

| Country | Model | Impact Multiplier $\beta_2$ | Long-run Multiplier $\theta$ | Error Correction Speed of Adjustment | | | Optimal Threshold $a_0*$ | Lags | |
|---|---|---|---|---|---|---|---|---|---|
| | | | | $\beta_{ECT,1}$ | $\beta_{ECT,2}$ | $\beta_{ECT,3}$ | | $k*$ | $n*$ |
| Austria | TAR* | –0.176<br>(–0.724) | 0.223 | –0.280<br>(–3.286) | –0.044<br>(–1.086) | | –1.577 | 8 | 0 |
| Belgium | M-TAR* | 0.771<br>(12.676) | 0.940 | –0.257<br>(–2.960) | –0.043<br>(–0.103) | | –0.233 | 1 | 1 |
| Denmark | M-TAR* | –0.149<br>(–1.869) | 0.499 | 0.023<br>(0.159) | –0.186<br>(–2.319) | | –0.312 | 12 | 12 |
| Finland | B-TAR* | 0.248<br>(8.257) | 1.010 | –0.006<br>(–0.141) | –0.134<br>(–3.551) | –0.168<br>(–5.165) | 0.402 | 5 | 8 |
| France | TAR* | 0.073<br>(1.177) | 0.509 | –0.074<br>(–0.586) | –0.057<br>(–0.497) | | 0.353 | 3 | 0 |
| Germany | SYM | 0.186<br>(1.790) | 1.008 | –0.070<br>(–2.305) | | | | 1 | 0 |
| Greece | STD | 0.502<br>(2.342) | 0.461 | | | | | 1 | 0 |
| Ireland | B-TAR* | 0.668<br>(13.470) | 0.911 | –1.257<br>(–4.624) | –1.167<br>(–2.914) | –1.079<br>(–4.846) | 0.096 | 1 | 1 |
| | STD | 0.723<br>(13.349) | 1.013 | | | | | 6 | 7 |

*(continued p. 182)*

*Table 6* (continued)

| Country | Model | Impact Multiplier $\beta_2$ | Long-run Multiplier $\theta$ | Error Correction Speed of Adjustment $\beta_{ECT,1}$ | $\beta_{ECT,2}$ | $\beta_{ECT,3}$ | Optimal Threshold $a_0*$ | Lags $k*$ | $n*$ |
|---|---|---|---|---|---|---|---|---|---|
| Italy | M-TAR* | 0.197 (7.027) | 0.866 | 0.053 (1.295) | –0.103 (–2.270) | | –0.065 | 11 | 5 |
| Luxembourg | M-TAR* | 0.084 (1.058) | 0.475 | –0.170 (–2.281) | –0.660 (–1.312) | | –0.368 | 1 | 10 |
| | STD | 0.089 (1.114) | 0.083 | | | | | 1 | 0 |
| Netherlands | SYM | 0.115 (0.934) | 0.983 | –0.222 (–2.447) | | | | 5 | 0 |
| Portugal | SYM | 0.168 (2.214) | 1.170 | –0.156 (–2.494) | | | | 2 | 1 |
| Spain | B-TAR* | 0.705 (6.169) | 1.074 | –0.071 (–0.438) | 0.715 (1.393) | 0.064 (0.241) | 0.141 | 1 | 1 |
| | SYM | 0.735 (6.543) | 1.074 | –0.103 (–0.893) | | | | 1 | 0 |
| Sweden | STD | 0.552 (2.589) | 0.698 | | | | | 5 | 1 |
| United Kingdom | M-TAR* | 0.272 (4.923) | 0.636 | –0.176 (–2.709) | –0.439 (–2.538) | | –0.203 | 3 | 4 |

Note: t-statistics are given in parentheses.

levels above 0.7 in Belgium and Spain. It is also interesting to note that our estimates fall by and large in the range indicated in the earlier studies reported in Table 3 with the only exception of the Netherlands. However, in this country we find a reasonably strong symmetric error correction mechanism at work that brings the lending rate back to its equilibrium relationship with the money market rate. The long-run multipliers are close to one (full pass-through) in most but not all cases. In Austria, Denmark, France, Greece, Luxembourg, Sweden, and the UK there is no full pass-through even in the long run.

