The Long Journey of Bank Competition: New Evidence on Italy from 1890 to 2013

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PRELIMINARY DRAFT

Abstract

The main contributions of this paper are represented by the analysis of bank competition over a period of 120 years and the use of the Boone indicator. We study the link between profits and average costs of banks regressing the former variable on the latter. According to Boone et al. (2013), the greater the negative value of the estimated coefficient, the higher the level of competition. We also analyze the sensitivity of loan market shares to average costs, since in a competitive market inefficient firms are supposed to lose market shares. Regressions show negative values of beta for most of the period. The results are more robust using market shares instead of net profits as measures of bank performance. We interpret the evolution of bank competition in Italy estimating annual values of beta and looking at the main regulatory changes and the developments of the business cycle. Competition was high between 1890 and the mid-1920s, when a sort of free banking was present. With the introduction of the two banking laws in 1926 and 1936, strong barriers to entry were introduced in the banking system. As a result, competition decreased until the 1950s and subsequently stabilized. A new rise of competition took place in the 1980s. Competition reached its maximum levels along the 120 years around the mid-1990s. Between the second half of the 1990s and the first years of the new millennium competition decreased probably as an effect of the wave of mergers and acquisitions. Such decrease of competition came to an end in the years following the global financial crisis.

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1 Introduction

Normally, competition is assessed on a short-term basis. This is a reasonable approach as antitrust authorities must evaluate mergers and acquisitions, possible abuses of dominant positions and cartels in order to make rapid policy decisions. Moreover, economists often evaluate the effects on competition deriving from exogenous shocks such as changes in regulations. This is common in banking literature. For instance, many contributions analyzed the effects of the creation of the common market in Europe (1993) or of the birth of the euro area (1999) on competition. These studies generally considered one or two decades.

On the contrary, the emphasis of our paper is on the long-run. Exploiting a unique data set which covers 120 years of balance-sheets and profits and loss accounts, we study banking competition in light of different regulatory regimes. We distinguish three conventional periods: the free banking era that characterized Italy from 1890 to the end of the 1920s, the strong prudential regulatory regime introduced in the 1930s and that remained virtually unchanged until the end of the 1970s, and the third period started in the 1980s with bank deregulation and liberalization. We analyze whether competition was different in these three periods devoting attention also to subperiods - such as the 1990s - characterized by structural breaks.

Among the available measures of competition, we chose to use the Boone et al. (2013) indicator. The idea behind such indicator is that the greater is the market power of firms, the more easily they can pass on higher costs to higher prices. The more competitive a market is, the more negative is the coefficient obtained from a regression of profits on costs. Furthermore, the stronger the competition, the more firms with higher average costs - i.e. the less efficient ones - lose market shares. Following this approach we study the elasticity of profits and of market shares with respect to average costs.

The paper is divided into five sections. After this introduction, section 2 summarizes the literature on competition indicators and briefly sketches our approach. Section 3 describes our methodology in detail and the data set, discussing the behaviour of bank profits in the long-run. Section 4 contains the econometric results deriving from the application of both parametric and non-parametric methods. Section 5 concludes.

2 On Measures of Banking Competition and our Approach

Competition plays a major role in economic theory but there is not wide consensus on how to measure it. A first strand of literature, the Structure-Conduct-Performance (SCP) paradigm, has focused for decades on assessing competition through market concentration. The intuition behind this method, originally developed by Bain (1951), is that collusion is hard to be achieved when the number of firms is large. According to the SCP hypothesis, market concentration (Structure) leads firms to behave non-competitively (Conduct) and to reach higher profitability (Performance). Testing the

SCP hypothesis requires to regress a measure of profitability on an index of concentration, like the Herfindahl index or a concentration ratio. A positive coefficient would justify the use of concentration as a measure of competition, by showing that high concentration allows firms to obtain higher profits. The Efficient-structure hypothesis (Demsetz, 1973; Peltzman, 1977) and the "contestable markets" theory (Baumol et al. 1982) rejected the SCP paradigm arguing that high market shares associated with high profitability could also be a signal of a competitive environment in which only the most efficient firms - in terms of better management or appropriate scale - survive. In the basic SCP regression, concentration cannot be considered as an exogenous regressor because efficiency is correlated both with profitability and concentration. After including market shares in the specification, Berger et al. (2004) find out that the relationship between concentration and profitability is very weak, which is consistent with the idea that best firms gain larger market shares.

Therefore concentration cannot be considered a reliable measure of competition. The New Empirical Industrial Organization (NEIO) approach tries to estimate competition through the derivation of conduct parameters. One method consists in computing the H-Statistic, proposed by Panzar and Rosse (1987), which is equal to the sum of the elasticities of revenues to production factors. In perfect competition, if input prices increase, revenues grow in the same extent, so that the H-Statistic is equal to one. In monopoly, since the price elasticity of demand is larger than 1, an increase in input prices determines a reduction in revenues, implying a negative H-statistic. In monopolistic competition, the indicator ranges between 0 and 1. Therefore, under the assumption of being in a long-run equilibrium, the H statistic allows discriminating among these three degrees of market competition, but it does not provide a measure of the evolution of competition.

Another competition indicator is the Lerner Index, which is equal to the difference between price and marginal cost, divided by the price itself. In perfect competition, where price and marginal cost coincide, the index is equal to zero; instead, a firm with market power is able to charge a price which is greater than marginal cost. Therefore, the larger the index the higher the market power in the industry. Whereas prices are directly observable, researchers have to compute marginal costs at the firm level by estimating a cost function or using a proxy like average costs (which is equivalent to assume a linear cost function). After having obtained estimates of the Lerner index at the firm level, market shares are used as weights to compute the index for the entire industry. Unfortunately, this measure cannot deal with the reallocation effect, which refers to the loss of market shares (and even the exit from the market) for the least efficient firms due to an increase of competition. In this case, the more efficient firms which survive are characterized by higher Lerner index because, given prices, they have lower marginal costs (this reminds the conclusions of the efficient structure hypothesis). Therefore an increase in competition could raise the aggregate Lerner index, which would turn out to be an incorrect indicator of the evolution of competition[1].

Boone (2008b) proposes a new way to measure the evolution of competition, the Relative Profit Difference (RPD), which theoretically captures changes in the intensity of competition due, for example, to a decrease in entry barriers, to closer product substitutability, to more aggressive interaction among firms. The basic idea of this indicator is that when competition increases, less efficient firms are punished in terms of profits more harshly than efficient ones. Boone et al. (2013) shows that the elasticity of profits to marginal cost (PE), which can be estimated through a regression of profits on marginal costs (both in log terms), is closely related to the RPD. The coefficient beta obtained by regressing profits on marginal costs should be negative. Efficient firms, thanks to lower marginal costs, tend to price lower but not to the extent that price to cost margin decreases; moreover, this margin tends to fall when competition increases. Under these two assumptions, which turn out to be quite reasonable, PE captures the evolution of competition. According to Boone et al. (2013), PE performs better than the Lerner index, especially when the reallocation effect is high. Furthermore Boone (2008a) shows that the Herfindahl index and the Lerner index are not monotone functions of competition. An increase of competition may correspond both to an increase or a decrease of these indicators whereas this is not the case for the Boone indicator (before the appearance of Boone's contributions other theoretical papers had already shown that more intense competition may lead to higher price-cost margins).

