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Impact of bank competition on the interest rate pass-through in the euro area

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Abstract in English

This paper analyses the impact of loan market competition on the interest rates applied by euro area banks to loans and deposits during the 1994-2004 period, using a novel measure of competition called the Boone indicator. We find evidence that stronger competition implies significantly lower spreads between bank and market interest rates for most loan market products, in line with expectations. Using an error correction model (ECM) approach to measure the effect of competition on the pass-through of market rates to bank interest rates, we likewise find that banks tend to price their loans more in accordance with the market in countries where competitive pressures are stronger. Further, where loan market competition is stronger, we observe larger bank spreads (implying lower bank interest rates) on current account and time deposits. This would suggest that the competitive pressure is heavier in the loan market than in the deposit markets, so that banks under competition compensate for their reduction in loan market income by lowering their deposit rates. We observe also that bank interest rates in more competitive markets respond more strongly to changes in market interest rates. These findings have important monetary policy implications, as they suggest that measures to enhance competition in the European banking sector will tend to render the monetary policy transmission mechanism more effective.

Key words: Monetary transmission, banks, retail rates, competition, panel data;

JEL code: D4, E50, G21, L10;

Abstract in Dutch

Deze paper analyseert het effect van concurrentie op de markt voor leningen voor het eurogebied. Wij analyseren de periode 1994-2004. Zoals verwacht vinden we voor de meeste leningsproducten dat sterkere concurrentie resulteert in significant kleinere opslagen op de rentes van banken ten opzichte van obligatiemarktrentes. Dit vinden we ook als we een andere benadering kiezen: een fouten-correctiemodel (ECM). Verder vinden we dat meer concurrentie op de markt van leningen leidt tot lagere rentes op deposito's. Dit komt waarschijnlijk doordat banken een zekere vorm van marktmacht hebben op de depositomarkt en hierdoor lagere rentes aan depositohouders kunnen vragen als de opbrengsten op de markt van leningen onder druk staan door concurrentie. Tenslotte observeren we dat rentes van banken op leningen in meer competitieve markten sneller reageren op rentes in de obligatiemarkt. Deze uitkomsten betekenen voor het monetaire beleid dat maatregelen ter versterking van competitie in het Europese bankwezen het monetaire transmissiekanaal effectiever maken.

Steekwoorden: Monetaire transmissie, banken, rente, competitie, panel data

Contents

Summary

In this paper, we investigate the effect of loan market competition on euro area banks' retail pricing behaviour and focus, in particular, on its effect on the adjustment of retail bank interest rates to changes in market interest rates. Given the prominent role of the banking sector in the euro area's financial system, it is of significant importance for the ECB to monitor the degree of competitive behaviour in the euro area banking market. A more competitive banking market is expected to drive down bank loan rates, adding to the welfare of households and enterprises. In addition, in a more competitive market, changes in the ECB's main policy rates supposedly will be more effectively passed through to bank interest rates.

We apply a novel measure of bank competition called the Boone indicator, which is based on the notion that in a competitive market, more efficient companies are likely to gain market shares. Hence, the stronger the impact of efficiency on market shares is, the stronger is competition. Furthermore, by analyzing how this efficiency-market share relationship changes over time, this approach provides a measure which can be employed to assess how changes in competition affect the cost of borrowing for both households and enterprises, and how it affects the pass-through of policy rates into loan and deposit rates.

We test three hypotheses concerning the impact of loan market competition on euro area banks' loan and deposit rates. First, we examine the effect of loan market competition on the level on bank loan and deposit rates; second, using a panel error-correction model (ECM) we estimate the effect of loan market competition on the long-run equilibrium pass-through of bank interest rates to changes in corresponding market interest rates; third, we also test the impact of competition in the loan market on the immediate adjustment of bank interest rates to changes in market interest rates.

Our results suggest that stronger competition implies significantly lower interest rate spreads for most loan market products, as we expected. This result implies that bank interest rates are lower and that the pass-through of market rates is stronger, the heavier competition is. We find evidence of the latter in our error correction model of bank interest rates. Furthermore, when loan market competition is stronger, we observe larger bank spreads (that is, lower bank interest rates) on current account and time deposits. Lower time deposits rates are confirmed by the estimates of the ECM. Apparently, the competitive pressure in the loan market is heavier than in the deposit markets, so that banks under competition compensate for their reduction in loan market income by lowering their deposit rates. Furthermore, in more competitive markets, bank interest rates appear to respond stronger and sometime faster to changes in market interest rates. These findings underline that bank competition has a substantial impact on the monetary policy transmission mechanism. More loan market competition enhances the strength and speed of transmission of monetary policy.

1 Introduction*

This paper discusses the effects of bank competition on bank loan and deposit rate levels as well as on their responses to changes in market rates and, hence, on the monetary policy transmission mechanism. Given the prominent role of the banking sector in the euro area's financial system, it is of significant importance for the ECB to monitor the degree of competitive behaviour in the euro area banking market. A more competitive banking market is expected to drive down bank loan rates, adding to the welfare of households and enterprises. Further, in a more competitive market, changes in the ECB's main policy rates supposedly will be more effectively passed through to bank interest rates.

This study extends the existing empirical evidence, which suggests that the degree of bank competition may have a significant effect on both the level of bank rates and on the pass-through of market rates to bank interest rates. Understanding this pass-through mechanism is crucial for central banks. However, most studies that analyse the relationship between competition and banks' pricing behaviour apply a concentration index such as the Herfindahl-Hirschman index (HHI) as a measure of competition. We question the suitability of such indices as measures to capture competition. Where the traditional interpretation is that concentration erodes competition, concentration and competition may instead increase simultaneously when competition forces consolidation. For example, in a market where inefficient firms are taken over by efficient companies, competition may strengthen, while the market's concentration increases at the same time. In addition, the HHI suffers from a serious weakness in that it does not distinguish between small and large countries. In small countries, the concentration ratio is likely to be higher, precisely because the economy is small.

The main contribution of this paper is that it applies a new measure for competition, called the Boone indicator (see also Boone, 2001; Bikker and Van Leuvensteijn, 2008; Van Leuvensteijn et al., 2007). The basic notion underlying this indicator is that in a competitive market, more efficient companies are likely to gain market shares. Hence, the stronger the impact of efficiency on market shares is, the stronger is competition. Further, by analyzing how this efficiency-market share relationship changes over time, this approach provides a measure which can be employed to assess how changes in competition affect the cost of borrowing for both households and enterprises, and how it affects the pass-through of policy rates into loan and deposit rates.

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Our study contributes also to the pass-through literature in the sense that it applies a newly-constructed data set on bank interest rates for eight euro area countries covering the January 1994 to March 2006 period. We include data for Austria, Belgium, France, Germany, Italy, the Netherlands, Portugal and Spain.[1] Further, we consider four types of loan products (mortgage loans, consumer loans and short and long-term loans to enterprises) and two types of deposits (time deposits and current account deposits). We apply recently developed dynamic panel estimates of the pass-through model. Our approach is closely related to that of Kok Sørensen and Werner (2006), on which it expands by linking the degree of competition directly to the pass-through estimates.

Against this background, we test the following three hypotheses:

1. Are loan interest rates lower, and are deposit interest rates higher, in more competitive loan markets than in less competitive loan markets?
2. Are long-run loan and deposit interest rate responses to corresponding market rates stronger in more competitive loan markets than in less competitive loan markets?
3. Do bank interest rates in more competitive markets adjust faster to changes in market interest rates than in less competitive markets?

This paper uses interest rate data that cover a longer period and that are based on more harmonised principles than those used by previous pass-through studies for the euro area. We find that stronger competition implies significantly lower interest rate spreads for most loan market products, as we expected. Using an error correction model (ECM) approach to measure the effect of competition on the pass-through of market rates to bank interest rates, we likewise find that banks tend to price their loans more in accordance with the market in countries where competitive pressures are stronger. Furthermore, where loan market competition is stronger, we observe larger spreads between bank and market interest rates (that is, lower bank interest rates) on current account and time deposits. Lower time deposit rates in countries with stronger bank competition are confirmed by the ECM estimates. Apparently, the competitive pressure is heavier in the loan market than in the deposit markets, so that banks under competition compensate for their reduction in loan market income by lowering their deposit rates. Furthermore, in more competitive markets, bank interest rates appear to respond more strongly and sometime more rapidly to changes in market interest rates.

The structure of the paper is as follows. Section 2 discusses the literature on both measuring competition and the bank interest rate pass-through. Section 3 describes the Boone indicator of competition and Section 4 the employed interest rate pass-through model of the error-correction

[1] For other euro area countries we had insufficient data to estimate the Boone indicator.

type and the applied panel unit root and cointegration tests. Section 5 presents the various data sets used. The results on the various tests and estimates of the spread model and the error correction model equations are shown in Section 6. Finally, Section 7 summarises and concludes.

2 Literature review

2.1 Measuring competition

Competition in the banking sector has been analysed by, amongst other methods, measuring market power (i.e. a reduction in competitive pressure) and efficiency. A well-known approach to measuring market power is suggested by Bresnahan (1982) and Lau (1982), recently used by Bikker (2003) and Uchida and Tsutsui (2005). They analyse bank behaviour on an aggregate level and estimate the average conjectural variation of banks. A strong conjectural variation implies that a bank is highly aware of its interdependence (via the demand equation) with other banks in terms of output and prices. Under perfect competition, where output price equals marginal costs, the conjectural variation between banks should be zero, whereas a value of one would indicate monopoly.

Panzar and Rosse (1987) propose an approach based on the so-called H-statistic which is the sum of the elasticities of the reduced-form revenues with respect to the input prices. In principle, this H-statistic ranges from $-\infty$ to 1. An H-value equal to or smaller than zero indicates monopoly or perfect collusion, whereas a value between zero and one provides evidence of a range of oligopolistic or monopolistic types of competition. A value of one points to perfect competition. This approach has been applied to all (old) EU countries by Bikker and Haaf (2002) and to 101 countries by Bikker et al. (2006).

A third indicator for market power is the Herfindahl-Hirschman Index, which measures the degree of market concentration. This indicator is often used in the context of the 'Structure Conduct Performance' (SCP) model (see e.g. Berger et al., 2004, and Bos, 2004), which assumes that market structure affects banks' behaviour, which in turn determines their performance.[2] The idea is that banks with larger market shares may have more market power and use that. Moreover, a smaller number of banks make collusion more likely. To test the SCP-hypothesis, performance (profit) is explained by market structure, as measured by the HHI. Many articles test this model jointly with an alternative explanation of performance, namely the efficiency hypothesis, which attributes differences in performance (or profit) to differences in efficiency (e.g. Goldberg and Rai, 1996, and Smirlock, 1985). As has been mentioned above, the Boone indicator can be seen as an elaboration on the assumptions underlying this efficiency hypothesis (EH). This EH test is based on estimating an equation which explains profits from both market structure variables and measures of efficiency. The EH

[2] Bikker and Bos (2005), pages 22 and 23.

assumes that market structure variables do not contribute to profits once efficiency is considered as cause of profit. As Bikker and Bos (2005) show, this EH test suffers from a multicollinearity problem if the EH holds.

