Journal of Banking & Finance 33 (2009) 1196–1210

Contents lists available at ScienceDirect

Journal of Banking & Finance

journal homepage: www.elsevier.com/locate/jbf

Market power in local banking monopolies

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ARTICLE INFO

Article history:
Received 31 March 2008
Accepted 27 December 2008
Available online 8 January 2009

JEL classification:
G21
L10
L13

Keywords:
Banking
Monopoly
Market structure
Conduct

ABSTRACT

By means of two NEIO techniques, this paper analyzes the conduct of a group of Italian single-branch banks operating as monopolists in small local areas (municipalities) in the years 1988–2005, in order to assess pricing behavior in highly concentrated banking markets. Both tests strongly reject the hypothesis of pure monopoly pricing: regardless the advantageous condition, these banks are able to exploit only partially their market power, principally by reason of the nearby competition, the latest banking consolidation trend and the local presence of big banks. Employing another sample, we also show that in duopolistic markets the conduct of single-branch banks is virtually competitive.

1. Introduction

For several countries, the most recent decades have been characterized by a dynamic process of consolidation within the banking industry, a vital sector for any economy because of its role in transmitting monetary policy and providing credit to firms and households. On one side, this process has been fastened by a worldwide deregulation of capital markets, a tendency to harmonized financial legislations and a generalized reduction of entry barriers; on the other side, it has represented an unavoidable answer to the notably increased competition that has forced domestic banks to search for higher levels of efficiency and offer diversified services to customers, both imposing the need of exploiting scale economies and therefore to look for bigger dimensions.

In Europe, this phenomenon has been amplified by the liberalization of capital flows among countries and the prospect of a common market, further aspects pushing banks to increase their size in order to cut costs and gain market shares at the same time.

As part of Europe, Italy has followed this path as well: in the last fifteen years the number of banks declined by a third, and their average size and their branch network more than doubled. In spite of this rapid consolidation, the dimension of the banking market still remains relatively small compared to other EU member countries: in 2005 the Herfindahl–Hirschman index for Italy has been the lowest in Europe after Germany, far from United Kingdom and France (ECB, 2006, p. 54).

Thus, the degree of market competition in the Italian banking sector becomes a foremost issue: if the number of credit institutions reduces, the resulting increased concentration might augment the market power of active banks. Previous studies conducted in Italy (but this is generally true also for other banking markets) have rejected this worry, yet the size and fragmentation of the market requires further investigations, with particular reference to small banks and local areas.

This paper focuses on single-branch banks (i.e. unit banks with only one branch) that operate as monopolists in local markets, namely in municipalities (in Italy, the smallest administrative territorial unit). We aim to assess the extent to which their conduct reflects this special condition, also controlling for factors like external competition and local market size. This is done by applying two different econometric techniques over a sample of 86 small credit institutions for the years 1988–2005, thus covering the entire span of time characterized by the consolidation trend. Finally, we compare the derived empirical results with those coming from an investigation on single-branch banks in duopolistic contexts.

The structure of the paper is the following. In Section 2 we briefly review the literature dealing with the assessment of the degree of competition in banking markets. Section 3 illustrates the first test, a simultaneous-equation model that allows to estimate a behavioral parameter measuring the degree of monopoly power that banks are actually able to exploit in their municipality, and discusses the results of different specifications of the model.

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doi:10.1016/j.jbankfin.2008.12.013

Section 4 describes the second test, the so-called Panzar–Rosse *H*-statistic, which proceeds from the estimation of a revenue equation, and examines the related results in parallel with those obtained from the previous model. Section 5 applies the *H*-test to a sample of banks that operate in local duopolies: in this way, a comparison between the conduct of similar banks in different contexts is made possible. Section 6 concludes.

2. A review of the literature

In order to measure the degree of competition in banking markets, empirical studies have usually made use of two different methodologies. The first is the structural approach, that stems from the traditional structure–conduct–performance (SCP) paradigm, and studies the possibility that highly concentrated markets generate a tendency to collusion among banks and hence allow them to enjoy higher profitability. The second is the so-called "new empirical industrial organization" (NEIO) approach, which relies on non-structural models that infer market power from the observation of banks' conduct, and requires the estimation of equations based on theoretical models of price and output determination.

According to the SCP approach (Mason, 1939; Bain, 1951), structure and performance are positively related, because firms in more concentrated markets are likely to earn higher profits than those operating in less concentrated environments, thanks to the exploitation of their greater market power or the ease of collusive agreements. However, the evidence deriving from empirical SCP studies for the banking industry does not fully support the theory.[1] On one side, the results by Berger and Hannan (1989), Hannan and Berger (1991) and Pilloff and Rhoades (2002) are in line with the SCP predictions. On the other side, Jackson (1992), Rhoades (1995) and Hannan (1997) find less robust evidence for the positive relationship between market concentration and banks' profitability.

Some authors have criticized the SCP paradigm by noting that a higher level of efficiency for one or more banks can give rise to higher profits irrespective of market concentration (Demsetz, 1973; Peltzman, 1977): as a consequence, the resulting positive link between market structure and profitability would be a spurious signal of the SCP hypothesis, unless one controls for efficiency.[2]

In order to overcome the above difficulties related to the structural approach, the NEIO models try to test conduct by directly addressing firms' behavior. This has been done in two ways. The first is to estimate a parameter that represents the behavior of firms (and therefore the degree of their market power): it can be interpreted as a conjectural variation coefficient (Iwata, 1974; Appelbaum, 1979, 1982; Roberts, 1984), or as the deviation of the perceived marginal revenue schedule of a firm in the industry from the demand schedule (Bresnahan, 1982, 1989; Lau, 1982; Alexander, 1988). Applications of this technique (and its variants) in banking have been performed by Shaffer (1989, 1993), Berg and Kim (1994), Shaffer and DiSalvo (1994), Coccorese (1998b, 2005), Neven and Roller (1999), Toolsema (2002), Angelini and Cetorelli (2003), Canhoto (2004) and Uchida and Tsutsui (2005). The advantage of this method is the direct analysis of firms' conduct (instead of focusing on the overall market structure), which helps to avoid indirect (and sometimes ambiguous) inferences about market power based on indicators of concentration; on the other side, it needs detailed information on costs and demand.

The second technique consists of a comparative statics analysis, and is based on the evaluation of the impact of variations in input prices on firms' revenues. The identification of market power is then obtained through an index (the Panzar–Rosse *H*-statistic), calculated as the sum of the elasticities of the reduced-form revenue with respect to all the factor prices (Rosse and Panzar, 1977; Panzar and Rosse, 1987). Empirical applications of the Panzar–Rosse test to the banking industry include those by Shaffer (1982, 2002, 2004), Nathan and Neave (1989), Molyneux et al. (1994), Coccorese (1998a, 2004), Hondroyiannis et al. (1999), De Bandt and Davis (2000), Bikker and Haaf (2002), Claessens and Laeven (2004), Gelos and Roldos (2004), Al-Muharrami et al. (2006), Casu and Girardone (2006), Staikouras and Koutsomanoli-Fillipaki (2006) and Matthews et al. (2007). The main advantage of the *H*-statistic is the need of firm-specific data on revenues and factor prices only; further information on costs and specific hypotheses on the definition of the (either product or geographic) market are not required, although the insertion of every variable shifting demand or cost (and possibly of firm-specific or also macroeconomic control variables) is most desirable.[3]

In what follows, we are going to test both NEIO techniques over the sample of Italian local monopolistic banks, trying to verify whether there is a great deal of market power in areas where only one bank is active. The Italian banking sector represents an important part of the overall European financial market, and shares many characteristics with a number of EU countries (e.g. France, Germany): a bank-oriented financial system; rather rigid labor markets that might impede thorough restructuring; a mix of large and small banks; a gradual shift from the traditional intermediation business to a more services-oriented industry.[4] Under this point of view, our analysis can therefore constitute a benchmark for analogous investigations regarding other European nations.

In Italy, apart few exceptions, banks usually have a fairly limited area of business: local banking markets are mainly oligopolies,[5] and the prevalence of small-size banks is generally widespread. In about 13% of provinces,[6] two banks are able to cover more than a half of the total number of branches; this figure grows to 39% of provinces when considering three banks (2005 figures). Few years ago, it was calculated that in more than half of Italian provinces two banks managed half of the deposits, while in another third of provinces only three were necessary.[7] Moreover, in 2005 the municipalities with at least one operating bank were 5924: of them, in 2219 there was only one branch (37%), two in 1309 municipalities (22%).

As a result, the conduct of local banks needs to be investigated. In what follows, we will implicitly assume that the municipality is the relevant banking market.[8] Of course, there is the possibility that close rivals exercise competitive pressures also on local monopolists or oligopolists, especially if they are big banks. On the other side, the competitors' lack of territorial roots and information

[1] Surveys of empirical SCP studies are provided by Gilbert (1984) and Weiss (1989).

[2] Investigations on the role of efficiency within the SCP paradigm for the banking sector have been proposed by Smirlock (1985), Berger (1995), Goldberg and Rai (1996) and Molyneux (2003).

[3] More considerations about advantages and disadvantages of the NEIO techniques can be found in Shaffer (2004).

[4] See Focarelli et al (2002, p. 1048). Sapienza (2002, pp. 336–337) emphasizes that the Italian banking industry has a number of similarities also with that of the United States.

[5] This pattern is again very much like the United States.

[6] In Italy, the province (*provincia*) is an administrative district that comprises a larger town or city and several neighbouring municipalities. By and large, it corresponds to a US county. Since 1995, the number of Italian provinces is 103.

[7] These data refer to mid-1990. See Coccorese (2004, p. 212). In the light of the growing concentration, they are likely to have been confirmed or even reinforced in the following years.

[8] This definition is narrower than what is usually supposed. However, the assumption that the relevant banking market is considered as geographically delimited within local areas (especially for deposits) has been widely recognized in the empirical literature on banking. About Italy, see for example Focarelli and Panetta (2002, p. 1156), Sapienza (2002, p. 335) and Angelini and Cetorelli (2003, p. 666). This is especially true for small banks, the focal point of this analysis.

about the local clientele, and also the presence of switching costs of any kind for the customers, may ease the exploitation of market power from incumbent banks. We can not even rule out that the significant market share (in terms of managed deposits and loans) may induce local leading banks to collude with nearby credit institutions so as to preserve their dominant position.

