The Long Journey of Bank Competition: New Evidence on Italy from 1890 to 2014

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PRELIMINARY DRAFT

Abstract

We study the evolution of bank competition in Italy over 125 years, from 1890 to 2014. We analyze three policy regimes: the "free banking era" from 1890 to the 1920s; the strong prudential regulatory regime introduced in the 1930s which remained virtually unchanged until the end of the 1970s; and the period of bank deregulation and liberalization that started in the 1980s. We assess competition using the Boone indicator, estimating the elasticity of profits to average costs. The basic idea behind this indicator is that when competition increases, less efficient firms are punished in terms of profits more harshly than efficient ones. According to Boone et al. (2013), the greater the absolute value of the elasticity, the higher the level of competition. We use a unique dataset of bank balance sheets from 1890 to 2014 with an average of around 400 institutions per year. We estimate the Boone indicator following both a parametric and a non-parametric approach. We also analyze the sensitivity of loan market shares to average costs, exploiting the idea that in a competitive market inefficient firms are supposed to lose market shares. We find that competition was high during the first decades of the 20th century, when a sort of free banking was present. The two banking laws in 1926 and 1936 introduced strong barriers to entry in the credit system. As a result, our estimates display a statistically significant decrease in competition: market power remained high until the 1970s. A new rise of competition took place in the 1980s, during the liberalization process. The Boone indicator shows that competition reached its maximum levels along the 125 years in the mid-1990s. In the second half of the 1990s competition decreased, but it soon stabilized during the first decade of the new millennium.

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1 Introduction

The study of competition in the banking industry raises important questions which do not emerge in other economic sectors. The central issue is the hypothesis that aggressive behaviors among banks may lead to instability in the financial system, with negative consequences to the real economy. The answer to this question changed over time and shaped the history of bank regulation, which displays similar patterns in many different countries. For this reason, the Italian case sheds lights on the experiences of other countries.

Until the 1920s, policy makers allowed banks to compete without any constraint. All around the world the Great Depression determined a policy U-turn, motivated by the belief that excessive bank competition causes financial instability. In Italy, the path towards stronger controls had already begun a couple of years before the Great Depression, with the banking law of 1926. A more severe legislation was introduced in 1936 and contributed in the following years to a reduction of competition. Between the mid-1930s and the beginning of the 1980s the banking system was characterized by undoubted stability but the lack of competition contributed to the rise of inefficiencies. For this reason, since the 1980s supervisory authorities inaugurated a liberalization process, up to the banking law in 1993.

Therefore, the emphasis of our paper is on analyzing competition in the long-run, in order to compare different regulatory regimes. We distinguish three regimes: 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. The objective of this research is to study the evolution of competition in these three periods, testing the changes of regime in the literature. Our hypothesis is that rules have influenced the behavior of banks and the degree of competition.

The work is innovative in two respects. First, it provides estimates of competition in the long-run by using a unique dataset of bank balance sheets covering 125 years of Italian history, from 1890 to 2014. Second, competition is measured by a new indicator, proposed by Boone et al. (2013). The idea behind this indicator is that the greater the market power is, the more easily the firms can pass on higher costs to higher prices. In other words, when competition is strong, inefficient firms – those with the highest marginal costs – are severely penalized in terms of profits. The indicator is obtained by calculating the elasticity of profits to marginal costs. Boone et al. (2013) show that this elasticity – expected to be negative – is a measure of competition: the higher the elasticity (in absolute value), the stronger the competition, because a change in efficiency is associated with a higher penalty in terms of profits. In the present work we follow also the approach of van Leuvensteijn et al. (2011) and Tabak et al. (2012) who study, on the basis of such reasoning, the elasticity of market shares to marginal costs: inefficient firms experience wider market shares reductions when competition increases. Because of the lack of detailed data on income statements, we use average costs rather than estimates of marginal costs in our analysis.

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 behavior 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 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 obtained by regressing log profits on log 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 to either an increase or a decrease of these indicators whereas this is not the case for the Boone indicator (before the ap-

[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.

pearance 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. Computing the Boone indicator requires three methodological choices: the definition of costs; the measure of profits; the estimation technique. First, the definition of the cost variable to use might be challenging. 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 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 (OLS, local regression, GMM).

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 interest 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 can coexist.

The main novelties of our paper are two. First, while previous studies took into account a few decades of statistics at best, our paper considers 125 years of observations for a large panel of Italian banks. Banking regulation has drastically changed over time 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. In the following section we illustrate the details of our approach and the main features 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 costs (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 a linear cost function, 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 a measure of competition. As highlighted by Tabak et al. (2012), a firm may use an efficiency improvement to raise profits in two ways: either it may 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:

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

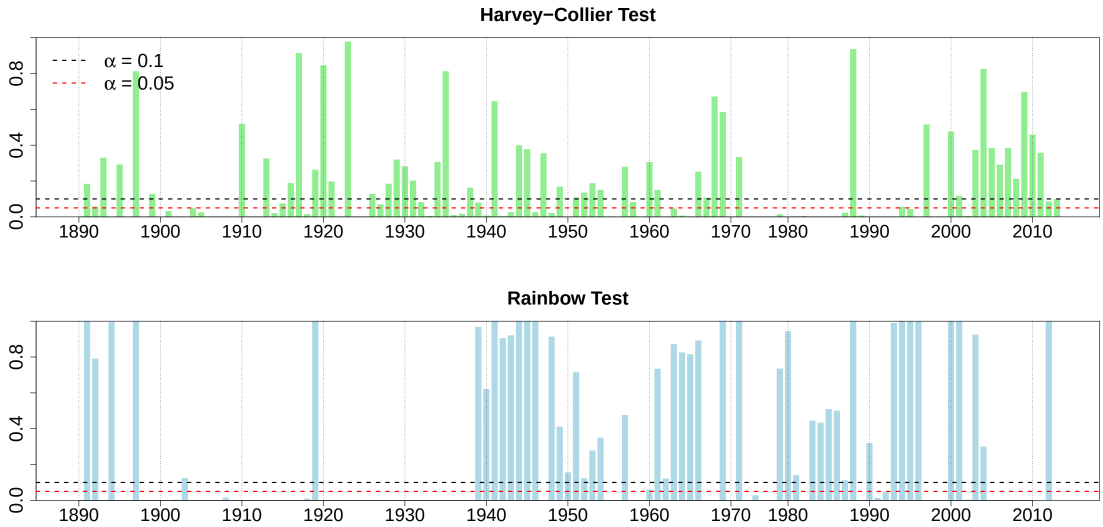
where $\pi_{i,t}$ is either the log of net profits of bank i at time t or the log of its loan market share, $c_{i,t}$ represents the log of the average costs and $\varepsilon_{i,t}$ is a vector of unobservables. $x_{i,t}$ is a vector of control variables, such as dummies for saving and cooperative banks and for geographic location. Our object of interest - the Boone indicator - is the partial derivative of $\pi_{i,t}$ with respect to $c_{i,t}$.