Market power may also be related to profits, in the sense that extremely high profits may be indicative of a lack of competition. A traditional measure of profitability is the price-cost margin (PCM), which is the output price minus marginal costs, divided by output price. The PCM is frequently used in the empirical industrial organization literature as an empirical approximation of the theoretical Lerner index.[3] In the literature banks' efficiency is often seen as proxy of competition. The existence of scale and scope economies has in the past been investigated thoroughly. It is often assumed that, under strong competition, unused scale economies would be exploited and, consequently, reduced.[4] Hence, the existence of non-exhausted scale economies is an indication that the potential to reduce costs has not been exhausted and, therefore, can be seen as an indirect indicator of (imperfect) competition (Bikker and Van Leuvensteijn, 2008). The existence of scale efficiency is also important as regards the potential entry of new firms, which is a major determinant of competition. Strong scale effects would place new firms in an unfavourable position.

A whole strand of literature is focused on X-efficiency, which reflects managerial ability to drive down production costs, controlled for output volumes and input price levels. X-efficiency of firm i is defined as the difference in cost levels between that firm and the best practice firms of similar size and input prices (Leibenstein, 1966). Heavy competition is expected to force banks to drive down their X-inefficiency, so that the latter is often used as an indirect measure of competition. An overview of the empirical literature is presented in Bikker (2004) and Bikker and Bos (2005).

2.2 Relationship between competition and monetary transmission

According to the seminal papers by Klein (1971) and Monti (1972) on banks' interest rate setting behaviour, banks can exert a degree of market pricing power in determining loan and deposit rates. The Monti-Klein model demonstrates that interest rates on bank products with smaller demand elasticities are priced less competitively. Hence, both the levels of bank interest rates and their changes over time are expected to depend on the degree of competition. With respect to the level of bank interest rates, Maudos and Fernández de Guevara (2004) show that

[3] The Lerner index derives from the monopolist's profit maximisation condition as price minus marginal cost, divided by price. The monopolist maximises profits when the Lerner index is equal to the inverse price elasticity of market demand. Under perfect competition, the Lerner index is zero (market demand is infinitely elastic), in monopoly it approaches one for positive non-zero marginal cost. The Lerner index can be derived for intermediary cases as well. For a discussion see Church and Ware (2000).

[4] This interpretation would be different in a market numbering only a few banks. It would also be different in a market where many new entries incur unfavourable scale effects during the initial phase of their growth path.

an increase in banks' market power (i.e. a reduction in competitive pressure) results in higher net interest margins.[5] In addition, Corvoisier and Gropp (2002) explain the difference between bank retail interest rates and money market rates by bank's product-specific concentration indices. They find that in concentrated markets, retail lending rates are substantially higher, while deposits rates are lower.

Regarding the effect of competition on the way banks adjust their lending and deposit rates, Hannan and Berger (1991) find that deposit rates are significantly more rigid in concentrated markets. Especially in periods of rising monetary policy rates, banks in more consolidated markets tend not to raise their deposit rates, which may be indicative of (tacit) collusive behaviour among banks. In a cross-country analysis, both Cottarelli and Kourelis (1994) and Borio and Fritz (1995) find a significant effect of constrained competition on the monetary transmission mechanism. Thus, lending rates tend to be stickier when banks operate in a less competitive environment, due to, inter alia, the existence of barriers to entry. This finding was confirmed in an Italian setting by Cottarelli et al. (1995). Reflecting the existence of bank market power and collusive behaviour as well as potential switching costs for bank customers (or other factors affecting demand elasticities), the degree of price stickiness is likely to be asymmetric over the (monetary policy) interest rate cycle.[6] Against this background, Mojon (2001) tests for the impact of banking competition on the transmission process related to euro area bank lending rates, using an index of deregulation, constructed by Gual (1999). He finds that higher competition tends to put pressure on banks to adjust lending rates quicker when money market rates are decreasing. Furthermore, higher competition tends to reduce the ability of banks to increase lending rates (although not significantly), when money market rates are moving up – and vice versa for deposit rates.[7] Similar findings of asymmetric pass-through effects have been found by Scholnick (1996), Heinemann and Schüler (2002), Sander and

[5] Of course, competition is not the only factor determining the level of bank interest rates. Factors such as credit and interest risk, banks' degree of risk aversion, operating costs, and bank efficiency are also likely to impact on bank margins. See, for example, Maudos and Fernández de Guevara (2004).

[6] See, for example, Neuwark and Sharpe (1992) and Mester and Saunders (1985) for empirical evidence of asymmetric interest rate pass-through effects among US banks.[6] In addition to bank competition, switching costs and other interest rate adjustment costs, bank rate rigidity may also be due to credit risk factors. For example, in a situation of credit rationing banks may decide to leave lending rates unchanged and to limit the supply of loans instead; see, for example, Winker (1999). Banks may also choose to provide their borrowers with 'implicit interest rate insurance' by smoothing bank loan rates over the cycle; see Berger and Udell (1992). Finally, sometimes banks give customers an interest rate option for a given period. These banks have to recoup the costs of their options which may reduce the speed of the interest rate pass-through for outstanding clients.

[7] In addition to bank competition, switching costs and other interest rate adjustment costs, bank rate rigidity may also be due to credit risk factors. For example, in a situation of credit rationing banks may decide to leave lending rates unchanged and to limit the supply of loans instead; see, for example, Winker (1999). Banks may also choose to provide their borrowers with 'implicit interest rate insurance' by smoothing bank loan rates over the cycle; see Berger and Udell (1992). Finally, sometimes banks give customers an interest rate option for a given period. These banks have to recoup the costs of their options which may reduce the speed of the interest rate pass-through for outstanding clients.

Kleimeier (2002, 2004) and Gropp et al. (2007).[8] Moreover, De Bondt (2005) argues that stronger competition from other banks and from capital markets has helped to speed up the euro area banks' interest rate adjustment's to changes in market rates.

A number of country-specific studies also provide evidence of sluggish pass-through from market rates into bank rates when competition is weak. For example, Heffernan (1997) finds that British banks' interest rate adjustment is compatible with imperfect competition whereas Weth (2002), by using various proxies for bank market power, provides evidence of sluggish and asymmetric pass-through among German banks. De Graeve et al. (2004) estimate the determinants of the interest rate pass-through on Belgian banks and find that banks with more market power pursue a less competitive pricing policy. In a microeconomic analysis of Spanish banks, Lago-González and Salas-Fumás (2005) provide evidence that a mixture of price adjustment costs and bank market power causes price rigidity and asymmetric pass-through. In a cross-country study, Kok Sørensen and Werner (2006) show that differences in the pass-through process across the euro area countries may to some extent be explained by national differences in bank competition. Finally, in another euro area based study, Gropp et al. (2007) provide evidence that the level of banking competition has a positive impact on the degree of bank interest rate pass-through.

[8] Sander and Kleimeier (2002, 2004) differ from others studies in that they also modelling asymmetries the severity of the interest rate shock (rather than merely its direction). This approach aims to take into account menu cost arguments implying that banks tend to pass on changes in market rates of a minimum size only.

3 The Boone indicator as measure of competition

Boone's indicator assumes that more efficient firms (that is, firms with lower marginal costs) will gain higher market shares or profits, and that this effect will be stronger the heavier competition in that market is. In order to support this intuitive market characteristic, Boone develops a broad set of theoretical models (see Boone, 2000, 2001 and 2004, Boone et al., 2004, and CPB, 2000). We use one of these models to explain the Boone indicator and to examine its properties compared to common measures such as the HHI and the PCM. Following Boone et al. (2004), and replacing 'firms' by 'banks', we consider a banking industry where each bank *i* produces one product *qi* (or portfolio of banking products), which faces a demand curve of the form:

$$p\ (q_i, q_{j\neq i}) = a - b\ q_i - d \sum_{j\neq i} q_j \tag{1}$$

and has constant marginal costs mc_i. This bank maximizes profits $\pi_i = (p_i - mc_i)\ q_i$ by choosing the optimal output level q_i. We assume that $a > mc_i$ and $0 < d \leq b$. The first-order condition for a Cournot-Nash equilibrium can then be written as

$$a - 2\ b\ q_i - d \sum_{i\neq j} q_j - mc_i = 0 \tag{2}$$

Where *N* banks produce positive output levels, we can solve the *N* first-order conditions (2), yielding:

$$q_i\ (c_i) = [(2\ b/d - 1)\ a - (2\ b/d + N - 1)\ mc_i + \textstyle\sum_j mc_j]/[(2\ b + d\ (N - 1))(2\ b/d - 1)] \tag{3}$$

We define profits π_i as variable profits excluding entry costs ε. Hence, a bank enters the banking industry if, and only if, $\pi_i \geq \varepsilon$ in equilibrium. Note that Equation (3) provides a relationship between output and marginal costs. It follows from $\pi_i = (p_i - mc_i)\ q_i$ that profits depend on marginal costs in a quadratic way. Competition in this market increases as the produced (portfolios of) services of the various banks become closer substitutes, that is, as d increases (with d kept below b). Further, competition increases when entry costs ε decline. Boone *et al*. (2004) prove that market shares of more efficient banks (that is, with lower marginal costs mc) increase both under regimes of stronger substitution and amid lower entry costs.

Equation (3) supports the use of the following model for market share, defined as $s_i = q_i / \sum_j q_j$:

$$\ln s_i = \alpha + \beta \ln mc_i \tag{4}$$

The market shares of banks with lower marginal costs are expected to increase, so that β is negative. The stronger competition is, the stronger this effect will be, and the larger, in absolute terms, this (negative) value of β. We refer to β as the *Boone indicator*. For empirical reasons, Equation (4) has been specified in log-linear terms in order to deal with heteroskedasticty. Moreover, this specification implies that β is an elasticity, which facilitates interpretation, particularly across equations.[9] The choice of functional form is not essential, as the log-linear form is just an approximation of the pure linear form.

The theoretical model above can also be used to explain why widely-applied measures such as the HHI and the PCM fail as reliable competition indicators. The standard intuition of the HHI is based on a Cournot model with homogenous banks, where a fall in entry barriers reduces the HHI. However, with banks that differ in efficiency, an increase in competition through a rise in d reallocates output to the more efficient banks that already had higher output levels. Hence, the increase in competition raises the HHI instead of lowering it. The effect of increased competition on the industry's PCM may also be perverse. Generally, heavier competition reduces the PCM of all banks. But since more efficient banks may have a higher PCM (skimming off the part of profits that stems from their efficiency lead), the increase of their market share may raise the industry's average PCM, contrary to common expectations.

We note that the Boone indicator model, like every other model, is a simplification of reality. First, efficient banks may choose to translate lower costs either into higher profits or into lower output prices in order to gain market share. Our approach assumes that the behaviour of banks is between these two extreme cases, so that banks generally pass on at least part of their efficiency gains to their clients. More precisely, we assume that the banks' passing-on behaviour, which drives Equation (4), does not diverge too strongly across the banks. Second, our approach ignores differences in bank product quality and design, as well as the attractiveness of innovations. We assume that banks are forced over time to provide quality levels that are more or less similar. By the same token, we presume that banks have to follow the innovations of their peers. Hence, like many other model-based measures, the Boone indicator approach focuses on one important relationship, affected by competition, thereby disregarding other aspects (see also Bikker and Bos, 2005). Naturally, annual estimates of β are more likely to be impaired by these distortions than the estimates covering the full sample period. Also, compared to direct measures of competition, the Boone indicator may have the disadvantage of being an estimate and thus surrounded by a degree of uncertainty. Of course, other model-based measures, such as Panzar and Rosse's H-statistic, suffer from the same disadvantage. The latter shortcoming affects the annual estimates β_t more strongly than the full-sample period estimate β.

