

Assessing the Competitive Conditions in the Italian Banking System: Some Empirical Evidence^{*}

PAOLO COCCORESE

1. Introduction

The European banking industry is currently facing the effects caused by the integration of the different national financial markets. This integration is a result of the elimination of the controls on capital movements agreed in 1989 by the European Union (EU) Council of Ministers, and the regulations introduced in 1992 in accordance with the Second Banking Directive, which allowed free access of banks in a given country to other domestic markets within the EU. In order to adapt to this new operational context caused by the adoption of the above legislative measures, the Italian financial system has undergone considerable transformation.

It is perhaps early to give a comprehensive evaluation of the impact of financial integration on each country. However, it is of prime importance to verify whether there have either been significant gains in efficiency for the consumer – derived from the enhancement of competition¹ – or that the imperfections of the financial markets suggest that stronger competition cannot simply result from unification² alone. In any case, such judgements presuppose the assessment of competitive conditions within each singular banking market before

□ Università di Salerno, Dipartimento di scienze economiche, Fisciano (Italy).

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¹ This is the view of Price Waterhouse (1988).

² On this point, see Branson (1990), Neven (1990) and Vives (1991).

integration, for which, in general, the studies draw attention to indicators such as the number of banks, the number of branches, and the concentration of assets and deposits.

This paper aims to evaluate the degree of competition in the Italian banking industry, analysing a sample of banks during the period 1988-96 thanks to the methodology first proposed by Rosse and Panzar. The next Section delineates the main features of the banking market in Italy, recalling the legislative and structural evolution over recent years. Section 3 is devoted to the theoretical description and the empirical implications of the Rosse-Panzar test. A brief review of some previous studies employing the same methodology to assess the degree of competitiveness within various banking industries is given in Section 4, while in Section 5 we present and discuss our application to the Italian banking system. Our conclusions are given in the last Section.

2. The structure of the Italian banking system: characteristics and evolution

Over the last twenty-five years there has been a gradual fall in the number of banks operating in Italy, from 1102 in 1973 to 1085 in 1982, finally reaching a value of 1043 in 1991. In contrast to this, the total number of branches has increased markedly from 11276 to 12853, and attaining 19080 for the respective years. These phenomena confirm the tendency to more concentrated markets already in action since several years. In particular, at the end of the last decade the financial markets underwent many significant changes, most notable in the demand for loans and the prospect of new services. These changes, coupled with the prospect of the elimination of intra-EU barriers, drove banks to undergo organisational rearrangements that could make them able to bear the challenges of their rivals in this new era. Consequently, many bank mergers have occurred, which were not confined to the smaller banks but also involved the larger banks. This was followed by the expansion of their territorial network, which is shown by the growth in the number of branches, and is linked to the liberalisation occurring in 1990.

These rearrangements were justified because of the characteristics traditionally shown by the Italian banking industry. In general these banks were small-size firms with higher operating costs than the other EU banks, and with differences between the credit and debit rates that were higher than the EU average. In fact, during the last decade the gross operating profit of the Italian banks was one of the highest amongst the industrialised countries, which was mainly due to the gross interest income, again greater than the other European credit institutions. In stark contrast, items such as the cost of labour and the number of branches – that can be regarded as good indicators of management efficiency – ranked the Italian banking system below that of nations like France, Germany and the United Kingdom.³

The reason for this situation is linked to the structure of industrial production in Italy, whose main feature is the existence of many small firms, and only a few large private/public firms. The smaller firms tend to have a family basis, and usually only have economic relations with a single bank, where they have limited contractual power. The larger firms have economic relations with several banks, making it impossible for the latter to have total control over such customers, and often producing insurance-type behaviour, which have a negative impact on the level of bank credit rates.⁴

In 1990 two important laws were issued in Italy: the first (no. 218/90) encouraged bank mergers, while the second (no. 287/90) designated the job of monitoring the impact of these mergers on the inter-bank competition to the central bank. However, some observers emphasised the difficulties for the Bank of Italy to reach simultaneously the objectives of concentration, which is needed to increase the average firm size in the banking sector and hence its efficiency, and competition, which constitutes a corollary to the abolition of intra-EU bounds. These points of view originate from the idea that there exists a trade-off between concentration and competition, which is itself based on the hypothesis that the banking industry follows the Stigler oligopolistic model.⁵ This important contribution supports the structure-conduct-performance (SCP) paradigm⁶ in its structuralist

³ See Parrillo (1989, pp. 45-46). See also Passacantando (1983) and Conti (1989).

⁴ See Desario (1995, pp. 121*-23**).

⁵ See Stigler (1964).

⁶ The structure-conduct-performance paradigm holds that the degree of competition in an industry (performance) can be explained in terms of conduct of firms (de-

version, where the performance of the market is strongly linked to its structure, and therefore to the market concentration. It is well known that the theory of contestability⁷ does not accept this paradigm, maintaining that, in order to have competition, it is sufficient to guarantee the possibility of both free entry to and exit from the industry, with no regard to the number of incumbent firms, since potential competition is able to reduce or even remove any monopoly power.