Unfortunately, the analysis of market power in narrow areas requires very disaggregated data that are hard to be found. Nonetheless, a way to overcome the problem is to refer to the balance sheets of single-branch banks that operate in local markets, particularly banks with a monopolistic position, whose data (especially those related to loans and deposits) can be also assumed as market data (although at a local level), at the same time allowing us to infer banks' behavior through the implementation of the NEIO tests.

3. The simultaneous-equation approach for monopolistic banks

3.1. The model

Let us suppose that bank i is monopolist in its reference market, the latter being considered as local and therefore of relatively narrow size. Assuming a price-setting model,[9] let the demand for loans it faces at time t be $q_{it} = q_{it}\,(p_{it},Z_t)$, where q_{it} is the quantity demanded, p_{it} is the price that bank i charges, and Z_t is a vector of exogenous factors influencing demand.

We therefore consider only one output (loans), implicitly assuming that credit intermediation represents the predominant activity of bank i. Let the related cost function be $C_{it} = C_{it}(q_{it}(\cdot),\omega_{it})$, where ω_{it} is the vector of the prices of input factors.

Hence, the profit function for bank i can be written as:

$$\pi_{it} = q_{it}(\cdot)p_{it} - C_{it}(q_{it}(\cdot),\omega_{it}). \tag{1}$$

Maximization implies that

$$\frac{\partial \pi_{it}}{\partial p_{it}} = q_{it} + \frac{\partial q_{it}}{\partial p_{it}}(p_{it} - MC_{it}(\cdot)) = 0, \tag{2}$$

where $MC_{it}(\cdot) = \frac{\partial C_{it}}{\partial q_{it}}$ represents the marginal cost function. Rearranging, we get:

$$L_{it} = \frac{p_{it} - MC_{it}}{p_{it}} = -\frac{1}{\varepsilon_{it}}. \tag{3}$$

In the above expression, L_{it} is the well-known Lerner index, measuring the extent to which the profit-maximizing price exceeds marginal cost (mark-up). Its value depends on the price elasticity of demand faced by bank i, ε_{it}, and is equal to 0 for a competitive firm, whose demand is perfectly elastic, implying $p = MC$.

In the empirical studies, the Lerner index has been often employed to evaluate the degree of market power of firms operating in a given industry.[10] Indeed, it is able to measure the extent to which market power is *actually* exercised: if the firm prices below the optimal price (to avoid antitrust investigations, to make rivals' entry unattractive, to avoid hit-and-run behavior from external competitors, etc.), its potential market power will not be captured by the index (Pindyck and Rubinfeld, 1995, p. 334).

In our framework, we are dealing with local banking monopolies. Hence, it is possible to assess the real market power of banks by slightly changing (3) into

$$L_{it} = \frac{p_{it} - MC_{it}}{p_{it}} = -\frac{\theta}{\varepsilon_t}. \tag{4}$$

The new parameter θ (with $0 \leqslant \theta \leqslant 1$) just measures the average deviation of banks' behavior from the monopolistic case and, if correctly identified in the estimation, expresses the true degree of market power exercised by banks. The full exploitation of the monopolistic market power coincides with $\theta = 1$, as (4) turns into (3). Conversely, $\theta = 0$ means that, in spite of the local monopolistic status, the bank is compelled to behave as a perfectly competitive firm (for which $p = MC$), therefore enjoying no market power. Note also that the price elasticity of demand has been now indicated as ε_t, without the subscript i, to stress that in this framework banks face the whole market demand curve.

In order to identify the behavioral parameter θ, we estimate a three-equation system. The first equation represents the market demand function, which is assumed to take the following form:

$$\ln q_{it} = a_0 + a_1 \ln p_{it} + a_2 \ln p_{PROVt} + a_3 \ln POP_t + a_4 \ln BR_{NEIt} + \tau_{it}. \tag{5}$$

Here q_i and p_i are the quantity and the price of the output (loans) of bank i, respectively; p_{PROV} is the calculated average value expressing the price set by the banks belonging to the same province of bank i,[11] and represents a proxy for the price set by nearby competitors[12]; POP measures the population of the municipality where bank i operates, and tries to take into account the size of the main reference market for the monopolistic bank; BR_{NEI} is the number of branches of all banks belonging to the neighboring municipalities (including bank i, in order to avoid possible zero values), and allows to consider the effect of bordering competition on the monopolist's demand; finally, τ_i is an error term.

The second equation is a total cost function. In line with the main studies on the banking industry, we consider a translog specification, which does not require particular strong assumptions about its functional form and is useful especially when dealing with both scale and scope economies in case of multi-product firms. Consistent with the intermediation approach,[13] the banks' output is here represented by loans[14]; regarding the inputs, we consider three factors (deposits, labor and physical capital). As a result, our cost function assumes the following form:

$$\begin{aligned}\ln C_{it} = {} & \beta_0 + b_0 \ln q_{it} + \frac{b_1}{2}(\ln q_{it})^2 + \sum_{r=1}^{3} \beta_r \ln \omega_{rit} \\ & + \ln q_i \sum_{r=1}^{3} b_{r+1} \ln \omega_{rit} + \frac{1}{2}\sum_{r=1}^{3} \beta_{r+3}(\ln \omega_{rit})^2 \\ & + \beta_7 \ln \omega_{1it} \ln \omega_{2it} + \beta_8 \ln \omega_{1it} \ln \omega_{3it} \\ & + \beta_9 \ln \omega_{2it} \ln \omega_{3it} + \varphi_{it},\end{aligned} \tag{6}$$

where φ_i is an error term.

We also impose the usual conditions on its coefficients. In particular, while the above specification makes possible to avoid the test for symmetry,[15] we constrain the cost function to be homogeneous of degree one in input prices, implying the restrictions

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9 Probably, prices represent a more appropriate strategic variable for banks. See Freixas and Rochet (1997, p. 65) and de Haan and Sterken (2006, p. 11).

10 For example, see Genesove and Mullin (1998) for the sugar industry, and Sheldon and Sperling (2003) for food retailing. Regarding the banking industry, recent papers are those by Angelini and Cetorelli (2003), and Maudos and Fernandez de Guevara (2004).

11 We refer to the province because we do not have more disaggregated information regarding the interest rates on loans for the sample period.

12 Similar studies employ the price of a product substitute rather than the average rival price (e.g. Shaffer, 1993). We opt for p_{PROV} in order to focus on rivals' choices. However, using the short-term government bond rate instead of p_{PROV} does not significantly change our results.

13 This approach is the most widely adopted in banking. For a general discussion, see Freixas and Rochet (1997, pp. 77–81).

14 In spite of the recognized multi-product nature of banks, our assumption that loans are the only output for banks seems by and large acceptable, given that credit intermediation is still the predominant activity of commercial banks, especially in the first part of the time interval that characterizes our sample.

15 Actually, symmetry in the coefficients of outputs is not necessary because we suppose banks produce only loans. Symmetry in the coefficients of input prices would be necessary if different parameters both for $\ln\omega_r \ln\omega_s$ and $\ln\omega_s \ln\omega_r$ are estimated ($r = 1,\ldots,m$; $s = 1,\ \ldots,m$), but in (6) we consider one coefficient for each pair of multiplications, what permits to implicitly impose such condition.

$\sum_{r=1}^{3}\beta_r = 1, \quad \sum_{r=1}^{3} b_{r+1} = 0$ and $\sum_{r=1}^{6}\beta_{r+3} = 0$.

We can now calculate the marginal cost function:

$$MC_{it} = \frac{\partial C_{it}}{\partial q_{it}} = \frac{C_{it}}{q_{it}}\left(b_0 + b_1 \ln q_{it} + \sum_{r=1}^{3} b_{r+1} \ln \omega_{rit}\right), \tag{7}$$

where ω_{rit} are the prices of input factors.

Substituting (7) in (4), we get:

$$\frac{p_{it} - \frac{C_{it}}{q_{it}}\left(b_0 + b_1 \ln q_{it} + \sum_{r=1}^{3} b_{r+1} \ln \omega_{rit}\right)}{p_{it}} = -\frac{\theta}{\varepsilon_t}, \tag{8}$$

Rearranging (8) through simple manipulations, we obtain:

$$p_{it} = \frac{\frac{C_{it}}{q_{it}}\left(b_0 + b_1 \ln q_{it} + \sum_{r=1}^{3} b_{r+1} \ln \omega_{rit}\right)}{1 + \frac{\theta}{\varepsilon_t}} + \phi_{it}, \tag{9}$$

where ϕ_{it} is an error term.

The system we are going to estimate is formed by Eqs. (5), (6), and (9). If the behavioral parameter θ can be identified through the estimation, it describes the degree of average market power exploitation characterizing the monopolistic banks in our price-setting game context: therefore, the presence of a conduct that is closer to monopoly (perfect competition) should give rise to higher (lower) values of θ.

According to Lau (1982), a necessary and sufficient condition for identification of the conduct parameter θ in a system of demand and cost functions is that the demand function must not be separable in at least one of the exogenous variables, which has also to be excluded from the marginal cost function. Our log-linear specification of demand ensures the satisfaction of this condition: actually, for example, it is $\frac{\partial^2 q_{it}}{\partial p_{it}\partial p_{PROVt}} = \frac{a_1 a_2 q_{it}}{p_{it} p_{PROVt}} \neq 0$.

Note also that the precision and efficiency of estimates are improved by the simultaneous estimation of total cost function (6) and price-cost margin function (9), and the related cross-equation restrictions (Angelini and Cetorelli, 2003, pp. 667–668; Coccorese, 2005, pp. 1087–1088). Besides, given the monopolistic nature of the considered local banking markets, our three-equation model allows to estimate an unbiased average behavioral parameter θ, as each observation spans at least one complete market (Shaffer, 2001, p. 88). Finally, in order to capture both bank-specific and time effects, two vectors of dummy variables have been inserted in the demand and cost equations[16] (as a result, they replace the constants).

