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START-UP BANKS' DEFAULT AND THE ROLE OF CAPITAL

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Abstract

Regulation requires banks to hold a minimum capital endowment upon their establishment. But what role does initial capital play in a bank's lifecycle? This paper addresses the issue for start-up banks. We use both survival-time and binary choice models for a sample of newly-established Italian banks in the period 1994-2006, controlling for a broad set of possible drivers of default, such as market, managerial and financial variables. Our results suggest that initial capital does play a leading role in explaining both the timing and the likelihood of a failure. Other important drivers are organisation and a balanced growth path, while market and management variables appear to play a minor role. We then turn to a quasi-experimental design: exploiting a regulatory shift in 1999 we run a counterfactual analysis of the impact of a regulatory tightening of initial capital, which affected only a subsample of banks. The set of results suggests that the effect on banks' survival may be significant.

JEL Classification: G21, G28.
Keywords: bank capital, survival analysis, probability of default analysis, start-up banks, counterfactual analysis.

Contents

* Bank of Italy, Banking and Financial Supervision.

1. Introduction[1]

Banking failures have been extensively analysed in the literature;[2] conversely, the default of start-up banks has drawn less attention. Embarking on this stream of research, DeYoung (2003a) identifies a cyclical pattern in their exit behaviour. As determinants of the exit, DeYoung (2003b) finds that intense competition, slow economic growth and urban location are good predictors of start-up banks' default. Evidence on the role of macroeconomic conditions is also available in Nuxoll et al. (2003) and in Porath (2006).

A small number of empirical studies examine start-up banks in the Italian banking market. By focusing on Italian mutual banks established during the 1990s, Maggiolini and Mistrulli (2005) find that their survival is positively related to a set of market characteristics, such as the market share of incumbent large banks and the absence of other mutual banks;[3] the local level of GDP also positively affects the survival probability. Santarelli (2000) tests the impact of a regulatory reform on the size of start-up banks: he runs a duration analysis on two cohorts of entrants in the Italian financial market in the years bridging the introduction of the Italian Banking Law of the early 1990s. His results suggest that the lifting of constraints on both branching and new bank formation fostered a pre-entry selection process.

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[2] See, among others, Cole and Gunther (1995), Demirguc-Kunt and Detragiache (1998), Wheelock and Wilson (2000).

[3] In a related work, Cocozza and Lozzi (2006) explore market characteristics which may promote the build up of a new bank; they find that new banks often spring from large, concentrated markets, particularly those affected by bank mergers in the previous years.

Our work adds to this strand of literature and explores in more depth the role played by initial capital endowment. In most countries, entrepreneurs wishing to establish a new bank must comply with a set of regulatory requirements in terms of corporate governance and financial soundness.[4] With respect to the latter, bank start-ups are commonly required to hold a minimum amount of capital.

From a market perspective, bank capital addresses the higher uncertainty a new bank is likely to face; indeed, it should preserve the new bank from potential losses during the start-up period, thus softening information asymmetries and increasing market confidence (Berger et al., 1995; Elizalde and Repullo, 2007). Furthermore, it has been broadly related to enhancing stability (Diamond and Rajan, 2000) and sound management (Hellmann et al., 2000).

From the regulators' perspective, capital is intrinsically linked to banks' stability, which is one of the main supervisory objectives. Thus, by assessing the relationship between capitalization and stability, a first result of our work is to provide banking regulators with a tool for measuring the potential benefit of different endowments of initial capital.

As a matter of fact, not only is capital the cornerstone of prudential regulation, but it is also a key regulatory parameter for licensing new banks. Within the European Union, for instance, the Capital Requirement Directive provides for a minimum initial capital of €5 million, with some exceptions for particular categories of credit institution.[5] Among EU Members some countries, such as Germany and France, have aligned their national regulations to the EU general threshold; others, such as Italy, have set a higher general

[4] See the survey by Barth et al. (2006).