Looking at the different optimal models it is striking how diverse the workings of the banking sector in Europe still are at the advent of the single currency. For example, in Greece and Sweden no error correction mechanism could be identified. In these countries the impact multipliers appear to be quite high but there is no full pass-through in the long run. On the other hand, we find symmetric adjustment processes in the cases of Germany, the Netherlands, Portugal, and Spain, although in the latter case the speed of adjustment was found insignificant.

A threshold autoregressive adjustment was found in five countries. Austria and France are characterised by a TAR* adjustment process. In Austria there is a strong indication that an error correction process is only invoked if the lending rates are above their equilibrium level by more than the 1.577 percentage points indicated by the optimal threshold. This "shielding" of the customers may have been responsible for not finding cointegration in the Engle-Granger model. By contrast, in France the threshold is lower and the test statistics indicate a close-to-non-existent adjustment process.

The B-TAR* models that have been selected for Finland and Ireland show significant error correction mechanisms. Since the B-TAR* was marginally significant for Spain, we additionally report this model. In Finland it surprisingly appears that even small deviations within the band will lead to adjustments (see $\beta_{ECT,2} = -0.134$) as well as situations where the lending rate is below its equilibrium level (see $\beta_{ECT,3} = -0.168$). Ireland exhibits for all three error correction mechanisms a full or even overshooting adjustment with coefficients close or even below –1. However one should recall that we had problems establishing the I(1) property for Irish interest rates. Therefore, we also report the STD model.

M-TAR* models have been advocated for cases in which increasing or decreasing deviations from equilibrium may induce strong and asymmetric reaction. For Belgium, positive discrepancies from long-run equilibrium resulting from a decrease of the money market rate or an increase of the lending rate such that the $\Delta u_t < -0.233$ are eliminated very quickly (see $\beta_{ECT,1} = -0.257$) whereas other changes display a large amount of persistence (see $\beta_{ECT,2} = -0.043$). A similar pattern can be found for Luxembourg, which could be expected under the monetary union between the two countries. The opposite appears to be the case for Denmark and Italy. In the UK the adjustment mechanism on both sides of the threshold is significant but the speed of adjustment is faster for negative discrepancies.

## IV. Summary and Conclusion

Our study extends the traditional pass-through literature by incorporating an error correction mechanism that is based on cointegration analyses allowing for symmetric as well as for a variety of asymmetric adjustment mechanisms. By and large the results of earlier pass-through studies are confirmed, in particular the finding that monetary policy in euro area is still to be conducted under the conditions of an "asymmetric EMU" of which the differences in the way the different banking systems in euro area countries work are arguably among the most important ones. However, not only is the speed of adjustment different across countries but as we find that on a more fundamental level the nature of the adjustment process itself is heterogeneous. Therefore, our analysis provides a deeper insight into the differential workings of the banking markets across Europe than previous studies. While optimists hope that the elimination of currency risks may contribute to an institutional harmonization within EMU the evidence provided here suggests that for the nearer future asymmetries will continue to influence the monetary mechanism within the euro area. However, with an increasing knowledge of the degree of heterogeneity in European banking markets, shooting at a moving target might become an easier task in the future.