3.2. Data and results

The sample considers 86 single-branch banks over the period 1988–2005. This implicitly means that our data refer to 86 municipalities, those where the above banks were the sole credit institutions, therefore acting as monopolists. All of them are cooperative credit banks (CCBs), small organizations fairly similar to US credit unions.[17] Angelini and Cetorelli (2003) argue that there exist several characteristics that put them in a "niche position" and therefore would potentially give them extra market power, but their empirical analysis allows one to conclude that CCBs' conduct is similar to commercial banks, and that the hypothesis that they operate in market niches sheltered from competition finds no support in the data.[18]

As Table 1a shows, the total number of sample observations amounts to 381. On average, in the reference period we consider about 32% of Italian monopolistic single-branch banks, but the sample coverage notably increases during years. The most represented areas are North-East and Center. The sample is unbalanced, especially because of mergers and acquisitions occurred in this time interval, which caused also a drop of data in the last years. Bank-level data come from Bankscope, the database managed by Fitch-IBCA, but have been combined with those published yearly by the Italian financial magazine "MF Milano Finanza", which collects the balance-sheet information of a large group of banks. Data on the distribution of branches come from the Bank of Italy, while the local population information is made available by ISTAT (the Italian National Institute of Statistics).

In the demand Eq. (5), the output, q, is measured by the value of the bank loans, and the price of output, p, is assumed as the interest rate earned on loans, calculated as the ratio between interest revenue and total loans. We expect the coefficient of p be negative, in conformity to a downward-sloping demand curve. Regarding the calculation of p_{PROV}, we followed an analogous procedure; particularly, starting from bank-level data, we computed it as the interest rate on loans earned by the "average" bank operating in the same province of bank i, set equal to the ratio between the interest revenue of all those banks whose business can be attributed to that province and their loans.[19] Its coefficient is expected to be positive if loans are substitutable across banks, on condition that customers are willing to go to a bank operating in another municipality to ask for a loan. We also expect that the cross-price elasticity is smaller than the own-price elasticity, if banks are able to soften price competition somehow.

The population of the municipality where banks are situated, *POP*, is supposed to positively affect their loans, because it represents the primary potential clientele. The variable BR_{NEI} should be negatively related to loans: if in the bordering municipalities there are several rival branches, the possibility of choice for customers is greater (since they enjoy a wider variety of services), with negative effects on the demand faced by the monopolistic bank.

As regards the cost Eq. (6), a correct specification of the behavioral parameter θ requires that the prices of the inputs (deposits, labor and physical capital) are exogenously given. This assumption can be acceptable in our framework: concerning labor and physical capital, banks compete with many other firms for the acquisition of these inputs, while it is hard to imagine that deposit interest rates are under full banks' control, because of the fierce competition coming from government bonds and other alternative investment options. In addition, those prices, being almost surely predetermined, are not likely to cause endogeneity bias in the estimates.

In Eqs. (6) and (9), we computed the price of deposits, ω_1, as the ratio between interest expenses and deposits, and the price of la-

[16] In an alternative specification, time dummies have been substituted by a linear trend that accounts for the evolution of the endogenous variables. As the results are virtually the same, we do not report them.

[17] For a description of this category of banks, see Angelini et al. (1998).

[18] See Angelini and Cetorelli (2003, p. 666 and p. 675). This evidence helps to solve a potential problem that derives from the fact that CCBs are non-profit organizations, namely the possibility that they do not behave according to the maximization problem we have postulated. However, Shaffer (2002) holds that, even if firms follow a different strategy than profit maximization, or if regulations constrain banks' behavior, it is always possible to compare a firm's price setting strategy with a competitive or socially optimal benchmark (p. 2002).

[19] In order to ensure that this provincial interest rate reflects local specificities, we considered only those banks with at least 50% of branches belonging to a given province, and calculated p_{PROV} only when there were at least three of such banks per province. This happened for 367 over 381 observations (96.3% of the sample). When the provincial banks were less than three (10 observations, 2.6%), we aggregated them to a contiguous province prior to the computation of p_{PROV}. Finally, for one bank (4 observations, 1.1%) it was impossible to have a proxy of the provincial rate, so we employed the regional value (still calculated at the bank level).

Table 1
Single-branch banks in Italian municipalities (1988–2005).

(a) Monopolistic single-branch banks				(b) Duopolistic single-branch banks			
	Italy	Sample	Sample coverage (%)		Italy	Sample	Sample coverage (%)
1988	197	18	9.1	1988	71	11	15.5
1989	191	30	15.7	1989	69	13	18.8
1990	171	45	26.3	1990	58	19	32.8
1991	124	33	26.6	1991	49	16	32.7
1992	97	37	38.1	1992	47	16	34.0
1993	77	24	31.2	1993	36	15	41.7
1994	66	30	45.5	1994	27	10	37.0
1995	56	29	51.8	1995	25	8	32.0
1996	45	26	57.8	1996	17	5	29.4
1997	37	22	59.5	1997	17	4	23.5
1998	29	18	62.1	1998	14	5	35.7
1999	23	10	43.5	1999	12	3	25.0
2000	20	7	35.0	2000	11	5	45.5
2001	18	14	77.8	2001	8	5	62.5
2002	15	12	80.0	2002	8	4	50.0
2003	10	10	100.0	2003	7	6	85.7
2004	9	8	88.9	2004	6	5	83.3
2005	9	8	88.9	2005	6	5	83.3
Total	1194	381	31.9	Total	488	155	31.8
North-West	151	23	15.2	North-West	45	4	8.9
North-East	401	239	59.6	North-East	164	107	65.2
Center	95	52	54.7	Center	62	23	37.1
South	535	63	11.8	South	123	9	7.3
Islands	12	4	33.3	Islands	94	12	12.8
Total	1194	381	31.9	Total	488	155	31.8

Source: our elaborations on Bank of Italy data.

bor, ω_2, as the ratio between total labor costs and the number of employees.[20] For the price of physical capital, ω_3, our idea is that in this kind of analysis a good measure can be obtained simply considering the value of all net operating costs different from those related to deposits and labor; since firms in our sample are single-branch banks, the above value constitutes a good proxy for the running costs of the local office.

All financial figures are expressed in 2000 euro using the Gross Domestic Product deflator. Table 2a reports summary statistics of the data regarding the monopolistic banks.

The system has been estimated simultaneously through non-linear three-stage least squares, in order to take into account both the endogeneity of the left-hand variables and the cross-equation restrictions. This technique also allows to cope with possible correlation of the right-hand side variables with the error terms, as well as the contemporaneous presence of heteroskedasticity and correlation in the residuals. The results are shown in the first column of Table 3 (only the coefficients regarding the dummy variables associated to time and firm effects are not reported).

The estimate of the price elasticity of demand, a_1, is highly significant and amounts to -0.82. This means that the (downward-sloping) demand for loans in these municipalities is inelastic. However, it is not a profit-maximizing behavior to fix the equilibrium quantity where the demand that firms face (here corresponding to the market demand) is perceived to be inelastic, because this would imply a negative marginal revenue, which can not be equated to marginal cost, the latter being always positive (Shaffer and DiSalvo, 1994, p. 1073).

Given that the first-order condition for static profit maximization can not be satisfied, this value is inconsistent with a monopoly conduct, and implies that behavior must be therefore closer to that of a firm operating in an environment characterized by a certain degree of competition.

The point estimate of θ for the whole sample is 0.6957, significantly different both from the monopoly value of 1 and from the competitive value of 0. This result conforms with the preceding discussion on the elasticity of demand for loans: our monopolistic banks are not able to fully exploit their market status, and enjoy a price-cost margin that is about 30% lower than that theoretically characterizing a monopoly.

The average level of provincial loan rates does not affect local demand, since the estimated coefficient a_2 is not significant, and therefore there is no cross-price effect. This outcome also helps to corroborate our assumption that the municipality can be regarded as the main reference market for this sample of banks. In contrast, a role is clearly played by rivals' closeness: as expected, the variable BR_{NEI} has a negative and significant coefficient, so that the mere presence of other banks in the neighborhood adversely impacts the loan demand of monopolistic banks. In general, this is an indicator of the pressure that potential competition exerts on banks, even when they operate alone in a market. Instead, the population size of the municipality does not appear to affect demand.

A question arises: why potential competition, whose role has been widely recognized especially in local banking markets, and should be even stronger in our narrow context (i.e. at the municipality level), is not able to lessen market power of incumbent firms up to a competitive level? One factor that could explain the surviving market power of our sample of banks is represented by information asymmetries. In the banking industry the interaction between market structure, competition and prices is heavily influenced by the nature of the loan contract. Throughout their activity, banks produce valuable information about financial characteristics of borrowers. If such information can not be transferred easily to new, potential lenders, then banks could acquire information-based market power on those borrowers who suffer most from asymmetric information (Sapienza, 2002, p. 332).

As known, some banks' assets constitute sunk costs as long as they can not be recovered in their entirety. Particularly, this

[20] We agree with Shaffer (2001), who notes that the construction of an accurate measure of wage rates requires data on the number of bank employees.

Table 2
Sample descriptive statistics.