[5] Member States may grant authorisation to particular categories of credit institution, but their initial capital shall be no less than € 1 million.

threshold (€ 6.3 million) while simultaneously allowing a particular category of banks (mutual banks) to hold a lower minimum initial capital.[6] Outside Europe we find slightly different fixed thresholds, such as for Japan (¥ 1 billion) and Canada (5 million Canadian dollars); in the U.S. initial capital requirements are expressed in relative terms, as add-ons to the solvency ratio to be held for the first years.[7]

Regulators face a clear trade-off in setting the floor for initial capital. Regarding bank stability, they should not raise an overly high entry barrier, thus preventing a sound source of competition from being introduced; an overly demanding threshold would also discourage investors from starting up new bank projects.

This paper aims at identifying a role for initial capital and at quantifying the benefit of different minimum levels for start-up banks.[8] We adopt two complementary approaches. First, we estimate the survival function and the probability of default of start-up banks.[9] While our estimation of default probability is primarily aimed at addressing the role of capital in avoiding "sudden death" (in our definition, a bank default in its first five years of life), the survival analysis is intended to highlight the role of capital endowment in postponing a bank failure. Second, we carry out a counterfactual analysis – by exploiting a regulatory shift for a subset of Italian start-up banks – in order to analyse the effectiveness of a regulatory tightening of initial capital.

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[6] The minimum capital for mutual banks was set at ITL 2 billion (about € 1 million) until the end of 1998 and it was raised to € 2 million thereafter.

[7] DeYoung (2003a) explains how U.S. Federal regulators imposed higher minimum capital requirements on newly-established banks in response to the numerous wind-ups during the late 1980s and early 1990s; the Federal Deposit Insurance Corporation (FDIC), for example, required all new banks to maintain a tier-1 capital ratio at 8 per cent for the first three years, doubling the 4 per cent tier-1 capital requirement requested for incumbent banks in order to be considered adequately capitalized.

[8] The relative costs and competition effects are beyond the scope of our work.

[9] See Lennox (1999) for a survey of default analysis methods; Lane et al. (1986) for survival time methods.

The approach used in our exercises generates a broad set of empirical results. Based on these findings, initial capital endowment plays a leading role in determining banks' probability and timing of default; the effectiveness of banks' organisation and a balanced expansion in the first years of life produce the same positive and significant effects, while directors' experience seems to be quite trivial. These results shed light on the efficacy of regulatory measures in addressing policy objectives (namely, start-up banks' stability). More generally, regulation seems to play a role: according to the counterfactual analysis, an increase in regulatory initial capital provides positive results (in terms of reducing the probability of default) for treated banks with respect to a control group.

The remainder of this paper is organised as follows: in Section 2 we present the dataset, the variables of interest and the statistical methods employed. Sections 3 and 4 show the results of survival and probability of default analyses, respectively. Section 5 presents a counterfactual analysis, conducted by a difference-in-difference model. Before concluding, Section 6 presents some robustness checks aiming at supporting the results of the previous sections.

2. Data

2.1. The sample

We use a unique dataset of start-up banks, established in Italy between 1994 and 2006. The reasons for choosing this sampling period are threefold: first, the homogeneity of the regulatory framework; second, the availability of data; and third, the need for a suitable period in which to observe a default event for those banks established in more recent years. We analyse the banks established immediately after the adoption of the Italian Banking

Law (*Testo Unico Bancario*) of 1993[10]; this time span allows us to observe the initial stage of activity for a large number of banks. We include in our sample banks established up to end-2006, in order to monitor them for at least five years.[11]

We use a narrow definition of "start-up bank" by excluding any new banks that carried on the business of previous existing subjects already operating in the financial system. In particular, we do not consider: i) branches of foreign banks; ii) banks resulting from extraordinary operations, such as transformations, mergers, spin-offs, sales of branches; and, iii) banks established within an already operative banking group.

The resulting sample is composed of 119 banks. Most of them (58 per cent) are mutual banks (*MUTUAL*, *banche di credito cooperativo*) while commercial banks (*COM*, *banche spa*) and cooperative banks (*COOP*, *banche popolari*) account for 23 per cent and 19 per cent of the total, respectively (Table 1); South of Italy attracted more mutual banks, while most commercial banks were established in the North.