The function f is indexed at time t since we want to estimate the evolution of the partial derivative f_{c_i}. In order to obtain time-dependent estimates, we might split our sample according to different subperiods, but this would require the determination of intervals' extremes. Alternatively, as in De Bonis et al. (2015), we might use rolling-windows estimations, running the regressions over sliding subperiods and attributing the average derivative to the central year. In this paper we opt for year by year regressions. This method does not require arbitrary choices about either the subperiods extremes or the rolling windows length. A major drawback is the estimates' instability due to the high sample size variability. However, we can apply the Hodrick-Prescott filter to the estimated elasticities in order to get a sense of the evolution of competition, disregarding annual changes which are mostly related to sample size issues.

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. Indeed, a visual inspection of the data by year[2] 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 through 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.

[2]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).

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 and the 10% black line indicate that the null hypothesis of linearity has to be rejected for that specific year.

In light of such evidence, we formulate the following parametric specification

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

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. The partial derivative for bank i and year t depends on the level of average costs:

$$\frac{\partial \pi_{i,t}}{\partial c_{i,t}} = \beta_1 + 2 \cdot \beta_2 c_{i,t} + 3 \cdot \beta_3 c_{i,t}^2 \tag{3}$$

We obtain the Boone indicator at the country level by computing the marginal effects for each single bank and then taking the average. We apply cluster techniques at the provincial level to adjust standard errors for geographical correlations.

Our non-parametric specification is

$$\pi_{i,t} = f_t\left(c_{i,t}, x_{i,t}, \varepsilon_{i,t}\right) \quad \forall t \in (1890, 1891, ..., 2014)$$

where we estimate a function f_t for each year in our sample through the local regression approach. 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[3] 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

$$\hat{\pi}_{j,t} = \sum_{i=1}^{n} W\left(\frac{c_{i,t} - c_j}{h}\right)\left[\pi_{i,t} - \theta_0 - \theta_1\left(c_{i,t} - c_j\right)\right]^2 \tag{4}$$

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

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

where $u = \left(c_i - c_j\right)/h(c_j)$. The choice of the values of the parameters we plug into the local regression estimation procedure is not arbitrary but is fully data-driven. The nearest neighbor component of the smoothing parameter is chosen on the basis of a constrained minimization of the generalized cross-validation (GCV) statistic defined as:

$$GCV_t = n\frac{\sum_{i=1}^{n}(\pi_{i,t} - \hat{f}(x_{i,t}))^2}{(n - tr(H))^2}$$

where n is the number of banks in year t and H is the hat matrix of the estimation. The optimal value of the smoothing parameter is then used in the estimation of the gradients for each observation through the robust estimation techniques proposed by Cleveland (1979), according to which outliers are iteratively identified and assigned a lower weight. The local regression returns a vector of fitted values for the gradients - one for each bank in the sample at time t - whose mean represents the estimated value of the Boone indicator for that particular year. Standard errors are derived from the distribution of the estimated coefficients generated through bootstrap techniques. For each year we

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

drew 100 samples with replacement and of the same dimension of the original sample of banks.

In the non-parametric setup, 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 2014. 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)[4];
- from 1977 to 2014 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; 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.[5] 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 1977-2014 are much more detailed than those of the 1890-1973 subperiod, especially for profit and loss accounts[6]. 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

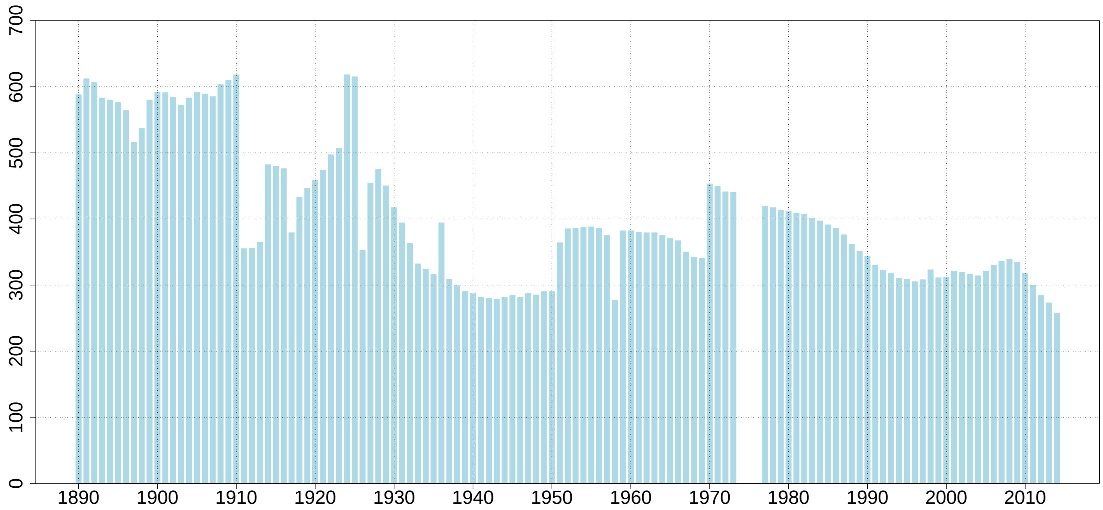
[4]See Natoli et al. (2014) for a full description of the data set.

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

[6]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 1983-2014 and 1995-2014, 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 around 400 institutions. Taking into account that we cover the interval 1890-2014 our dataset include around 50,000 observations.