Several recent studies use the Boone approach investigating the nexus between profits and marginal costs. Estimation methods vary. As Amador and Soares (2012) and Schiersch and Schmidt-Ehmcke (2010) point out, PE is highly sensitive to econometric specifications, estimation methodologies, the set of firms considered and potential non-linearities. First, the definition of variables might be challenging, especially for marginal costs. Many studies, like Maliranta et al. (2007), Schaeck and Cihák (2014), Peroni and Ferreira (2012), Amador and Soares (2012), Kick and Prieto (2013), use average variable costs as a proxy of marginal costs. This is the road we follow. Second, the specification suggested by Boone et al. (2013) applies a log-transformation to profits, dropping all firms with losses. This choice appears unreasonable as the natural outcome of competition is to drive less efficient firms out of the market. Instead, Clerides et al. (2013) transform profits following a method suggested by Bos and Koetter (2011), which avoids the loss of observations. We will also use the BK method. Third, estimation techniques range from OLS, fixed effects models and local regressions (Delis, 2012). Schaeck and Cihák (2014) use a two-step GMM estimator with one year lagged values

[1] A further measure, introduced by Iwata (1974), Bresnahan (1982) and Lau (1982), is the conjectural-variations method. It requires the simultaneous estimation of a system of demand and supply equations to obtain a conduct parameter, which represents the perceived response of industry output to a change in one firms output. It can be shown that the conduct parameter corresponds to a Lerner index adjusted by the elasticity of demand and therefore it cannot capture the reallocation effect. Clearly, the data requirement for estimating the supply function can be stringent when only balance sheets data are available and scholars do not have information on prices.

because endogeneity issues may arise if performance and cost are jointly determined (but Boone et al. (2013) shows that even in case of endogeneity, PE is still a good measure of the evolution of competition). We will also resort to several econometric methods[2].

There are several applications to banking. Schaeck and Cihák (2014) use a panel dataset for European banks between 1995 and 2005 and data on US banks in 2005 aiming to evaluate the effects of competition on financial stability. Delis (2012) investigates the relation between bank competition and financial reforms with balance sheets data from Bankscope over the 1987-2005 period. Clerides et al. (2013) observe a correlation between banking competition and business cycle using Bankscope and covering the 1997-2010 period for 148 countries. Kick and Prieto (2013) use data on German banks in the 1994-2010 period and find out that an increase in competition, if measured through PE, lowers the riskiness of banks. van Leuvensteijn et al. (2011) and Tabak et al. (2012) consider the elasticity of market shares to marginal costs as indicator of competition in the banking sector. The main advantage of using market shares is that no observation is dropped while the main disadvantage is that the national market shares do not correspond to the relevant local markets. van Leuvensteijn et al. (2011) uses an extended Bankscope database covering the 1992-2004 period for the US, Japan and the major Euro area countries and show that the US have the most competitive loan market. Tabak et al. (2012) investigates the relationship between competition and financial stability, focusing on banks in Latin America between 2003 and 2008.

The literature on banking competition in Italy has focused on the effects of the liberalization process that took place in the 1980s and in the 1990s. Angelini and Cetorelli (2003) estimate the Lerner index through the conjectural-variations method over the 1984-1997 period; they find an increase in competition after 1992, and provide evidence that it was due to the deregulation process. Focarelli and Panetta (2003) show that the mergers which took place in Italy in the 1990s determined a decrease in deposit interest rates in the short term; however, this increment has been only temporary because in the long-run rates have increased, so that the consolidation process has turned to be beneficial for banks clients. The Italian deregulation process, and the related expansion of banks branches, has also provided the opportunity to test the multimarket hypothesis, which postulates that more contacts among firms facilitate collusion. De Bonis and Ferrando (2000) and Coccorese and Pellecchia (2013) obtain opposite results: the former study rejects the multimarket hypothesis, the latter accepts it. The controversial result is related to different measures of competition, different estimation strategies and different periods of analysis (the former 1990-1996, the latter 1997-2009). Coccorese (2009), using data between 1988 and 2005, shows that banks operating as monopolists in local markets did not entirely exploit their market power (measured through both Lerner index and H-statistic), suggesting that concentration and competition in banking

[2] The empirical works estimating the link between profits and marginal costs consider many industries (Peroni and Ferreira (2012); Boone et al. (2013); Amador and Soares (2012)) or focuses on specific sectors. Boone and van Leuvensteijn (2010), for example, compute PE for American sugar industry between 1890 and 1914.

can coexist.

The main novelties of our paper are three. First, while previous studies took into account a few decades of statistics at best, our paper considers more than 120 yearly observations for a large panel of Italian banks. Banking regulation has drastically changed in the last 120 years and we try to capture the effects of regime-switching on competition. Second, while the Boone indicator has been already estimated for the Italian banking market, previous contributions took into account only very short time periods. Third, to study competition we regress not only profits on average costs but also loan market shares on average costs. In the following section we illustrate the details of our approach and the main characteristics of our data set.

3 Methodology and Data

In this section we first describe our methodology (3.1) and then we summarize the characteristics of our data set (3.2).

3.1 Profits equation

As already mentioned, Boone et al. (2013) show that the elasticity of profits to marginal cost (PE) is closely related to the Relative Profit Difference. This elasticity can be approximated through the derivative of the log of profits to the log of marginal costs. Under the assumption of constant average costs, average and marginal costs coincide. Therefore we study the relationship between the log of profits and the log of average costs.

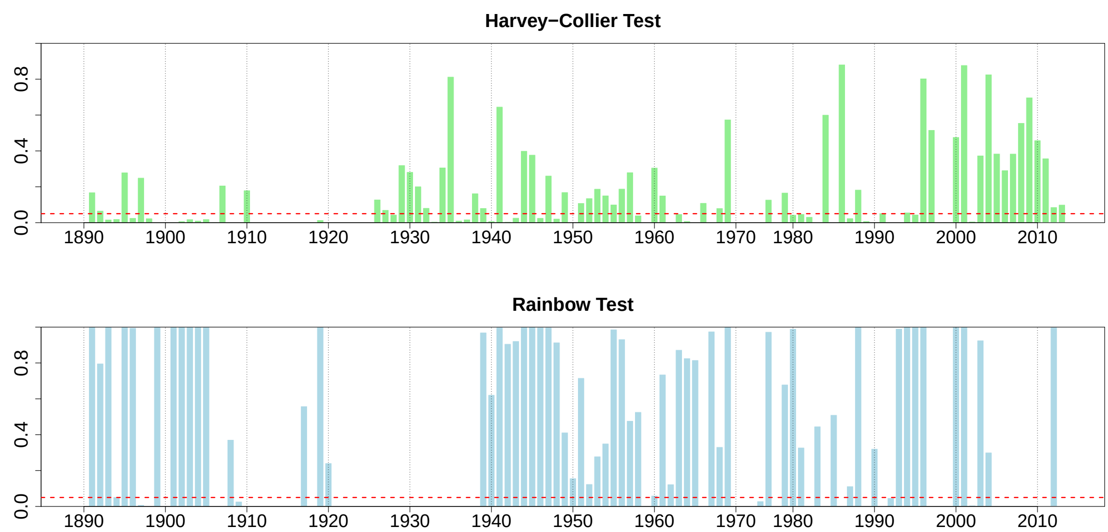
We also follow the approach proposed by van Leuvensteijn et al. (2011), considering the elasticity of market shares to average costs as measure of competition. As highlighted by Tabak et al. (2012), a firm may use an efficiency improvement to raise profits in two ways: it may either charge the same price and keep the previous volume of revenues unchanged or it may lower the price and increase its market share. If we assume that all banks always pass their efficiency gains to the consumers, at least partially, we can study the evolution of competition through the elasticity of market shares to average costs.