Next we identify the subset of banks gone out of business for reasons assimilated to default. The concept of mortality is interpreted in a restrictive manner. We consider among the causes of default: i) liquidation (voluntary or compulsory); ii) special administration; iii) the sale of assets and liabilities; and, iv) some cases of mergers and acquisitions.[12] As

[10] Gobbi and Lotti (2004) underline that a first wave of liberalisation hit the banking system in 1989, when the First Banking Directive was implemented. New legislation allowed the establishment of new banks, notwithstanding issues relating to by-laws, minimum capital endowments and personal conduct for both founders and management. These solutions represented a major change to the previous system, where by the establishment of new banks was limited to mutual institutions and the subsidiaries of foreign banks.

[11] We observe bank default taking place within the first semester of 2011.

[12] In the definition of start-up bank default, we considered only the following cases of mergers and acquisitions: i) when a mutual bank merged with a bank of a different nature: according to the Italian Banking Law, this is subject to the approval of the Bank of Italy, only admissible in the interests of creditors and having regard to the need for stability, or; ii) when, before the merger, the target bank displayed tensions in its

summarized in Table 1, 60 per cent of our sample defaulted; more than a half of the exits were within the bank's fifth year of life. Of this 60 per cent, only 25 per cent exited the market following or within the context of an official crisis procedure (special administration or compulsory liquidation). The remaining 34 per cent of bank exits ended in a market solution, mainly in M&A.[13]

2.2. *Initial capital and other relevant variables*

Regardless of the approach we use, we investigate the relation between a bank default and its initial capital endowment. Table 2 shows that the average initial capital for the whole sample is € 5.7 million: € 1.9 million for mutual banks, € 8.6 million for cooperative banks and € 12.9 million for commercial banks.[14]

The regulatory minimum level of initial capital differs among bank categories: since 1999, it has been set at € 2 million for mutual banks and € 6.3 million for both cooperative and commercial banks.

Admittedly, estimating a simple model of the form *Default* = f(*initial capital endowment*) would lead to a spurious regression: the estimated relation between capital and default could be due to unobserved characteristics. Thus, in order to insulate the causal effect of capital on default, we complement initial capital with a broad set of information. These additional variables are meant to capture the effect of banks' management, organisational effectiveness, market characteristics and financial features in determining a default. Controlling for these additional variables addresses a key challenge of our research:

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capital base and/or in profitability, that is, reported a capital ratio about to breach the regulatory minimum and/or losses in its income statement.

[13] M&As have sometimes been recommended by the Supervisory Authority as a resolution tool.

[14] The observed values refer only to banks that have successfully passed the licensing process by the Bank of Italy and thus to a sample of selected banks. Accordingly this sample has better features than the universe of new bank projects (for which data are unavailable).

that is, how to disentangle the effect of initial capital from that of other aspects likely to be related to the success of a start-up bank. A list of all the variables used, and their definition, is reported in Table 3.

Before introducing the controls, we need to clarify further the timing of their observation period. Most of the variables are observed at the time of the start-up (or close after), thus mimicking the information set at the supervisor's disposal when the bank applies to be established. Two of them (i.e. *BUSINESS_MIX* and *DEV_ASST_GWT*) can only be observed during the first years of life.

Bivariate statistics reported in both Table 1 and Table 2 confirm that mutual banks have a lower average initial capital (mainly due to their lower licensing requirement) and a higher default rate. Evidently, other reasons than lower initial capital may account for their higher default rate; it may be that some unobserved heterogeneity feeds the capital effect even if it is unrelated to capitalization. This is why we introduce the variable *MUTUAL*. This dummy variable captures all the unobservable characteristics associated with mutual banks and allows us to check whether being a mutual bank has some independent effect on default, uncorrelated with the (lower) capital level.

Conventional wisdom suggests that a greater level of board competence and experience allows for more effective action and improves firm performance.[15] We gathered information about management by exploiting the Bank of Italy's "*Organi Sociali*" (Or.So.) database, which collects data on Italian banks' board members for supervisory purposes. Relevant variables include previous experiences in the financial sector and education. Table 3 shows, respectively, the percentage of board members with previous experience in the

[15] Hermalin and Weisbach (2003) provide a comprehensive survey on the topic. For the banking industry, see Berger et al. (1993).

banking industry and university academic training. In our analyses these data take the form of dummy variables, signalling higher-than-average expertise (*EXPER*) and education (*EDUC*) of the bank's board.