# Appendix

*Table A-1*

**Data Description**

| Country | Primary Data Source | Secondary Data Source | Rate | Type | Maturity |
|---|---|---|---|---|---|
| Central Bank Policy Rates | | | | | |
| Austria | IFS line 60 | Central Bank | discount rate | end of period | |
| Belgium | IFS line 60 | Central Bank | discount rate | end of period | |
| Denmark | IFS line 60 | Central Bank | discount rate | end of period | |
| Finland | IFS line 60 | Central Bank | central bank rate | end of period | |
| France until 7/89 | IFS line 60 | Central Bank | discount rate | end of period | |
| France after 7/89 | IFS line 60a | | repurchase rate | average | |
| Germany | IFS line 60 | Central Bank | discount rate | end of period | |
| Greece | IFS line 60 | Central Bank | discount rate | end of period | |
| Ireland | IFS line 60 | Central Bank | short term facility rate | end of period | |
| Italy | IFS line 60 | Central Bank | discount rate | end of period | |
| Luxembourg | IFS line 60 for Belgium | Central Bank | discount rate | end of period | |
| Netherlands until 1/94 | IFS line 60 | Central Bank | discount rate | end of period | |
| Netherlands after 1/94 | IFS line 60a | | rates on advances | end of period | |

(continued p. 186)

*Table A-1* (continued)

| Country | Primary Data Source | Secondary Data Source | Rate | Type | Maturity |
|---|---|---|---|---|---|
| Central Bank Policy Rates | | | | | |
| Portugal | IFS line 60 | IFS | discount rate | end of period | |
| Spain | IFS line 60 | IFS | central bank rate | end of period | |
| Sweden | IFS line 60 | Statistical offices' publication | discount rate | end of period | |
| Money Market Rates | | | | | |
| Austria | IFS line 60b | | interbank loan rate | average | 1 day |
| Belgium | IFS line 60b | Central bank | interbank loan rate (until 12/90), call money rate (as of 1/91) | average | 3 months (as of Jan 1991) |
| Denmark | IFS line 60b | Central bank | interbank rate | average | 3 months |
| Finland | IFS line 60b | Central bank | interbank (Helibor) rate | average | 3 months |
| France | DS series FFRMM1M | | money market rate | | 1 month |
| Germany | IFS line 60b | Central bank | call money rate | average | 1 day |
| Ireland | IFS line 60b | Central bank | interbank deposit rate | average | 1 month |
| Italy | IFS line 60b | IMF Treasurer's Department | interbank rate | average | 3 months |
| Luxembourg | IFS line 60b | | interbank rate | average | |

| Country | Primary Data Source | Secondary Data Source | Rate | Type | Maturity |
|---|---|---|---|---|---|
| Netherlands | IFS line 60b | | call money rate | average | |
| Portugal | IFS line 60b | Central bank | interbank deposit rate (until 6/88), overnight interbank rate (as of 7/88) | average | up to 5 days (until 6/88), 1 day (as of 7/88) |
| Spain | IFS line 60b | | interbank rate | average | |
| Sweden | IFS line 60b | Central bank | interbank rate | average | 1 day |
| United Kingdom | IFS line 60b | Central bank | interbank overnight deposit rate | average | 1 day |
| Lending Rates | | | | | |
| Austria | DS series AUSCBPL | The Economist | prime lending rate | | |
| Belgium | IFS line 60p | Central bank | prime lending rate | average | short-term (liquidity) loans |
| Denmark | DS series DENCBPL | | prime lending rate | | |
| Finland | IFS line 60p | Central bank | | average | |
| France | IFS line 60p | Central bank | lending rates for best customers | average | short term loans |
| Germany | IFS line 60p | Central bank | current account lending rate | average | |
| Greece | IFS line 60p | Central bank | | average | |

(continued p. 188)

*Table A-1* (continued)

| Country | Primary Data Source | Secondary Data Source | Rate | Type | Maturity |
|---|---|---|---|---|---|
| Lending Rates | | | | | |
| Ireland | IFS line 60p | Central bank | commercial loan rate | average | short term loans |
| Italy | IFS line 60p | Central bank | lending rate | average | short term loans |
| Luxembourg | IFS line 60p | | minimum mortgage rate | average | |
| Netherlands | IFS line 60p | Central bank | rate on current account advances | average | |
| Portugal | IFS line 60p | IFS | commercial loan rate | average | 91 to 180 days |
| Spain | IFS line 60p | Central bank | commercial bill discount rate | average | 3 months |
| Sweden | DS series SWECBPL | | prime lending rate | | |
| United Kingdom | IFS line 60p | | clearing banks' minimum base lending rate | average | |

Note: The primary data source is the database from which the interest rates are obtained. IFS stand for IMF's International Financial Statistics, DS for Datastream. The secondary data source presents the source from which either the IFS or DS have obtained the data.