Figure 2: Number of banks in the sample

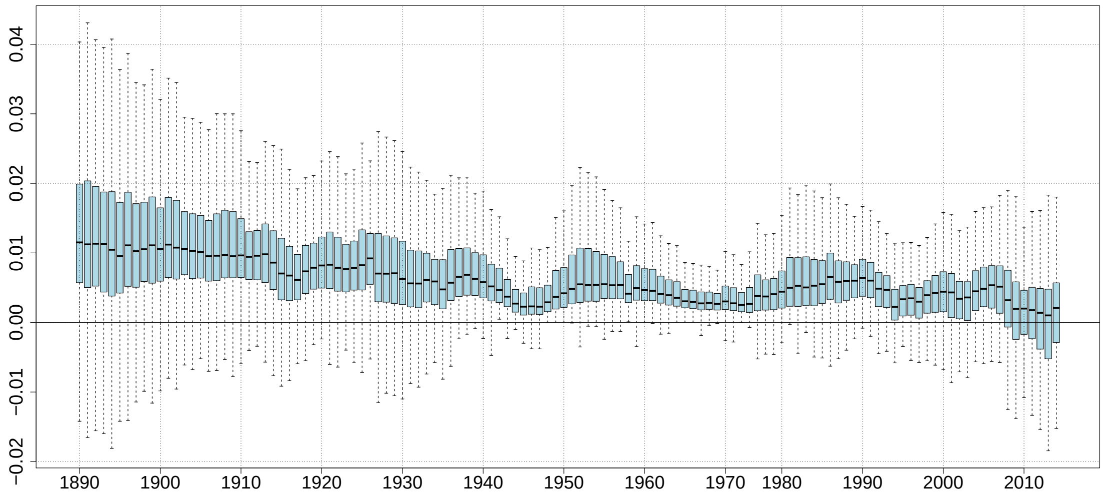


Source: authors' elaborations on Bank of Italy data. Number of banks in the sample of the present paper.

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 2014 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 2013, because of the two recessions that hit the Italian economy after 2008, ROA has reached the lowest level 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.

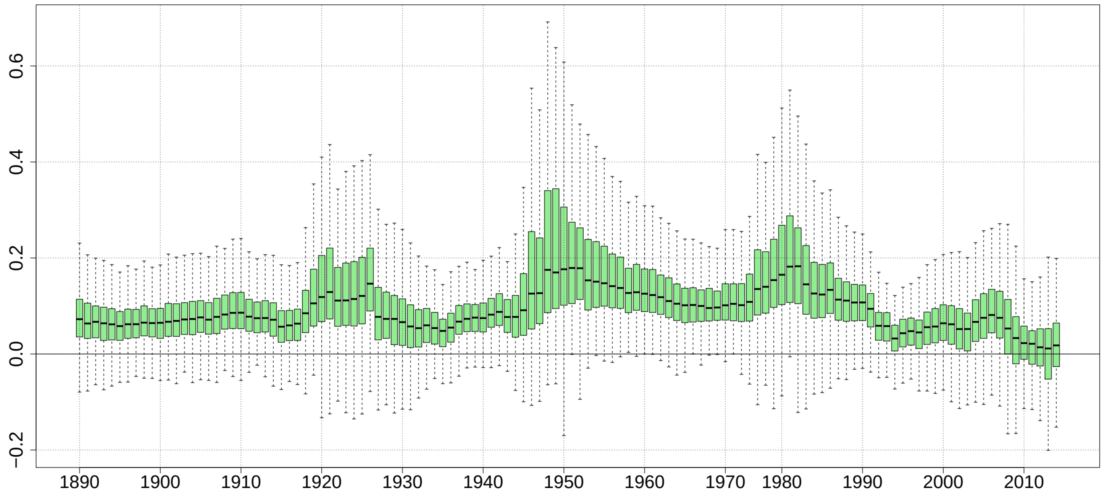
Figure 3: Banks' return on assets (ROA)



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.

In figure 4 we show the evolution of ROE, the ratio of net profits over capital and reserves. Even if high levels of profits do not necessarily imply strong market power, nonetheless ROE offers interesting insights to study competition. 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[7].

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 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. In subsection 4.3 we describe the results using a GMM estimator to deal with potential endogeneity.

[7]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

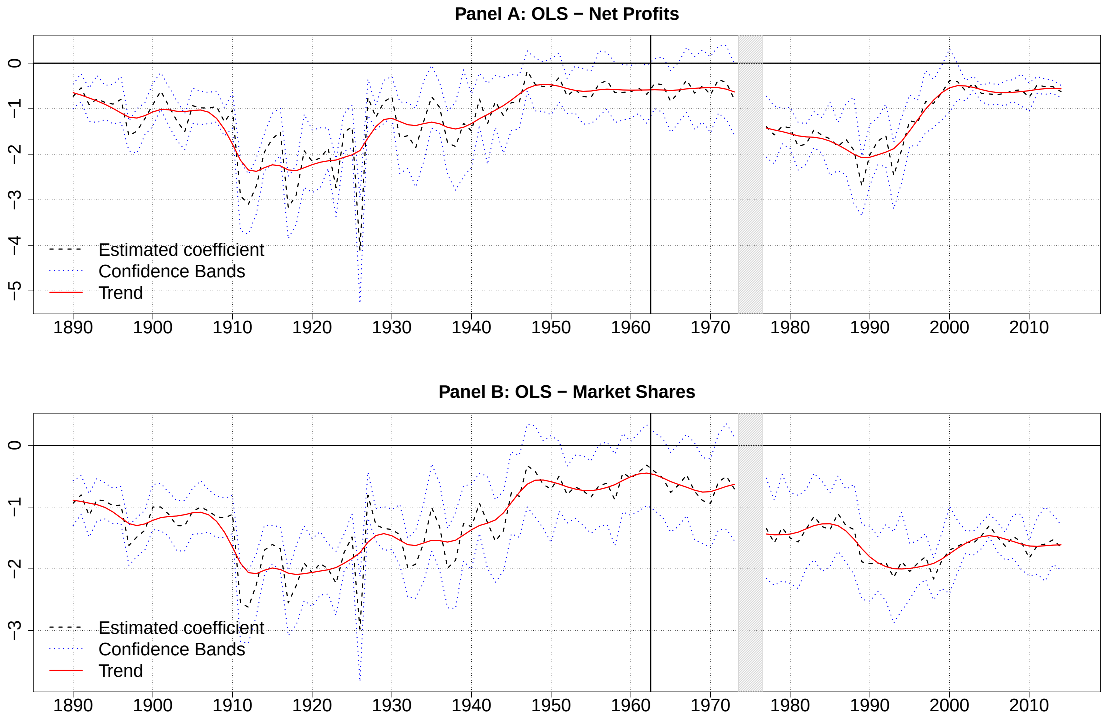
Tables A.3 and A.4 report the estimates of the Boone indicator using, respectively, net profits and loan market shares without adding any control variables. The elasticities are always negative, as expected: an increase in costs is associated to a reduction in profits and market shares. The estimates are generally significant at the 99% level, except between mid-1940s and the beginning of the 1970s. In that period, therefore, the Boone indicator is not statistically different from zero, implying an extremely low degree of competition.