The performance of a bank's organisation – in terms of process design, internal control, information flows etc. – is also assessed by the Italian supervisor. *ORG_SCORE* is the score assigned by the Bank of Italy to the bank's organisational profile[16] in the context of its annual Supervisory Review and Evaluation Process.[17] The score is based on off-site information available to line-supervisors as well as meetings held with the management and on-site examination results. We refer to the score assigned to each bank of our sample at its first year of life.

Market characteristics can also affect a bank start-up.[18] Accordingly, we consider market concentration, proxied by the Hirschman-Herfindahl index of branches (*HHI*), and urbanisation (*URB*), as captured by an index ranging from 1 (less urbanised) to 3 (highly urbanised). Additionally, we use GDP per capita *(GDP_PC)* as a measure of competitiveness and economic development at local level. To take into account all the externalities related to geographical position properly, the dummy variable *SOUTH* indicates the banks established in Southern Italy. Furthermore, to take into account the effects of the economic cycle, *GDP_GWT* controls for GDP growth over the two-year time span before the start-up.

[16] Up to 2007 the score was related to the "organisation" profile and was based on a 4-level scale (the higher the score, the worse the situation). Since 2008 the score has been related to the "governance and control systems" profile and is based on a 6-level scale; for reasons of homogeneity, extreme values (1-2 and 5-6) are kept together, thus turning to a 4-level scale.

[17] This variable is equivalent to the managerial decision-making component of the U.S. CAMEL rating system.

[18] See DeYoung (2003a), Maggiolini and Mistrulli (2005).

Finally, a set of balance-sheet indicators aims at capturing the bank's strategy. *DEV_ASST_GWT* indicates the distance (in absolute value) of the first five years asset growth rate from the sample median value; we expect an unbalanced growth (either too fast or too slow) to be detrimental to the bank. On the one hand, an exuberant asset development may jeopardize credit quality; on the other hand, by delaying the break-even point excessively slow growth can be economically unsustainable. *BUSINESS_MIX* is intended to describe the business model of a bank: it captures the weight of retail lending (i.e., lending to households and small-medium enterprises) on total assets.

3. Survival analysis

As a first step in examining the role of capital in a start-up bank's life cycle, we run a broad set of survival analysis regressions. These analyses examine how initial capital endowment may affect the timing of a default event. Duration models are based on the probability of observing a default, conditional on a bank's survival up to that point in time.[19] Given that a bank is alive at time *t*, survival models give the probability it will remain alive during a subsequent time span, whereas the hazard function gives the probability of failure in that period.[20] Therefore these models differ from binary outcome

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[19] The survival function is a cumulative distribution function (cdf): defining the time until death as *T*, the survival function can be defined as $S(t) = \Pr(T > t)$, with $0 \le S(t) \le 1$, $S(0) = 1$ and $dS(t)/dt \le 0$. It describes the probability that the variable *T* will be greater than any chosen value of *t*. In this setting, the survival function describes the probability of surviving beyond *t*.

[20] Hazard is defined as $h(t) = \lim_{\Delta t \to 0} \frac{\Pr(t \le T \le t + \Delta t \mid T \ge t)}{\Delta t}$ and can be expressed as a function of the pdf of *S(t)*: $h(t) = -\frac{d}{dt} \log S(t)$.

For a given *T*, the hazard function quantifies the instantaneous risk that an event (in this context, a bank default) will take place at time *t*, conditional on the individual having survived up to time *t*. A convenient way of representing it is by taking the reciprocal of its value, *1/h(t)*: this represents the length of time until a given event (default) occurs, by assuming the hazard remains constant over this time span.

models, focused on the unconditional probability of default within a pre-defined time horizon.

As a descriptive analysis, the Kaplan-Maier Survival curve[21] (Chart 1) shows that the empirical survival rate's estimate faces a first downward jump just before the 10^{th} semester; then it recovers and the curve flattens accordingly, up to the 20^{th} semester, when the survival probability is close to 35 per cent. Afterwards, the survival probability remains fairly constant until the end, with a final value of 28 per cent.