Variable	Mean	Std. dev.	Minimum	Maximum	Median
(a) Monopolistic banks[a]					
Total assets	23702.0	15304.9	2316.6	113977.8	19957.8
Loans	9925.6	9658.7	379.8	82715.0	7484.0
Risk capital	3074.5	2480.1	232.0	18444.0	2303.0
Total cost	1894.1	1062.2	200.2	6508.4	1572.9
Employees (∗)	6.77	3.60	1	22	6
Population (municipality) (∗)	2156.8	1624.1	326	8378	1484
Neighboring municipalities (∗)	5.70	1.87	2	11	6
Neighboring branches (∗)	40.84	154.33	0	1122	11
HHI (province) (∗)	962.15	679.25	316.73	3456.79	773.74
p_i (∗∗)	0.2743	0.1217	0.0486	0.5861	0.2658
p_{PROV} (∗∗)	0.2086	0.0774	0.0493	0.3957	0.2296
ω_1 (∗∗)	0.0568	0.0206	0.0089	0.0970	0.0629
ω_2 (∗∗∗)	60.52	10.51	18.57	117.50	59.55
ω_3 (∗∗∗)	384.45	190.91	75.54	1196.63	344.99
(b) Duopolistic banks[b]					
Total assets	34824.7	22721.4	3539.1	107435.7	29095.8
Loans	15348.8	13736.4	1150.0	74935.0	10378.4
Risk capital	4820.9	4729.1	258.0	24071.0	3202.0
Total cost	2658.2	1563.3	259.0	6596.6	2203.1
Employees (∗)	9.45	5.13	2	29	9
Population (municipality) (∗)	3870.3	2740.2	992	20872	2811
Neighboring municipalities (∗)	5.87	2.08	2	10	6
Neighboring branches (∗)	25.78	75.16	3	677	13
Rival bank's branches (∗)	137.16	198.62	1	1194	70
HHI (province) (∗)	1005.75	613.47	316.73	3560.09	785.28
p_i (∗∗)	0.2510	0.1096	0.0526	0.5442	0.2436
p_{PROV} (∗∗)	0.1965	0.0682	0.0493	0.4316	0.2132
ω_1 (∗∗)	0.0570	0.0211	0.0095	0.0911	0.0647
ω_2 (∗∗∗)	59.41	11.03	14.00	97.53	58.60
ω_3 (∗∗∗)	496.25	259.43	109.27	1263.67	456.46

Total assets = customer deposits + interbank deposits + shareholders' equity + reserves.
p_i = interest revenue/total loans.
p_{PROV} = interest revenue/total loans (for banks operating in the same province).
ω_1 = interest expenses/total deposits.
ω_2 = labor costs/number of employees.
ω_3 = other operating costs/number of branches ≡ other operating costs.
All variables are expressed in millions of 2000 euro, except: (∗) units; (∗∗) ratios; (∗∗∗) thousands of 2000 euro.
[a] Number of observations: 381.
[b] Number of observations: 155.

depends on the type of loan: supposing that a bank wants to leave the market, assets like government bonds, interbank loans and credits to large-size firms may not be regarded as sunk costs, because the debtor's degree of solvency is known to the whole market; in contrast, credits linked to a guaranty as well as specific loans (especially those to small firms) are sunk costs to a much larger extent, because they imply a personal relationship between creditor and debtor whose specificity makes its transfer to other credit institutions, who do not know the exact associated degree of risk, much more difficult and costly (Coccorese, 2004, p. 212). Now, there is often a "size effect" in lending: large banks tend to lend to medium and large companies, while small banks usually specialize in lending to small businesses.[21] As a result, the information asymmetries, and the related costs of entry or exit, allow little local banks – no matter their size – to count on a limited market power, notwithstanding possible competitive pressures from banks operating in the bordering municipalities or even outside them (especially if competitors are large-size and can therefore rely on the possibility of enjoying scale economies, which may balance their poor territorial roots).

Another factor preventing a more substantial competitive behavior is the inertia of banks' clientele. Some surveys show that consumers rely on nearby institutions for their banking services (close to home or the workplace) and tend to cluster several financial products with their deposit account.[22] This pattern of behavior makes the local presence a crucial need for banks. On the other hand, even if a widespread branch network helps to increases banks' business (and also to reduce the risk of banking activity thanks to the geographic diversification), it nevertheless imposes higher costs. As a result, in those areas where other credit institutions decide not to open new offices, existing banks will benefit from the survival of their market power.[23]

Referring to the cost function, most coefficients are highly significant. Their estimated values indicate that banks are operating in a region of strong economies of scale: considering the sample means, the calculated average total cost (0.24) is higher than the marginal cost (0.04). This finding, which is consistent with the results of previous studies on cost banking functions, represents a plausible reason explaining why in many of these small markets additional banks chose not to enter.

[21] For an interesting discussion on this point, see Sapienza (2002, pp. 333–335).

[22] See Focarelli and Panetta (2002, p. 1169). This may also explain why the number of branches in many countries has continued to increase notwithstanding a decline of the number of banks.

[23] A general discussion on why locally owned banks may behave differently from non-locally owned banks, especially those operating in economically small areas, is provided by Collender and Shaffer (2003, pp. 29–31).

Table 3
Simultaneous equation model: estimation results.

	Whole sample		Years 1988–1992		Years 1993–2005		$HHI_{PROV}<773.74$		$HHI_{PROV}\geqslant 773.74$		North-Centre		South-Islands	
	Coeff.	*t*-Value	Coeff.	*t*-Value	Coeff.	*t*-Value	Coeff.	*t*-Value	Coeff.	*t*-Value	Coeff.	*t*-Value	Coeff.	*t*-Value
Demand equation (dependent variable: $\ln q_i$)														
$\ln p_i$	−0.8188	-13.68^{***}	−0.9332	-12.84^{***}	−0.6600	-9.23^{***}	−1.0791	-18.88^{***}	−0.4521	-5.02^{***}	−0.7851	-12.17^{***}	−0.9741	-4.23^{***}
$\ln p_{PROV}$	−0.0157	−0.13	0.0262	0.30	0.3544	1.84^{*}	0.0559	0.43	0.0695	0.43	−0.2639	-1.70^{*}	0.1881	0.69
$\ln POP$	0.3767	1.32	1.1697	29.55^{***}	1.4855	2.54^{**}	0.1647	0.78	−0.4291	−0.62	1.1454	9.11^{***}	1.1270	10.69^{***}
$\ln BR_{NEI}$	−0.2817	-2.08^{**}	−0.1238	−1.04	−0.2036	−0.97	0.2479	1.67^{*}	−0.2464	−1.44	−0.1989	−1.27	−0.7584	-2.36^{**}
Adj. R^2	0.9254		0.9616		0.9473		0.9619		0.9331		0.9190		0.9198	
Cost equation (dependent variable: $\ln C_i$)														
$\ln q_i$	0.0717	2.76^{***}	0.1454	2.94^{***}	0.0218	1.24	0.1380	3.72^{***}	0.0330	0.67	0.2506	4.86^{***}	−0.0073	−0.19
$(\ln q_i)^2$	0.0221	6.03^{***}	0.0051	1.76^{*}	0.0187	3.16^{***}	0.0039	1.63	0.0433	4.78^{***}	0.0181	3.88^{***}	0.0253	1.34
$\ln \omega_{1i}$	3.0954	6.48^{***}	−1.2982	-8.76^{***}	2.9878	2.92^{***}	2.1590	5.14^{***}	1.4229	1.89^{*}	−0.7998	-7.31^{***}	−1.3609	-5.67^{***}
$\ln \omega_{2i}$	−0.8167	-1.83^{*}	3.1708	5.30^{***}	0.1211	0.09	−0.5114	−1.35	1.3419	1.71^{*}	2.0697	6.27^{***}	3.5657	3.15^{***}
$\ln \omega_{3i}$	−1.2786	-5.00^{***}	−0.8726	-1.82^{*}	−2.1088	-4.16^{***}	−0.6476	-2.83^{***}	−1.7649	-2.97^{***}	−0.2699	−1.03	−1.2048	−1.33
$(\ln q_i)(\ln \omega_{1i})$	0.0094	2.88^{***}	0.0070	1.41	0.0066	2.20^{**}	0.0032	1.11	0.0159	2.24^{**}	0.0170	2.99^{***}	0.0076	1.01
$(\ln q_i)(\ln \omega_{2i})$	0.0085	2.48^{**}	−0.0045	−0.94	0.0128	2.75^{***}	−0.0010	−0.34	0.0240	2.90^{***}	−0.0037	−0.62	0.0171	1.23
$(\ln q_i)(\ln \omega_{3i})$	−0.0179	-5.32^{***}	−0.0025	−0.97	−0.0194	-3.08^{***}	−0.0022	−0.94	−0.0399	-4.40^{***}	−0.0133	-2.92^{***}	−0.0247	−1.32
$(\ln \omega_{1i})^2$	0.2800	5.97^{***}	−0.2512	-4.33^{***}	0.3274	5.13^{***}	0.2416	6.15^{***}	0.0366	0.48	−0.0463	-1.68^{*}	−0.2441	-3.70^{***}
$(\ln \omega_{2i})^2$	−0.2272	-2.42^{**}	−0.2440	-3.76^{***}	−0.5386	-1.92^{*}	−0.1950	-2.60^{***}	−0.2158	−1.18	−0.3675	-3.44^{***}	−0.0385	−0.14
$(\ln \omega_{3i})^2$	0.0465	1.16	−0.0081	−0.21	0.1333	1.78^{*}	−0.0168	−0.39	0.2390	2.91^{***}	−0.0207	−0.46	0.9419	4.01^{***}
$(\ln \omega_{1i})(\ln \omega_{2i})$	−0.1330	−1.52	0.8396	4.39^{***}	−0.0460	−0.24	−0.1120	-1.72^{*}	0.2464	2.08^{**}	0.4208	7.55^{***}	−0.2584	−1.52
$(\ln \omega_{1i})(\ln \omega_{3i})$	−0.2041	-6.16^{***}	−0.3721	-2.68^{***}	−0.2035	-5.19^{***}	−0.0993	-4.12^{***}	−0.3414	-4.01^{***}	−0.1256	-3.60^{***}	0.2997	2.76^{***}
$(\ln \omega_{2i})(\ln \omega_{3i})$	0.2377	4.92^{***}	0.0358	1.14	0.3273	3.26^{***}	0.1815	4.59^{***}	0.0352	0.41	0.1393	2.74^{***}	−0.7005	-3.79^{***}
Adj. R^2	0.9723		0.9874		0.9782		0.9822		0.9734		0.9650		0.9715	
Price-cost margin equation (dependent variable: p_i)														
θ	*0.6957*	13.67^{***}	*0.8208*	13.19^{***}	*0.5980*	9.19^{***}	*0.9401*	19.79^{***}	*0.3589*	4.94^{***}	*0.5981*	12.04^{***}	*0.8749*	4.20^{***}
Adj. R^2	0.9530		0.9209		0.9606		0.9720		0.9340		0.9647		0.8952	
H_0: $\theta = 0$ (χ^2 test)		186.80^{***}		173.86^{***}		84.45^{***}		391.64^{***}		24.38^{***}		145.04^{***}		17.67^{***}
H_0: $\theta = 1$ (χ^2 test)		35.74^{***}		8.28^{***}		38.16^{***}		1.59		77.77^{***}		65.51^{***}		0.36
Observations	381		163		218		189		192		314		67	

Systems have been estimated with three-stage least squares.
In the demand and cost equations a full set of dummy variables capturing time effects and bank effects is also added (coefficient estimates are not reported).
The instruments used are: levels and logs of p_{PROV}, POP, BR_{NEI}, HHI_{PROV}, ω_1, ω_2, ω_3, total assets, number of employees, equity, deposits, and average total cost; time dummies; bank dummies.
* Significance for the parameter estimates = 10% level.
** Significance for the parameter estimates = 5% level.
*** Significance for the parameter estimates = 1% level.