To have a visual representation of the dynamics of the Boone indicator along the 125 years of our sample, we plot the estimates in Figure 5. The dashed line is the estimated elasticity of either net profits (Panel A) or market shares (Panel B) with respect to costs. The dotted lines indicate the extremes of the confidence interval at the 95% level. The solid line smoothes the series of the Boone indicator applying the Hodrick-Prescott filter in order to provide a better intuition of the evolution of competition. The high variation in sample size, especially before 1951, makes the estimates not very stable. Therefore, in order to provide a better intuition of the evolution of competition, the solid line smoothes the series of the Boone indicator applying the Hodrick-Prescott filter. The vertical line signals a break in the average costs variable, which is equal to average total costs until 1962 and to average operating costs as from 1963. We remind that an increase in the elasticity (in absolute values) indicate a reduction of market power.

In both panels we can observe the three regulatory regimes we briefly introduct in section 1: a period of high competition until the 1930s; the decades following the 1936 banking reform characterized by strong market power; the rise of competition after the 1970s, when the liberalization process started. Since the mid-1990s we observe a reduction in competition, which is less pronounced using market shares. This result, which we will discuss below, is in line with the literature on the Boone indicator for the 1990s and 2000s (van Leuvensteijn et al., 2011). Let us analyse the evolution of competition more in details.

In Panel A of Figure 5, the elasticity of net profits to average costs is stable at a value of nearly -1 for the first two decades of our sample. Then it decreases to nearly -2 until the mid-1920s, but the annual estimates are volatile due to sample size variation which is particularly serious after 1911. In that period barriers to entry were virtually inexistent: banks were free to constitute and to open branches as the supervisory controls were absent(Gigliobianco and Giordano, 2012). The number of banks and branches rose: banks fiercely competed for deposits, both offering higher interest rates and opening new branches. This situation raised concerns on banking stability and led to the 1926 bank law, which entrusted the Bank of Italy a discretion on the operations of incorporation, merger, acquisition, and opening branches. The estimates show a slight reduction of competition at the end of the 1920s, which might be the result of the introduction of the 1926 law.

Figure 5: OLS results without control



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 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.

The Boone indicator signals a rapid increase of market power at the beginning of the 1940s: the elasticity is not statistically different from zero for most of the years until the 1970s, which means that the profits did not react to changes in costs. This can be associated to the severe banking regulation that was introduced in 1936 in reaction to bank failures during the Great Depression[8] (Toniolo, 1995). With the banking law, 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. There were constraints on the geographic destination of the loans and their duration. The Bank of Italy used his discretion to restrict competition between banks, favoring the stability of the system. The authorizations were granted mainly to the opening of branches in small banks and in places with poor financial services. Banking regulation underwent changes after the Second World War; nevertheless banking competition remained strongly restricted until the 1970s. For in-

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

stance, 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 Boone indicator is able to identify the low level of competition that characterized that period in which inefficient banks - with high average costs - could remain on the market, thanks to the protection given by the entry barriers and the bank cartel.

Figure 5 shows an increase of competition since the mid-1970s. The index of Boone drops to -1.5 and mantains low levels until the mid-1990s. Several factors contributed to this process. European integration favored a pro-competition climate. The Bank of Italy removed constraints on geographical expansion of loans. The bank cartel was first weakened and then deleted. Through the so-called "Piani Sportelli" in 1978, 1982 and 1986 clear rules for branch openings were introduced, anticipating the full liberalization of branches in 1990. The transformation of state-owned savings banks into joint-stock banks, sanctioned by "Amato" law in 1990, laid the foundations for privatization. The liberalization came to fruition with the Banking Law of 1993.