The preliminary assessment through a graphical inspection of the Kaplan-Maier survival curve reveals that hazard is non-monotonic: it starts by rising and then it declines. To test some hypotheses regarding the shape of the hazard function directly, we start by using parametric regression models (3.1). The parametric regression models come with a cost, in the form of estimation inefficiency in case of misidentification of distributional form. To circumvent this additional issue, we then move to a semi-parametric estimation, by using a Cox regression model (3.2).[22]

3.1. Parametric model

We begin our analysis of duration by fitting a set of parametric models to our sample. We conduct this analysis through accelerated failure time (*AFT*) models. Besides analysing

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[21] The Kaplan-Meier estimator of the survival function can be expressed as $\hat{S}(t)=\prod_{j\leq t}\frac{r_j-d_j}{r_j}$ where r_j indicates the number of individuals at risk and d_j the number of exits at each time t. Considering the absence of regressors, this method is often referred to as non-parametric estimation. Calendar time is ignored: a different establishment date plays no role in this scenario.

[22] In analysing the hazard function other approaches are possible. For example, DeYoung (2003a) analyses this phenomenon by estimating basic logit regression over rolling intervals of a specified length (e.g. 3 years). The exit probabilities generated in these models are interpreted as hazard rates as they are conditional on a bank surviving up to the start of the exit window.

the effect of covariates on survival time, parametric models allow us to check how closely the model fits the estimation sample. Unlike semi-parametric proportional hazard models (such as the Cox model used in the next section), *AFT* models have the advantage of allowing us to explicitly test the shape of the hazard function.

We estimate an *AFT* model in the form:

$$\begin{aligned}\ln[T_b] &= \alpha + \beta_1 CAP_b + \beta_2 MUTUAL + \sum_{i=1}^{n} \gamma_i (management_b) + \\ &+ \sum_{i=1}^{n} \delta_i (market_b) + \sum_{i=1}^{n} \varphi_i (balance\_sheet_b) + \sum_{y=1}^{Y} \phi_y (dummy\_year_n) + \sigma\varepsilon_b\end{aligned} \quad [1]$$

where T_b is the time until the default of bank *b*; *CAP* – the log of capital at the start-up – is our variable of interest; *MUTUAL* is a dummy variable identifying mutual banks; *management, market* and *balance_sheet* are vectors of bank- and market-specific regressors; *dummy_year* is a vector of 12 dummy variables, valorised from 1994 to 2005, whose aim is twofold: taking into account both the different time at risk (deriving from a different date of establishment) and the business condition; σ the shape parameter that allows for different distributions of ε, besides the normal one, while still retaining the assumption of constant mean and variance over *b*.

We estimate model [1] under four alternative hypothetical distributions: gamma, Weibull, exponential and log-normal (see Appendix *A* for details).[23] A different distributional form affects the hazard function; in turn, this yields different substantive interpretations related to start-up banks' lifecycle (DeYoung, 2003a).[24] A constant hazard

[23] In the text, we refer to the distribution of *T* rather then to the distribution of the disturbance term.

[24] An exponential distribution implies that the hazard of default is constant throughout time. A Weibull distribution would imply a hazard of default that can follow different paths over time, according to a scale

would indicate that the risk of default is independent of a bank age; otherwise, a non-monotonic hazard function would imply that start-up banks experience a period of financial fragility during which they are more vulnerable to failure.

To test the distributional form, we run a Likelihood Ratio (LR) test on nested models (Weibull, exponential and log-normal) from the baseline gamma model of the form:

$$LR = -2 \log(L_{\text{Baseline\_Gamma}}) - (-2 \log(L_{\text{Nested\_model}}))$$

that is distributed as a χ^2 under the null hypothesis of the best fitting of the sample.

The bottom row of Table 4 shows the LR test results.