We focus on the following equation:

$$\pi_{i,t} = f\left(c_{i,t}, x_{i,t}, \eta_t, \varepsilon_{i,t}\right) \tag{1}$$

where $\pi_{i,t}$ is either the log of net profits of bank i at time t or the market share of loans, $c_{i,t}$ represents the log of the average costs, η_t a time fixed effect and $\varepsilon_{i,t}$ is a vector of unobservables. $x_{i,t}$ is a vector of control variables, such as dummies for saving and cooperative banks, for geographic location, bad debts to loans ratio and the leverage ratio. We estimate the profit equation in both a parametric and a non-parametric setup. The twofold approach is dictated by the idea that the relation between profits and costs may turn out to be non-linear. Thus the parametric and, obviously, the non-parametric specification are devised to account for a departure from the standard assumption of linearity. Indeed, a visual inspection of the data by year[3] suggests that the relation may be non-linear in many years. We tested such hypothesis with the Harvey-Collier and the Rainbow linearity tests for each year in the sample. Figure 1 reports the p-values of the tests with the null hypothesis being the linearity of profits with respect to average costs. While the Harvey-Collier test can be considered a proper linearity test, on the other hand the Rainbow test verifies that, even when a non-linear relation exists, a subset of the observations can be nonetheless used to obtain a good linear fit. The subset of the observations is controlled though a proportion which is usually set to 0.5. Table A.2 reports the matrix of results for the two tests depending on the choice of the significance level. The results suggest that non-linearity is a valid hypothesis for at least half of the 120 years of our sample period.

Figure 1: Linearity Tests



Source: authors' elaborations on Bank of Italy data. The Harvey-Collier and the Rainbow tests for linearity assume linearity as null hypothesis. The figure reports the p-values of the two linearity tests run on the log of net profits with respect to the log of average costs on each year of the sample period. Bars that are lower than the 5% red line indicate that the null hypothesis of linearity has to be rejected for that specific year.

In light of the above-reported evidence, we formulate the following parametric spec-

[3] We do not report the scatterplots of the log of net profits with respect to the log of average costs for each of the 120 years of the sample period for conciseness reasons (the figures are available upon request from the authors).

ification

$$\pi_{i,t} = \beta_0 + \beta_1 c_{i,t} + \beta_2 c_{i,t}^2 + \beta_3 c_{i,t}^3 + \sum_{k=1}^{K} \gamma_k x_{k,i,t} + \eta_t + \varepsilon_{i,t} \tag{2}$$

where the log of the average costs $c_{i,t}$ enter also with a quadratic and a cubic term, in order to allow for non-linearities. In order to avoid using a truncated data set due to the application of the logarithmic transformation to negative profits, we use the Bos and Koetter (2011) transformation. The log of negative values is set to zero and a control variable $\zeta_{i,t}$ is added to the model:

$$\zeta_{i,t} = \begin{cases} 0 & \text{if } \pi_{i,t} \geq 0; \\ \log(-\pi_{i,t}) & \text{if } \pi_{i,t} < 0. \end{cases}$$

Instead, market shares cannot assume negative values thus overcoming the issue of the log transformation of negative profits.

Our non-parametric specification is

$$\pi_{i,t} = f_t\left(c_{i,t}, x_{i,t}, \varepsilon_{i,t}\right) \quad \forall t \in (1890, 1891, ..., 2013)$$

where we estimate a function f_t for each year in our sample through the local regression approach. The vector of control variables $x_{i,t}$ includes the leverage ratio and the bad debts to loans ratio. No *a priori* restriction is imposed on the functional form except for the fact that it has to be smooth. The function is estimated for each observation c_j[4] considering its *neighbourhood*, i.e. the remaining closest observations which are weighted according to their distance from the mentioned c_j observation. The smoothing window is defined as $\left[c_j - h\left(c_j\right), c_j + h\left(c_j\right)\right]$ where h is the bandwidth parameter that determines the smoothness of the fit and the width of the window. A fitting point $\hat{\pi}_{j,t}$ is derived as

$$\hat{\pi}_{j,t} = \sum_{i=1}^{n} W\left(\frac{c_{i,t} - c_j}{h}\right)\left[\pi_{i,t} - \theta_0 - \theta_1\left(c_{i,t} - c_j\right)\right]^2 \tag{3}$$

where n is the total number of observations in year t and the weight function W is represented by:

$$W(u) = \begin{matrix} \left(1 - |u|^3\right)^3 & if \quad |u| < 1 \\ 0 & if \quad |u| \geq 1 \end{matrix}$$

where $u = \left(c_i - c_j\right)/h(c_j)$. The choice of the values of the parameters to plug into the local regression estimation procedure is fully data-driven. The three main parameters, namely the nearest neighbour component and the constant component of the smoothing

[4] For brevity reasons, we illustrate the univariate case with one predictor, i.e. average costs.

parameter and the penalty for adaptive fitting are all chosen on the basis of a constrained minimization of the globally cross-validated indicator.

In this case, the Bos-Koetter transformation cannot be used and our sample suffers from truncation of negative values of profits. Clearly, this is not an issue when using market shares as dependent variable.

3.2 Data

Our database goes from 1890 to 2013. For each year we collected information on banks profit and loss accounts and on balance sheets. The data set may be split into two parts:

- from 1890 to 1973 statistics are taken from the historical banking archive of the Bank of Italy (ASCI)[5];
- from 1977 to 2013 information is taken from the electronic database of prudential statistics of the Bank of Italy.

Unfortunately, data on balance sheets are missing for the years 1974-1976.

The ASCI collects around 40,000 balance sheets which have been turned homogeneous through a unique common scheme for the entire period. The scheme includes 14 variables for the assets side, 9 for the liabilities side; it also provides total costs and total revenues. The number of banks per year is not constant as it mainly depends on the availability of historic data. Almost the universe of saving banks is included in the sample; more than 70% of commercial banks is included too (except for 1926); the number of cooperative banks instead is highly volatile, generally larger than 30% (50% since 1951) but almost null between 1911 and 1935. Overall, the banks in the dataset cover more than 80% of both total deposits and total assets using the estimates of banking aggregates present in the literature (Cotula and Raganelli (1996), Garofalo and Colonna (1999) and De Bonis et al. (2012)). The lack of detailed banks profit and loss accounts in the ASCI is a major issue, since a better estimate of the Boone indicator would require either average operating costs or marginal costs estimated through a cost function. Moreover, because of the lack of profit and loss accounts, we could not use better measures of performance, such as gross profits or operating revenues.[6] For the 1963-1973 period we were able to add more details on bank profit and loss accounts using an unpublished dataset of the Bank of Italy.