[9] The few existing empirical studies based on the Boone indicator all use a log linear relationship. See, for example, Bikker and Van Leuvensteijn (2007).

As the Boone indicator may be time dependent, reflecting changes in competition over time, we estimate β separately for every year (hence, β_t). An absolute benchmark for the level of β is not available. We only know that more negative betas reflect stronger competition. Comparing the indicator across countries or industries helps to interpret estimation results. For that reason, Boone and Weigand in CPB (2000) and Boone *et al.* (2004) apply the model to different manufacturing industries. Since measurement errors – including unobserved country or industry specific factors – are less likely to vary over time than across industries, the time series interpretation of beta is probably more robust than the cross-sector one (that is, comparison of β for various countries or industries at a specific moment in time). Therefore, Boone focuses mainly on the *change* in β_t over time within a given industry, rather than comparing β between industries.

We improve on Boone's approach in two ways. First, we calculate marginal costs instead of approximating this variable with average costs. We are able to do so by estimating a translog cost function, which is more precise and more closely in line with theory. An important advantage is that these marginal costs allow focussing on segments of the market, such as the loan market, where no direct observations of individual cost items are available. Second, we use market share as our dependent variable instead of profits. The latter is, by definition, the product of market shares and profit margin. We have views with respect to the impact of efficiency on market share and its relation with competition, supported by the theoretical framework above, whereas we have no *a priori* knowledge about the effect of efficiency on the profit margin. Hence, a market share model will be more precise. An even more important advantage of market shares is that they are always positive, whereas the range of profits (or losses) includes negative values. A log-linear specification would exclude negative profits (losses) by definition, so that the estimation results would be distorted by sample bias, because inefficient, loss-making banks would be ignored.

In order to be able to calculate marginal costs, we estimate, for each country, a translog cost function (TCF) using individual bank observations. This function assumes that the technology of an individual bank can be described by a single one multiproduct production function. Under proper conditions, a dual cost function can be derived from such a production function, using output levels and factor prices as arguments. A TCF is a second-order Taylor expansion around the mean of a generic dual cost function with all variables appearing as logarithms. It is a flexible functional form that has proven to be an effective tool in explaining multiproduct bank services. Our TCF has different marginal costs for different types of banks, resulting in the following form:

$$\ln c_{ith} = \alpha_0 + \sum_{h=1,..,(H-1)} \alpha_h d_{ih} + \sum_{t=1,..,(T-1)} \delta_t d_t + \sum_{h=1,..,H} \sum_{j=1,..,K} \beta_{jh} \ln x_{ijt} d_{ih} + \sum_{h=1,..,H} \sum_{j=1,..,K} \sum_{k=1,..,K} \gamma_{jkh} \ln x_{ijt} \ln x_{ikt} d_{ih} + v_{it} \quad (5)$$

where the dependent variable c_{it}^h reflects the production costs of bank i (i = 1, .., N) in year t (t = 1, .., T). The sub-index h (h = 1, .., H) refers to the type category of the bank (commercial, savings or cooperative bank). The variable d_i^h is a dummy variable, which is 1 if bank i is of type h and otherwise zero. Another dummy variable is d_t,, which is 1 in year t and otherwise zero. The explanatory variables x_{ikt} represent three groups of variables (k = 1, .., K.). The first group consists of (K_1) bank output components, such as loans, securities and other services (proxied by other income). The second group consists of (K_2) input prices, such as wage rates, deposit rates (as price of funding) and the price of other expenses (proxied as the ratio of other expenses to fixed assets). The third group consists of (K-K_1-K_2) control variables (also called 'netputs'), *e.g*. the equity ratio. In line with Berger and Mester (1997), the equity ratio corrects for differences in loan portfolio risk across banks. The coefficients α_h, β_{jh} and γ_{jkh}, all vary with h, the bank type. The parameters δ_t are the coefficients of the time dummies and v_{it} is the error term.

Two standard properties of cost functions are linear homogeneity in the input prices and cost-exhaustion (see *e.g*. Beattie and Taylor, 1985, and Jorgenson, 1986). They impose the following restrictions on the parameters, assuming – without loss of generality – that the indices j and k of the two sum terms in Equation (5) are equal to 1, 2 or 3, respectively, for wages, funding rates and prices of other expenses:

$$\beta_1 + \beta_2 + \beta_3 = 1,\ \gamma_{1,k} + \gamma_{2,k} + \gamma_{3,k} = 0 \text{ for k = 1, 2, 3, and } \gamma_{k,1} + \gamma_{k,2} + \gamma_{k,3} = 0 \text{ for } k = 4,.., K \qquad (6)$$

The first restriction stems from cost exhaustion, reflecting the fact that the sum of cost shares is equal to unity. In other words, the value of the three inputs is equal to total costs. Linear homogeneity in the input prices requires that the three linear input price elasticities (β_i) add up to 1, whereas the squared and cross terms of all explanatory variables ($\gamma_{i,j}$) add up to zero. Again without loss of generality, we also apply symmetry restrictions $\gamma_{j,k} = \gamma_{k,j}$ for j, k = 1, .., K.[10] As Equation (5) expresses that we assume different cost functions for each type of banks, the restrictions (6) likewise apply to each type of bank.

The marginal costs of output category $j = l$ (of loans) for bank i of category h in year t, mc_{ilt}^h are defined as:

$$mc_{i1t}^h = \partial\, c_{it}^h / \partial\, x_{i1t} = (c_{it}^h./\, x_{i1t})\, \partial\, ln\, c_{it}^h / \partial\, ln\, x_{ilt} \qquad (7)$$

[10] The restrictions are imposed on Equation (5), so that the equation is reformulated in terms of a lower number of parameters.

The term $\partial \ln c_{it}^h / \partial \ln x_{ilt}$ is the first derivative of Equation (5) of costs to loans. We use the marginal costs of the output component 'loans' only (and not for the other K_1 components) as we investigate the loan markets. We estimate a separate translog cost function for each individual sector in each individual country, allowing for differences in the production structure across bank types within a country. This leads to the following equation of the marginal costs for output category loans (l) for bank i in category h during year t:

$$mc_{i1t}^h = c_{it}^h / x_{i1t} \left(\beta_{1h} + 2\,\gamma_{1lh}\,\ln x_{ilt} + \sum_{k=1,..,K;\,k \neq l} \gamma_{1kh} \ln x_{ikt}\right) d_i^h \qquad (8)$$

4 The interest rate pass-through model

Our analysis of the pass-through of market rates to bank interest rates takes into account that economic variables may be non-stationary.[11] The relationship between non-stationary but cointegrated variables should preferably be based on an error-correction model (ECM), which allows disentangling the long-run co-movement of the variables from the short-run adjustment towards the equilibrium. Accordingly, most of the pass-through studies conducted in recent years apply an ECM, as it allows testing for both the long-run equilibrium pass-through of bank rates to changes in market rates and the speed of adjustment towards the equilibrium.[12] Using a panel-econometric approach, we test for the impact of banking competition (measured by the Boone indicator) on the long-run bank interest rate pass-through.

4.1 Estimation of the long-run relationship

If bank interest rates and their corresponding market rates are cointegrated, we may analyse their long-run relationship in an error-correction framework. Hereby, we test for the three hypotheses by estimating the following two equations for each of the six considered interest rates:[13]

$$BR_{i,t} = \alpha BI_{i,t} + \beta_i MR_{i,t} + \gamma BI_{i,t} MR_{i,t} + \delta_i D_i + u_{i,t} \tag{9a}$$

$$\Delta BR_{i,t} = \theta_i u_{i,t-1} + \eta_i \Delta MR_{i,t} + \varphi BI_{i,t} \Delta MR_{i,t} + v_{i,t} \tag{9b}$$

Equation (9.a) reflects the long-run equilibrium pass-through, while Equation (9.b) presents the short-term adjustments of bank interest rates to their long-run equilibrium. $BR_{i,t}$ and $MR_{i,t}$ are the bank interest rate and the corresponding market rate, respectively, in country i (for $i = 1, \ldots, N$) at time t (for $t = 1, \ldots, T$), observed at a quarterly basis. $BI_{i,t}$ is the Boone indicator of country i at time t. For convenience's sake, the Boone indicator is redefined in positive terms, so that an increase in the Boone indicator reflects stronger competition (hence $BI = -\beta$). In all estimations, we include the market interest rates for the different countries separately ($\beta_i\, MR_{i,t}$ and $\eta_i\, \Delta MR_{i,t}$, respectively, in the long and short run), in order to observe country-specific effects, as well as multiplied by the Boone indicator ($\gamma\, BI_{i,t}\, MR_{i,t}$ and $\varphi\, BI_{i,t}\, \Delta MR_{i,t}$, respectively, in the long and short run), in order to capture the (overall) impact of competition on the pass-through. Furthermore, in the long-run model we account for country effects, by using country dummies (D_i). The short-run model includes the error-correction term

[11] In order to avoid spurious results, see Granger and Newbold (1974).

[12] See, for example, Mojon (2001), De Bondt (2002, 2005), Sander and Kleimeier (2004), and Kok Sørensen and Werner (2006).

[13] Namely, four types of loan products (mortgage loans, consumer loans and short and long-term loans to enterprises) and two types of deposits (time deposits and current account deposits).

($\theta_i\, u_{i,t-1}$), the effects of competition on short-term adjustments in market rates ($\varphi\, BI_{i,t}\, \Delta MR_{i,t}$) for all countries simultaneously and the change in the market interest rate for each country separately ($\eta_i\, \Delta MR_{i,t}$).

In Equations (9.a) and (9.b), we estimate European-wide (or panel) parameters for the various competition effects (α, γ and φ), because the Boone indicator varies insufficiently over time to estimate reliable country-specific effects. The other parameters (β_i, η_i and θ_i) remain country-specific, unless restrictions that these parameters are equal across all countries considered would be accepted by a Wald test.

The three hypotheses to be tested are:

1. Are loan interest rates lower, and are deposit interest rates higher, in more competitive loan markets than in less competitive loan markets? H_0: $\alpha + \gamma\, MR_{i,t} < 0$ and H_1: $\alpha + \gamma\, MR_{i,t} \geq 0$;[14] (and H_0: $\alpha + \gamma\, MR_{i,t} > 0$ and H_1: $\alpha + \gamma\, MR_{i,t} \leq 0$, respectively, for deposit rates).
2. Are long-run loan and deposit interest rates responses to the corresponding market rates stronger in more competitive loan markets than in less competitive loan markets?H_0: $\gamma > 0$ and H_1: $\gamma \leq 0$.
3. Do more competitive markets adjust faster, in the short run, to changes in market interest rates than in less competitive markets? H_0: $\varphi > 0$ and H_1: $\varphi \leq 0$.

As we measure competition on the loan market, the competition effects on the deposit-rate pass-through may be less reliable. Loan market competition might have a positive impact on deposit markets also, implying $\alpha_1 + \gamma_1\, MR_{i,t} > 0$. Alternatively, banks may try to compensate for strong loan market competition by exploiting their market power in the deposit market, in which case $\alpha_1 + \gamma_1\, MR_{i,t} < 0$.

4.2 Unit root and panel cointegration tests

Unit root tests

As a first preparatory step, we investigate the unit root properties of the variables.[15] We apply two types of tests based on two different null hypotheses. The Im, Pesaran and Shin (2003) test (henceforth the IPS test) is a panel version of the Augmented Dickey Fuller (ADF) test on unit roots. It is based on the following regression equation:

[14] Note that competition causes a downwards shift to the level of bank interest rates (that is, $\alpha_1 < 0$) as well as a change in the relationship between market rates and bank rates (expressed by $\gamma_1\, MR_{i,t}$).