While the tests reject the assumption of both monotonic and constant hazard, there is (limited) evidence of non-monotonic, twisted U-shaped hazard, as implied by a log-normal distribution. An analysis of probability plots (Chart 2) for each distribution confirms this finding: estimated values for the empirical distribution closely fit the plot of theoretical log-normal distribution; the fitting for all the other alternative distributions is less adequate.

These results suggest that newly-established banks face an initial period of financial vulnerability during which they are more exposed to the risk of failure. This finding is consistent with DeYoung's life-cycle theory of start-up bank failure (2003a): the probability of default increases at first and then declines the longer the new bank survives.

We then examine the role played by a set of variables in determining the time of default. For each distributional form we use two specifications: the first one (a) controls for the whole set of information; the second one (b) includes only the information available at the start-up. Table 4 shows the results.

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parameter. Finally, log-normal distribution would imply a non-monotonic hazard function: it increases in the first years; then, after reaching a peak, it declines.

Estimates of log-normal version (column IV.a) reveal that, among the information available at the start-up, capital endowment plays a major role in determining survival time. A one million increase in initial capital endowment *CAP* extends survival by 4.3 per cent.[25] Similarly, an improvement in the organisation (*ORG_SCORE* one level lower) determines an increase of 31.7 per cent in estimated survival time. The local economic environment, as captured by both *SOUTH* and *GDP_PC* affects estimated survival time as well: starting up a bank in Southern Italy reduces the estimated survival time by 36.2 per cent; after controlling for the *SOUTH* dummy, choosing a less developed area as headquarter (with lower per capita GDP), has a statistically significant but negligible influence. Finally, among balance-sheet variables, only *DEV_ASST_GWT* shows significant effects: a one-unit deviation from median asset growth jeopardizes survival time by 38.3 per cent.

3.2. Semi-parametric model

AFT models used in Section 3.1 provide efficient estimates only if the underlying distribution is correctly identified. Although the results seem to indicate that log-normal distribution allows a close fit of our sample, some caution is needed.

To circumvent all the shortcomings of distributional misidentification, we shift to a Cox regression model. This approach circumvents distributional issues by estimating a semi-parametric model through partial likelihood.[26]

We estimate Cox regressions under the same set of covariates of model [1]. The results are reported in the first two columns of Table 5.

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[25] The estimated coefficients of *AFT* models are not informative; in order to arrive at an intuitive interpretation, we need to compute a simple algebraic transformation $100*(e^{\beta}-1)$ to work out the effect on survival time (see Appendix A).

[26] See Cox (1972) and Appendix *B* for more details.

The estimates are largely in line with the log-normal ones of the previous section, thus confirming the adequacy of the functional form adopted.

Five out of twelve variables have a significant effect on survival time; remarkably, all the covariates with a statistically significant coefficient in the log-normal model – including our variable of interest *CAP* – confirm their significance in the Cox model. Both *ORG_SCORE* and *DEV_ASST_GWT* confirm their major role in affecting the timing of a failure.

We use the Cox model estimates for a graphical inspection of our hypothesis about the shape of the hazard function: a graph of the log-survival function flattening after a peak would add more evidence of non-monotonic hazard. The plot of the fitted model (Chart 3) confirms our previous findings: the hazard rate – the first derivative of the plotted function – climbs until the 10^{th} semester and, for a while, between the 16^{th} and the 18^{th} semester; it flattens thereafter. This finding is consistent with DeYoung (2003a), who highlights a non-linear pattern in start-up banks' hazard of failure. It is worth noting that the fifth year of life appears to represent a sort of watershed in a start-up bank's life cycle.

4. Probability of default analysis

The previous section has investigated the role of initial capital in affecting the duration of a start-up bank. In this section, we complement this analysis through a binary choice model, estimating how the level of initial capital, along with other variables, affects the probability of a start-up bank to default within the first years of life.

Drawing on the evidence of the previous section, we first set the time horizon at five years, with default within the fifth year of life *(DEF_5)* as our dependent variable.[27]

[27] Maggiolini and Mistrulli (2005) also show that some balance-sheet indicators of newly-established mutual banks converge with those of incumbent ones after ten semesters of life.

Besides this, since this choice might be considered arbitrary and affect our results, we will use two alternative time-spans in Section 6.1 as a robustness check.