Statistics for the time span 1977-2013 are much more detailed than those of the 1890-1973 subperiod, especially for profit and loss accounts[7]. Again, data cover the majority of Italian banks, reaching a market share of approximately 80 per cent for the main balance sheet items. To avoid statistical breaks our data do not include mutual

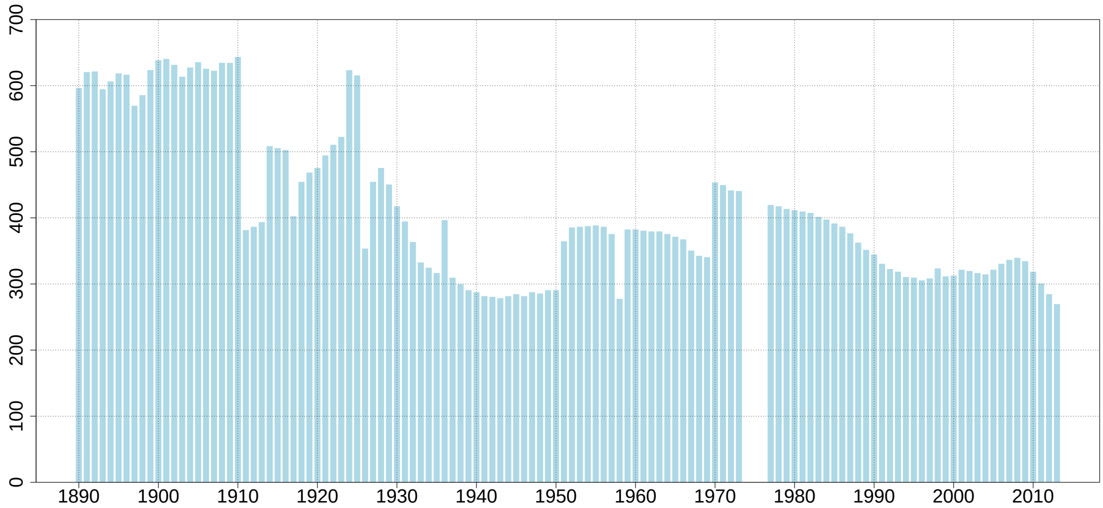
[5]See Natoli et al. (2014) for a full description of the data set.

[6]However, the Boone indicator seems to be more sensitive to the definition of costs than that of the performance measure.

[7]We thank Carlo Mauri and his colleagues for making the data available for the time span 1977-1982.

cooperative banks and special credit institutions. Actually, the former started to report complete statistics only in 1983 while the latter in 1995. In other words, even if data on these categories are available respectively for the time spans 1983-2013 and 1995-2013, we chose to exclude them in order to preserve continuity of the time series. In order to reduce the high variability of the sample, we have deleted all observations that are not present for at least two consecutive years. As shown in Figure 2, our average yearly number of banks is more than 400 institutions. Taking into account that we cover the interval 1890-2013 our dataset include around 50,000 observations.

Figure 2: Number of banks in the sample

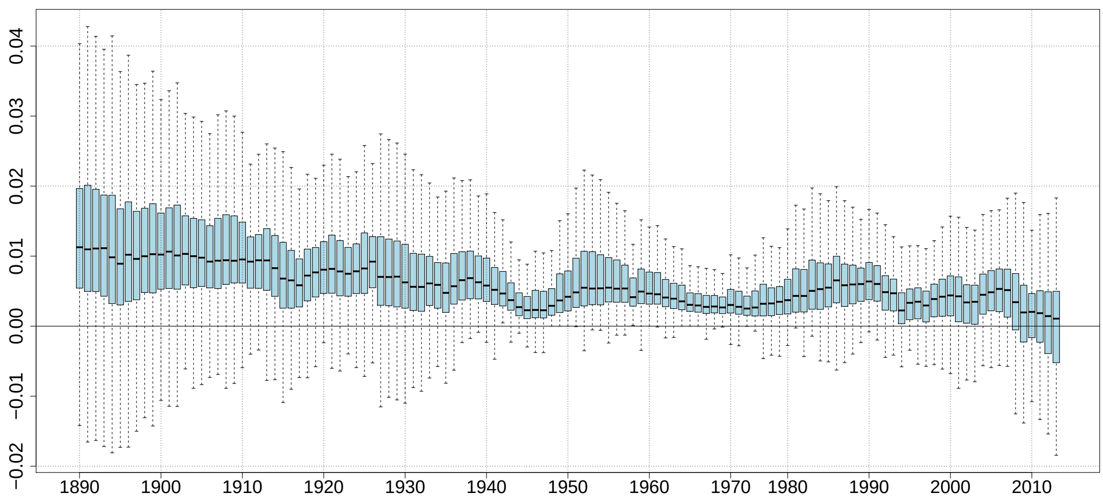


Source: authors' elaborations on Bank of Italy data. Number of banks in the sample of the present paper.

Following the approach of most of the papers reviewed in the previous section, our initial dependent variable is the log of net profits. As an alternative we also use the log of loan market shares. Turning to independent variables, the main indicator is the ratio of total costs to total assets as a proxy for average costs. Unfortunately, as we have already mentioned, for the time span 1890-1962 we have information only on banks' total costs and not on operating costs that are the key indicator to measure efficiency. Therefore in the regression we chose to include the average total costs from 1890 to 1962 and the more appropriate average operating costs in the following years.

Table A.1 reports the main descriptive statistics of the variables in our data set. Figure 3 shows the trend and the dispersion of the ratio of net profits to total assets (ROA) from 1890 to 2013 in our sample. ROA reaches its highest levels at the beginning of the period taken into account, and then shows a negative trend until the 1970s, with some relevant fluctuations. Disregarding the two World Wars periods, it is important to notice the drop between mid-1920s and mid-1930s, which is also due to the Great Depression. From the beginning of the 1970s ROA started rising, reaching again high levels in the 1980s. After a reduction associated with the 1992-93 recession of the Italian economy, ROA had a new increase until the break out of the global financial crisis. In the last years, because of the two recessions that hit the Italian economy after 2008, ROA has reached the lowest levels of the 120 years of banking history we are studying. It's worth noting that dispersion has also declined since the 1890s, with some fluctuations: it increased at the end of the 1920s, during the 1950s, at the beginning of the 1980s and finally from the mid-1990s.

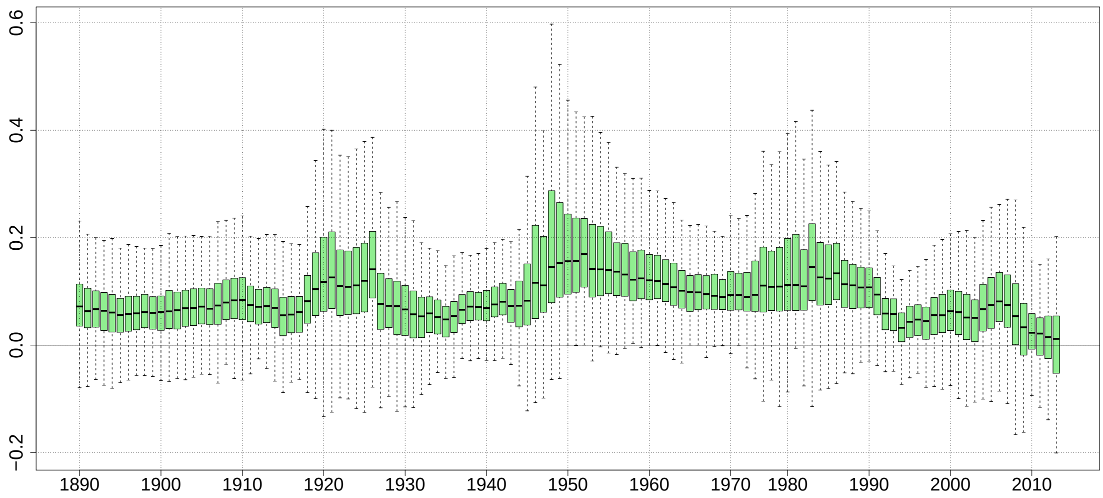
Figure 3: Banks' return on assets (ROA)



Source: authors' elaborations on Bank of Italy data. The evolution of the distribution of the variable over time is represented through a boxplot where the line in the middle of the box is the median; the lower and the upper side of the box, in blue, represent the first and the third quartile respectively; and the notches extend to $\pm 1.58 \cdot IQR/\sqrt{n}$ so as to give rougly a 95% confidence interval.