Both the above laws of 1990, and the adoption of the Second Banking Directive⁸ of 1993 – authorising commercial banks to act as ‘universal banks’ operating freely in almost every branch of the financial market – has hastened the search for more efficient industrial configurations. The result of this reorganisation of the Italian credit system has been a sharp growth in the number of concentration processes to obtain economies of both scale and scope. In fact, from 1984 to 1986 the average number of operations per annum aiming toward concentration was 7 (without considering those involving co-operative credit banks), which increased to 14 in the period 1987-89, and then to 26 for the period 1990-94.⁹ This means that the adoption of the new national legislation – followed by Italian banks also when operating abroad – has imposed the improvement in the efficiency of the banks. It should be noted that the current level of bank credit rates is closer to government bond rates than it was ten years ago: this change can be interpreted as a signal of increased competition within the Italian banking system, as well as proof that the removal of some of the restraints to competition has implied a stronger incentive to concentration. In other words, the increase in efficiency that a higher size guarantees seems to have produced a fall in prices and a better quality supply for customers. This result should not be a surprise, be-

gree of collusion, innovation, advertising, etc.) and that in turn the conduct depends on the structural characteristics of the market (number and size of firms, cost and demand conditions). In the last twenty years there has been a general abandonment of the paradigm in industrial economics literature, due mainly to its low flexibility (a crucial role is played by the structural variables that are considered exogenous, so that the behaviour of firms cannot modify them) and to its insufficient explanatory capability from the empirical point of view.

⁷ See Baumol, Panzar and Willig (1982).

⁸ The First Directive on credit institutions was adopted by the European Commission in 1977. For further details, see Molyneux, Lloyd-Williams and Thornton (1994).

⁹ See Desario (1995, pp. 125*-26*).

cause this market is also open to foreign banks, which contribute to higher competition.

Given this continuously evolving picture – especially in the last period – it seems to be appropriate to try an evaluation of the degree of competition that typifies the Italian banking market. In particular, it would be useful to assess whether the increased concentration of recent years implies a move in the market performance towards a less competitive configuration (according to the SCP paradigm), or if larger sizes are compatible with an increase in competition (as the theory of contestability asserts). These questions can be answered by estimating the Rosse-Panzar H-statistic, which is able to offer a measure of competitive conditions within the Italian banking market.¹⁰

3. Empirical assessment of competitive conditions

The comparative statics analysis suggested by Panzar and Rosse¹¹ for the identification of market power is based upon the estimation of a reduced form revenue equation, when considering that the total revenue is easily observable, unlike the price and quantity. For a single firm, the equilibrium total revenue is given by the equilibrium quantity times the equilibrium price. Both the equilibrium quantity and price depend on the cost, demand and conduct, and therefore in the revenue function all the shifters of cost and demand must be included, with particular attention given to the factor prices. For the i -th firm, the following reduced form revenue equation can be written:

$$R_{it} = f(W_{it}, Z_{it}, Y_t, E_t),$$

where W_{it} represents the factor prices, Z_{it} are the other variables which shift the cost function, Y_t are the variables that shift the demand function, and E_t is the error term.

If R_{wk} is the derivative of the total revenue with respect to the k -th input, the Rosse and Panzar H-test can be written as:

¹⁰ An exhaustive survey on the econometric models that allow the identification of market power is given by Bresnahan (1989).

¹¹ See Rosse and Panzar (1977), and Panzar and Rosse (1987). See also Bresnahan (1989, pp. 1034-39).

$$H = \sum_k (R_w W_{it} / R_{it}),$$

i.e. the sum of the elasticities of the reduced form revenue with respect to all the factor prices. Hence, for each firm the calculation of the H-statistic requires only specific data on the revenues and factor prices (the information on costs are not required), although it requests the insertion of *all* the variables shifting demand or cost. In particular, the H-statistic is linked to the variable R_{it} , which – as we have already seen – can be estimated in many cases when the structural equations of the model (supply and demand) cannot.

The Rosse and Panzar test often has a clear interpretation when applied to the study of the markets. For example, when the market under study is a monopoly, then the value of the H-statistic is less than 0. Intuitively, H represents the percentage variation of the equilibrium revenue derived from a unit percent increase in the price of all factors used by the firm. Clearly, an increase of 1% in the price of each factor produces a 1% increase in marginal costs. Economic theory tells us that revenues in monopoly fall when there is an increase of marginal costs, and H measures the percentage fall in the equilibrium revenues associated with the 1% increase in costs. The authors have also stressed that it is not only the sign of the H-index which is important, but also its magnitude.¹² The above result can also be generalised to the case when a monopolist has more than one choice of variable, e.g. it produces more than one good or makes use of advertising. For similar reasons, H is less than 0 also for the case of a perfectly colluding oligopoly or a homogeneous conjectural variations oligopoly.¹³

If we now consider the case when the observed firm is in a symmetric perfectly competitive market in long-run equilibrium – where all firms produce a quantity equal to Q^* , corresponding to the minimum point of the long-run average cost curve – then $H = 1$. This should also not be a surprise. If all factor prices rise by 1%, the average cost will shift upward by also 1%, leaving its minimum point unchanged. In the long-run equilibrium the price p^* must be always equal to the minimum level Q^* of average cost, which remains the long-run equilibrium quantity, and therefore it is necessary that the price – and hence the total revenue – increases by the same percent-

¹² See Panzar and Rosse (1987, p. 446).

¹³ These demonstrations can be found in Panzar and Rosse (1987).

age, so that the H value remains unity. Shaffer (1982) has shown that the H-statistic is also equal to one for a natural monopoly operating in a market which is perfectly contestable, as well as for a sales-maximising firm that is subject to break-even constraints.

Concerning the equilibrium in a symmetric monopolistic competition market,¹⁴ we know that firms set their output where perceived marginal revenue equals marginal cost, and that the possibility of entry and exit related to the existence of positive or negative short-run profits causes zero economic profits in the long run. Given these circumstances, a general factor price increase shifts upward the average and marginal cost curves and reduces the optimal level of production. This results in losses for the operating firms, induces the exit of some producers and, because of the reduction in global supply, shifts upward the demand curve of the other firms until a new tangency occurs between the price and average cost curves. It follows that the sum of elasticities of the total revenue with respect to factor prices, i.e. the H value, is less than or equal to one.