3.3. Sub-samples

An interesting investigation concerns the likely change of banks' conduct during time. Particularly, we aim to verify whether there have been modifications in their behavior after 1993, when Italy implemented the Second Banking Coordination Directive, after a long succession of steps in earlier years that intended to gradually deregulate the internal banking contexts. As a result, banks from European Union countries have been able to branch freely into other EU countries. This chance, combined with the outlook of a unique European capital market, has certainly affected the strategic policy of domestic banks throughout Europe, causing a vigorous concentration wave in banking markets: the bigger size has been regarded as the best option for better facing the toughened competition, achieving both scale and scope economies, and by and large improving overall efficiency.

In Italy, the number of banks has dropped from 1100 in 1988 to 784 in 2005, while in the same period the number of branches has increased from 15,363 to 31,501. Both figures drive to envisage an intensification of the competitive pressures on banks, as well as a widespread lessening of their market power. To check for this possibility with reference to our monopolistic banks, essentially by means of θ, we have split them in two sub-samples according to years: the first considers data from 1988 to 1992, the second spans from 1993 (when the Second Banking Directive became effective) to 2005.

The results are shown in the second and third columns of Table 3. The behavioral parameter drops from 0.8208 (years 1988–1992) to 0.5980 (years 1993–2005), both statistically different from 0 and 1. Hence, even if the banks included in our sample have remained monopolist in their municipality, they were forced to noticeably reduce the exploitation of market power. This is another signal of the effective potential competition that seems to characterize the Italian banking industry, in spite of the ongoing consolidation process.[24]

The comparison of the results also shows that there has been a fall in the price elasticity of demand, which was expected since the lower market power moves banks farther from the profit-maximizing equilibrium. Besides, while the coefficient related to BR_{NEI} is not significant and p_{PROV} has the predicted positive impact on q only for the most recent period of time (however at the 10% level), the variable *POP* shows now the expected positive coefficient, significantly different from 0 at least at the 5% level. It is worth to further notice that there is a remarkable drop in the estimated value of both the marginal cost (from 0.04 to 0.02) and the average cost (from 0.29 to 0.21), which corresponds to the fall in the average rate on loans (from 0.33 to 0.23), to be ascribed not only to the generalized reduction of nominal interest rates, but also to the strengthened competition. In general, our little monopolistic banks have answered the environmental changes with a more efficient cost structure.

It seems also interesting to evaluate the impact of local market concentration on monopolistic banks' behavior, and hence the role of nearby rival banks' dimension. We expect that, when concentration is high, it is likely that there are big banks in the neighborhood, which can count on their size to gain customers even at expense of smaller credit institutions. For the banks in our sample, the only way to react to this stronger competition is to behave in a more competitive fashion as well, so that the parameter θ should be smaller.

To assess the effects of market concentration, we split again the sample in two sections, one under and one above the median value of the Herfindahl–Hirschman index (*HHI*) that characterizes the province in which banks are situated[25] (this figure being equal to 773.74). The estimation results of the two systems are presented in the fourth and fifth columns of Table 3.

For banks that have operated as local monopolists in provinces with a lower *HHI* than the median value (i.e. in less concentrated markets), the behavioral parameter θ is 0.9401, not statistically different from 1: these banks have been therefore able to almost fully exploit their monopolistic position. Quite to contrary, where the provincial *HHI* has been higher than the median (i.e. in more concentrated markets), θ amounts to 0.3589, significantly different from both 0 and 1, and also much lower than before. So our evidence is that bigger banks, and (in a wider sense) the wave of mergers and acquisitions in the banking industry, appear to constitute a deterrent for the conduct of incumbents, even if the latter do not share with them the same market and are the only banks in the municipality. This result is consistent with the previous remark regarding the reduction of monopoly rents after 1993, and further emphasizes the pro-competitive role that the concentration process in banking can play, if properly supervised.[26]

About the other estimated parameters, we again note a conspicuous drop of the own-price demand elasticity, passing from −1.08 (greater than 1 in modulus, which is now compatible with a profit-maximizing behavior and hence with a value of θ close to 1) to −0.45, while the other coefficients are not significant (except for BR_{NEI}, which is positive in the first specification, even if significant only at the 10% level, surprisingly showing that in less concentrated areas the demand of loans for monopolistic banks grows with the number of rivals' neighboring branches).

Our framework allows the investigation of another important issue that characterizes Italy, namely the financial backwardness of Southern regions, where notable differences exist compared to the rest of the country, especially with reference to the thinness and the competitiveness of markets, the efficiency of intermediaries, and the cost and quality of credit provided (Galli and Onado, 1990). One striking feature is the interest rate differentials between areas: in the South and Islands they are constantly some points above those prevailing in the North and the Centre. Besides the difference in risk conditions, this gap is often attributed to the lack of competition among banks, which can be due to the imperfect and asymmetric information of firms in favor of banks and the excessive market segmentation (Faini et al., 1993; Messori and Silipo, 1997). However, in the 1990s the spread between Northern and Southern rates has markedly reduced. Some authors ascribe this phenomenon to the consolidation process of banks, which helped to improve their efficiency without harming Southern Italy's economy (Panetta, 2003), also if the loan dynamics may not have been able to adequately support Southern firms' growth (Mattesini and Messori, 2004, pp. 164–179).

To get some insights about local bank competition, we estimate our model for two different sub-samples: North-Centre and South-Islands (Table 3, sixth and seventh columns). The results show that θ is higher for Southern regions (0.8749 vs. 0.5981), also if it is statistically different from 1. Hence, in line with the bulk of the literature, there is evidence that in the period 1988–2005 bank competition has been tighter in Northern regions.[27] In addition, the average loan rate for the sample banks confirms to be higher in Southern regions (0.29 vs. 0.27), as is average total cost (0.26

[24] Our conclusion conforms with the evidence by Angelini and Cetorelli (2003, p. 669).

[25] Because of the lack of data on local deposits and loans, we have computed the HHI values on the basis of the geographical distribution of branch networks, conforming to the standard approach of the literature when disaggregated balance-sheet data are missing (e.g. Carbo Valverde et al., 2003; Degryse and Ongena, 2005). Relevant statistics for provincial *HHI* are given in Table 2a.

[26] Hannan and Prager (2009) also find that in the US banking industry small single-market banks operating in more highly concentrated rural banking markets earn higher rates of return.

[27] Coccorese (2004) finds a similar result for a shorter period of time (1997–1999).

vs. 0.24). Note that, as the coefficient of BR_{NEI} is negative and significant for the South and Islands, here bank loans appear to particularly suffer from neighboring competition, which is nonetheless unable to guarantee a degree of competition at least comparable with Northern regions. Finally, the variable *POP* has the expected positive coefficient, highly significant for both areas and slightly higher for the North and the Centre.

4. The Panzar–Rosse *H*-statistic for local monopolistic banks

4.1. Methodology

In order to assess the reliability of previous results, we use the same dataset to perform our second empirical test, the Panzar–Rosse *H*-statistic. As noted above, it is based on a reduced-form revenue equation to be estimated on cross-sectional data, and consists in the estimation of the elasticities of gross revenue with respect to each input price. Their sum, denoted as *H* and ranging between $-\infty$ and +1, assumes different values according to the structure of the market: it is negative for a monopolist, collusive oligopolist, or conjectural-variations short-run oligopolist[28]; is 1 for a perfectly competitive firm[29]; and lies between 0 and 1 for the case of monopolistic competition.

From an economic point of view, *H* represents the percentage variation of the equilibrium revenue originating from a unit percent variation in the price of all inputs of the firm. In case of monopoly, an increase of the factor prices moves the marginal cost curve upwards, also causing a reduction of the optimal output and hence of firm's revenues (since the profit-maximizing monopolist operates in the elastic part of the market demand), what means $H < 0$. For a firm in a perfectly competitive market, a given percentage growth in input prices must match with an equivalent percentage growth in revenues in order to pass along the entire cost increase, otherwise it would be forced to exit the market: hence $H = 1$.

Earlier studies have estimated the Panzar–Rosse statistic in the banking sector, both on single countries and on groups of nations.[30] To our knowledge, the only attempt to apply this test to a structural monopoly is performed by Shaffer (2002), who analyzes a single US bank with no office in other markets during a period of fifteen years. He is able to reject the hypotheses of static monopoly pricing and perfectly contestable pricing, while the results are consistent with a form of monopolistic competition. Since the market structure under his exam literally precludes this outcome, Shaffer maintains that the true explanation must lie in a form of behavior not yet formally linked to specific ranges of the Panzar–Rosse *H*-parameter.

Our context is very close to Shaffer's (2002), except the fact that we are dealing with panel data, and therefore with more banks (and also more observations), for which the estimation of the *H*-statistic can help to assess the significance of the results obtained so far.

For bank *i* at time *t*, the revenue equation to be estimated is the following:

$$\begin{aligned}\ln TR_{it} = {} & \alpha_1 \ln \omega_{1it} + \alpha_2 \ln \omega_{2it} + \alpha_3 \ln \omega_{3it} + \alpha_4 \ln TOTASS_{it} \\ & + \alpha_5 \ln LOANASS_{it} + \alpha_6 \ln CAPASS_{it} + \alpha_7 \ln MIX_{it} + \alpha_8 \\ & \times \ln POP_t + \alpha_9 \ln BR_{NEIt} + \lambda_i + \gamma_t + \varepsilon_{it}. \end{aligned} \tag{10}$$

[28] Shaffer (1983) shows that $H < 0$ also characterizes any profit-maximizing firm facing a fixed demand curve, even in a short-run but not long-run competitive equilibrium.

[29] It has been demonstrated that $H = 1$ describes a natural monopoly situation in a contestable market, or a firm that maximizes sales subject to a breakeven constraint. See Shaffer (1982).

[30] Some of them have been recalled in Section 2.

The dependent variable, *TR*, is total revenue, given by the sum of interest revenue and net non-interest income. This measure appears suitable so as to account for the remarkable increase of the revenue flow coming from fee-based products and off-balance-sheet activities, especially in recent years.