More interesting insights to study competition can be drawn by looking at the evolution of ROE, the ratio of net profits over capital and reserves (figure 4). In contrast with ROA, ROE was quite low before the 1930s, except for a temporary increase in the mid-1920s. In this period there were no official restrictions to competition and the banking cartel was not binding. In the mid-1930s, at the same time of the 1936 Banking Act which strongly restricted bank competition, ROE started rising and it kept to increase until the beginning of the 1950s. Of course, this rise has been driven also by inflation (especially in the post-war years), but it is worth noting that ROE has remained on high levels until the 1980s: this coincides with a period of strong barriers to competition. ROE decreased since the mid-1980s until the first half of the 1990s when profitability decreased sharply due to the strong recession that hit the Italian economy in 1992-1993 and the virtual defaults of large Southern banks that were taken over by Central and Northern banks. A new rise of profitability occurred in the second half of the 1990s but without coming back to the levels observed in the 1980s. The eruption of the global financial crisis, two recessions and the euro area sovereign debt crisis led to a fall in profitability. In the last years ROE has been on the lowest levels in the entire sample period[8].

Figure 4: Banks' return on equity (ROE)



Source: authors' elaborations on Bank of Italy data. The evolution of the distribution of the variable over time is represented through a boxplot where the line in the middle of the box is the median; the lower and the upper side of the box, in red, represent the first and the third quartile respectively; the notches extend to $\pm 1.58 \cdot IQR/\sqrt{n}$ so as to give rougly a 95% confidence interval.

4 Econometric Results

The sensitivity of profits or of loan market shares to average costs may be analyzed using parametric or non-parametric methods. In subsection 4.1 we illustrate the results obtained with the classic parametric approach. In subsection 4.2 we discuss the evidence obtained using the non-parametric approach, particularly the local regression method.

[8]The behaviour of ROE is influenced by the attitude of bank supervisors towards capital and reserves. Capital requirements were introduced in industrial countries at the end of the 1980s while these measures were rarely used in previous years.

4.1 Parametric Model Results

We start with a simple regression where the log of net profits is regressed on banks' average costs for the whole sample period (1890-2013); results are reported in Table A.3. The OLS coefficient is negative and statistically significant. On the contrary, if we allow for bank fixed effects and perform first difference estimations (FD) we obtain a positive coefficient. Results are more consistent when we use the individual market share of loans as dependent variable. In this case the coefficient of the average costs is negative and statistically significant using both the OLS and FD estimators.

Competition and regulation have drastically changed over the 120 years of our sample[9]. We assume that this long period may be split into three phases: 1890-1930, 1931-1979, 1980-2013. Here we provide some rationale behind our choice. A first period ranging from 1890 to 1930 may be labeled as the "free banking" era when barriers to entry were virtually inexistent. Banks were free to constitute and to open branches as the supervisory controls were absent until the approval of the first banking law in 1926 (Gigliobianco and Giordano, 2012).

A second period started in the 1930s when a severe banking regulation was introduced in reaction to bank failures during the Great Depression[10] (Toniolo, 1995). With the banking law introduced in 1936, the Bank of Italy was entitled with the supervision of the banking sector. The constitution of new banks, the opening of branches and mergers and acquisitions were subject to the authorization of the Bank of Italy. Banking regulation underwent changes after the Second World War; nevertheless banking competition remained strongly restricted until the 1970s. For instance, a bank cartel, sponsored by the Government, had been created in 1919 but it had strengthened only in 1932 when it had become compulsory for most of the banking sector. The regulation included caps on deposit rates and had to be enforced by the Bank of Italy. After 1952, the cartel became voluntary but remained effective until 1974 (for a discussion on the evolution of bank competition in Italy in the 1950s see Albareto, 1999).

The third period started in the 1980s. Following autonomous initiatives of the Bank of Italy and directives/regulations from the European Union, barriers to entry in the banking market were progressively lowered. Old limitations to the geographic destination of loans were canceled. The authorization to the creation of new banks was simplified. Branch opening was liberalized in 1990 but, at the same time, mergers and acquisitions increased in the 1990s.

Our working assumption is that competition has decreased in the second phase with respect to the first one while it has increased in the third period. In other words, we expect that banking competition was lower in the interval 1930-1979 than in the period 1890-1929. Moreover, we expect that competition rose since the 1980s even if the

[9]See Battilossi et al. (2013) for a general discussion on this topic.

[10]On the link between acceleration of credit and bank crises see De Bonis and Silvestrini (2014).

strong increase of mergers and acquisitions in the 1990s and in the first years of the new millennium must be scrutinized.

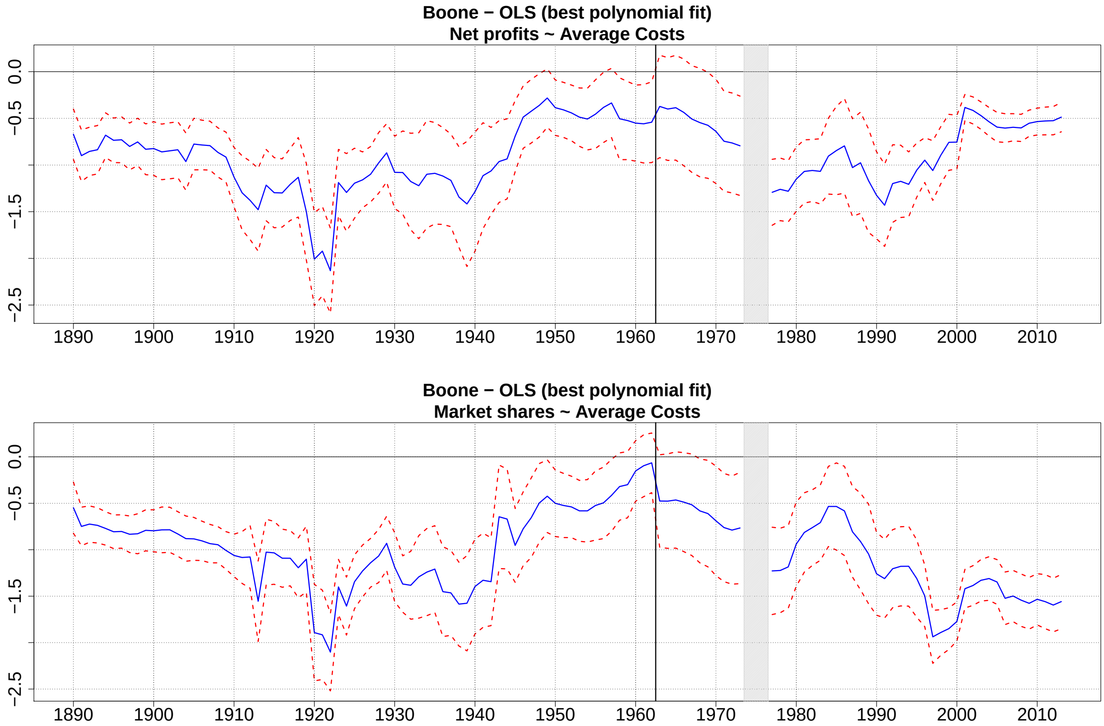
As already mentioned in the Section on descriptive statistics, only total costs are available from 1890 to 1962; from 1963 onwards statistics refer to the more appropriate operating costs aggregate. Therefore the interval 1930-1979 is split into two subperiods: 1930-1962 and 1963-1979. Table A.4 presents the econometric results splitting the sample in the above-mentioned subperiods: 1890-1929, 1930-1962, 1963-1979, 1980-2013. The OLS estimates using net profits as dependent variable obtain a negative coefficient. Competition is higher in the time span 1890-1930 than in the other periods. However competition is not higher in the interval 1981-2013 than in the previous periods. The FD estimates with net profits as dependent variable do not confirm our *a priori*. On the contrary, results are closer to our expectations when loan market shares are employed as dependent variable. The OLS shows a high level of competition in 1890-1930, a lower level of competition in the 1930-1979 subperiod and the highest level of competition in the more recent subperiod. The FD estimates are similar to the OLS ones.