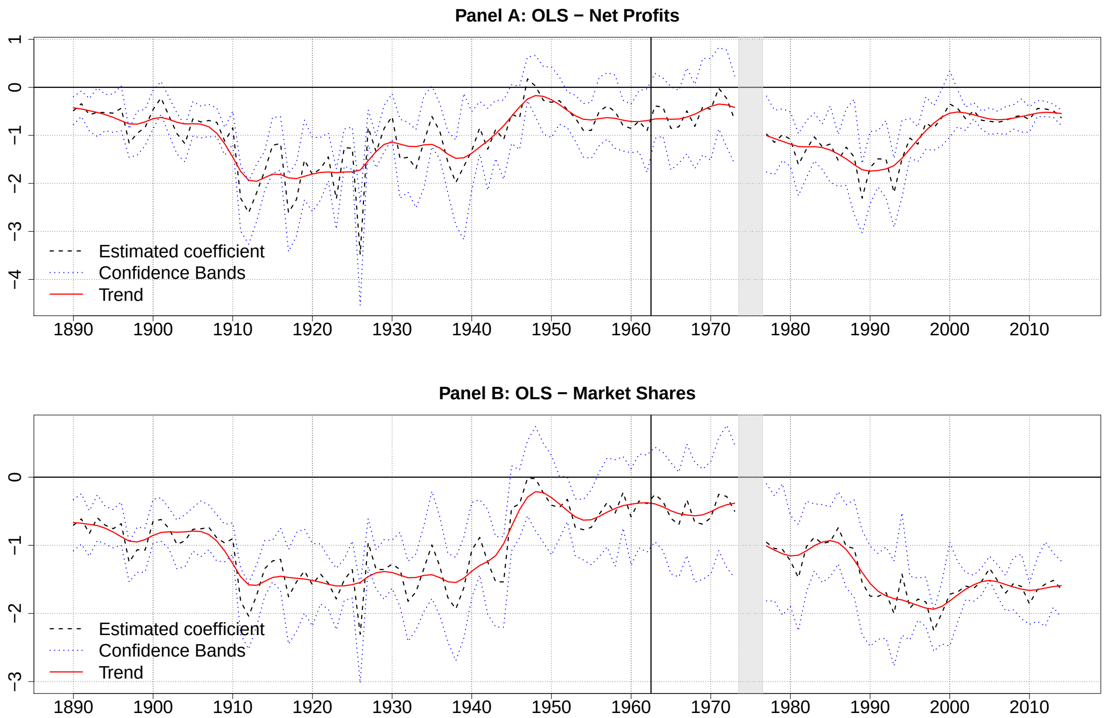
From the second half of the 1990s, the Boone indicator returns on higher levels, like in the 1940-1970s period. However, in contrast with that interval of time, the estimates are statistically different from zero (except in 2000), indicating a significant negative relationship between profits and costs. The reduction of the degree of competition in Italy at the end of the 1990s also emerges in other works (van Leuvensteijn et al., 2011; Clerides et al., 2013; Schaeck and Cihák, 2014) and can be attributed to three factors. First, the wave of bank mergers and acquisitions - which in the 1990s was associated with increased competition because allowed to weaken the old local oligopolies - slowed down towards the end of the decade and stopped in the early millennium. Between 1990 and 2001 occurred more than 400 merger and acquisition, which had a pro-competitive effect, as claimed by Grillo (2006). Between 2002 and 2014 they have dropped to about half of the value of the interval 1990-2001. Second, competition benefited the liberalization of branches in 1990. The annual flow of new openings were intense until the end of the 1990s, before declining towards the end of the decade, when the behavior of the banks became less aggressive. Between 1989 and 2001 bank branches increased by an average of 5.4 percent per year, while between 2001 and 2014 the average growth rate was 0.7 percent. Third, a positive trend of bank profitability in those years may also have influenced a less aggressive interaction among banks. Thanks to the "Dot-Com" bubble between 1995 and 2000 and the development of assets management[9], much of the banking system reached a high profitability level, as shown in Figure 4 by the performance of ROE. The Boone indicator suggests that in those years the most efficient firms were satisfied with their results. They put aside profits as a buffer stock, without exploiting

[9]Saving collected by investment funds in Italy skyrocketed from 65 billions of euros in 1995 to 450 billions of euros in 2000.

their efficiency gains to subtract profit shares to competitors.

In Panel B of Figure 5, the estimates of the Boone indicator using the market shares of loans mostly confirm the results of Panel A. However, there is a relevant departure in the period after 1990s. In fact, competition is particularly strong for the entire decade. Then it slightly weakens in the 2000s but it does not return to the low levels recorded in the 1940-1970s period.

Figure 6: OLS results with control variables

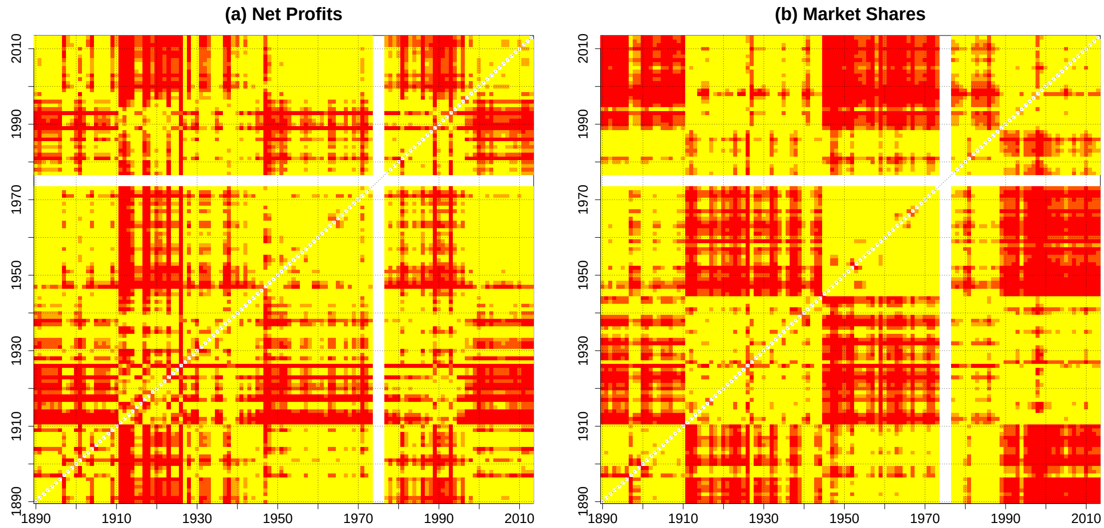


Source: authors' elaborations on Bank of Italy data. The estimated parametric model includes a third degree polynomial specification for the independent variable, net profits and market shares. The vertical black line indicates the break in the average costs variable which is equal to average total costs 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[10]. The path is consistent with what we have already seen in Figure 5. There are few differences looking at the elasticity of market shares with respect to average costs. in particular, it does not display a reduction of competition after the 1926 banking law, which would confirm the thesis that it was not very effective

[10]Results are very similar also including the total assets to equity ratio and the bad debts to loans ratio among control variables. However, we prefer excluding these controls because they might introduce endogeneity in the estimates.

Figure 7: Statistical significance of variations in the Boone indicator



Source: authors' elaborations on Bank of Italy data. Matrices of statistical significance (p-values) of variations in the Boone indicator for each pair of years in the sample.

(Cotula and Garofalo, 1996; Gigliobianco and Giordano, 2012). Moreover, the increase of market power in the 2000s is very small.

After describing the evolution of competition according to the Boone indicator, we check now if the observed variations are statistically significant. The graphs in Figure 7 use different colors to identify the significance level of the elasticity changes for any combination of two years in the sample. For a given pair of years, when the color is yellow, the difference between the elasticities computed in the two years is not statistically significant. The closer the color to red, the more significant the difference is. For example, in Panel A we can focus on the period 1890-1910 and see that it is associated with the 1910-1920s through red color, whereas the area in correspondence with the 1950-1960s is yellow. This means that in general the Boone indicator computed for the 1890-1910 period is significantly different from the one computed for the 1910-1920s, whereas the difference is not statistically significant with respect to the 1950-1960s.