[15] For a survey of panel unit root tests, see Banerjee (1999). For a more detailed description and application to a similar set of data, see also Kok-Sørensen and Werner (2006).

$$\Delta y_{i,t} = \alpha_i + \rho_i y_{i,t-1} + \sum_{j=1}^{p_j} \beta_{i,j} \Delta y_{i,t-j} + \varepsilon_{i,t} \tag{10}$$

The interest rate series under investigation is $y_{i,t}$ and it must be observable for each country i and each month t. The autoregressive parameter ρ_i is estimated for each country separately, which allows for a large degree of heterogeneity. The null hypothesis is, H_0: $\rho_i = 0$ for all i, against the alternative hypothesis H_1: $\rho_i > 0$ for some countries. The test statistic Z_{t_bar} of the IPS test is constructed by cross-section-averaging the individual t-statistics for ρ_i. Rejection of the null hypothesis indicates stationarity.

As a cross-check, we add results based on Hadri's (2000) test, which is a panel version of the Kwiatkowski, Phillips, Schmidt, and Shin (KPSS) test, testing the null hypothesis of stationarity. The model underlying the Hadri test can be written as:

$$y_{i,t} = \alpha_i + \sum_{\tau=1}^{t} u_{i,\tau} + \varepsilon_{i,t} \tag{11}$$

The time series $y_{i,t}$ are broken down into two components, a random walk component $\Sigma_\tau u_{i,\tau}$ and a stationary component $\varepsilon_{i,t}$. The test statistic Z_τ:is based on the ratio of the variances $\sigma^2_u / \sigma^2_\varepsilon$. The null hypothesis of the test assumes that this ratio is zero, which implies that there is no random walk component. Rejection of this test's null hypothesis indicates the presence of unit root behaviour of the variable under investigation. Both panel series test statistics are asymptotically normal.

Cointegration tests

In a second preliminary step, we test for cointegration using panel cointegration tests by Pedroni (1999, 2004) which are based on the following regression models:

$$y_{i,t} = \alpha_i + \sum_{j=1}^{K} \beta_{j,i} x_{j,i,t} + \varepsilon_{i,t} \tag{12}$$

The long-run coefficients $\beta_{i,j}$ may be different across the euro area countries. We use the group mean panel version of the Pedroni test. The null hypothesis of this test assumes a unit root in the residuals of the cointegration regression, which implies absence of cointegration. The alternative hypothesis assumes a root less than one, but allows for different roots in different countries.[16] We use three different types of test statistics: an ADF type which is similar to the

[16] In the panel versions of the tests the alternative hypothesis assumes a root which is less than one but is identical between the countries. Hence, the group mean versions allow for stronger heterogeneity. As a result, we focus on the test's group mean version.

ADF statistic used in univariate unit-root tests, a nonparametric Phillips-Perron (PP) version, and a version which is based directly on the autoregressive coefficient (ρ-test).

5 The Data

5.1 The Boone indicator

This paper uses the Bankscope database of banks from eight euro area countries during 1992-2004, namely Austria, Belgium, France, Germany, Italy, the Netherlands, Portugal and Spain. Our choice of countries was limited by the availability of (usable) data. For countries such as Finland, Greece and Ireland not enough data are available. Luxembourg is excluded from our sample because its figures presumably do not reflect local market conditions due to the high international profile of its banks. We focus on commercial banks, savings banks, cooperative banks and mortgage banks, ignoring the 25% more specialized institutions such as investment banks, securities firms, long-term credit banks and specialized governmental credit institutions. An exception is made for Germany in order to achieve a more adequate coverage of the national banking systems: specialized German governmental credit institutions, comprising mainly the major Landesbanken, are included. In addition to certain public finance duties, the Landesbanken also offer banking activities in competition with private sector banks, and thus should be included to ensure adequate cover of the competitive environment in the German banking system (see Hackethal, 2004). The appendix provides a detailed description of the data; see also Van Leuvensteijn *et al*. (2007). Table 5.1 presents summary statistics of the estimated Boone indicator.[17] Over the 1994-2004 period we observe that, on average, banking competition is heaviest in Spain, Germany and Italy. Competition appears to be less strong in Belgium, the Netherlands and Austria, and is found to be weakest in France and Portugal. At the same time, Boone indicators for many countries vary considerably over time.[18]

Table 5.1 Summary statistics of the Boone indicator (1994-2004)

	AT	BE	DE	ES	FR	IT	NL	PT
Average	– 1.5	– 2.6	– 4.0	– 4.8	– 0.6	– 4.0	– 2.5	– 0.9
Standard deviation	2.3	0.7	1.5	1.8	0.5	1.8	1.5	1.2
Maximum	4.3	– 1.5	– 2.5	– 2.7	0.3	– 1.6	1.0	1.6
Minimum	– 4.0	– 3.4	– 7.1	– 9.6	– 1.3	– 7.3	– 4.4	– 2.4

[17] The Boone indicator results in this paper may seem different from those in Van Leuvensteijn et al. (2007). However, both working papers use identical estimates of the Boone indicator. The estimates in the appendix of the present paper are exactly equal to the estimates in Table 5.4 in Van Leuvensteijn et al. (2007). However, the presentation of the results differs in two respects from Table 5.3 in Van Leuvensteijn et al. (2007). First, in this paper we present three additional euro-area countries, namely Austria, Belgium and Portugal. Second, in Table 5.3 in Van Leuvensteijn et al. (2007) we compare the average Boone indicator across the European countries by estimating a single parameter for each country over the entire sample period. In this way, we obtain a weighted average of the Boone indicator over the entire period instead of an unweighted average of the annually (time dependent) estimates as in Table 5.1. See the appendix for the yearly estimates of the Boone indicator.

[18] For more details, see Van Leuvensteijn *et al*. (2007).

5.2 Bank interest rates and market rates

Our bank loan interest rates are from the ECB's MFI Interest Rate (MIR) statistics, which since January 2003 have been compiled on a harmonised basis across all euro area countries. Prior to January 2003 the series have been extended backwards to January 1994 using the non-harmonised national retail interest rate (NRIR) statistics compiled by the national central banks of the (later) Eurosystem.[19] The MIR statistics consist of more detailed breakdowns than the NRIR statistics, particularly with respect to the size of loans and the rate fixation periods. In order to link the two sets of statistics, the MIR series have been aggregated (using new business volumes as weights) to the broader product categories of the NRIR statistics, which include rates on mortgage loans, rates on consumer loans, rates on short-term loans to non-financial corporations (≤1 year), rates on long-term loans to non-financial corporations (>1 year), rates on current account deposits and rates on time deposits. The data period covers 147 monthly observations ranging from January 1994 to March 2006.

Table 5.2 Availability of bank interest rates and corresponding market rates

	Mortgage loans	Consumer loans	Short-term enterprise loans	Long-term enterprise loans	Current account deposits	Time deposits
AT	April 1995	April 1995	April 1995		April 1995	April 1995
	3M MR	3M MR	3M MR		3M MR	3M MR
BE	Jan. 1994	Jan. 1994	Jan. 1994	Jan. 1994		Jan. 1994
	3M MR	5Y MR	3M MR	5Y MR		3M MR
DE	Jan. 1994	Jan. 1994	Jan. 1994	Nov. 1996		Jan. 1994
	10Y MR	5Y MR	3M MR	5Y MR		3M MR
ES	Jan. 1994	Jan. 1994	Jan. 1994	Jan. 1994	Jan. 1994	Jan. 1994
	3M MR	3M MR	3M MR	3M MR	3M MR	3M MR
FR	Jan. 1994	Jan. 1994	Jan. 1994	Jan. 1994		Jan. 1994
	10Y MR	5Y MR	3M MR	5Y MR		3M MR
IT	Jan. 1995		Jan. 1994	Jan. 1995	Jan. 1994	Feb. 1995
	3M MR		3M MR	3M MR	3M MR	3M MR
NL	Jan. 1994		Jan. 1994		Jan. 1994	Jan. 1994
	10Y MR		3M MR		3M MR	3M MR
PT	Jan. 1994	Jan. 1994	Jan. 1994			Jan. 1994
	3M MR	3M MR	3M MR			3M MR

Sources: ECB and Bloomberg.
Note: Date indicates: 'available since'; '3M MR' is the 3-month money market rate (MR). '5Y MR' is the 5-year government bond yield. '10Y MR' is the 10-year government bond yield, all for the respective country.

We select market rates which correspond to these bank interest rates in terms of the rate fixation period. Hence, a three-month money market rate is selected to correspond with bank rates that are either floating or fixed for short periods (below one year), while longer-term government

[19] For some bank products in some countries, it is not possible (due to insufficient data being available) to extend interest rates series all the way back to 1994. Hence, we use unbalanced samples for some bank products.

bond yields are selected for long-term fixed bank rates.[20] Table 5.2 presents the data availability of bank interest rates in each country and for each product category together with the corresponding market rates. Note that there is strong variation in interest rate fixation periods across both products and countries. For instance, in many of the considered euro area countries the predominant fixation period for mortgages is rather short, proxied by three months. For Germany and France, however, the typical fixation period on consumer loans is quite long, approximated here by five years.

Table 5.3 Summary statistics of the various bank interest rates (1994-2004; in %)

	AT	BE	DE	ES	FR	IT	NL	PT
	Mortgage rates							
Average	5.6	5.9	6.4	6.6	6.1	7.0	5.7	7.6
Standard deviation	1.0	1.2	1.1	2.7	1.5	3.2	1.0	3.5
Maximum	7.9	8.8	9.1	11.5	8.9	13.0	8.0	14.5
Minimum	3.8	3.8	4.5	3.1	3.9	3.7	3.8	3.4
	Consumer lending rates							
Average	6.6	8.1	7.5	10.4	8.8			13.1
Standard deviation	1.1	0.5	1.0	2.8	1.7			3.6
Maximum	9.5	9.1	10.2	16.2	12.1			19.6
Minimum	5.0	7.3	6.3	7.1	6.2			8.6
	Rates on short-term loans to enterprises							
Average	4.8	4.6	4.0	5.9	4.5	6.7	4.2	8.8
Standard deviation	1.0	1.1	0.7	2.2	1.5	2.8	1.0	3.8
Maximum	7.2	7.6	5.8	10.5	7.8	11.7	6.5	16.8
Minimum	2.9	2.9	3.1	3.2	2.6	3.3	2.8	4.4
	Rates on long-term loans to enterprises							
Average		5.1	5.2	5.7	5.9	6.3		
Standard deviation		1.1	0.5	2.4	1.4	2.7		
Maximum		8.2	6.1	10.4	8.8	11.8		
Minimum		3.4	4.2	3.0	4.0	3.1		
	Current account deposit rates							
Average	1.3			1.8		2.6	1.7	
Standard deviation	0.2			1.2		1.8	0.3	
Maximum	1.7			4.6		5.7	2.0	
Minimum	1.0			0.5		0.7	1.1	
	Time deposit rates							
Average	3.5	3.4	4.4	3.8	4.0	3.3	4.1	3.4
Standard deviation	1.0	0.9	2.1	1.3	2.3	0.9	2.2	0.8
Maximum	6.3	5.4	8.9	8.0	9.1	5.4	8.7	5.1
Minimum	1.9	2.0	1.9	2.0	1.6	2.0	1.8	2.0

[20] The market rates have been chosen to best match bank interest rates on the basis of information from the Methodological Notes for the NRIR statistics and from the volume weights of the MIR statistics.