In summary: in the case of monopoly H is non-positive; for the (symmetric) long-run competitive equilibrium $H = 1$; the (symmetric) Chamberlinian equilibrium is identified by $H \leq 1$.

From an econometric point of view, the rejection of the $H \leq 0$ hypothesis rules out the monopoly model; the rejection of the hypothesis $H \leq 1$ excludes all the above three models; and the rejection of both the $H \leq 0$ and the $H = 1$ hypotheses (but not the $H \leq 1$ hypothesis) implies that of the models so far examined only the monopolistic competition model could be consistent with the data.¹⁵

It should be further noted that the results concerning both the perfect and Chamberlinian competition models rely crucially on the assumption that firms are observed in long-run equilibrium,¹⁶ whereas the monopoly case does not. To test this hypothesis, one can suppose that competitive markets equalise rates of return across firms, so that in equilibrium these rates of return should not be significantly correlated with input prices. Therefore to test for equilibrium it is sufficient to calculate the Rosse-Panzar H-statistic using the return on as-

¹⁴ See Chamberlin (1962).

¹⁵ See Panzar and Rosse (1987, p. 453).

¹⁶ See Panzar and Rosse (1987, p. 447).

sets as the dependent variable in place of the total revenue in the regression equation. A value of $H < 0$ would show non-equilibrium, whereas $H = 0$ would prove equilibrium.¹⁷

In any case, it is useful to remember that the adoption of such a methodology for the study of a given industry, which is based on a partial equilibrium analysis, requires that given variations in the equilibrium price or quantity are not significantly influenced by and have not a significant influence on the price levels in other markets. However, this condition appears to be rather unrealistic, as an economic system is characterised by the division of labour, and so each industry is connected to many other markets for the acquisition of the necessary inputs. In an ideal situation it would therefore be better to consider the exchange values between the involved industries.¹⁸

4. Previous studies

The first application of the H-test was by Rosse and Panzar (1977) using linear regression on a cross-section of data in order to estimate the H-statistic for the newspaper firms in the local media markets. In the analysis they rejected the hypothesis that newspapers were monopolies even when they were the only newspaper in the market. In fact, the empirical findings showed that the industry behaved as if it were competitive, and the authors ascribed this to the role and importance of competition from other media.¹⁹

In the banking market there has only been a few sporadic applications of the Rosse-Panzar methodology. Shaffer (1982) used this approach to examine the behaviour of a sample of banks in New York. In his analysis, the total revenue was explained by variables such as the unit price of labour, capital and funds, together with other variables which were supposed to affect long-run equilibrium bank reve-

¹⁷ See Shaffer (1982).

¹⁸ Sraffa (1926) was one of the first to stress this problem. He suggested to solve it through the simultaneous determination of the relative prices and the distributive variables for the whole economic system.

¹⁹ In Sullivan (1985) and Ashenfelter and Sullivan (1987), an extension of the Rosse-Panzar analysis is offered in connection with the possibility that variables other than revenue are observable.

nues, e.g. total bank assets. His estimation gave the value of H to be 0.318. The test to verify the long-run equilibrium – performed by substituting the total revenues by the return on assets (ROA) – produced a negative value of H, but not significantly different from zero. Therefore, these results suggested that the banks behaved neither as monopolists nor as perfectly competitive firms in long-run equilibrium, and that the forces preventing monopolistic conduct were primarily potential rather than actual.

Nathan and Neave (1989) used a similar procedure to analyse Canadian banks in the years between 1982 and 1984, and also trust companies and mortgage companies.²⁰ Concerning the commercial banks, for 1982 they estimated a value of $H = 1.058$ – not significantly different from unity – while for 1983 and 1984 the index values were 0.680 and 0.729 respectively, both of which were significantly different from 0 and 1. They concluded that it was possible to reject both the monopoly hypothesis, and – except in 1982 – the perfect competition hypothesis: hence, the banking revenues behaved as if earned under monopolistic competition.

Another application to the banking industry was made by Molyneux, Lloyd-Williams and Thornton (1994), based on a model similar to those of Shaffer and Nathan and Neave. Here the Rosse-Panzar test was performed on a sample of German, British, French, Italian and Spanish banks for the period 1986-89, in order to offer an appraisal on the degree of integration of the banking markets within the EU. The results showed values of H which were significantly different from zero and one for Germany (except for 1986), France, Spain, and United Kingdom. In contrast, the H-statistic for the Italian banks between 1987 and 1989 was always negative and significantly different from zero – the 1986 data were missing. Hence, it was impossible to reject the monopoly or conjectural variations short-run oligopoly hypotheses for the sample of Italian commercial banks under consideration. It should also be stressed that the 1988 data for Italy did not represent long-run equilibrium values.

²⁰ On this estimation, see also the note by Perrakis (1991) and the reply by Nathan and Neave (1991).

5. Application to the Italian banking market: method, results and discussion

In this paper we aim to evaluate the degree of competition in the Italian banking industry using the same approach as employed by the aforementioned studies. Table 1 shows the number of commercial banks in the sample for each year. Their accounting data have been observed during the period 1988-96. We can therefore follow the progress of the H-index over the nine years under inspection: this enables a rather homogeneous comparison between the years and allows for an evaluation of the trend of the data over time. Note that the analysis by Molyneux, Lloyd-Williams and Thornton – the only comparable study – also considered a sample of Italian banks whose dimension varied from year to year.