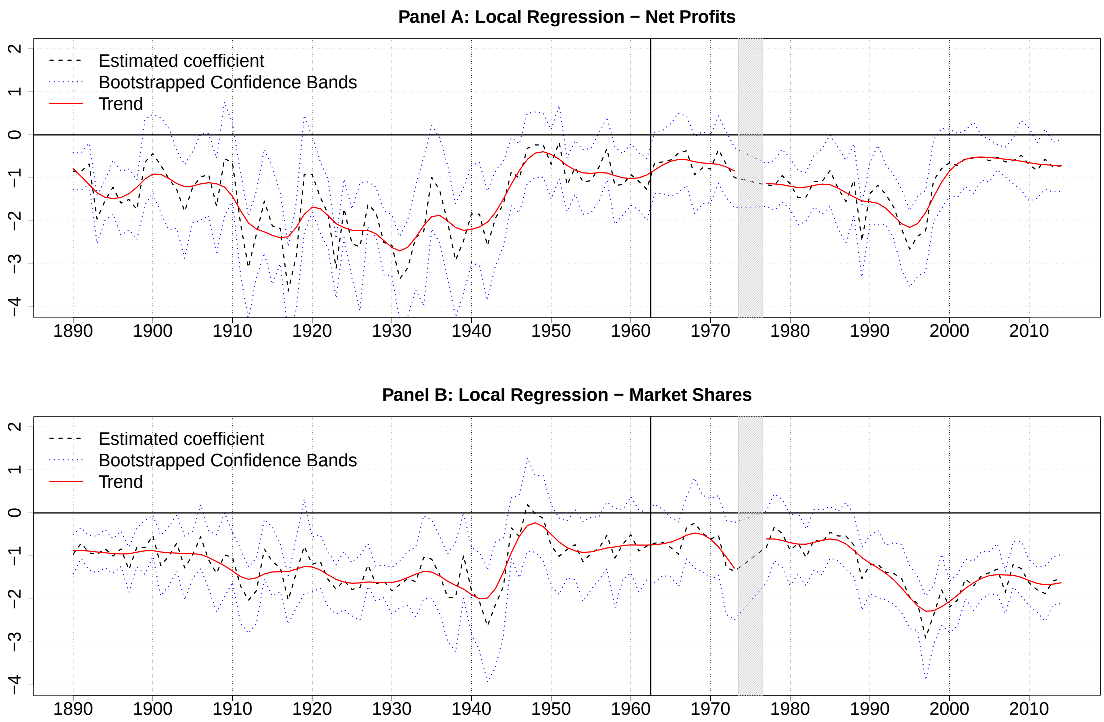
The interpretation is easier when inspecting Panel B of Figure 7, which relates to the Boone indicator computed using market shares of loans in Panel B of Figure 6. After 1910 there is a significant drop in the degree of market power. In the 1950-1970s period, competition is significantly lower than before the banking reform of the 1930s. Market power began to decrease during the 1970s until reaching, at the end of the 1980s, a significantly lower level than in the 1950-1970 period. Differently from the results in Figure 7, competition has never come back to the low levels experienced during the

1950-1960s, and, instead, it is comparable with the degree of the 1910-1930s.

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 2014 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.

Figure 8: 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.

The upper panel of Figure 8 reports the evolution of the elasticity from 1890 to 2014 taking into account the log of net profits as dependent variable. The figure shows that competition increased until the late 1920s, during the period of overbanking and overbranching. Then from the early 1930s until the 1950s competition decreased because of the new emphasis on bank stability and did not change substantially until the mid-1980s. Since then, competition increased 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 stabilization of the level of competition starting from early years of the new millennium [11].

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 8 is roughly similar to that represented in the previous graph. Competition increased more gradually between 1890 and the early 1920s. With the introduction of the two banking laws competition decreased until the 1950s and stabilized until the 1970s. A sharp increase of competition took place from the late 1980s but this trend came to an end in the mid-1990s. After a decade of gradual contraction, competition stabilized again in the years following the financial crisis[12].

4.3 Endogeneity issues

Estimating the elasticity of profits to average costs may raise endogeneity issues. Boone et al. (2013) prove that the dynamics of the indicator is still correct if endogeneity is driven by the dependence of costs on competition. For example, an increase in competition may negatively affect profits and induce managers to reduce operating costs through layoffs. In this case, a change in the elasticity is still interpretable as a variation of competition. However, there are other drivers of endogeneity since in general performance and costs are jointly determined. Tabak et al. (2012) and van Leuvensteijn et al. (2011) apply a GMM methodology, instrumenting costs through their lagged values. We follow the same approach, using the first lag of average costs as instrument.