Table 5.3 shows summary statistics of the bank interest rate data. Bank interest rates differ substantially across countries, across products and over time. On average, over the 1994-2004 period, mortgage rates and consumer lending rates were highest (lowest) in Portugal (Austria). Regarding short-term loans to enterprises rates were on average highest (lowest) in Portugal (Germany), whereas regarding long-term loans to enterprises rates were highest (lowest) in Italy (Belgium). On the deposit side, current account deposit rates were lowest (highest) in Austria (Italy), while time deposit rates were lowest (highest) in Italy (Germany). Regarding developments over time, it may be noted that the variation of bank interest rates was highest in the Mediterranean countries reflecting the particular strong decline in the overall level of interest rates in those countries.

Table 5.4 details the market interest rates for the considered countries. We find that Italy has, on average, the highest three-month money market rate and the Netherlands the lowest. The same picture arises for the 5-year government bond yield. The minima for the three-month money market rates and the two government bond yields with, respectively, a 5 and 10 year fixation period are very similar across all countries: these minima where reached after the introduction of the euro in 1999.

Table 5.4 Summary statistics of the various market rates (1994-2004; in %)

	AT	BE	DE	ES	FR	IT	NL	PT
	3-month money market rate							
Average	3.6	3.6	3.6	4.9	3.9	5.4	3.5	5.3
Standard deviation	0.9	1.1	1.0	2.3	1.4	2.8	1.0	2.9
Maximum	5.5	7.0	5.9	9.7	8.1	11.0	5.4	12.7
Minimum	2.0	2.0	2.0	2.0	2.0	2.0	2.0	2.0
	5-year government bond yield							
Average	4.7	4.8	4.5	5.7	4.8	6.1	4.6	5.9
Standard deviation	1.1	1.2	1.0	2.6	1.3	2.9	1.1	2.7
Maximum	7.3	8.0	7.1	12.2	7.9	13.4	7.3	12.2
Minimum	2.8	2.9	2.8	2.7	2.7	2.9	2.8	2.7
	10-year government bond yield							
Average			5.2		5.4		5.3	
Standard deviation			1.0		1.2		1.0	
Maximum			7.6		8.2		7.7	
Minimum			3.6		3.6		3.6	

Table 5.5 presents the spreads between the various bank and market rates. We present the spreads on deposits as a negative number as the market interest rates are higher than the bank lending rates on these products. On average, the spreads are narrow ranging from 0.5% to 2.0%,

with the notable exception of consumer loans where bank interest rates often include very high risk premiums.

Table 5.5 Summary statistics of the various bank-rate spreads (1994-2004; in %)

	AT	BE	DE	ES	FR	IT	NL	PT
	Mortgage rates							
Average	2.1	2.2	1.8	1.6	1.3	1.9	1.1	2.2
Standard deviation	0.6	0.6	0.3	0.5	0.7	0.7	0.2	1.0
Maximum	3.6	3.5	2.4	2.9	3.8	3.7	1.7	4.5
Minimum	0.8	0.3	1.0	0.8	0.1	0.7	0.6	0.5
	Consumer lending rates							
Average	3.2	4.2	3.1	5.5	4.0			7.7
Standard deviation	0.7	0.9	0.8	0.6	0.9			1.3
Maximum	5.1	6.5	5.2	7.2	7.0			10.2
Minimum	2.1	2.6	1.4	4.2	2.3			4.4
	Rates on short-term loans to enterprises							
Average	1.3	1.0	0.5	1.0	0.6	1.3	0.7	3.4
Standard deviation	0.6	0.2	0.6	0.2	0.8	0.5	0.3	1.1
Maximum	2.9	1.5	1.6	2.0	2.8	2.5	1.3	6.7
Minimum	0.4	0.4	– 0.4	0.5	-1.8	– 0.4	– 0.1	1.9
	Rates on long-term loans to enterprises							
Average		0.4	1.1	0.9	1.1	1.3		
Standard deviation		0.4	0.2	0.4	0.7	0.4		
Maximum		1.2	1.8	1.8	2.2	3.3		
Minimum		– 0.3	0.5	0.1	– 0.4	– 0.5		
	Current account deposit rates							
Average	– 2.0			– 2.9		– 2.7	– 1.7	
Standard deviation	0.7			1.2		1.1	0.8	
Maximum	– 1.0			– 1.4		– 1.3	– 0.8	
Minimum	– 3.8			– 5.9		– 6.0	– 3.5	
	Time deposit rates							
Average	– 0.4	– 0.1	– 0.2	– 0.5	– 0.1	– 0.9	– 0.2	– 1.1
Standard deviation	0.4	0.2	0.2	0.3	0.1	0.5	0.4	0.9
Maximum	0.6	0.2	0.2	0.1	0.2	– 0.2	0.6	– 0.1
Minimum	– 1.5	– 0.7	– 0.6	– 1.1	– 0.3	– 2.6	– 1.1	– 4.7

6 Empirical results

Estimates of the Boone indicator for the loan markets in the euro area countries are presented in the appendix. This approach is similar to the procedure applied in Van Leuvensteijn *et al*. (2007). We obtain annual estimates of the Boone indicator. As the regressions in this section are based on monthly data, we calculate 'smoothed' Boone indicator values using moving averages over six months.

6.1 Unit roots and cointegration

Table 6.1 reports the panel unit root tests for the bank and market interest rate series of the considered eight euro area countries simultaneously. The outcomes indicate non-stationarity at the 5% significance level for all the bank and market interest rate series used. The IPS test on the null hypothesis of a unit root cannot be rejected at the 5% significance level for either the bank rates or the market rates, suggesting non-stationary interest rates. While the IPS test indicates stationarity of the Boone indicator, the null hypothesis of non-stationarity cannot be rejected at the 5% significance level for the product of the Boone indicator and the market rates for three of the six categories, namely mortgage loans, consumer loans and time deposits. However, the Hadri-test on the null hypothesis of stationarity is clearly rejected in all cases. Furthermore, we apply the panel unit root tests for the first differences in interest rates to test on second order non-stationarity. The results reject I(2) and, hence, support the conclusion that the interest rate series are integrated of order 1, so that I(1) holds. Given these findings, we proceed to test on cointegration between bank interest rates and the corresponding market rates.

Table 6.1 Panel unit root tests on model variables applied to all countries

	Im, Pesaran and Shin test		Hadri test	
	Z_{t_bar}[a]	p-value	Z_{τ}	p-value
Boone-indicator	– 2.16	0.02	10.67	0.00
	Bank interest rates			
Mortgage loans	0.98	0.84	18.78	0.00
Consumer loans	– 0.89	0.19	16.59	0.00
Short-term loans to enterprises	– 0.68	0.25	18.83	0.00
Long-term loans to enterprises	0.40	0.66	13.10	0.00
Current account deposits	1.64	0.95	13.86	0.00
Time deposits	– 0.72	0.24	16.03	0.00
	Market interest rates [b]			
Mortgage loans	0.04	0.52	17.08	0.00
Consumer loans	0.34	0.64	15.21	0.00
Short-term loans to enterprises	– 0.68	0.25	17.23	0.00
Long-term loans to enterprises	0.94	0.83	13.39	0.00
Current account deposits	0.38	0.65	12.60	0.00
Time deposits	– 1.56	0.06	16.46	0.00
	Boone indicator times market interest rates [a]			
Mortgage loans	– 2.16	0.01	15.76	0.00
Consumer loans	– 1.88	0.03	12.64	0.00
Short-term loans to enterprises	– 1.44	0.08	17.46	0.00
Long-term loans to enterprises	– 1.38	0.08	13.74	0.00
Current account deposits	– 1.60	0.06	12.65	0.00
Time deposits	– 2.46	0.01	15.70	0.00

[a] The test statistics are explained in Section 4.2.
[b] Market rates are approximated according to Table 5.2.

Table 6.2 shows the results for Pedroni's three panel cointegration tests as applied to the long-run models of the six bank rates.[21] For bank interest rates on consumer loans and current account deposits, the null hypothesis of no cointegration cannot be rejected. Apparently, therefore, the adjustment of interest rates on consumer loans and current account deposits to changes in market rates is so sluggish that even a long-run relationship cannot be detected in our sample.[22] Consequently, the results of the error-correction model on consumer loans and current account deposits, presented in Section 6.2 below, have to be interpreted with caution. For the other four long-run bank rate models, the null hypothesis of no cointegration has been

[21] P-values of the various test statistics have been derived using the standard normal distribution, which is a valid assumption for cointegration tests; see Pedroni (1999).
[22] Data on interest rates on consumer loans and current account deposits prior to January 2003 are only available for six and four countries, respectively, which somewhat limits the analysis of these rates.

rejected (for two of the three tests), indicating a long-run equilibrium relationship between bank rates, market rates and the Boone indicator.

Table 6.2 Pedroni cointegration tests on the six long-run bank interest rates models

Bank interest rates	Group mean panel cointegration tests [a]		
	P-statistic	PP-statistic	ADF-statistic
Mortgage loans	– 3.19 (0.00)	– 3.56 (0.00)	– 0.07 (0.53)
Consumers loans	0.73 (0.77)	0.19 (0.57)	0.05 (0.52)
Short term loans to enterprises	– 5.79 (0.00)	– 4.75 (0.00)	– 1.50 (0.07)
Long term loans to enterprises	– 2.68 (0.00)	– 2.91 (0.00)	– 0.75 (0.22)
Current account deposits	1.14 (0.87)	1.29 (0.90)	0.66 (0.75)
Time deposits	– 8.28 (0.00)	– 7.08 (0.00)	– 0.43 (0.33)

[a] P-values in parentheses.

6.2 Competition and the bank interest-rate pass-through

As a first investigation into the impact of competition on the bank interest rate pass-through, we analyse the effect of competition on the various spreads between bank and market interest rates (see Table 6.3). The main finding is that competition tends to keep bank loan rates more closely in line with the corresponding market rates (implying that they are lower). Moreover, the results in Table 6.3 show that competition significantly diminishes the bank rate spreads for three out of four loan products, namely for mortgages, consumer loans and short-term loans to enterprises. No significant effect is found for long-term loans to enterprises. The Boone indicator's elasticities of the first three loan products indicate that mortgage loans are least affected by competition while short-term loans to enterprises are influenced most strongly.

For the two deposit categories, competition in the loan market seems to increase the (negative) spread between bank and market rates. Hence, deposit rates become lower where there is fierce competition in the loan market. This could reflect that the competitive pressure is heavier in the loan market than in the deposit markets, so that banks under competitive pressure compensate for their decline in loan market income by lowering their deposit rates.