*Table A-2*

**Cointegration of Lending Rates and Money Market Rates from 1994 to 1998**

| Country | Cointegration Regression (t-statistics) | | | |
|---|---|---|---|---|
| Austria | L = | 8.86 | + 0.22 | M |
| | | (13.41) | (1.32) | |
| Belgium | L = | 3.95 | + 0.94 | M |
| | | (41.57) | (42.35) | |
| Denmark | L = | 4.31 | + 0.50 | M |
| | | (17.37) | (10.09) | |
| Finland | L = | 2.16 | + 1.01 | M |
| | | (10.49) | (21.61) | |
| France | L = | 4.80 | + 0.51 | M |
| | | (42.10) | (21.40) | |
| Germany | L = | 6.14 | + 1.01 | M |
| | | (23.90) | (15.86) | |
| Greece | L = | 1.15 | + 1.17 | M |
| | | (1.15) | (21.58) | |
| Ireland | L = | 1.00 | + 0.91 | M |
| | | (6.07) | (32.30) | |
| Italy | L = | 3.81 | + 0.87 | M |
| | | (15.32) | (28.43) | |
| Luxembourg | L = | 3.90 | + 0.47 | M |
| | | (25.79) | (13.47) | |
| Netherlands | L = | 3.16 | + 0.98 | M |
| | | (24.77) | (29.39) | |
| Portugal | L = | 2.72 | + 1.17 | M |
| | | (6.28) | (21.03) | |
| Spain | L = | 0.36 | + 1.07 | M |
| | | (2.27) | (48.50) | |
| Sweden | L = | 0.59 | + 1.12 | M |
| | | (2.45) | (29.67) | |
| United Kingdom | L = | 2.53 | + 0.64 | M |
| | | (7.91) | (12.18) | |

Note: Results for Greece are based on the policy rate.