Every year, the banks of the sample have been classified into three different groups, according to the funds under management: large-size banks (8 of these banks can be regarded as major, since the funds under their management are always notably larger than the other banks), medium-size banks and small-size banks. The reader can refer to Table 1 for further details.

The role of these banks within the Italian banking system can be deduced from Table 2. The sample of banks under study holds approximately a 75% share of total deposits, and 80% of total loans. The figures also show that from 1993 these percentages grew substantially as a consequence of a series of mergers and acquisitions which involved the national credit system from that year; this concentration is to be regarded as one of the main results of the new laws aiming to enhance the European integration.

Accounting data were provided by the Italian weekly *Milano Finanza*.

In the calculation of the Rosse-Panzar H-statistic, the sample data were used to estimate the following revenue equation:

$$\ln TR = a + b \ln PF + c \ln PL + d \ln PK + e \ln DEP + f \ln ASS + g \ln CAPASS + h \ln LNASS + i \ln BR + j D8,$$

TABLE 1

CHARACTERISTICS OF THE SAMPLES
(size in thousand billion lire)

Year	Large banks		Medium banks		Small banks		Total
	Size	Number	Size	Number	Size	Number	
1988	F > 4.1	41	1.25 < F < 4.1	64	F < 1.25	61	166
1989	F > 4.5	41	1.37 < F < 4.5	62	F < 1.37	66	169
1990	F > 5.0	41	1.50 < F < 5.0	64	F < 1.50	62	167
1991	F > 5.5	38	1.65 < F < 5.5	56	F < 1.65	70	164
1992	F > 5.7	42	1.70 < F < 5.7	58	F < 1.70	70	170
1993	F > 6.2	42	1.85 < F < 6.2	58	F < 1.85	68	168
1994	F > 6.3	44	1.87 < F < 6.3	44	F < 1.87	69	157
1995	F > 6.4	43	1.89 < F < 6.4	54	F < 1.89	57	154
1996	F > 6.6	43	1.94 < F < 6.6	52	F < 1.94	57	152

TABLE 2

MARKET CONCENTRATION IN THE ITALIAN BANKING SECTOR
(figures in %)

Year	Deposits			Loans		
	5-firm	10-firm	Sample	5-firm	10-firm	Sample
1988	23.4	38.7	79.0	29.1	46.1	81.8
1989	22.9	38.0	78.6	29.4	46.2	84.3
1990	22.7	37.8	75.4	29.7	48.0	84.8
1991	23.8	37.4	69.4	29.4	46.0	77.6
1992	22.8	35.6	68.8	28.5	40.8	78.7
1993	23.7	36.7	72.0	30.5	45.4	77.7
1994	26.3	40.5	76.5	32.4	49.1	83.0
1995	28.1	42.8	81.4	36.6	52.8	89.6
1996	26.9	42.1	81.9	34.8	50.9	88.4

where:

ln	= natural logarithm
TR	= total revenue
PF	= interest expenses/total deposits (proxy for unit price of funds)
PL	= personnel expenses/number of employees (proxy for unit price of labour)
PK	= other operating costs/number of branches (proxy for unit price of capital)
DEP	= total deposits
ASS	= administered funds
CAPASS	= risk capital/administered funds
LNASS	= loans/administered funds
BR	= number of branches/total number of branches
D8	= dummy variable (related to the amount of administered funds): 1 for the eight largest banks, 0 for all the other banks.

The nature of the estimation of the H-statistic means that we are especially interested in understanding how the total revenue reacts to variations in the cost figures, and for this reason the dependent variable is given by the sum of all the revenues, including the interest revenues.

The independent variables introduced to explain the variations in total revenue are similar to those used in other studies. In particular, the unit price of funds is calculated by considering the interest expenses for each lira received as a deposit, and the unit price of labour is computed as the labour cost for each employee. The latter proxy is the same as used by Shaffer (1982) and Nathan and Neave (1989), but different to the one employed by Molyneux, Lloyd-Williams and Thornton (1994), where the unit price was taken as the personnel expenses divided by the bank assets. In determining the unit price of capital, we have taken into consideration the value of all the operating costs minus those related to funds and labour, ensuring that the resulting amount is a good proxy for the general costs; this figure has then been divided by the number of branches, obtaining the average cost per branch. This ratio (already used by Nathan and Neave) appears to be a more appropriate indicator of the unit price for the capital input

rather than the ratio between capital costs and the value of fixed assets (as used by Shaffer and Molyneux, Lloyd-Williams and Thornton) particularly because premises and equipment may be frequently either rented or leased.

Besides the variables representing the prices of the various inputs to a commercial bank, some additional variables have been included in the estimation to take account of other characteristics. Total deposits (*DEP*) are a proxy for the aggregate demand, and should positively affect total revenues. The administered funds (*ASS*) are included to identify possible scale economies; its sign will be either positive or negative depending on whether the differences between the banks, due to the size of the funds under management, lead to higher or lower revenues. The risk capital to administered funds ratio (*CAPASS*) and the loans to administered funds ratio (*LNASS*) are included to account for firm risk: the coefficient of the former is expected to be negative as a lower level of risk capital should lead to higher bank revenues, while the coefficient of the latter is expected to be positive because a higher fraction of loans on the total funds under management also envisages greater revenues. The ratio between number of branches of a single bank and the global number of branches (*BR*) represents another proxy useful for evaluating the effect of the bank size on its revenues. Finally, the dummy variable *D8* is added to distinguish Italy's eight largest banks from the others: if administered funds and branches are sufficient to explain any size effect, its coefficient should not be significant; otherwise, if there exists an oligopoly power associated with their large size, *D8* should be significant.