Table 6.3 Effect of competition on the spreads between bank and market lending rates

	Mortgage loans		Consumer loans		Short term loans to enterprises	
	Parameter	Z-value[1)]	Parameter	Z-value	Parameter	Z-value
Boone indicator	– 0.030	– 2.12**	– 0.075	– 3.03***	– 0.128	– -6.72***
Constant	1.357	5.54***	5.818	16.91***	.736	3.02***
Country dummies [2)]	$\chi^2(7)=498$		$\chi^2(5)=3095$		$\chi^2(7)=911$	
Monthly dummies[2)]	$\chi^2(119)=693$		$\chi^2(119)=766$		$\chi^2(119)=223$	
R-squared, centred	0.687		0.907		0.793	
Number of observations	957		717		957	
	Long term loans to enterprises		Current account (sight) deposits		Time deposits	
	Parameter	Z-value	Parameter	Z-value	Parameter	Z-value
Boone indicator	0.003	0.15	– 0.154	– 8.26***	– 0.036	– 3.06***
Constant	1.114	4.26***	– 3.496	– 12.30***	– 0.655	– 2.80***
Country dummies	$\chi^2(4)=240$		$\chi^2(3)=141$		$\chi^2(7)=640$	
Monthly dummies	$\chi^2(119)=1084$		$\chi^2(119)=1499$		$\chi^2(119)=389$	
R-squared, centred	0.670		0.832		0.691	
Number of observations	578		477		956	

Two and three asterisks indicate a level of confidence of 95% and 99%, respectively. [1)] The z-value indicates whether the parameter significantly differs from 0 under the normal distribution with mean zero and standard deviation one. [2)] Chi-squared distributed Wald tests on H_0 'all country dummy coefficients are zero' and 'all monthly time dummy coefficients are zero', respectively. The null hypotheses are rejected for all loan and deposit types.

Table 6.4 presents the estimated long-run relationship of the error-correction model (ECM) described in Section 4.1 (Equation (9.a)), in order to test the three hypotheses mentioned in that section. This model explains bank interest rates from the Boone indicator and the market interest rates. We use Newey-West's kernel-based heteroskedastic and autocorrelation consistent (HAC) variance estimations to correct for heteroskedasticity and autocorrelation, where the bandwidth has been set on two periods. We observe that the impact of market rates on bank interest rates is highly significant for all six interest rates considered and in all eight euro area countries. Moreover, in line with the existing literature, we find that the country-specific long-run pass-through coefficients (β_i) differ considerably across product categories (and across countries) for both the long and short term. The adjustment of bank interest rates to changes in market rates is highest for mortgage loans, loans to enterprises and time deposits.[23]

The first hypothesis is: are loan interest rates lower, and are deposit interest rates higher, in more competitive loan markets than in less competitive loan markets? Contrary to the

[23] See also Mojon (2001), De Bondt (2005) and Kok Sørensen and Werner (2006).

estimations of the spreads presented above, the ECM long-run equation does not assume full pass-through of market rates within one month. Table 6.4 shows that the effect of the interaction terms with the Boone indicator of competition and the market rate is (slightly) positive for all four considered loan products.[24] But the Chi-squared distributed Wald tests on H_0: $\alpha + \gamma\, MR_{i,t} = 0$ also shows that the combined effects of $\alpha + \gamma\, MR_{i,t}$ on bank rates are not significant. This outcome does not confirm our earlier finding of significantly lower loan market spreads under competition. Apparently, the simple spread model is a more successful tool to observe the competition effect than the more complicated ECM.[25]

The second hypothesis is: do bank interest rates in more competitive markets show stronger long-run responses to the corresponding market rates compared to less competitive markets? Our results suggest that all four bank loan rates do indeed respond significantly more strongly to market rates when competition is high, as reflected by the significant positive coefficient γ of the product terms of indicator and market rates for all loan categories. We find that competition in the loan market contributes also to a more complete pass-through of interest rates on current accounts.[26] All in all, we observe that, generally, competition does make for stronger long-run bank rate responses to corresponding market rates.

The third hypothesis is: do more competitive markets adjust faster in the short run to changes in market interest rates than in less competitive markets? To test this hypothesis, we estimate Equation (9.b). The results in Table 6.5 indicate that the immediate responses of banks' interest rates on loans to changes in market rates tend indeed to be higher in more competitive markets (see the coefficient φ of the product terms).[27] However, the effect is not statistically significant. All in all, we find only limited evidence to support the third hypothesis.

[24] When tested, one single EU-wide parameter for market interest rates was rejected in favour of separate country-specific parameters for market interest rates.

[25] We have tested on a single EU-wide parameter for market interest rates in the long-run ECM model. This null hypothesis was rejected for all loan and deposit categories in favour of separate country-specific parameters for market interest rates.

[26] As mentioned in Section 4, the estimated long-run relationship between interest rates on consumer loans and current account deposits and corresponding market rates may be spurious owing to the lack of a statistically significant cointegration relationship.

[27] We have tested on one single EU-wide parameter for market interest rates and for one single EU-wide parameter for residuals in the short-run ECM model. The null hypotheses of a single EU-wide parameter were rejected for most loan and deposit categories in favour of separate country-specific parameters.

Table 6.4 Estimates of the long-run ECM models for the six bank interest rates

	Mortgage loans		Consumer loans		Short-term loans to enterprises	
	Parameter	Z-value	Parameter	Z-value	Parameter	Z-value
Boone indicator (α)	– 0.198	– 3.32***	– 0.196	– 2.39**	– 0.153	– 3.39**
Market interest rate AT	0.843	8.02***	0.824	6.15***	0.937	***8.76***
Market interest rate BE	0.913	2.26***	1.000	5.98***	0.892	23.05***
Market interest rate DE	0.923	14.88***	0.312	2.41**	0.325	6.22***
Market interest rate ES	0.777	10.89***	0.785	7.63***	0.725	10.90***
Market interest rate FR	0.989	12.85***	1.093	13.38***	0.877	13.04***
Market interest rate IT	0.870	16.07***			0.807	16.90***
Market interest rate NL	0.784	18.11***			0.879	20.11***
Market interest rate PT	1.274	24.63***	1.336	23.06***	1.344	37.41***
Market interest rate*Boone ind. (γ)	0.053	4.29***	0.057	3.21***	0.039	3.47***
Constant	1.951	9.74***	5.679	11.21**	2.813	13.62***
R-squared, centred	0.940		0.927		0.952	
Number of observations	957		717		957	
$\alpha + \gamma MR_{i,t}$	0.034		0.055		0.002	
χ^2 H_0: $\alpha + \gamma MR_{i,t} = 0$ 1)	2.92, p-value = 0.09		2.39, p-value =0.12		0.01, p-value = 0.92	

	Long term loans to enterprises		Current account (sight) deposits		Time deposits	
	Parameter	Z-value	Parameter	Z-value	Parameter	Z-value
Boone indicator (α)	– 0.181	– 3.59***	– 0.146	– 5.75***	– .001	– 0.60
Market interest rateAT			0.063	2.28***	0.616	10.17***
Market interest rate BE	0.808	16.79***			0.921	39.45***
Market interest rate DE	0.615	11.48***			0.894	33.03***
Market interest rate ES	0.691	10.89***	0.259	6.75***	0.925	26.99***
Market interest rate FR	0.982	14.42***			0.997	137.37***
Market interest rate IT	0.745	18.84***	0.433	18.09***	0.856	*26.99***
Market interest rate NL			0.083	2.19***	0.831	12.41***
Market interest rate PT					0.798	38.33***
Market interest rate*Boone-ind. (γ)	0.046	4.48***	0.037	5.86***	– 0.015	– 0.60
Constant	2.591	11.58***	1.457	10.43***	0.302	3.15**
R-squared, centred	0.956		0.966		0.972	
Number of observations	578		477		956	
$\alpha + \gamma MR_{i,t}$	0.028		0.005		– 0.024	
χ2 H_0: $\alpha + \gamma MR_{i,t} = 0$ 1)	2.26, p-value=0.13		0.53, p-value=0.47		4.29, p-value =0.04	

Note: One, two and three asterisks indicate levels of confidence of 90%, 95% and 99%, respectively. Country dummies are included but not shown. 1)Chi-squared distributed Wald tests on H0 '$\alpha + \gamma MRi,t = 0$'. The null hypothesis is not rejected for any of the loan and for current account deposits.

Table 6.5 The short-term ECM model of bank interest rates

	Mortgage loans		Consumer loans		Short term loans to enterprises	
	Parameter	Z-value	Parameter	Z-value	Parameter	Z-value
ΔMarket interest rate AT	0.2272	3.15***	0.203	1.84*	0.275	3.41***
ΔMarket interest rate BE	0.207	1.73*	0.358	1.32	0.408	2.49***
ΔMarket interest rate DE	0.511	4.33***	– 0.267	– 1.30	0.159	1.20
ΔMarket interest rate ES	0.217	1.75*	0.041	0.10	0.573	3.36***
ΔMarket interest rate FR	– 0.025	– 0.58	– 0.005	– 0.09	0.079	0.73
ΔMarket interest rate IT	0.156	1.11			0.066	0.42
ΔMarket interest rate NL	0.262	2.79***			0.464	3.01***
ΔMarket interest rate PT	0.173	1.88*	0.001	0.00	0.159	0.87
ΔMarket interest rate*Boone-ind. (φ)	0.020	0.86	0.071	1.52	0.050	1.66*
						***-
Residual AT (-1) [a]	– 0.005	– 3.10***	– 0.004	– 2.89***	– 0.005	3.00***
Residual BE (-1)	– 0.007	– 2.20**	– 0.003	– 1.09	– 0.005	– 1.52
Residual DE (-1)	– 0.003	– 1.56	– 0.003	– 2.07**	– 0.001	– 0.23
Residual ES (-1)	– 0.006	– 2.80***	– 0.003	– 0.86	– 0.000	– 0.03
Residual FR (-1)	– 0.006	– 3.45***	– 0.004	– 3.25***	– 0.003	– 0.44
Residual IT (-1)	– 0.006	– 1.96**			– 0.004	– 1.64*
Residual NL (-1)	– 0.004	– 1.63			– 0.000	– 0.10
Residual PT (-1)	– 0.009	– 3.89***	– 0.006	– 1.50	– 0.011	– 2.28**
R-sq centred	0.19		0.03		0.19	
Number of observations	949		711		949	

	Long term loans to enterprises		Current account (sight) deposits		Time deposits	
	Parameter	Z-value	Parameter	Z-value	Parameter	Z-value
ΔMarket interest rate AT			0.107	3.05***	0.229	2.90***
ΔMarket interest rate BE	0.987	6.97***			0.532	6.02***
ΔMarket interest rate DE	0.657	3.56***			0.587	6.27***
ΔMarket interest rate ES	0.994	3.67***	0.374	3.90***	0.344	2.09**
ΔMarket interest rate FR	0.162	1.47			0.972	38.82***
ΔMarket interest rate IT	0.744	3.34***	0.312	3.68***	0.146	1.28
ΔMarket interest rate NL			0.099	2.45**	0.463	4.95***
ΔMarket interest rate PT					0.281	3.37***
ΔMarket interest rate*Boone-ind. (φ)	0.070	1.41	– 0.033	–2.47**	0.020	0.92
Residual AT (-1)			– 0.004	–2.16**	– 0.004	– 1.69*
Residual BE (-1	0.001	0.31			– 0.004	– 1.58
Residual DE (-1)	– 0.001	– 0.80			– 0.001	– 0.64
Residual ES (-1)	– 0.005	– 1.51	– 0.010	–2.13**	– 0.006	– 2.03**
Residual FR (-1)	– 0.004	– 1.36			0.000	0.24
Residual IT (-1)	– 0.004	– 1.33	– 0.007	–1.41	– 0.009	– 2.33**
Residual NL (-1)			– 0.003	–2.18**	– 0.005	– 1.46
Residual PT (-1)						***-
					– 0.009	3.39***
R-squared centred	0.27		0.18		0.63	
Number of observations	573		473		948	

Note: One, two and three asterisks indicate a level of confidence of, respectively, 90%, 95% and 99%.
[a] See equation (9b).