The factor prices coincide with those already used for the system estimation: hence, the price of deposits (ω_1) is the ratio between interest expenses and deposits, the price of labor (ω_2) is the ratio between labor costs and employees, and the price of physical capital (ω_3) is the value of net operating costs other than those associated to deposits and labor.

Additional control variables have been inserted to consider bank-specific factors as well as local features. Total assets (*TOTASS*) are usually regarded as a good proxy for the size of firms, and are positively correlated to revenues as long as larger banks enjoy scale economies in their activity. Unfortunately, we were not able to collect this figure for all banks in every year, so we have chosen to proxy total assets with the funds under management, calculated as the sum of customer deposits, interbank deposits, shareholders' equity and reserves.[31] The coefficient of the ratio between loans and total assets (*LOANASS*) is also expected to be positive: even if loans represent the most risky asset for banks, they guarantee the highest gross yield, so a higher proportion of loans should generate greater revenues. The ratio of risk capital to total assets (*CAPASS*) could exhibit either a negative or positive sign: on one side, less equity reflects a more risk-taking behavior, greater leverage and so higher interest revenues; on the other side, banks could want to set aside more risk capital just to balance more risky loans and investments, and this would imply a positive coefficient.

The variable *MIX* is calculated as the share of interest revenues over total revenues, and is added to control for the business mix. We suppose it is negatively related to gross revenue: if banks can not have a full control over loan rates, those with a greater portion of interest revenues (hence a lower share of non-interest income) will enjoy less total revenue, as a consequence of their poor ability to develop kinds of business different from credit intermediation. The municipality population (*POP*) is added as a proxy for the size of local market demand, and should therefore be positively correlated with revenues. Finally, BR_{NEI} is again the number of branches belonging to the neighboring banks: being an indicator of outside competition, we expect a negative coefficient.

All firm-specific and time-varying factors that could affect the level of total revenues, but not explicitly addressed in our specification, are captured through the insertion of dummy variables associated to banks and years, and denoted as λ_i and γ_t, respectively.

4.2. Results

Table 2a again contains summary statistics for the relevant variables, while Table 4 reports the results of our panel estimations. Since we are interested in comparing the value of the *H*-statistic with the level of the behavioral parameter θ, we have performed seven regressions, each corresponding to the sample of observations already considered for the estimation of the simultaneous-equation model. To account for local level shocks, we estimate robust Huber–White standard errors clustering observations by province: this allows to control for possible within-zone correlation of error terms over time.

For the total sample of monopolistic banks (first column), a positive (and significant) correlation between revenues and input prices (ω_1, ω_2 and ω_3) emerges. Particularly, the coefficient on

[31] Referring to the available observations, the mean of the funds under management was only 0.07% lower than total assets.

Table 4
Panzar–Rosse *H*-statistic for the monopolistic banks: estimation results.

	Whole sample		Years 1988–1992		Years 1993–2005		HHI_{PROV} < 773.74		$HHI_{PROV} \geqslant$ 773.74		North-Centre		South-Islands	
	Coeff.	*t*-Value	Coeff.	*t*-Value	Coeff.	*t*-Value	Coeff.	*t*-Value	Coeff.	*t*-Value	Coeff.	*t*-Value	Coeff.	*t*-Value
$\ln\omega_{1i}$	0.4398	8.78***	0.3901	2.13**	0.5063	8.92***	0.3132	17.63***	0.5136	7.44***	0.4044	8.38**	0.3865	4.31***
$\ln\omega_{2i}$	0.0702	2.64**	0.0389	0.92	0.1779	3.63***	0.0898	125.39*	0.0477	0.61	0.0620	2.01*	0.1345	1.79
$\ln\omega_{3i}$	0.1058	3.06***	0.0571	1.08	0.1229	3.94***	0.0804	1.60	0.0738	1.34	0.0701	1.99*	0.1360	2.26*
$\ln TOTASS_i$	0.8277	26.35***	0.8518	6.56***	0.7582	24.15***	0.8227	23.73***	0.8708	17.48***	0.8781	28.50***	0.7241	25.87***
$\ln LOANASS_i$	0.1286	2.49**	0.0577	1.66	0.0943	1.16	0.2036	4.43***	0.1118	2.22**	0.1355	2.59**	0.1957	4.90***
$\ln CAPASS_i$	0.0367	0.68	0.1853	1.97*	−0.0071	−0.17	0.1503	5.38***	−0.0450	−1.32	0.1213	2.91***	−0.0915	−1.28
$\ln MIX_i$	−0.7405	−15.25***	−0.0275	−0.62	−0.9669	−16.11***	−0.6401	−153.17***	−0.7486	−10.45***	−0.7046	−22.25***	−0.7839	−12.51***
$\ln POP$	−0.1361	−1.89*	−0.0637	−0.32	−0.2646	−1.39	−0.0797	−0.97	0.0393	0.15	−0.0649	−0.44	−0.2944	−0.63
$\ln BR_{NEI}$	−0.0986	−1.74*	−0.0403	−0.54	−0.1046	−1.47	−0.0609	−0.32	−0.0863	−1.05	−0.1635	−1.77*	−0.0387	−0.81
H	0.6158		0.4861		0.8071		0.4834		0.6351		0.5365		0.6570	
H_0: *H* = 0 (*F*-test)		77.75***		6.46**		144.47***		208.05***		29.50***		109.62***		88.35***
H_0: *H* = 1 (*F*-test)		30.27***		7.22**		8.26*		237.56***		9.74***		81.81***		24.07**
Adj. R^2	0.9856		0.9945		0.9861		0.9823		0.9908		0.9833		0.9938	
DW statistic	1.84		2.85		1.73		1.61		2.23		1.73		2.45	
Observations	381		163		218		189		192		314		67	

Dependent variable: $\ln TR_i$.
t-Statistics are based on robust Huber–White standard errors (with observations clustered by province).
All regressions include bank and year intercepts (coefficient estimates are not reported).
* Significance for the parameter estimates = 10% level.
** Significance for the parameter estimates = 5% level.
*** Significance for the parameter estimates = 1% level.

the price of deposits is considerably larger than the other two. This result may depend on the fact that interest expenses usually represent the main part of total costs: in our sample, in the period 1988–2005 they amounted to 59% of overall expenses (21% for labor, and 20% for the other inputs), so that a stronger association with revenues is acceptable. Furthermore, interest rates on both loans and deposits usually move jointly, with the result that the coefficient of ω_1 should be positive in the revenue regression. The sign could be negative only if the demand for loans were elastic (as an increase in the factor price would reduce both the provision of loans and the level of revenues, in spite of the concurrent increase of the price of loans), but the previous results indicate that for our sample monopolistic banks it was actually inelastic over the sample period.[32]

Turning to the Panzar–Rosse *H*-statistic, its point estimate amounts to 0.6158. The *F*-tests strongly reject (at the 1% level) the hypothesis that it is non-positive as well as the hypothesis that it equals unity. Thus, we can rule out the possibilities of monopoly conduct, conjectural-variations equilibrium, pure sales maximization, perfect competition or perfect contestability for this sample, and get a confirmation of the fact that banks in our sample are unable to sustain a profit-maximizing monopolistic position in static equilibrium. Along with the theory underlying the Panzar–Rosse statistic, the estimated value of *H* might result from a form of monopolistic competition, but our banks do not face any direct competitor (at least in the municipality), so we have to look for another explanation of the outcome.

One possibility is that banks adopt a limit pricing behavior to deter entry. It is likely that this policy entails a conduct lying between monopoly and competitive pricing. Actually, the need to hold their prices below the monopoly level forces banks to pass along higher input prices only partially to their customers as higher output prices. Thanks to the inelastic demand curve, loans fall of a lower percentage, so that revenues increase, implying $H > 0$. All the above requires that input prices are exogenous to the bank, while output prices and revenues adjust endogenously.

In other words, limit pricing can give reasons for both $\theta < 1$ and $0 < H < 1$. Our behavioral parameter θ reproduces the actual difference between price and marginal cost for a given set of input prices: in order to deter entry, this gap must be small, because a limit price must lie near enough to marginal cost. Furthermore, since the *H*-value measures how revenues change according to modifications in input prices, it is possible that an equilibrium limit price approximately varies with input prices, what guarantees the desired differential over marginal cost (Shaffer and DiSalvo, 1994, p. 1077).

According to Rosse and Panzar (1977), we can rule out the hypothesis of cost-plus pricing behavior, as it requires a unit value of *H*. Besides, our scenario is not even compatible with a contestable market, a situation that would still generate $H = 1$: this proves that potential competition, although existing in our framework (as the negative coefficient of the variable BR_{NEI} in the above demand equation confirms, even though significant at the 10% level), cannot be fully effective. Again, we ascribe this to the informational problems that new entrants could face, as well as to the customers' inertia.[33]

Concerning the other explanatory variables, *TOTASS* and *LOANASS* take a positive and significant coefficient. As expected, a bigger bank size and a higher share of loans play an important role in gen-

[32] See also Shaffer (2002, pp. 228–230). It should be noted that other previous banking studies employing the Panzar–Rosse test have found a similar pattern for the coefficients of input prices.

[33] In his analysis regarding one single monopolistic bank over time, Shaffer (2002) discusses other hypotheses that can explain a behavior different from profit maximization: expense preference behavior, monopsony power, declining market.

Table 5
Panzar–Rosse H-statistic for the monopolistic banks: equilibrium test.