The Rosse-Panzar tests were performed cross-sectionally thanks to separate estimations on the sample of data for the Italian commercial banks during the years 1988 through to 1996. The results are shown in Table 3a.

The sign of the coefficients for the price of the input factor proxies are always positive (with the exception of *PL* in 1993, which is negative but not significantly different from zero, and in 1995, which is also negative but different from zero at 5% level) and statistically significant. The data shows that the estimated values of H for the nine years are always significantly non-negative. They are also significantly different from unity, with the exclusion of 1992 and 1994, when this statistic does not differ significantly from unity. Therefore, the results indicate that the Italian banks have essentially

TABLE 3A
REGRESSION RESULTS FOR THE ITALIAN COMMERCIAL BANKS (1988-96)

Year	Constant	lnPF	lnPL	lnPK	lnDEP	lnASS	lnCAPASS	lnLNASS	lnBR	D8	R ²	H
1988	2.3052 (6.40) ^a	0.5358 (15.69) ^a	0.1822 (4.30) ^a	0.1488 (8.79) ^a	0.5351 (15.45) ^a	0.2506 (6.69) ^a	0.0971 (6.90) ^a	0.0224 (1.16)	0.2139 (12.03) ^a	-0.0308 (-1.21)	0.998	0.8668 285.74 ^{ab} 6.7 ^{bc}
1989	1.8472 (5.50) ^a	0.4925 (14.30) ^a	0.2074 (5.78) ^a	0.1385 (8.47) ^a	0.4701 (12.58) ^a	0.3264 (8.14) ^a	0.1023 (7.06) ^a	0.0342 (2.08) ^b	0.1941 (11.38) ^a	-0.0327 (-1.48)	0.999	0.8384 292.5 ^{ab} 10.9 ^{bc}
1990	2.3557 (7.82) ^a	0.4838 (14.34) ^a	0.1908 (5.57) ^a	0.1780 (11.15) ^a	0.4447 (11.82) ^a	0.3125 (7.89) ^a	0.0522 (3.73) ^a	0.0430 (2.27) ^b	0.2361 (14.20) ^a	-0.0145 (-0.65)	0.999	0.8526 271.8 ^{ab} 8.1 ^{bc}
1991	2.9520 (9.31) ^a	0.4984 (13.35) ^a	0.0518 (1.39)	0.1641 (8.06) ^a	0.4800 (11.78) ^a	0.2844 (6.57) ^a	0.0729 (4.43) ^a	0.0079 (0.50)	0.2242 (11.40) ^a	0.0071 (0.27)	0.998	0.7142 152.5 ^{ab} 24.4 ^{bc}
1992	3.1747 (10.67) ^a	0.7526 (32.43) ^a	0.0612 (2.81) ^a	0.1526 (6.10) ^a	0.7559 (23.95) ^a	0.0291 (1.09)	0.0657 (3.24) ^a	-0.0412 (-1.50)	0.1933 (8.37) ^a	-0.0320 (-0.91)	0.997	0.9664 1081.9 ^{ab} 1.3 ^c
1993	2.7509 (7.73) ^a	0.4344 (11.99) ^a	-0.0407 (-1.11)	0.1789 (8.87) ^a	0.5555 (12.08) ^a	0.2323 (4.91) ^a	0.0613 (4.86) ^a	0.0429 (2.23) ^b	0.2117 (11.10) ^a	-0.0103 (-0.47)	0.999	0.5726 97.2 ^{bc} 54.1 ^{bc}
1994	2.7392 (5.90) ^a	0.6129 (17.53) ^a	0.1241 (2.28) ^b	0.1991 (7.36) ^a	0.6510 (18.06) ^a	0.1121 (2.67) ^a	0.0902 (4.98) ^a	-0.0059 (-0.24)	0.2236 (8.66) ^a	-0.0025 (-0.08)	0.997	0.9361 216.1 ^{ab} 1.0 ^c
1995	1.4778 (2.86) ^a	0.2662 (4.99) ^a	-0.0575 (-2.12) ^b	0.1117 (4.35) ^a	0.2584 (4.80) ^a	0.5826 (9.80) ^a	0.0431 (3.15) ^a	0.0288 (1.38)	0.1414 (5.64) ^a	0.0094 (0.33)	0.998	0.3204 31.0 ^{bc} 139.5 ^{bc}
1996	1.0840 (2.71) ^a	0.3091 (6.58) ^a	0.1442 (3.49) ^a	0.1742 (9.00) ^a	0.3211 (6.57) ^a	0.4825 (8.72) ^a	0.0594 (4.54) ^a	0.0401 (2.09) ^b	0.1836 (10.42) ^a	-0.0411 (-1.66) ^c	0.998	0.6275 91.6 ^{bc} 32.3 ^{bc}

Dependent variable: lnTR.

t-statistics for the parameter estimates in parentheses.

^a coefficient estimates significant at 1% level.

^b coefficient estimates significant at 5% level.

^c coefficient estimates significant at 10% level.

^d F-statistics for testing the hypothesis H=0 (* = significantly different from 0 at 5% level).

^e F-statistics for testing the hypothesis H=1 (* = significantly different from 1 at 5% level).

operated under monopolistic competition between 1988 and 1996, rejecting both the monopoly hypothesis and (except in 1992 and 1994) the perfect competition hypothesis.²¹ This is compatible with the hypothesis that the market is contestable when it can be assumed that potential competition is able to guarantee a level of price set by the incumbent firms, which is close to the competitive level – given that higher prices would generate *hit-and-run* behaviour from potential firms.