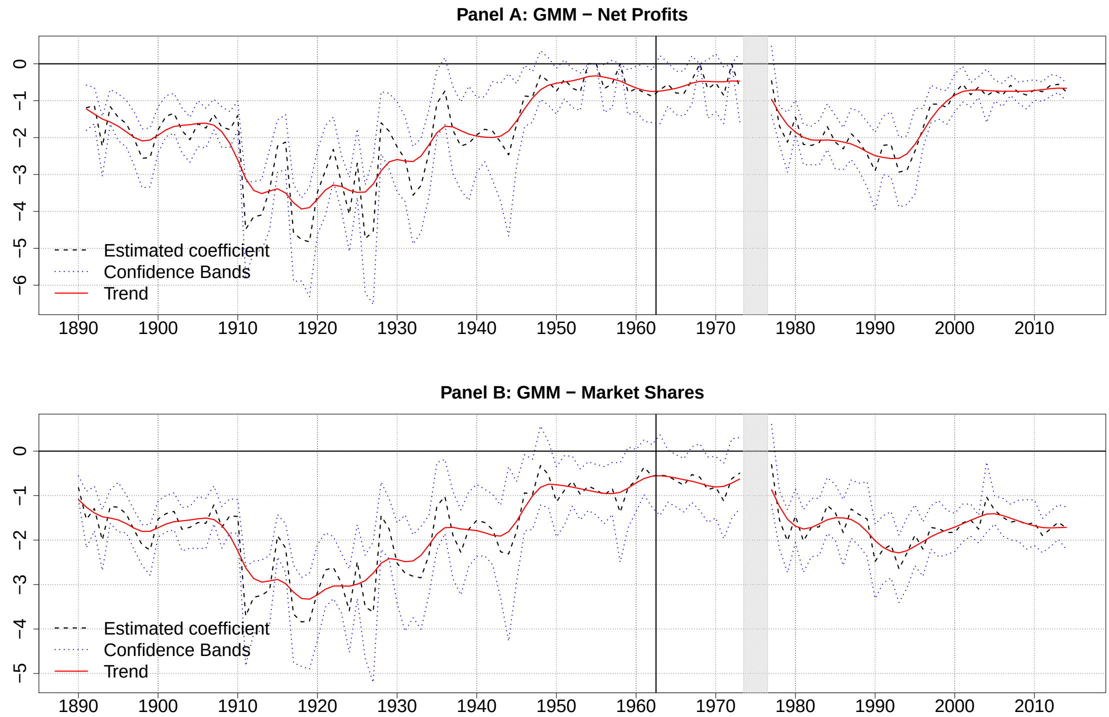
Figure 9 reports the GMM estimates, which are robust with respect to the OLS ones. The range is larger, both computing the Boone indicator using net profits (Panel A) or market shares (Panel B): for example, the elasticities of net profits exceed 4 in absolute values for several years during the 1910-1930 period, whereas the correspondent estimates in Figure 5 are always below 3 (except in 1926). However, the GMM estimates display similar dynamics with respect to the OLS ones (Figure 5). Competition reaches the highest peak during the mid-1920s and then starts decreasing, almost gradually: in the 1950-1960s the Boone indicator is not statistically significant, like in the OLS estimates. The pattern is also comparable after the 1970s, with an increase of competition until

[11] As a caveat, such results may be influenced by the left-truncation of the sample due to the log transformation of net profits when these are negative. However, the parametric estimates show similar patterns both using log profits and the Bos and Koetter transformation.

[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.

the mid-1990s, followed by a remarkable rise of market power using net profits and a slight increase using market shares. Again, there is a divergence in terms of levels: the elasticities for the 1920s are higher in absolute terms than those for the 1990s using the GMM approach, whereas they are almost equal using the OLS methodology. As a caveat, the variability in sample size between 1911 and 1951 may introduce stronger distortion in the GMM estimates than in the OLS ones because it requires dropping the banks which are not present for two consecutive years.

Figure 9: GMM results without 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 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.

5 Conclusions

This paper contributes to the literature in two ways. First, it provides annual estimates of bank competition over more than one century (from 1890 to 2014). A long-run perspective allows to consider different regulation regimes and their influence on the evolution of competition. Second, competition is analyzed using a measure recently introduced, the indicator of Boone, which has never been calculated for the Italian case over such a long period. The Boone indicator is sensitive to the econometric methods and to the linearity hypothesis. For this reason, we use both parametric and non-parametric methods and we check for the impact of endogeneity issues.

At the beginning of our research, we identified three regimes, characterized by varying degrees of competition: an initial period of substantial free banking until the 1920s, when banks were free to compete; a second period, from the 1930s to the 1970s, in which strict constraints to banking were placed in response to the 1920s banks' failures and the Great Depression; a third period, started at the end of the 1970s, when independent initiatives of the Bank of Italy and Community directives led to the elimination of barriers to entry in the credit markets.

The main conclusions of the paper are five. First, taking into account the entire time span 1890-2014, an increase in banking costs is negatively associated with net profits and loan market shares. Second, our estimates show that banking competition was high during the first decades 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. The Boone indicator displays a statistically significant rise of market power at the end of the 1930s, which remained stable until the 1960s. Fourth, since the 1970s – and more remarkably since the mid-1980s in the estimates with market shares and in non-parametric estimates – competition between banks increased: in this period barriers to competition were phased out. The highest levels of competition were observed in the mid-1990s, when the flow of new branches peaked and banking mergers produced pro-competitive effects. Finally, from the second half of the 1990s to the years preceding the global financial crisis, the estimates show a reduction of competition, especially when considering the elasticity of net profits to average costs. The decline of branch openings as well of M&As since the early 21st century – so that their pro-competitive effects reduced – possibly explains the lowering of competition. Moreover, the high profitability, due to the development of asset management and the growth in revenues from services especially during the "Dot-Com" bubble, may have limited the incentive for more efficient banks to use their competitive advantage to subtract profit shares from competitors. Such decrease of competition came to an end in the years following the global financial crisis.

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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	0.41	0.06	0.20	5.29
(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
	1980-2014	0.68	16.93	3.43	12.83	212.09
	1890-2014	0.03	4.63	0.19	1.31	102.09
Loan Market Shares	1890-1929	0.01	0.17	0.03	0.09	0.89
(%)	1930-1979	0.01	0.29	0.03	0.13	1.15
	1963-1979	0.01	0.25	0.04	0.13	0.86
	1980-2014	0.02	0.29	0.06	0.18	0.88
	1890-2014	0.01	0.24	0.04	0.12	0.95
Average Operating	1963-1979	2.31	2.83	2.77	3.29	0.97
Costs (%)	1980-2014	2.09	4.08	2.81	3.60	14.76
	1963-2014	2.18	3.70	2.80	3.49	12.39
Average Total	1890-1929	3.26	4.69	4.34	5.53	5.76
Costs (%)	1930-1979	3.72	5.33	4.95	6.40	3.20
	1963-1979	5.19	6.67	6.32	8.07	2.38
	1980-2014	4.25	8.79	7.57	10.27	16.15
	1890-2014	3.72	6.15	5.10	7.19	9.41
Bad Debts	1890-1929	0.00	3.23	0.11	2.04	9.45
to Loans Ratio	1930-1979	0.00	1.04	0.05	0.74	3.39
(%) (2)	1963-1979	0.55	2.83	1.52	3.37	4.46
	1980-2014	1.39	5.86	4.01	7.37	8.17
	1890-2014	0.00	3.34	0.61	3.59	7.87
Leverage Ratio	1890-1929	4.33	11.32	8.00	13.63	12.02
(3)	1930-1979	11.07	31.72	19.22	37.67	37.55
	1963-1979	22.94	41.41	34.76	52.42	26.16
	1980-2014	9.99	40.91	15.38	25.65	216.71
	1890-2014	6.98	26.13	13.55	25.98	108.17