Table A.5 reports the estimations of the profit elasticity to average costs adding some control variables. We introduce geographic and bank type dummies, the ratio of total assets to capital (a rough measure of leverage) and the ratio of bad debts to loans. Results and interpretation of the coefficients are consistent with the baseline results reported in Table A.4.

Our conventional splitting of 120 years of the evolution of banking competition may be considered arbitrary as it may hide a more complex dynamics of competition. For this reason we perform the estimations with five-year rolling windows (Figure 5). This method allows to study the evolution of competition without imposing restrictions on the definition of the subperiods.

Using net profits as dependent variable we find that competition increases from the beginning of the 20th century to the 1930s. There is consensus on the strong growth in the number of banks and branches until the mid-1920s and on the concomitant banking crises in Italy. Then, following the introduction of the severe banking regulation, competition stabilized and subsequently declined from the end of the 1930s. It remained on low levels until the 1970s. Competition increased in the second half of the 1980s while it fell in the 1990s. Figure 5 shows a similar behaviour of competition looking at loan market shares except. The strengthening of competition continues during the 1990s. A subsequent fall in competition followed between the end of the 1990s and the first years of the new millennium. We attribute this decrease to two factors. First, a wave of mergers and acquisitions characterized those years. Second, between 1995 and 2000 the *dot-com* bubble induced a sharp rise of stock prices. Italian banks took advantage of the economic developments by increasing the profits deriving from asset management (in Italy the asset management industry is dominated by banks). A reduction of banking competition between 1994 and 2004 is also obtained by van Leuvensteijn et al. (2011).

Figure 5: OLS results



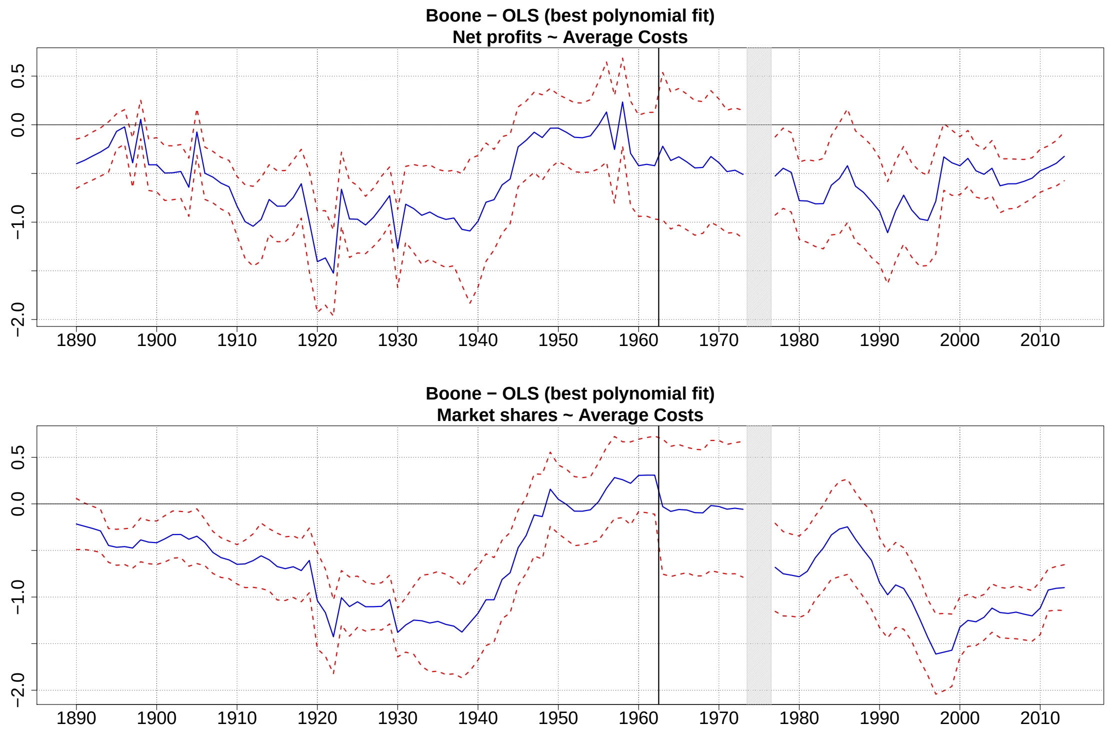
Source: authors' elaborations on Bank of Italy data. The estimated parametric model includes a third degree polynomial specification for the independent variable, net profits and market shares. The degree of the polynomial used for the estimation of the gradients and of the standard errors is chosen on the basis of the significance of the coefficients of the polynomial terms, i.e. the non-significant coefficients are dropped in light of a parsimonious specification. The vertical black line indicates the break in the average costs variable which is equal to average total costs until 1962 and to average operating costs as from 1963.

Figure 6 replicates the exercises in Figure 5 adding control variables such as geographic and bank type dummies, the total assets to equity ratio and the bad debts to loans ratio. The path is consistent with what we have already seen in Figure 5. An increase of competition - shown by a greater negative coefficient - took place in the first years of the 20th century. This corresponded to the entry of new banks in the market. With the introduction of entry barriers and the limits to competition of the banking law, competition decreased. The Boone indicator remains negative but is closer to zero and not statistically significant until the 1970s. As in the previous estimates, competition started increasing again from the mid-1980s. In the second half of the 1990s and the years preceding the global financial crisis competition declined. As already pointed out, the possible explanations are the wave of mergers and acquisitions and the boom of asset management[11]. This evidence is consistent with the increase in ROE commented

[11]Saving collected by investment funds in Italy skyrocketed from 65 billions of euros in 1995 to 450

in Figure 4: the return on equity increased from 1994 to 2006.