The previous results seem to be highly compatible within the Italian context, remembering that in Italy the role and strength of a bank must be evaluated with reference to the markets where it works. Actually, there are few banking services in Italy where the reference market is the whole national area. In the majority of cases, the reference market for banks is much more limited. The data show that the local banking markets are mainly oligopolies, where the most powerful firms are generally small-size banks. For example, in more than a half of the Italian provinces only two banks are sufficient to concentrate one half of the deposits, while in another third of provinces only three are necessary.²²

The local forms of oligopoly can be explained by considering the possibility that some assets could not be recovered in their entirety, and hence are sunk costs.²³ In fact, it can be verified that a bank credit represents a sunk cost depending on its category. For example, if 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, since the debtor's degree of solvency is known to the whole market; in contrast, the credits linked to a guaranty as well as the specific loans – especially if they have been given to small firms – can be considered as 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 – very difficult and costly. We can conclude that there is strong competition between banks for

²¹ For 1988 and 1989, our result sharply contrasts with the one by Molyneux, Lloyd-Williams and Thornton (1994), whose H-statistic is negative and significant. This is probably due to the proxies used for the price of labour and capital, which have been computed quite differently (see before).

²² See Padoa-Schioppa (1995, pp. 55*-56*).

²³ See Di Battista and Grillo (1988, pp. 436-57).

the credits of large firms due to the absence of sunk costs, while on the other hand the information asymmetries, and related costs of exit, persuade a bank operating in a local market – no matter its size – to reduce its competitive pressure on the other rivals, or even to come to a collusive agreement with them.

In any case, it is believable that the competitive pressures are also quite strong in local concentrated markets, especially those from the largest banks because of their dimension and the resulting possibility of enjoying scale economies, which are often able to balance their little territorial roots.²⁴

The picture emerging from the previous discussion on local oligopolies could induce an expectation of a negative value of H , since we have already seen that in these conditions an upward shift of the marginal cost curve produces a fall in the equilibrium output and therefore a fall in the total revenues. But the comparative statics approach underlying the Rosse-Panzar test helps to make clear that, in spite of a group of banks which behave as monopolists in their local operating area, there exist competitive forces determining a situation where the Chamberlinian competition prevails on a national basis, and therefore contributing to qualify the Italian banking industry as a monopolistic competitive market (or as a contestable market, if *hit-and-run* behaviours were possible due to the absence of sunk costs).²⁵

Concerning the other variables, we can observe that the variable DEP has the expected positive sign, which is highly significant in all regressions. The sign of ASS is always positive and statistically significant (with the only exception of 1992), allowing us to state that the differences between banks based on the size of administered funds lead to bigger revenues for the large-size banks. The coefficient of $CAPASS$ has a positive sign, and therefore contrasts with our expectation. The variable $LNASS$ has a positive coefficient – except in 1992 and 1994 – and confirms the direct relationship between loans and revenues (also if it is not significantly different from zero in five of the nine estimations). Another significant variable is BR , whose positive sign proves that the wide diffusion over the country has an im-

²⁴ Starting from these considerations, Di Battista and Grillo (1988) use the theory of contestability to analyse the role and extent of competition in Italian banking industry. On related arguments, see also Coccoresse (1995).

²⁵ This scenario partially recalls the analysis of the American daily newspaper industry made by Rosse and Panzar (1977).

portant role on the level of revenues. The estimation also indicates that the dummy variable $D8$, which is never significant, with the exception of 1996, produces no additional explanatory power, i.e. for the eight largest banks there are no significant revenue effects, other than those linked to their size (measured in terms of administered funds and relative number of branches).

Finally, the long-run equilibrium test for the value of H , performed by using the return on assets (ROA) as the dependent variable (see Table 3b), shows that for four years (1988, 1989, 1990 and 1992) the data are in long-run equilibrium, since it is not possible to reject the hypothesis that $H = 0$: therefore, for these years the Rosse-Panzar test can be meaningfully interpreted.

In Table 4 some tests on the regressions results are shown. For the revenue equations they prove the absence of heteroscedasticity in the data, with the exception of the years 1992, 1993 and 1994; furthermore, the residuals appear normally distributed in four of the nine regressions, and the results of the test aiming to verify the correct specification for the functional form has fully supported our choice of the linear model. The same tests have given worse results when estimating ROA.

6. Conclusions

This paper has tried to assess the latest tendency of competitive conditions in the Italian banking industry by utilising the Rosse-Panzar test for the years between 1988 and 1996, a period of time characterised by noticeable changes in banking legislation and therefore in the internal organisation of credit institutions.

The empirical evidence has indicated that Italian banks earn revenues as if they were under conditions of monopolistic competition. This result is compatible with the limited average dimension of banks, but appears to be in opposition with the features of the local banking markets, where situations of concentrated oligopoly prevail. A possible explanation for this contrast could be due to the potential competition of the large-size banks – with a low presence in the territory, but able to enjoy large scale economies – towards the small-size

TABLE 3B

REGRESSION RESULTS FOR THE ITALIAN COMMERCIAL BANKS (1988-96)