## References

*Andrews,* D. W. K. (1993): Tests for parameter instability and structural change with an unknown change point, Econometrica 61, 821–856. – *Banerjee,* A., *Lumsdaine,* R. L. and *Stock,* J. H. (1992): Recursive and sequential test of the unit-root and trend-break hypotheses: Theory and international evidence, Journal of Business & Economic Statistics 10, 271–287. – *Balke,* N. S. and *Fomby,* T. B. (1997): Threshold cointegration, International Economic Review 38 (3), August, 627–645. – Bank for International Settlements (BIS) (1994): National differences in interest rate transmission, CB394, Basle. – Bank for International Settlements (BIS) (1995): Financial structure and the monetary transmission mechanism, Basle. – *Barran,* F., *Coudert,* V. and *Mojon,* B. (1997): The transmission of monetary policy in European countries, European Monetary Policy, Pinter, London. – *Baum,* C. F. and *Karasulu,* M. (1998): Modelling Federal Reserve discount policy, Computational Economics 11, 53–70. – *Bernanke,* B. S. and *Blinder,* A. S. (1992): The Federal Funds rate and the channels of monetary transmission, The American Economic Review 82(4), 901–921. – *Bernanke,* B. S. and *Gertler,* M. (1995): Inside the black box: the credit channel of monetary policy transmission, Journal of Economic Perspectives 9, Fall, 27–48. – *Borio,* C. E. V. and *Fritz,* W. (1995): The response of short-term bank lending rates to policy rates: A cross country perspective, in: BIS (1995). – *Britton,* E. and *Whitley,* J. (1997): Comparing the monetary transmission mechanism in France, Germany, and the UK: Some issues and results, Bank of England 37(2), May. – *Cecchetti,* S. G. (1999): Legal Structure, Financial Structure, and the Monetary Transmission Mechanism, Federal Reserve Bank of New York Economic Policy Review, July. – *Chan,* K. S. (1993): Consistency and limiting distribution of the least squares estimator of a threshold autoregressive model, The Annals of Statistics 21, 520–533. – *Cottarelli,* C., *Ferri,* G. and *Generale,* A. (1995): Bank lending rates and financial structure in Italy: A case study, IMF Working Paper 95/38. – *Cottarelli,* C. and *Kourelis,* A. (1994): Financial structure, bank lending rates, and the transmission mechanism of monetary policy, IMF Staff Papers 41, No.4. – *Diebold,* F. S. and *Chen,* C. (1996): Testing structural stability with endogenous breakpoint – A size comparison of analytic and bootstrap procedures, Journal of Econometrics 70, 221–241. – *Dornbusch,* R., *Favero,* C. A. and *Giavazzi,* F. (1998): A red letter day?, CEPR Discussion Paper 1804, February 1998. – *Enders,* W. and *Granger,* C. W. J. (1998): Unit root tests and asymmetric adjustment with an example using the term structure of interest rates, Journal of Business and Economic Statistics 16(3), 304–11. – *Enders,* W. and *Siklos,* P. I. (2000): Cointegration and threshold adjustment, working paper. – *Engle,* R. F. and *Granger,* C. W. J. (1987): Co-integration and error correction: Representation, estimation, and testing, Econometrica 55, 251–276. – European Central Bank (2000): The euro area one year after the introduction of the Euro: key characteristics and changes in the financial structure, Monthly Bulletin January. – *Gregory,* A. W., *Nason,* J. M. and D. G. *Watt,* (1996): Testing for structural breaks in cointegrated relationships, Journal of Econometrics 71, 321–341. – *Gual,* J. (1999): Deregulation, integration and market structure in European banking, Journal of the Japanese and International Economies 13, 372–396. – *Hansen,* B. E. (1992): Tests for parameter instability in regressions with I(1) processes, Journal of Business & Economic Statistics 10, 321–341. – International Monetary Fund, (1996): World Economic Outlook, October. – *Kashyap,* A. N. and *Stein,* J. C. (1993): Monetary policy and bank lending, NBER Working Paper 4317, April. – *Kieler,* M. and *Saarenheimo,* T. (1998): Differences in monetary policy transmission? A case not closed, European Commission Directorate-General for Economic and Financial Affairs, Economic Papers 132. – *Kleimeier,* S. and *Sander,* H. (2002): Consumer credit rates in the Eurozone, European Credit Research Institute (ECRI), Brussels, ECRI Research Report No. 2. – *Kleimeier,* S. and *Sander,* H. (2000): Regionalisation versus globalisation in European financial market integration: Evidence from co-integration analyses, Journal of Banking and Finance 24(6), 1005–1043. – *Kremers,* J. J. M., *Ericsson,* N. R. and *Dolado,* J. J. (1992): The power of cointegration tests, Oxford Bulletin of Economics and Statistics 54(3), 325–348. – *Modigliani,* F. (1963): The monetary transmission mechanism and its interaction with real phenomena, Review of Economics and Statistics 45(1), 79–107. – *Mojon,* B. (2000): Financial structure and the interest channel of the ECB monetary policy, ECB working paper 40, November. – *Quandt,* R. (1960): Tests of the hypothesis that a linear regression system obeys two separate regimes, Journal of the American Statistical Association 55, 324–330. – *Ramaswamy,* R. and *Sloek,* T. (1997): The real effects of monetary policy in the European Union: What are the differences?, IMF Working Paper 97/160. – *Sander,* H. and *Kleimeier,* S. (2000): Asymmetric adjustment of commercial bank interest rates in the euro area: Implications for monetary policy, working paper presented at the 3rd Conference of the Swiss Society For Financial Market Research, Zurich, April 7, 2000 and at the "Financial Structure, Bank Behaviour and Monetary Policy in the EMU" conference, Groningen University, October 5–6, 2000, (mimeo). – *Scholnick,* B. (1996): Asymmetric adjustment of commercial bank interest rates: evidence from Malaysia and Singapore, Journal of International Money and Finance 15, 485–496. – *Scholnick,* B. (1999): Interest rate asymmetries in long term loan and deposit markets, Journal of Financial Services Research 16(1), 5–26. – *Stiglitz,* J. and *Weiss,* A. (1981): Credit rationing in markets with imperfect information, American Economic Review 69, 339–345. – *Tong,* H. (1983): Threshold models in non-linear time series analysis, Springer Verlag, New York. – *Toolsema,* L. A. and *de Haan,* J. (2000): Convergence of monetary transmission in EMU. New Evidence, Department of Economics, University of Groningen (mimeo). – *Toolsema,* L. A., *Sturm,* J.-E. and *de Haan,* J. (2001): Convergence of monetary transmission in EMU. New Evidence, CESifo Working Paper No. 465, Munich, April. – *Wickens,* M. R. and *Breusch,* T. S. (1988): Dynamic specification, the long-run and estimation of transformed regression models, The Economic Journal 98, 189–205.