Figure 6: OLS results with control variables



Source: authors' elaborations on Bank of Italy data. The estimated parametric model includes a third degree polynomial specification for the independent variable, net profits and market shares. The degree of the polynomial used for the estimation of the gradients and of the standard errors is chosen on the basis of the significance of the coefficients of the polynomial terms, i.e. the non-significant coefficients are dropped in light of a parsimonious specification. The vertical black line indicates the break in the average costs variable which is equal to average total costs until 1962 and to average operating costs as from 1963.

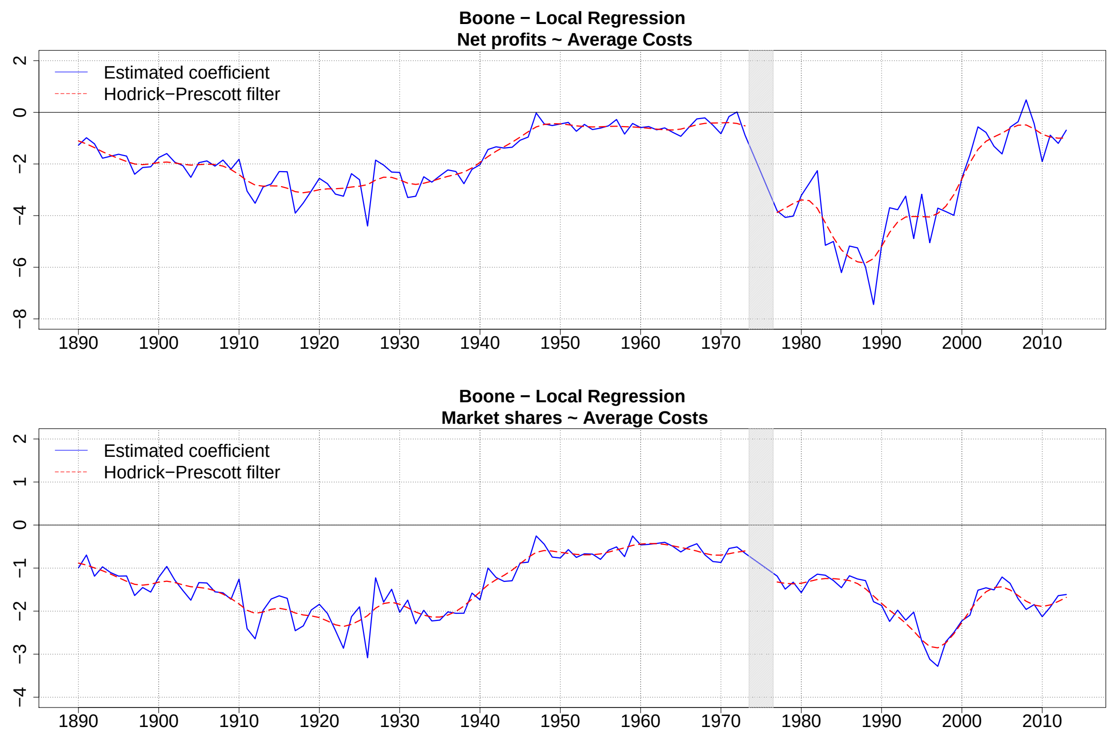
4.2 Non-parametric Model Results

In the previous exercises we used a parametric approach where an *a priori* functional form, be it linear, quadratic or cubic, is imposed. Now we apply a non-parametric approach analogously to Delis (2012). We estimate the Boone indicator using, as in the previous paragraph, the log of net profits and loan market shares as dependent variables. Average costs are the main independent variable. Also in this case we use the average total costs from 1890 to 1962. From 1963 to 2013 we use the more appropriate concept of average operating costs. The non-parametric regression allows to estimate the gradients for each bank in a given year. In order to summarize the level of competition in each year, we consider the mean gradient estimated through the local regression procedure.

billions of euros in 2000.

The upper panel of Figure 7 reports the evolution of the elasticity from 1890 to 2013 taking into account the log of net profits as dependent variable. The figure shows that competition increased until the early 1920s, during the period of overbanking and overbranching. Then from the second half of the 1920s until the 1950s competition decreased because of the new emphasis on bank stability. Competition did not change substantially until the beginning of the 1970s. Since then, competition increased especially in the 1980s and reached the highest levels of the entire sample period in the mid-1990s. Then competition decreased in the second part of the 1990s and in the first years of the new millennium: this matches the evidence reported in subsection 4.1. The estimation results of the more recent years seem to support a moderate expansion of competition starting from the global financial crisis. Such result has to be considered with caution as it may be heavily influenced by the left-truncation of the sample due to the log transformation of net profits when these are negative. The proportion of banks with negative profits has dramatically increased in the more recent years of financial crisis and has reached the value of one third in our sample.

Figure 7: Non-parametric results



Source: authors' elaborations on Bank of Italy data. The figure reports the yearly mean gradients estimated through the local regression procedure and the corresponding trend obtained by applying the Hodrick-Prescott filter.

As in the parametric estimates, we also regressed the log of loan market shares on average costs. The path of the Boone indicator shown in the lower panel of Figure 7 is very similar to that represented in the previous graph. Competition increased between 1890 and the early 1920s. With the introduction of the two banking laws competition decreased until the 1950s, then it stabilized until the 1970s. A sharp increase of competition took place from the early 1980s but this trend came to an end in the mid-1990s. After a decade of contraction, competition has started increasing again in the years following the financial crisis[12].

5 Conclusions

Economists do not agree on the best method to measure competition. This caveat applies also to the banking sector. This paper contributes to the literature in three ways. First, we study banking competition taking into account 120 years (from 1890 to 2013). Regulation and the business cycles have been very different over time and we study how these changes have influenced the evolution of competition. Second, while other studies applied the Boone measure to the Italian banking sector for 10-20 years, we provide a long-term analysis of this indicator. Third, we study not only the response of profits to costs but also the sensitivity of loan market shares to costs.

The paper studies the evolution of bank competition over 120 years starting from the following hypotheses. Competition should have been higher in the period of substantial free banking observed between 1890 and 1926. It is expected to have decreased between the 1930s and the end of the 1970s, when strong entry barriers and limits to competition were introduced as a reaction to bank failures observed in the 1920s and during the Great Depression. An increase of competition should have occurred in the 1980s when the Italian banking system was liberalized by autonomous initiatives of the Bank of Italy and by European directives. However a great wave of mergers and acquisitions took place from the mid-1990s until the years preceding the eruption of the financial crisis. The Boone indicator is sensitive to the econometric methods and to the linearity hypothesis. For this reason, we used different methods and introduced non-linearities in our specification.

The main conclusions of the paper are four. First, taking into account the entire time span 1890-2013, an increase in banking costs is generally negatively associated with net profits. The estimated elasticity appears more robust when we consider the relation between loan market shares and costs. Second, our estimates show that banking competition was high between the end of the 19th and the beginning of the 20th century characterized by overbanking and overbranching. Third, the two banking laws of 1926 and 1936 introduced severe entry barriers and supervisory controls that ensured stability but contributed to a reduction of the degree of competition until the 1970s. Finally, an

[12] Additionally, we run the local regression including the leverage ratio and the bad debts to loans ratio as controls. Results are substantially the same and are available from the authors.

increase in competition took place since the 1980s, following banking deregulation, such as the abolition of credit ceilings and of the limits to territorial expansion of bank loans. The non-parametric estimates show the highest elasticity - in absolute terms - of the entire sample period in the years 1990-1995, during the first phase of branching liberalization. From the second half of the 1990s to the years preceding the global financial crisis a reduction of competition occurred: this probably derived from the wave of mergers and acquisitions and the revenues obtained by banks from the boom in asset management linked to the *dot-com* bubble. A small increase of competition was observed during the years of the global financial crisis and of the last two recession that struck the Italian economy.