Year	Constant	lnPF	lnPL	lnPK	lnDEP	lnASS	lnCAPASS	lnLNASS	lnBR	D8	R ²	H
1988	0.8474 (0.47)	0.2773 (1.62)	-0.1567 (-0.74)	0.0699 (0.82)	0.5940 (3.41) ^a	-0.6810 (-3.62) ^a	0.7319 (10.35) ^a	-0.0944 (-0.97)	0.1095 (1.23)	-0.1685 (-1.31)	0.807	0.1905 0.6 ^d
1989	-1.9796 (-0.94)	-0.1266 (-0.58)	0.0269 (0.12)	0.0157 (0.15)	0.1507 (0.64)	-0.1736 (-0.69)	0.8279 (9.09) ^a	-0.2486 (-2.40) ^b	0.0574 (0.53)	-0.3500 (-2.52) ^b	0.690	-0.0840 0.1 ^d
1990	-2.0113 (-1.22)	-0.2386 (-1.29)	0.0738 (0.39)	0.1039 (1.19)	-0.1792 (-0.87)	0.0718 (0.33)	0.5794 (7.55) ^a	-0.1242 (-1.19)	0.1169 (1.28)	-0.2994 (-2.45) ^b	0.536	0.4163 0.1 ^d
1991	1.6941 (0.77)	-0.0176 (-0.07)	-0.7796 (-3.01) ^a	-0.1930 (-1.36)	-0.1169 (-0.41)	0.1712 (0.57)	0.8801 (7.69) ^a	-0.0165 (-0.15)	-0.0321 (-0.23)	-0.2828 (-1.55)	0.460	-0.9902 6.1 ^{d*}
1992	-3.4464 (-2.60) ^b	-0.3809 (-3.69) ^a	0.0288 (0.30)	0.0984 (0.88)	-0.2824 (-2.01) ^b	0.2501 (2.10) ^b	0.7003 (7.77) ^a	0.3791 (3.10) ^a	0.0026 (0.03)	-0.1914 (-1.23)	0.677	-0.2537 3.8 ^d
1993	-2.1117 (-1.15)	-0.3146 (-1.68) ^c	-0.9302 (-4.92) ^a	-0.1993 (-1.92) ^c	0.2079 (0.88)	0.0023 (0.01)	0.4183 (6.43) ^a	0.1344 (1.36)	-0.1923 (-1.96) ^b	-0.3494 (-3.09) ^a	0.479	-1.4441 23.2 ^{d*}
1994	-5.8314 (-1.60)	-0.7577 (-2.78) ^a	-0.8330 (-1.96) ^b	-0.3951 (-1.83) ^b	-0.5440 (-1.96) ^b	0.9078 (2.79) ^a	0.7409 (4.94) ^a	0.0614 (0.32)	-0.3483 (-1.71) ^c	-0.5165 (-2.08) ^b	0.410	-1.9858 16.1 ^{d*}
1995	-7.1682 (-2.35) ^b	-0.5976 (-1.92) ^b	-0.1144 (-0.71)	-0.2473 (-1.59)	-0.3588 (-1.13)	0.6056 (1.73) ^c	0.7258 (8.09) ^a	0.1793 (1.41)	-0.3024 (-2.02) ^b	-0.1723 (-0.99)	0.600	-0.9593 7.8 ^{db*}
1996	-7.8041 (-3.43) ^a	-0.8358 (-3.10) ^a	-0.2625 (-1.11)	-0.2132 (-1.91) ^c	-0.8027 (-2.86) ^a	1.0610 (3.32) ^a	0.5727 (6.60) ^a	0.2882 (2.67) ^a	-0.2853 (-2.80) ^a	-0.3451 (-2.36) ^b	0.502	-1.3142 11.8 ^{d*}

Dependent variable: lnROA.

t-statistics for the parameter estimates in parentheses.

^a coefficient estimates significant at 1% level.

^b coefficient estimates significant at 5% level.

^c coefficient estimates significant at 10% level.

^d F-statistics for testing the hypothesis H=0 (* = significantly different from 0 at 5% level).

TABLE 4

RESIDUALS AND SPECIFICATION TESTS: RESULTS

Year	Heteroscedasticity (Ramsey test) ^a		Normality test (Jarque-Bera statistic) ^b		Functional form test (Ramsey's RESET test) ^c	
	lnTR	lnROA	lnTR	lnROA	lnTR	lnROA
1988	0.059*	48.187	12.310	170.110	0.569*	24.110
1989	1.221*	71.013	1.333*	135.040	0.001*	0.342*
1990	0.278*	0.604*	0.305*	20.012	0.003*	1.657*
1991	3.749*	68.508	7.023	2664.71	0.078*	3.252*
1992	4.472	1.871*	7.455	40.818	0.147*	12.167
1993	5.233	2.289*	5.691*	2.292*	0.470*	3.354*
1994	5.909	5.357	5.413*	21.824	0.197*	2.044*
1995	0.030*	4.766	8.581	14.229	0.844*	3.158*
1996	3.076*	3.957	8.908	7.489	0.261*	2.803*

^a To verify the presence of heteroscedasticity in the errors, we employ the Ramsey test (Ramsey 1969). Square residuals are regressed on the squared fitted values, then an F test (with 1 and $N-2$ degrees of freedom) is performed to check for the significance of the coefficient of this regression. If $F(1, N-2)$ is lower than the corresponding theoretical value $F_{0.95}$ (the cases satisfying this condition are marked with an asterisk) we are unable to reject the hypothesis H_0 that the errors are homoscedastic at the 5% significance level.

^b The Jarque-Bera (JB) statistic tests whether the residuals of a series are normally distributed. Its value is calculated in the following way: $JB = \frac{N}{6} \left[S^2 + \frac{(K-3)^2}{4} \right]$, where N is the number of observations, S is the skewness and K is kurtosis. Under the null hypothesis of normality of errors, JB has a chi-square distribution with 2 degrees of freedom. Thus, we reject this hypothesis at the 5% level if $JB > 5.99$ (the values compatible with this occurrence are marked with an asterisk).