## Summary

### Asymmetric Adjustment of Commercial Bank Interest Rates in the Euro Area: An Empirical Investigation into Interest Rate Pass-Through

Our study extends the traditional pass-through literature by incorporating an error correction mechanism that is based on cointegration analyses allowing for structural breaks and symmetric as well as for a variety of asymmetric adjustment mechanisms. While some results of earlier pass-through studies regarding a symmetric monetary transmission mechanism within the euro area are confirmed, our study provides additional evidence that not only the speed of adjustment differs but that the nature of the adjustment process itself is heterogeneous across countries. Therefore, our analysis provides a deeper insight into the differential workings of the banking markets across Europe. (JEL E43, E52, E58, F36)

## Zusammenfassung

### Asymmetrische Anpassung von Geschäftsbankenzinsen im Eurogebiet: Eine empirische Analyse des Zins-Pass-Through

Unsere Studie erweitert die traditionelle Literatur des Zins-Pass-Through durch die Berücksichtigung von Kointegrationsbeziehungen und der korrespondierenden Fehlerkorrekturmechanismen. Dabei erlaubt unser Ansatz, sowohl Strukturbrüche als auch symmetrische und verschiedene asymmetrische Anpassungsmechanismen zu analysieren. Während einige Ergebnisse früherer Pass-Through-Studien bezüglich des Transmissionsmechanismus innerhalb des Eurogebiets bestätigt werden, zeigt unsere Studie darüber hinaus, daß sich im Ländervergleich nicht nur die Anpassungsgeschwindigkeiten unterscheiden, sondern auch, daß der Charakter der Anpassungsprozesse selbst heterogen ist. Unsere Ergebnisse erlauben daher einen tieferen Einblick in die unterschiedlichen Funktionsweisen der Bankenmärkte Europas.

## Résumé

### Ajustement asymétrique des taux d'intérêt des banques commerciales dans la zone euro: une investigation empirique sue la répercussion des taux d'intérêt

Notre étude porte plus loin la littérature traditionnelle en incorporant un mécanisme de correction d'erreurs basé sur des analyses de cointégration, tenant compte de ruptures structurelles ainsi que de mécanismes d'ajustements symétriques et asymétriques. Alors que certains résultats d'études passées sur un mécanisme de transmission monétaire symétrique au sein de la zone euro sont confirmés, notre étude montre une évidence supplémentaire: non seulement la vitesse d'ajustement diffère, mais aussi la nature du processus d'ajustement lui-même est hétérogène selon les pays. Notre analyse offre donc un aperçu plus profond sur les fonctionnements différentiels des manchés bancaires européens.