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	26	35	21	31
	H_1	24	36	28	41

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: Boone indicator using net profits (OLS)

Year	Coeff.	Year	Coeff.	Year	Coeff.	Year	Coeff.
1890	−0.726***	1921	−2.082***	1952	−0.720***	1985	−1.659***
1891	−0.541***	1922	−1.868***	1953	−0.578**	1986	−1.830***
1892	−0.911***	1923	−2.736***	1954	−0.731**	1987	−1.671***
1893	−0.785***	1924	−1.514***	1955	−0.754**	1988	−1.914***
1894	−0.861***	1925	−1.399***	1956	−0.455	1989	−2.697***
1895	−0.898***	1926	−4.116***	1957	−0.381	1990	−2.010***
1896	−0.795***	1927	−0.750***	1958	−0.648*	1991	−1.713***
1897	−1.595***	1928	−1.187***	1959	−0.639**	1992	−1.587***
1898	−1.491***	1929	−0.851***	1960	−0.626**	1993	−2.467***
1899	−1.218***	1930	−0.734***	1961	−0.549*	1994	−1.896***
1900	−0.912***	1931	−1.634***	1962	−0.685**	1995	−1.263***
1901	−0.611***	1932	−1.592***	1963	−0.446	1996	−1.309***
1902	−0.939***	1933	−1.855***	1964	−0.472	1997	−0.843**
1903	−1.259***	1934	−1.279***	1965	−0.842**	1998	−0.882***
1904	−1.504***	1935	−0.736**	1966	−0.643*	1999	−0.653**
1905	−0.935***	1936	−0.972***	1967	−0.367	2000	−0.382
1906	−0.981***	1937	−1.751***	1968	−0.657	2001	−0.402**
1907	−0.985***	1938	−1.830***	1969	−0.499	2002	−0.606***
1908	−0.952***	1939	−1.320**	1970	−0.694	2003	−0.441***
1909	−1.289***	1940	−1.488***	1971	−0.356	2004	−0.663***
1910	−1.034***	1941	−0.788***	1972	−0.428	2005	−0.676***
1911	−2.913***	1942	−1.337***	1973	−0.800**	2006	−0.689***
1912	−3.095***	1943	−0.842***			2007	−0.663***
1913	−2.690***	1944	−1.151***	1977	−1.394***	2008	−0.600***
1914	−1.953***	1945	−0.865***	1978	−1.574***	2009	−0.587***
1915	−1.659***	1946	−0.851***	1979	−1.384***	2010	−0.749***
1916	−1.513***	1947	−0.170	1980	−1.421***	2011	−0.479***
1917	−3.168***	1948	−0.465	1981	−1.822***	2012	−0.515***
1918	−2.891***	1949	−0.513*	1982	−1.776***	2013	−0.518***
1919	−1.924***	1950	−0.517	1983	−1.460***	2014	−0.625***
1920	−2.155***	1951	−0.314	1984	−1.585***		

Source: authors' elaborations on Bank of Italy data. The coefficients displayed in this table are the marginal effects obtained through annual regressions. The controls included in each regression are: the Bos and Koetter control variable; region and bank type dummies.

Table A.4: Boone indicator using market shares (OLS)

Year	**Coeff.**	**Year**	**Coeff.**	**Year**	**Coeff.**	**Year**	**Coeff.**
1890	−0.935***	1921	−1.912***	1952	−0.796***	1985	−1.366***
1891	−0.801***	1922	−2.000***	1953	−0.671**	1986	−1.105***
1892	−1.133***	1923	−2.234***	1954	−0.735**	1987	−1.299***
1893	−0.879***	1924	−1.738***	1955	−0.837***	1988	−1.363***
1894	−0.902***	1925	−1.483***	1956	−0.652*	1989	−1.888***
1895	−0.978***	1926	−2.979***	1957	−0.612*	1990	−1.917***
1896	−0.967***	1927	−0.805***	1958	−0.884**	1991	−1.917***
1897	−1.621***	1928	−1.287***	1959	−0.444	1992	−1.899***
1898	−1.483***	1929	−1.345***	1960	−0.532*	1993	−2.140***
1899	−1.376***	1930	−1.363***	1961	−0.430	1994	−1.879***
1900	−0.999***	1931	−1.447***	1962	−0.319	1995	−2.042***
1901	−0.998***	1932	−1.983***	1963	−0.425	1996	−1.911***
1902	−1.116***	1933	−1.924***	1964	−0.526	1997	−1.802***
1903	−1.302***	1934	−1.580***	1965	−0.762**	1998	−2.165***
1904	−1.302***	1935	−1.006***	1966	−0.638*	1999	−1.876***
1905	−1.070***	1936	−1.320***	1967	−0.479	2000	−1.695***
1906	−1.000***	1937	−1.988***	1968	−0.744*	2001	−1.639***
1907	−1.062***	1938	−1.861***	1969	−0.893**	2002	−1.565***
1908	−1.156***	1939	−1.264***	1970	−0.939**	2003	−1.583***
1909	−1.174***	1940	−1.317***	1971	−0.592	2004	−1.465***
1910	−1.116***	1941	−0.938***	1972	−0.501	2005	−1.299***
1911	−2.552***	1942	−1.246***	1973	−0.704	2006	−1.480***
1912	−2.622***	1943	−1.552***			2007	−1.633***
1913	−2.252***	1944	−1.386***	1977	−1.343***	2008	−1.473***
1914	−1.697***	1945	−0.768**	1978	−1.578***	2009	−1.562***
1915	−1.605***	1946	−0.825**	1979	−1.335***	2010	−1.827***
1916	−1.674***	1947	−0.328	1980	−1.501***	2011	−1.626***
1917	−2.552***	1948	−0.421	1981	−1.565***	2012	−1.598***
1918	−2.270***	1949	−0.632*	1982	−1.358***	2013	−1.529***
1919	−1.909***	1950	−0.709	1983	−1.149***	2014	−1.649***
1920	−2.062***	1951	−0.497*	1984	−1.315***		

Source: authors' elaborations on Bank of Italy data. The coefficients displayed in this table are the marginal effects obtained through annual regressions. The controls included in each regression are region and bank type dummies.