^c The Ramsey's RESET test (Ramsey 1969) is used in this context to detect errors in the specification of the functional form. Actually, their presence produces a non-zero mean value for the errors. The test, based on an augmented regression that considers also the square of the fitted values, has an F distribution with 1 and $N-k-2$ degrees of freedom. Here, if $F(1, N-k-2)$ is greater than the corresponding theoretical value $F_{0.95}$, we must reject the null hypothesis that the functional form is correctly specified at the 5% significance level (the asterisk signals the cases when this hypothesis cannot be rejected).

banks, which have a much higher average market share at a local level. The tendency to concentration, which was traditionally considered with worry, following the liberalisation of European financial markets, appears not to have had a significant influence on the Italian banking industry conduct, since the data reject the hypotheses of both monopolistic and oligopolistic behaviour.

However, it must be underlined that the years which have been considered in this analysis refer to an interval of time when the proc-

ess of integration had only just started and therefore could not be regarded as anywhere near concluded at all. In the examined years the banking industry was still far from the equilibrium, which it was moving towards after the initiation of the European integration process, an equilibrium that requires more time in order to be attained.

REFERENCES

- ASHENFELTER, O. and D. SULLIVAN (1987), "Nonparametric tests of market structure: an application to the cigarette industry", *Journal of Industrial Economics*, vol. 25, pp. 483-98.
- BAUMOL, W.J., J.C. PANZAR and R.D. WILLIG (1982), *Contestable Markets and the Theory of Industry Structure*, Harcourt Brace Jovanovich, San Diego.
- BRANSON, W.H. (1990), "Financial market integration, macroeconomic policy and the EMS", in C. Bliss and J. Braga de Macedo eds, *Unity with Diversity in the European Community: The Community's Southern Frontier*, CEPR/Cambridge University Press, Cambridge.
- BRESNAHAN, T. (1989), "Empirical studies of industries with market power", in R. Schmalensee and R.D. Willig eds, *Handbook of Industrial Organisation*, vol. 2, North-Holland, Amsterdam.
- CHAMBERLIN, E.H. (1962), *The Theory of Monopolistic Competition*, Harvard University Press, Cambridge.
- COCCORESE, P. (1995), "Contendibilità e industria bancaria", *Giornale degli Economisti e Annali di Economia*, vol. 14, pp. 229-42.
- CONTI, V. (1989), *Margini, costi e prospettive strategiche per le aziende di credito: alcuni confronti internazionali*, Banca Commerciale Italiana, Milano.
- DESARIO, V. (1995), "La concentrazione del sistema bancario", *Bollettino Economico*, n. 25, Banca d'Italia.
- DI BATTISTA, M.L. and M. GRILLO (1988), "La concorrenza nell'industria bancaria italiana", in F. Cesarini et al., a cura di, *Banca e mercato*, Banca Commerciale Italiana, il Mulino, Bologna.
- MOLYNEUX, P., D.M. LLOYD-WILLIAMS and J. THORNTON (1994), "Competitive conditions in European banking", *Journal of Banking and Finance*, no. 18, pp. 445-59.
- NATHAN, A. and E.H. NEAVE (1989), "Competition and contestability in Canada's financial system: empirical results", *Canadian Journal of Economics*, vol. 22, pp. 576-94.
- NATHAN, A. and E.H. NEAVE (1991), "Reply to Perrakis", *Canadian Journal of Economics*, vol. 24, pp. 733-35.
- NEVEN, D.J. (1990), "Structural adjustment in European retail banking: some views from industrial organisation", in J. Dermine ed., *European Banking in the 1990s*, Basil Blackwell, Oxford.

- PADOA-SCHIOPPA, T. (1995), "Global banking Italia", *Bollettino Economico*, n. 24, Banca d'Italia, pp. 49*-58*.
- PANZAR, J.C. and J.N. ROSSE (1987), "Testing for monopoly equilibrium", *Journal of Industrial Economics*, vol. 25, pp. 443-56.
- PARRILLO, G. (1989), "Il sistema finanziario italiano e il mercato unico europeo: evoluzione normativa e strutturale", *Quaderno di Ricerca*, Editoriale Lavoro, luglio.
- PASSACANTANDO, F. (1983), "Costi e margini del sistema bancario italiano: un'analisi comparata", *Temi di Discussione*, n. 24, Banca d'Italia.
- PERRAKIS, S. (1991), "Assessing competition in Canada's financial system: a note", *Canadian Journal of Economics*, vol. 24, pp. 727-32.
- PRICE WATERHOUSE (1988), "The cost of non-Europe in financial services", *Research of the Cost of Non-Europe*, no. 9, Commission of the European Communities, Brussels.
- RAMSEY, J.B. (1969), "Tests for specification errors in classical linear least squares regression analysis", *Journal of Royal Statistical Society*, series B, vol. 31, pp. 350-71.
- ROSSE, J.N. and J.C. PANZAR (1977), "Chamberlin vs. Robinson: an empirical test for monopoly rents", *Bell Laboratories Economic Discussion Papers*, no. 90.
- SHAFFER, S. (1982), "A non-structural test for competition in financial markets", in *Bank Structure and Competition*, Conference Proceedings, Federal Reserve Bank of Chicago, Chicago.
- SRAFFA, P. (1926), "The laws of returns under competitive conditions", *Economic Journal*, vol. 36, pp. 535-50.
- STIGLER, G. (1964), "A theory of oligopoly", *Journal of Political Economy*, vol. 72, pp. 44-61.
- SULLIVAN, D. (1985), "Testing hypotheses about firm behaviour in the cigarette industry", *Journal of Political Economy*, vol. 93, pp. 586-98.
- VIVES, X. (1991), "Banking competition and European integration", in A. Giovanni and C. Mayer eds, *European Financial Integration*, Cambridge University Press, Cambridge.