	Whole sample		Years 1988–1992		Years 1993–2005		HHI_{PROV} < 773.74		$HHI_{PROV} \geqslant 773.74$		North-Centre		South-Islands	
	Coeff.	t-Value	Coeff.	t-Value	Coeff.	t-Value	Coeff.	t-Value	Coeff.	t-Value	Coeff.	t-Value	Coeff.	t-Value
$\ln \omega_{1i}$	−0.0014	−0.29	−0.0001	−0.01	0.0026	0.37	0.0048	1.51	−0.0016	−0.23	0.0043	0.87	0.0074	0.53
$\ln \omega_{2i}$	−0.0032	−0.92	−0.0003	−0.05	−0.0064	−1.03	0.0009	0.39	−0.0020	−0.24	−0.0035	−0.89	−0.0133	−2.74**
$\ln \omega_{3i}$	−0.0245	−3.38***	−0.0113	−1.53	−0.0303	−3.50***	−0.0116	−4.90***	−0.0343	−5.06***	−0.0145	−4.52***	−0.0362	−4.98***
$\ln TOTASS_i$	0.0232	4.15***	−0.0044	−0.35	0.0320	3.27***	0.0264	8.15***	0.0224	3.70***	0.0120	3.19***	0.0470	5.45***
$\ln LOANASS_i$	0.0046	1.89*	−0.0020	−0.38	0.0008	0.26	0.0002	0.39	0.0074	1.36	0.0024	1.68	0.0141	0.92
$\ln CAPASS_i$	0.0361	3.26***	0.0228	2.07**	0.0498	4.68***	0.0245	5.34***	0.0471	6.61***	0.0164	3.31***	0.0642	14.22***
$\ln MIX_i$	−0.0187	−2.81***	−0.0073	−1.37	−0.0230	−2.71**	−0.0250	−59.23***	−0.0140	−0.84	−0.0257	−6.43***	−0.0260	−2.18*
$\ln POP$	0.0280	1.40	−0.0317	−1.18	0.0311	1.34	0.0046	0.74	0.0037	0.10	0.0272	1.52	−0.0092	−0.14
$\ln BR_{NEI}$	−0.0090	−1.33	−0.0182	−2.02*	0.0165	2.26**	−0.0043	−1.64	−0.0160	−1.41	−0.0105	−1.06	0.0280	2.54**
H	−0.0291		−0.0117		−0.0342		−0.0059		−0.0379		−0.0137		−0.0421	
H_0: H = 0 (F-test)		8.23***		0.46		14.39***		3.74		7.83***		3.26*		6.32**
Adj. R^2	0.6399		0.6710		0.7410		0.6689		0.6961		0.6018		0.8559	
DW statistic	1.93		2.69		1.83		2.10		2.16		1.84		2.07	
Observations	381		163		218		189		192		314		67	

Dependent variable: $\ln(1 + ROA_i)$.
t-Statistics are based on robust Huber–White standard errors (with observations clustered by province).
All regressions include bank and year intercepts (coefficient estimates are not reported).
* Significance for the parameter estimates = 10% level.
** Significance for the parameter estimates = 5% level.
*** Significance for the parameter estimates = 1% level.

erating revenues. In contrast, the ratio of equity to total assets (*CAPASS*) does not influence banks' income, as its coefficient is not statistically different from zero. The negative sign of the variable *MIX* (significant at the 1% level) is also in line with our expectations, and proves that banks that are able to diversify their business earn more revenue.

The signs of the last two variables, population (*POP*) and neighboring branches (BR_{NEI}), are significant only at the 10% level. Hence, the evidence is that the size of local market demand does not particularly affect revenues (moreover, contrary to our conjecture, the negative sign would indicate that higher revenues are associated to a smaller population), and that the stronger outside competition appears to be only partially harmful for local monopolistic banks' income.

Shaffer (1982) has demonstrated that, if the sample is not in long-run equilibrium, it can negatively skew the Panzar–Rosse H-statistic. To test this hypothesis, he has proposed to estimate the value of H by using *ROA* (return on assets) as the dependent variable instead of revenues: the sum of elasticities of *ROA* with respect to input prices will be zero for a sample in long-run equilibrium, but not otherwise. Here, so as to adjust return on assets for possible small negative values due to banks' losses in any year, we compute it as $1 + ROA$ (Claessens and Laeven, 2004; Casu and Girardone, 2006).

The results of the new estimation are reported in Table 5 (first column). The value of H_{ROA} is negative and significantly different from zero. Hence, there is evidence of disequilibrium, with $H_{ROA} < 0$ meaning that an exogenous increase in input prices reduces banks' *ROA*: credit institutions are not able to pass along their higher costs fully to its borrowers and other asset-side customers without losing some patronage (Shaffer, 2004, p. 307). The lack of long-run equilibrium should be not surprising, given the dynamic nature of the banking industry and especially the structural and regulatory environment that has characterized the sample period. However, this indicates that the rejection of $H = 1$ might be spurious for these banks. On the other hand, the rejection of $H < 0$ in the revenue equations is valid also when the sample is not in long-run equilibrium.[34] Accordingly, our empirical rejection of the monopoly power for local monopolistic banks is robust to equilibrium considerations. There could be the possibility that their conduct is approximately competitive (or contestable), even in presence of the statistical rejection of $H = 1$, but the empirical evidence of the simultaneous-equation model seems to discard this conclusion.

4.3. Sub-samples

In analogy with the previous analysis, to examine how banks' conduct changed in the sample period we estimate the Panzar–Rosse statistic for the sub-intervals 1988–1992 and 1993–2005. In Table 4, the second and third columns exhibit the estimation results for the two sub-samples, where H grows from 0.4861 to 0.8071. The first value is significantly different from both 0 and 1, while the second is indistinguishable from 1 at the 5% level. The clear indication is that the environment has become much more competitive, again corroborating the analogous evidence coming from the estimation of the behavioral parameter θ: the enhanced potential competition has constrained banks not to take advantage from their territorial market power.

It is worth to note also that for the period 1988–1992 the variable *MIX* was not significant: banks did not appear to care about business mix. Later, in coincidence with a more competitive landscape, they have been forced to look for other sources of in-

[34] Only the results concerning the perfect and the monopolistic competition models rely crucially on the assumption that firms are observed in long-run equilibrium. See Panzar and Rosse (1987, p. 447).

come in addition to the flow originated from credit intermediation.[35]

We can make use of the Panzar–Rosse test also to evaluate how our single-branch monopolistic banks undergo the level of market concentration characterizing the province where they are located. For the purpose, and in analogy with a previous inspection, we perform two estimations of (10): one considers the observations for which $HHI \leqslant 773.74$ (the median), the other those for which $HHI > 773.74$. Table 4 displays the results of the regressions (fourth and fifth column).

In less concentrated markets, it is $H = 0.4834$, a lower value than more concentrated provinces, where $H = 0.6351$. It has been recognized that H is a decreasing function of the Lerner index, and therefore of market power, for a pure monopoly (Panzar and Rosse, 1987, p. 446), and also that, under stronger assumptions (particularly, a constant price elasticity of demand across markets), a continuous interpretation of H may be correct.[36] According to that, we discover that once more (but not surprisingly), the outcome is in line with the estimated behavioral parameter, and also with our expectations: higher concentration implies bigger credit institutions in the neighborhood, which surely impose a more competitive conduct to smaller banks, reflected by a higher H-statistic. However, now the difference between the competition index of the two sub-samples is less marked, and especially there is evidence of an average (rather than strong) exploitation of market power when $HHI \leqslant 773.74$. We also note that, with the only exception of *CAPASS*, the estimated coefficients of the control variables do not show particular differences across the samples.[37]

Finally, we have estimated the H-statistic also for the two macroregions, North-Centre and South-Islands (see Table 4, sixth and seventh columns). In this case, the Panzar–Rosse competition index is slightly higher for Southern areas (0.6570 vs. 0.5365, both significantly different from zero and one), an evidence that contrasts with what we got from the system estimation. However, the equilibrium test for the same groups of banks (Table 5, sixth and seventh columns) shows that for South-Islands $H_{ROA} < 0$: hence, this sub-sample is not in long-run equilibrium, implying that the Panzar–Rosse statistic could yield spurious indication of market power (Shaffer, 2004, p. 298).

5. The Panzar–Rosse *H*-statistic for local duopolistic banks

We have dealt with local monopolistic single-branch banks so far, and their behavior seems nonetheless far from a full exploitation of market power. What happens when two banks operate in local markets? If monopolies are characterized by behaviors that mimic a sort of monopolistic competition, one would expect that two banks are enough to give rise to an almost fully competitive environment.

Our last inspection has consisted in finding, for the same time interval, a set of banks similar to those already analyzed (hence, small single-branch credit institutions) but operating in municipalities where another branch was present, no matter if the competitor was a small or big bank. As Table 1b makes clear, the sample coverage in the reference period, its time pattern and the territorial characteristics are comparable to those of monopolistic banks. The new sample comprises 155 observations regarding 44 duopolistic banks all over Italy. The summary statistics for this group (shown in Table 2b) are useful for drawing a picture of their average characteristics, and allow also a comparison with the monopolistic banks. Particularly, we observe that on average their total assets and loans are about 50% higher, and that they employ more workers: therefore, duopolistic banks are generally bigger. Besides, they operate in municipalities where population is about twice as much, and where their rival is normally a big bank (its overall branches being 137 on average, 70 if we consider the median value). Finally, their loan rates are lower than monopolistic banks, with whom they share a comparable level of unit deposit and labor costs, while the cost of capital is higher.

To assess the degree of competition of this supplementary sample, we cannot rely on a system estimation anymore. Actually, we should measure a conjectural variation (rather than a behavioral) parameter now, because we are dealing with a duopoly; for the purpose, we would need municipal disaggregated data also for each rival of our banks, what is impossible to infer from their balance sheets, since they are essentially multi-branch institutions.

On the other hand, we can apply the Panzar–Rosse procedure, employing only the balance data concerning the single-branch banks. Clearly we are going to consider only a portion of market supply, but we trust that our sample banks are nevertheless able to disclose the average degree of market competition. Moreover, we are especially interested in evaluating their own conduct, in order to make a close parallel with the monopolistic single-branch banks, the foremost focus of our analysis, and to figure out whether conduct in local banking monopolies is less competitive than in markets with more than one firm.

For the duopolistic bank i, the revenue equation we are going to estimate is the following:

$$\begin{aligned}\ln TR_{it} = {} & \alpha_1 \ln \omega_{1it} + \alpha_2 \ln \omega_{2it} + \alpha_3 \ln \omega_{3it} + \alpha_4 \ln TOTASS_{it} \\ & + \alpha_5 \ln LOANASS_{it} + \alpha_6 \ln CAPASS_{it} + \alpha_7 \ln MIX_{it} + \alpha_8 \\ & \times \ln POP_t + \alpha_9 \ln BR_{NEIt} + \alpha_{10} \ln BR_{RIVt} + \lambda_i + \gamma_t + \varepsilon_{it}. \qquad (11)\end{aligned}$$

It reflects (10), plus a new variable, BR_{RIV}, representing the total number of branches managed in the whole country by the rival bank. We regard this as a variable measuring the size of the competitor, and expect that it is negatively correlated to revenues: a bigger rival can count on its dimension, possible scale economies, and also a more widely known name, to capture more customers at expense of bank i.