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A Appendix

Table A.1: Summary statistics over time

Variable	Sub-period	1st quartile	Mean	Median	3rd quartile	Std. Dev.
Net Profits	1890-1929	0.01	0.36	0.05	0.19	5.25
(1)	1930-1979	0.04	0.88	0.14	0.50	3.49
	1963-1979	0.24	2.66	0.67	1.93	7.80
	1980-2013	0.68	16.93	3.43	12.84	212.14
	1890-2013	0.03	4.52	0.18	1.26	101.14
Loan Market Shares	1890-1929	0.01	0.17	0.02	0.09	0.87
(%)	1930-1979	0.01	0.29	0.03	0.13	1.15
	1963-1979	0.01	0.25	0.04	0.13	0.86
	1980-2013	0.02	0.29	0.06	0.18	0.87
	1890-2013	0.01	0.23	0.03	0.12	0.94
Average Operating	1963-1979	0.02	0.03	0.03	0.03	0.01
Costs	1980-2013	0.02	0.04	0.03	0.04	0.14
	1890-2013	0.02	0.04	0.03	0.03	0.11
Average Total	1890-1929	0.03	0.06	0.04	0.06	0.09
Costs	1930-1979	0.04	0.05	0.05	0.06	0.03
	1963-1979	0.05	0.06	0.06	0.08	0.02
	1980-2013	0.04	0.09	0.08	0.10	0.15
	1890-2013	0.04	0.06	0.05	0.07	0.10
Bad Debts	1890-1929	0.00	5.28	0.15	2.49	15.80
to Loans Ratio (%)	1930-1979	0.00	1.04	0.05	0.74	3.39
(2)	1963-1979	0.56	2.85	1.53	3.38	4.47
	1980-2013	2.32	6.56	4.57	7.92	8.33
	1890-2013	0.00	4.34	0.72	3.91	11.74
Leverage Ratio	1890-1929	4.14	10.94	7.72	13.28	11.47
(3)	1930-1979	10.73	28.49	18.60	35.48	30.31
	1963-1979	22.21	37.69	32.51	47.14	22.31
	1980-2013	10.15	39.54	15.45	24.92	216.93
	1890-2013	6.77	24.25	13.19	24.89	105.92

Source: authors' elaborations on Bank of Italy data. (1) Net profits are expressed in millions of euros and are divided by the price index estimated by the Italian Statistical Office (Istat). - (2) Bad debts data suffer from some statistical breaks in the period 1929-35 due to the lack of the corresponding data for some of the banks in the sample. - (3) The leverage ratio is calculated as the ratio of total assets to capital and reserves and was winsorized at 1%.

Table A.2: Linearity Tests

		Rainbow Test			
		$\alpha = 0.05$		$\alpha = 0.1$	
		H_0	H_1	H_0	H_1
Harvey-Collier Test	H_0	30	25	24	23
	H_1	33	33	37	37

The table reports the number of years for which the linearity/non-linearity hypothesis is verified through the Harvey-Collier and the Rainbow test at the $\alpha = 0.05$ and 0.1 significance levels. The null hypothesis H_0 of the Harvey-Collier is that the relation between profits and costs is linear while the null of the Rainbow test is that, even in presence of non-linearity, a subset of the observations can be used to achieve a good linear fit. The fraction parameter of the Rainbow test was set to 0.5.

Table A.3: Profit elasticity: 1890-2013

	Net Profits (BK)	Market shares (loans)
OLS	-.553 ***	-1.004 ***
	(.0751)	(.061)
FD	.36 ***	-.251 ***
	(.0336)	(.0212)

The estimated parametric model includes a third degree polynomial specification for the independent variable, net profits and market shares. The degree of the polynomial used for the estimation of the gradients and of the standard errors is chosen on the basis of the significance of the coefficients of the polynomial terms, i.e. the non-significant coefficients are dropped in light of a parsimonious specification.

Standard errors in parentheses

Time dummies included in all regressions.

BK=Bos Koetter transformation (2011). Total costs over assets until 1962.

* p¡0.10, ** p¡0.05, *** p¡0.001

Table A.4: **Profit elasticity: subperiods**

	OLS	
	Net Profits (BK)	Market shares (loans)
1890-1930	-.886 ***	-.976 ***
	(.0944)	(.072)
1931-1962	-.798 ***	-.922 ***
	(.1519)	(.1544)
1963-1980	-.693 ***	-.623 ***
	(.1754)	(.2001)
1981-2013	-.552 ***	-1.341 ***
	(.0535)	(.0895)

	FD	
	Net Profits (BK)	Market shares (loans)
1890-1930	.247 ***	-.16 ***
	(.0513)	(.0223)
1931-1962	-.119	-.26 ***
	(.0776)	(.0314)
1963-1980	-.138	-.118 *
	(.0969)	(.0639)
1981-2013	.125 **	-.624 ***
	(.0546)	(.0713)

The estimated parametric model includes a third degree polynomial specification for the independent variable, net profits and market shares. The degree of the polynomial used for the estimation of the gradients and of the standard errors is chosen on the basis of the significance of the coefficients of the polynomial terms, i.e. the non-significant coefficients are dropped in light of a parsimonious specification.

Standard errors in parentheses

Time dummies included in all regressions.

BK=Bos Koetter transformation (2011). Total costs over assets until 1962.

* p¡0.10, ** p¡0.05, *** p¡0.001

Table A.5: Profit elasticity adding controls: subperiods

	OLS	
	Net Profits (BK)	Market shares (loans)
1890-2013	-.46 ***	-.688 ***
	(.0716)	(.0565)
1890-1930	-.597 ***	-.664 ***
	(.0798)	(.073)
1931-1962	-.626 ***	-.724 ***
	(.1401)	(.1326)
1963-1980	-.468 **	-.552 ***
	(.2071)	(.1978)
1981-2013	-.539 ***	-.972 ***
	(.0905)	(.1009)

	FD	
	Net Profits (BK)	Market shares (loans)
1890-2013	.316 ***	-.164 ***
	(.039)	(.0166)
1890-1930	.203 ***	-.14 ***
	(.0538)	(.0215)
1931-1962	-.111	-.255 ***
	(.079)	(.0321)
1963-1980	.04	-.078 **
	(.0999)	(.0389)
1981-2013	.056	-.283 ***
	(.0834)	(.0469)

The estimated parametric model includes a third degree polynomial specification for the independent variable, net profits and market shares. The degree of the polynomial used for the estimation of the gradients and of the standard errors is chosen on the basis of the significance of the coefficients of the polynomial terms, i.e. the non-significant coefficients are dropped in light of a parsimonious specification.

Standard errors in parentheses

Controls: time, geographic and bank type dummies; assets over capital; bad credits over loans.

BK=Bos Koetter transformation (2011). Total costs over assets until 1962.

* p¡0.10, ** p¡0.05, *** p¡0.001