7 Conclusion

This paper analyses the effects of loan market competition on bank interest rates on loans and deposits, measuring competition by a new approach, called the Boone indicator. Our results show that, in the euro area countries, bank interest rate spreads on mortgage loans, consumer loans and short-term loans to enterprises are significantly lower in more competitive markets. This result implies that bank loan rates tend to be lower under heavier competition, thus improving social welfare. Banks compensate for stronger loan market competition by lowering their deposit rates. Furthermore, evidence is found for all four loan categories that, in the long run, bank loan rates are closer in line with market rates where competition is higher. These results show that stronger loan market competition reduces bank loan rates while changes in market rates are transmitted more rapidly to bank rates. These findings underline that bank competition may have a substantial impact on the monetary policy transmission mechanism.

References

Banarjee, A., 1999, Panel data unit roots and cointegration: an overview, *Oxford Bulletin of Economics and Statistics*, Special Issue, pp. 607-629.

Beattie, B.R. and C.R. Taylor, 1985, *The Economics of Production*, John Wiley & Sons.

Berger, A. and G. Udell, 1992, Some evidence on the empirical significance of credit rationing, *The Journal of Political Economy*, 100, pp. 1047-1077.

Berger, A.N, A. Demirgüç-Kunt, R. Levine and J.G. Haubrich, 2004, Bank concentration and competition: An evolution in the making, *Journal of Money, Credit, and Banking*, 36(3), pp. 433–451.

Berger, A.N., and L.J. Mester, 1997, Inside the black box: What explains differences in the efficiencies of financial institutions?, *Journal of Banking and Finance*, 21(7), 895-947.

Bikker, J.A. and K. Haaf , 2002, Competition, concentration and their relationship: an empirical analysis of the banking industry, *Journal of Banking & Finance*, 26, pp. 2191-2214.

Bikker, J.A., 2003, Testing for imperfect competition on the EU deposit and loan markets with Bresnahan's market power model, *Kredit und Kapital*, 36, pp. 167-212.

Bikker, J.A., 2004, *Competition and efficiency in an unified European Banking market*, Edward Elgar.

Bikker, J.A. and J.W.B. Bos, 2005, Trends in Competition and Profitability in the Banking Industry: A Basic Framework, *SUERF Studies*, 2005/2.

Bikker, J.A., L. Spierdijk, and P. Finnie, 2006, Misspecification in the Panzar-Rosse model: assessing competition in the banking Industry, DNB Working Paper 114, Dutch Central Bank.

Bikker, J.A. and M. van Leuvensteijn, 2008, Competition and efficiency in the Dutch life insurance industry, *Applied Economics*, forthcoming.

Bos, J.W.B., 2004, Does market power affect performance in the Dutch banking market? a comparison of reduced form market structure models, *De Economist*, 152, pp. 491-512.

Boone, J., 2000, Competition, CEPR Discussion Paper Series, 2636, December.

Boone, J., 2001, Intensity of competition and the incentive to innovate, *International Journal of Industrial Organization*, 19, pp. 705-726.

Boone, J., 2004, A New Way to Measure Competition, CEPR Discussion Paper Series, 4330, March.

Boone J., R. Griffith and R. Harrison, 2004, Measuring Competition, presented at the Encore Meeting 2004: Measuring competition.

Borio, C. and W. Fritz, 1995, The response of short-term bank lending rates to policy rates: a cross-country perspective, BIS Working Paper27, May.

Bresnahan, T.F.,1982, 'The oligopoly solution concept is identified', *Economics Letters*, 10, 87-92.

Church, J. and R. Ware, 2000, Industrial Organization: A strategic approach, Irwin McGraw-Hill.

Corvoisier, S. and R. Gropp, 2002, Bank concentration and retail interest rates, *Journal of Banking and Finance*, 26, pp. 2155-2189.

Cottarelli, C. and A. Kourelis, 1994, Financial structure, bank lending rates and the transmission of monetary policy, IMF Staff Paper, 42, pp. 670-700.

Cottarelli, C., G. Ferri, and A. Generale, 1995, Bank lending rates and financial structure in Italy: A case study, IMF Working Paper 38 (April).

CPB, 2000, Measuring competition; how are cost differentials mapped into profit differentials, CPB Working Paper, CPB Netherlands Bureau for Economic Policy Analysis.

De Bondt, G.J., 2002, Retail bank interest rate pass-through: new evidence at the euro area level, ECB Working Paper Series 136.

De Bondt, G.J., 2005, Interest rate pass-through: empirical results for the euro area, *German Economic Review* , 6, pp.37-78.

De Graeve, F., O. De Jonghe, and R. Vander Vennet, 2004, The determinants of pass-through of market conditions to bank retail interest rates in Belgium, mimeo.

Goldberg, L.G. and A. Rai, 1996, The structure-performance relationship for European banking, *Journal of Banking and Finance*, 20, pp.745-771.

Granger, C.W.J. and P. Newbold, 1974, Spurious regressions in econometrics, *Journal of Econometrics*, 2, pp.111-120.

Gropp, R., C. Kok Sørensen and J. Lichtenberger, 2006, The dynamics of bank spreads and financial structure, European Central Bank Working Paper Series 714.

Gual, J., 1999, Deregulation, integration and market structure in European banking, *Journal of the Japanese and International Economies*,13, pp.372-396.

Hackethal, A., 2004, German banks and banking structure, in: J.P. Krahnen and R.H. Schmidt (eds.), *The German Financial System*, Oxford University Press, pp. 71-105.

Hadri, K., 2000, Testing for stationary in heterogeneous panel data, *Econometrics Journal*, 3, pp. 148-161.

Hannan, T., and A. Berger, 1991, The rigidity of prices: Evidence from the banking industry, *American Economic Review*, 81, pp. 938-945.

Hayashi, F., 2000, *Econometrics*, Princeton University Press.

Heffernan, S.A., 1997, Modelling British interest rate adjustment: An error-correction approach, *Economica*, 64, pp. 211-231.

Heinemann, F., and M. Schüler, 2002, Integration benefits on EU retail credit markets – Evidence from interest rate pass-through, ZEW, mimeo.

Hefferman, S.A., 1997, Modelling British Interest Rate Adjustment: An Error Correction Approach, *Economica*, 64 (254), pp. 211-231.

Hughes, J.P. and L.J. Mester, 1997, Bank capitalization and cost: Evidence of scale economies in risk management and signalling', Federal Reserve Bank of Philedelphia Working Paper 96-2/R.

Im, K. S., M. H.Pesaran and Y. Shin, 2003, Testing for unit roots in heterogeneous Panels, *Journal of Econometrics*, 115, pp. 53-74.

Jorgenson, D.W., 1986, Econometric methods for modelling producer behavior, in: *Handbook for Econometrics*, vol. III, eds. E. Grilliches and M. Intrilligator, the Netherlands, Elsevier Science Publishers.

Klein, M., 1971, A theory of the banking firm, *Journal of Money, Credit, and Banking*, 3, pp. 205-218.

Kok Sørensen, C. and T. Werner, 2006, Bank interest rate pass-through in the euro area: a cross-country comparison, ECB Working Paper Series 580.

Lago-González, R. and V. Salas-Fumás, 2005, Market power and bank interest rate adjustments, Banco de España Working Paper 0539.

Lau, L., 1982, On Identifying the Degree of Competitiveness from Industry Price and Output Data, *Economics Letters*,10, pp. 93-99.

Leibenstein, H., 1966, Allocative efficiency versus X-efficiency, *American Economic Review*, 56, pp. 392-415.

Maudos, J. and J. Fernández de Guevara, 2004, Factors explaining the interest rate margin in the banking sectors of the European Union, *Journal of Banking and Finance*, 28, pp. 2259-2281.

Mester, L.J. and A. Saunders, 1995, When does the prime rate change?, *Journal of Banking and Finance*,19, pp. 743-764.

Mojon, B., 2001, Financial structure and the interest rate channel of ECB monetary policy, *Economie et Provision*, 147, pp. 89-115.

Monti, M., 1972, Deposit, credit, and interest rate determination under alternative bank objectives, in *Mathematical Methods in Investment and Finance*, G.P. Szego and K. Shell (eds.), Amsterdam, North-Holland.

Neuwark, D. and S. Sharpe, 1992, Market structure and the nature of price rigidity: evidence from the market of consumer deposits, *Quarterly Journal of Economics*, 107, pp. 657-680.

Panzar, J., and J. Rosse, 1987, Testing for ’monopoly’ equilibrium, *Journal of Industrial Economics*, 35, pp. 443-456.

Pedroni, P.,1999, Critical values for cointegration tests in heterogeneous panels with multiple regressors, *Oxford Bulletin of Economics and Statistics*, Special Issue, pp. 653-670.

Pedroni, P.,2004, Panel cointegration: asymptotic and finite sample properties of pooled time series tests with an application to the PPP hypothesis, *Econometric Hypothesis*, 20, pp. 597-625.

Sander, H. and S. Kleimeier, 2002, Asymmetric adjustment of commercial bank interest rates in the euro area: an empirical investigation into interest rate pass-through, *Kredit und Kapital*, 35, pp. 161-192.

Sander, H. and S. Kleimeier, 2004, Convergence in euro zone retail banking?, *Journal of International Money and Finance*, 23, pp. 461-492.

Scholnick, B., 1996, Asymmetric adjustment of commercial bank interest rates: Evidence from Malaysia and Singapore, *Journal of International Money and Finance*, 15, pp. 485-496.

Smirlock, M., 1985, Evidence of the (non)relationship between concentration and profitability in banking, *Journal of Money, Credit and Banking*, 17, pp. 69-83.

Uchida, H. and Y. Tsutsui, 2005, Has competition in the Japanese banking sector improved?, *Journal of Banking and Finance*, 29, pp. 419-439.

Van Leuvensteijn, M., J.A. Bikker, A. van Rixtel and C. Kok-Sørensen, 2007, A new approach to measure competition in the loan markets of the euro area, ECB Working Paper Series 768.

Weth, M.A., 2002, The pass-through from market interest rates to bank lending rates in Germany, Deutsche Bundesbank Discussion Paper 11/02.

Winker, P., 1999, Sluggish adjustment of interest rates and credit rationing: an application of unit root testing and error correction modelling, *Applied Economics*, 31, pp. 267-277.

Appendix The estimation of the Boone Indicator Model

Description of the data used

The Boone indicator model uses Bankscope data of banks from eight euro area countries during 1992-2004.[28] This model is based on marginal costs which are derived from a translog cost function with output components and input prices. In order to exclude irrelevant and unreliable observations, banks are incorporated in our sample only, if they fulfilled the following conditions: total assets, loans, deposits, equity and other non-interest income should be positive; the deposits-to-assets ratio and loans-to-assets ratio should be less than, respectively, 0.98 and 1; the income-to-assets ratio should be below 0.20; personnel expenses-to-assets and other expenses-to-assets ratios should be between 0.05% and 5%; and, finally, the equity-to-assets ratio should be between 0.01 and 0.50. As a result, our final data set totals 520 commercial banks, 1506 cooperative banks, 699 savings banks, 28 special governmental credit institutions (Landesbanken) and 62 real estate banks (see Table A.1).