We have run five regressions: one contemplates all observations, two consider the sample according to different time intervals (as before, 1988–1992 and 1993–2005, respectively), and the last two split the single-branch banks according to the size of the other bank in the market, measured through the variable BR_{RIV}: in this case, the first sub-sample is composed by those observations for which it is $BR_{RIV} \leqslant 70$ (i.e. the median), the second by the remaining data.[38]

The regression results are presented in Table 6. For the whole sample (first column), $H = 0.8076$, not significantly different from 1 at the 5% level. In spite of the local duopoly, our banks behaved as if under perfect competition, or even in a contestable market. It follows that, while little monopolistic banks enjoy some market power, this is not possible when there is even one competitor in the same municipality. Of course, we do not mean that in all areas with more than one bank perfect competition prevails (especially when dealing with large-size banks, this occurrence may depend on factors that could have no direct relation with the specific territorial context), but only that small banks are unable to exercise all the potential market power, and that the latter seems to disappear at all when they cannot act as monopolists.

[35] The test on *ROA* shows the presence of long-run equilibrium only for the period 1988–1992 (see Table 5).

[36] See Vesala (1995, p. 56) and Bikker and Haaf (2002, p. 2203). Shaffer (1983) has shown that H is an increasing function of the Bresnahan-Lau conduct parameter for an oligopolist facing a fixed demand curve.

[37] Table 5 makes clear that the long-run equilibrium characterizes only banks for which the Herfindahl-Hirschman index is below the median value.

[38] Due to lack of sufficient observations for Southern regions, we are not able to estimate the macroregional H-statistic for duopolistic banks.

Table 6
Panzar–Rosse *H*-statistic for the duopolistic banks: estimation results.

	Whole sample		Years 1988–1992		Years 1993–2005		$BR_{RIV} \leqslant 70$		$BR_{RIV} > 70$	
	Coeff.	*t*-Value	Coeff.	*t*-Value	Coeff.	*t*-Value	Coeff.	*t*-Value	Coeff.	*t*-Value
$\ln \omega_{1i}$	0.5874	6.04***	0.4711	12.55***	0.5591	4.42***	0.5002	9.16***	0.5631	2.79**
$\ln \omega_{2i}$	0.0594	0.89	−0.0368	−1.83*	0.2188	1.65	0.0375	0.68	0.2227	0.95
$\ln \omega_{3i}$	0.1609	1.26	0.0130	0.43	0.3367	1.92*	−0.0232	−0.59	0.3405	1.61
$\ln TOTASS_i$	0.7873	7.50***	0.9968	10.62***	1.0564	4.90***	0.8281	34.08***	0.9833	4.45***
$\ln LOANASS_i$	0.1082	0.73	−0.0458	−0.87	0.1891	0.98	0.0357	1.47	0.3056	1.03
$\ln CAPASS_i$	0.0855	0.99	0.3068	4.53***	0.5922	3.69***	0.1203	1.77*	0.4463	1.95*
$\ln MIX_i$	−1.0753	−8.89***	−0.5237	−12.18***	−1.0722	−7.48***	−0.9616	−6.51***	−1.3699	−3.26***
$\ln POP$	−0.4736	−0.50	−0.5822	−3.72***	−0.9611	−0.72	−0.1093	−0.48	−1.4447	−0.60
$\ln BR_{NEI}$	0.1776	0.63	−0.0915	−1.04	0.4293	0.57	0.0994	0.62	0.7949	1.28
$\ln BR_{RIV}$	−0.0936	−2.87***	−0.0073	−0.64	−0.0589	−0.97	−0.0006	−0.04	−0.0606	−0.79
H	0.8076		0.4472		1.1145		0.5145		1.1263	
H_0: *H* = 0 (*F*-test)		17.90***		61.75***		17.39***		28.89***		14.37***
H_0: *H* = 1 (*F*-test)		1.02		94.33***		0.18		25.72***		0.18
Adj. R^2	0.9880		0.9988		0.9841		0.9969		0.9833	
DW statistic	2.23		3.34		2.14		2.86		2.12	
Observations	155		75		80		78		77	

Dependent variable: $\ln TR_i$.
t-Statistics are based on robust Huber–White standard errors (with observations clustered by province).
All regressions include bank and year intercepts (coefficient estimates are not reported).
* Significance for the parameter estimates = 10% level.
** Significance for the parameter estimates = 5% level.
*** Significance for the parameter estimates = 1% level.

Table 7
Panzar–Rosse *H*-statistic for the duopolistic banks: equilibrium test.

	Whole sample		Years 1988–1992		Years 1993–2005		$BR_{RIV} \leqslant 70$		$BR_{RIV} > 70$	
	Coeff.	*t*-Value	Coeff.	*t*-Value	Coeff.	*t*-Value	Coeff.	*t*-Value	Coeff.	*t*-Value
$\ln \omega_{1i}$	0.0161	2.99***	0.0058	0.31	0.0190	2.73**	0.0118	1.33	0.0165	1.65
$\ln \omega_{2i}$	−0.0042	−0.81	−0.0123	−3.06***	−0.0043	−0.63	−0.0053	−1.07	−0.0030	−0.32
$\ln \omega_{3i}$	−0.0038	−0.63	−0.0016	−0.23	−0.0035	−0.40	−0.0096	−1.84*	−0.0021	−0.21
$\ln TOTASS_i$	0.0110	1.29	0.0063	0.51	0.0479	3.18***	−0.0018	−0.47	0.0350	1.89*
$\ln LOANASS_i$	0.0070	1.08	−0.0162	−0.99	0.0097	1.62	−0.0024	−0.27	0.0096	0.73
$\ln CAPASS_i$	0.0048	0.47	0.0382	3.55***	0.0515	2.89**	0.0184	2.02*	0.0315	1.29
$\ln MIX_i$	−0.0833	−11.73***	−0.0760	−3.20***	−0.0733	−10.41***	−0.1134	−4.60***	−0.0638	−2.56**
$\ln POP$	−0.0581	−1.09	−0.0735	−1.85*	−0.1132	−1.55	−0.0355	−1.01	−0.1526	−1.10
$\ln BR_{NEI}$	0.0071	0.71	−0.0135	−1.12	0.0357	1.00	0.0025	0.11	0.0187	0.54
$\ln BR_{RIV}$	−0.0035	−2.36**	−0.0054	−0.83	−0.0034	−0.83	−0.0051	−2.50**	−0.0013	−0.38
H	0.0081		−0.0080		0.0113		−0.0032		0.0114	
H_0: *H* = 0 (*F*-test)		0.85		0.21		0.46		0.07		0.47
Adj. R^2	0.6889		0.7834		0.7694		0.8124		0.5600	
DW statistic	2.28		3.08		2.30		2.49		2.27	
Observations	155		75		80		78		77	

Dependent variable: $\ln(1 + ROA_i)$.
t-Statistics are based on robust Huber–White standard errors (with observations clustered by province).
All regressions include bank and year intercepts (coefficient estimates are not reported).
* Significance for the parameter estimates = 10% level.
** Significance for the parameter estimates = 5% level.
*** Significance for the parameter estimates = 1% level.

Regarding the individual variables, the coefficient of ω_1 is still positive, significant and quite large, while ω_2 and ω_3 are not statistically different from zero. The variables *TOTASS* and *MIX* confirm also their sign and significance, the latter being much greater in magnitude, what implies that the revenue of duopolistic banks is more sensitive to the business mix, perhaps for the need to respond to competition by searching for new sources of income. This could also be the reason why *LOANASS* loses its significance in this sample, while *CAPASS* and *POP* are still indistinguishable from zero at the 5% level. Finally, BR_{RIV} shows a negative and highly significant coefficient: as expected, single-banks' revenues are lower when their competitor in the municipality is a big bank.

Comparing the sub-samples before and after 1993 (second and third columns of Table 6), we find that *H* = 0.4472 in the former, a value statistically different from both 0 and 1, while *H* = 1.1145 in the latter, not significantly different from 1. Competition in our local markets has increased during time, in coherence with the conclusion drawn from the analysis of monopolistic banks.

Grouping banks according to the number of the rival's branches (fourth and fifth columns of Table 6), we get *H* = 0.5145 and *H* = 1.1263 for those under and above the median value of BR_{RIV}, respectively. Both are different from zero, so we rule out any monopoly equilibrium, but in the second sub-sample banks appear to behave as perfectly competitive firms, since *H* is not distinguishable from 1. Yet again, we get evidence that sizeable competitors force small banks to set lower prices.

In Table 7 we report the results of the equilibrium test for the various samples of duopolistic banks (with 1 + *ROA* as dependent variable). In all of them we cannot reject the hypothesis that *H* = 0, so now the Panzar–Rosse statistics can be always meaningfully interpreted.

6. Conclusions

Latest empirical studies have shown that market concentration in the banking industry does not necessarily reduce the level of competition, thus contrasting with conventional industrial organization models based on SCP paradigm, for which in concentrated industries operating firms can exercise a significant market power. This is an important indication, given the key function of banks in the economy.

This paper has applied two NEIO techniques to a sample of Italian single-branch banks that operated as monopolists in small local areas during the period 1988–2005, with the aim to assess pricing conduct in highly concentrated banking markets. The findings provide evidence that market power may be small even in markets with only one bank, confirming that in this sector concentration and competition can coexist. Particularly, we have found that, in spite of the advantageous condition, monopolistic banks are able to exploit only partially their market power, whose estimated level is far from the typical monopoly conduct. Significant factors that favor this outcome appear to be the nearby competition, the recent wave of banking consolidation and the local presence of big banks. Moreover, the same elements compel banks to have a significantly competitive conduct when local markets are duopolies.

Acknowledgements

I am grateful to Jacob Bikker, Jaap Bos, Santiago Carbo Valverde, Barbara Casu, Nicola Cetorelli, Claudia Girardone, Joaquin Maudos, Alfonso Pellecchia, Sherrill Shaffer, an anonymous referee and the participants at the Rome XVII Tor Vergata Conference for helpful hints and conversations. Any remaining errors are my responsibility.

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