Table A.1 Number of banks by country and by type

Country	Commercial banks	Cooperative banks	Real estate banks	Savings banks	Specialized governmental credit institutions	Total
AT	52	54	10	65	0	181
BE	24	6	0	5	0	35
DE	130	867	44	501	28	1570
ES	61	17	0	43	0	121
FR	115	83	2	30	0	230
IT	105	476	1	52	0	634
NL	24	1	4	1	0	30
PT	9	2	1	2	0	14
Total	520	1506	62	699	28	2815

Table A.2 provides a short description of the model variables. To grasp the relative magnitude of the key variables, such as costs, loans, security investment and other services, we present them as shares of corresponding balance sheet items. Total costs are defined as total expenses. They vary between 6.3% and 8.6% of total assets, whereas market shares in the loan market vary between 0.06% and 5.8%. Loans and securities are in the range of, respectively, 35%-60% and 4%-37% of total assets. One of the output components we distinguish is other services. For lack of direct observations, this variable is proxied by non-interest income. Non-interest income ranges from 12%-20% of total income. Wage rates are proxied as the ratio of personnel expenses and total assets, since for many banks the number of staff is not available. Wages vary across countries between 0.9% and 1.7% of total assets. The input price of capital is proxied by

[28] See also Van Leuvensteijn *et al.* (2007), where a similar approach has been used.

the ratio of other expenses and fixed assets. Finally, interest rates are proxied by dividing interest expenses by total funding and range from 3.2% to 5.9%.

Table A.2 Mean values of key variables for various countries (in %)

Country Code	Boone model: Average loans market shares in %	Translog cost function: Total costs as % of total assets	Loans as % of total assets	Securities as % of total assets	Other services as % of total income	Other expenses as % of fixed assets	Wages as % of total assets	Interest expenses as % of total funding
AT	0.87	6.34	56	22	20	229	1.4	3.2
BE	2.27	6.49	35	37	16	594	1.0	4.5
DE	0.06	6.44	60	22	12	227	1.5	3.7
ES	0.98	6.63	58	14	16	167	1.5	4.1
FR	0.41	7.42	54	4	20	537	1.5	4.8
IT	0.22	6.67	53	26	16	261	1.7	3.5
NL	3.02	6.59	54	15	13	340	0.9	5.4
PT	5.83	8.62	52	8	18	191	1.3	5.9

Estimation results for marginal costs

We estimate a translog cost function for each separate country and take the first derivative of loans to derive the marginal costs of lending, see Equations (5) and (8), respectively.[29] Table A.3 shows the marginal costs of loans across countries and over time. Marginal costs decline over time, reflecting the significant decreases in funding rates during 1992-2004 and possibly also technological improvements. Germany, France and Spain have relatively high marginal costs compared to the Netherlands and Belgium. Apart from differences in funding rates, this may be explained also by lower efficiency in the former countries.[30]

[29] See also Section 3.1 in Van Leuvensteijn *et al.* (2007).

[30] Another explanation is lower population density in the former countries. Low population density may raise operating costs, as it makes retail distribution of banking services more costly.

Table A.3	Marginal costs of loans across countries and over time (in %)							
	AT	BE	DE	ES	FR	IT	NL	PT
1992	10.3	7.1	10.2	15.9	13.8	13.2	9.2	21.3
1993	9.4	6.9	9.4	17.2	13.4	12.0	8.1	18.8
1994	7.1	6.4	9.2	14.3	11.9	12.2	7.4	16.6
1995	7.3	5.8	8.9	15.4	11.7	11.8	7.1	15.4
1996	7.1	5.2	8.5	14.3	10.9	11.3	6.3	13.4
1997	6.1	4.6	7.4	11.7	10.9	9.7	6.4	12.3
1998	6.0	3.6	7.1	11.1	11.2	7.5	7.4	9.4
1999	5.5	3.2	6.4	8.8	10.0	6.7	6.4	6.1
2000	6.1	3.3	7.1	9.9	11.2	6.7	6.5	6.3
2001	6.1	3.1	7.3	9.6	11.7	6.6	6.4	5.9
2002	5.7	3.1	7.1	7.8	10.7	6.1	5.7	5.2
2003	5.5	2.7	6.4	5.9	8.9	5.3	4.9	5.3
2004	5.2	2.5	6.0	4.8	7.9	4.9	4.6	5.5

Estimation results for the Boone indicator

Table A.4 shows the estimates of the Boone indicator across countries and over time (usually 1994-2004, depending on the respective country). The results are based on the following model:

$$\ln ms_{i,t} = \alpha + \sum_{t=1,..,T} \beta_t \ln mc_{i,t} + \sum_{t=1,..,(T-1)} \gamma_t\, d_t + u_{i,t} \qquad \text{(A.1)}$$

explaining loans market shares of bank *i* in year *t* ($ms_{i,t}$) by marginal costs ($mc_{i,t}$) and country dummies (d_t). Note that the Boone indicator, β_t, is time dependent. The estimations are carried out using the Generalized Moment Method (GMM) with as instrument variables the one-, two- or three-year lagged values of the explanatory variable, marginal costs, or average costs. To test on overidentification of the instruments, we apply the Hansen J-test for GMM (Hayashi, 2000). The joint null hypothesis is that the instruments are valid as such, *i.e.* uncorrelated with the error term. Under the null hypothesis, the test statistic is chi-squared with the number of degrees of freedom equal to the number of overidentification restrictions. A rejection would cast doubt on the validity of the instruments. Furthermore, the Anderson canonical correlation likelihood ratio is used to test for the relevance of excluded instrument variables (Hayashi, 2000). The null hypothesis of this test is that the matrix of reduced form coefficients has rank K-1, where K is the number of regressors, meaning that the equation is underidentified. Under the null hypothesis of underidentification, the statistic is chi-squared distributed with L-K+1 degrees of freedom, where L is the number of instruments (whether included in the equation or excluded). This statistic provides a measure of instrument relevance, and rejection of the null hypothesis indicates that the model is identified. We use kernel-based heteroskedastic and autocorrelation consistent (HAC) variance estimations. The bandwidth in the estimation is set at two periods and the Newey-West kernel is applied. Where the instruments are overidentified, 2SLS is used instead of GMM. For this 2SLS estimator, Sargan's statistic is used instead of the Hansen J-test.

Over the sample period, the Boone indicator for Belgium, Germany, and Italy are highly significant, except for one or two years, suggesting stronger loan market competition then elsewhere in the euro area.[31] The Dutch and Spanish loan markets take up an intermediate position with significant Boone indicators for at least a number of years. For France, the degree of competition declined over the years, where the reverse development is observed for Austra and Portugal. If, for each country, we had estimated only one beta for the full-sample period instead of annual ones (that is, $\beta_t = \beta$ for all t), we would have obtained significant values for all countries (except Portugal), reflecting a certain degree of competition in the whole area (see Van Leuvensteijn *et al.*, 2007).

Table A.5 The Boone indicator over time and across various countries[2)]

	Germany[1)]		France		Italy[1)]	
	β_t	Z-value	β_t	Z-value	β_t	Z-value
1993					– 5.90	– 1.18
1994					– 7.25**	– 3.24
1995	– 4.47	– 1.40	– 1.28**	– 3.36	– 4.51**	– 3.53
1996	– 7.09**	– 2.92	– 1.28**	– 3.56	– 5.58**	– 3.98
1997	– 4.64**	– 3.41	– 1.11**	– 3.55	– 5.89**	– 4.08
1998	– 5.10**	– 3.97	– 0.79*	– 1.99	– 4.60**	– 6.08
1999	– 2.60**	– 4.04	– 0.7*	– 2.30	– 4.05**	– 4.39
2000	– 2.50**	– 4.60	– 0.46	– 1.34	– 3.32**	– 4.39
2001	– 3.31**	– 7.02	– 0.68	– 1.67	– 2.66**	– 3.62
2002	– 4.53**	– 4.71	– 0.40	– 0.78	– 1.59	– 1.82
2003	– 2.73**	– 5.62	0.27	0.39	– 2.42**	– 3.69
2004	– 2.66((	– 4.15	0.10	0.12	– 1.81**	– 2.79
F-test	10.70			5.01	13.23	
Anderson canon corr. LR-test	185.20			1023.66	300.34	
Hansen J-test (p-value)	0.00			19.69 (0.48)	0.00	
Number of observations	14 534			918	4918	

Notes: Asterisks indicate 95% (*) and 99% (**) levels of confidence. Coefficients of time dummies have not been shown.
[1] 2SLS is used and the equation is exactly identified, so that the Hansen J-test is 0.00.
[2] Equation (A.1) is estimated with the GMM.

[31] Most likely, the favourable result for Germany hinges in part on the special structure of its banking system, being built on three pillars, *i.e.* the commercial banks, the publicly-owned savings banks and the cooperative banks (see Hackethal, 2004).

Table A.5 The Boone indicator over time and across various countries[2)] (continued).

	Spain[1)]		Netherlands		Belgium	
	β_t	Z-value	β_t	Z-value	β_t	Z-value
1993	– 4.21*	– 2.49				
1994	– 4.80*	– 2.28	– 1.92	– 1.42		
1995	– 5.20	– 1.92	– 4.42*	– 2.42	– 1.48	– 1.59
1996	– 9.61	– 0.67	– 2.09**	– 2.58	– 1.74**	– 2.93
1997	– 4.36	– 1.78	– 3.57	– 1.70	– 2.02((	– 3.78
1998	– 5.40	– 0.86	1.04	0.38	– 1.98**	– 3.19
1999	– 5.46*	– 2.21	– 1.44	– 0.85	– 2.62**	– 4.65
2000	– 3.44	– 1.93	– 3.26**	– 3.00	– 3.41**	– 6.10
2001	– 4.38**	– 2.55	– 3.91**	– 4.71	– 3.00**	– 4.51
2002	– 3.88*	– 2.09	– 2.45*	– 2.44	– 3.42**	– 4.34
2003	– 3.42	– 1.20	– 2.22	– 1.80	– 2.79**	– 3.18
2004	– 2.69**	– 5.62	– 3.09**	– 2.85	– 3.12**	– 4.02
F-test	3.33		3.90		6.35	
Anderson canon corr. LR-test	38.78		31.71		178.10	
Hansen J-test (p-value)	0.00		20.5 (0.039)		8.34 (0.60)	
Number of observations	1015		241		269	

	Austria		Portugal	
	β_t	Z-value	β_t	Z-value
1994	11.2	1.01	0.05	0.05
1995	– 4.03	– 0.94	1.57	0.91
1996	– 2.31*	– 1.93	0.09	0.16
1997	4.25	0.93	– 0.04	– 0.08
1998	– 0.91	– 0.52	– 0.55	– 0.76
1999	– 2.98	– 0.73	– 1.51	– 1.40
2000	– 2.31	– 0.50	– 2.43**	– 4.03
2001	– 0.96	– 1.30	– 1.92**	– 3.77
2002	– 1.49*	– 1.97	– 2.16**	– 7.33
2003	– 1.26**	– 3.52	– 1.74*	– 2.05
2004	– 2.99**	– 2.23	– 1.53	– 1.69
F-test	2.21		3.94	
Anderson canon corr. LR-test	28.89		77.92	
Hansen J-test, (p-value)	9.308 (0.59)		11.71 (0.38)	
Number of observations	988		134	

Notes: Asterisks indicate 95% (*) and 99% (**) levels of confidence. Coefficients of time dummies have not been shown.

[1] 2SLS is used and the equation is exactly identified, so that the Hansen J-test is 0.00.

[2] Equation (A.1) is estimated with the GMM.