Accordingly, we start by running the following logistic regression:

$$\ln\left[\frac{p(DEF\_5_b)}{1-p(DEF\_5_b)}\right] = \alpha + \beta_1 CAP_b + \beta_2 MUTUAL + \sum_{i=1}^{n}\gamma_i(management_b) + \\ + \sum_{i=1}^{n}\delta_i(market_b) + \sum_{i=1}^{n}\varphi_i(balance\_sheeet_b) + \varepsilon_b \qquad [2]$$

where *DEF_5* is a dummy variable taking value 1 if the bank defaulted within the first five years, 0 otherwise, and *CAP* – the log of capital at the start-up – is our variable of interest.

As outlined in previous sections, mutual banks show higher default rates as well as lower average initial capital, whose level is mostly driven by their lower licensing requirements. Indeed, it may be the case that some unobserved heterogeneity feeds the capital effect, even if it is unrelated to capitalization; this effect is captured by the dummy variable *MUTUAL*.

As in Model [1], we introduce management information, market-related variables and balance-sheet figures as controls. In line with the previous section, we estimate two specifications of the model. The first one (a) mimics the information set at the time of licensing, including all the variables known at that stage; the second specification (b) incorporates additional information available after a few years. The results are reported in Table 6.

Regardless of the specification used, a higher initial capital endowment is negatively associated with the five-year probability of default: the higher the bank's initial capital, the lower the probability of exiting the market within this time frame. The results for initial capital (*CAP*) – in terms of both the magnitude and significance of estimated coefficients – are stable (ranging from -2.118 to -3.422) and robust to both specifications presented.[28]

The use of a logit framework implies that the effect on the probability of default due to an increase of a given covariate is nonlinear and depends on the value of the other variables. Thus, to interpret the effect of a regressor on the dependent variable we compute marginal effects.[29]

In the unrestricted specification (a), the average marginal effect of a one million increase in initial capital is -3.4 per cent (-3.2 per cent in the restricted specification).[30]

Among control variables, the dummy *MUTUAL* is not statistically significant, thus being a mutual bank has no independent effect on default. As for management characteristics, neither board experience (*EXPER*), nor education (*EDUC*), shows a sizeable degree of significance, besides the fact that neither of them has the expected sign. Conversely, the supervisor's assessment of bank organisation at the first year of life (*ORG_SCORE*) looks like an effective predictor of default: a higher score (i.e. a worse assessment) is associated with a higher probability of exiting the market; a one-notch increase results in a marginal effect ranging from 23 to 24 per cent, according to the specification used.

[28] As a robustness check (not shown) we take into account the price level at the start-up date, by using the deflated value of initial capital, with broadly similar results.

[29] Because of the non-linearity of the logistic model, marginal effects vary depending on the starting level of covariates and the magnitude of the change. Table 6 shows average marginal effects whose values are calculated by first computing marginal effects at each observation level and then by averaging all the individual marginal effects. For small samples, averaging the individual marginal effects is preferred to computing the marginal effect at the sample means of the data (Greene, 1997).

[30] As *CAP* is expressed in logarithmic terms, its marginal effect has been computed accordingly.

Business context controls display a multifaceted behaviour. While market concentration and economic development (*HHI, GDP_PC, GDP_GWT*) turn out not to be statistically significant,[31] urbanisation (*URB*) does play some role in reducing the probability of default. Headquartering in a highly populated Italian municipality reduces the probability of default for start-up banks; a one-unit increase in the urbanisation index yields a drop in probability of default ranging from 9 to 11 per cent.

Moving to balance-sheet indicators, in column (b), it's worth noting the importance of a balanced asset growth during the start-up period (*DEV_ASST_GWT* is significant and shows the expected sign), while business-mix (*BUSINESS_MIX*) does not seem to significantly affect short-term default. Admittedly, the lack of pure high-risk-high-return seeker banks in our sample limits the magnitude of this finding.

To complement the above findings, in model [2] we introduce the interaction variable *CAP*MUTUAL*. This interaction variable is meant to test if the effect of higher initial capital is reinforced for mutual banks. Column I.c of Table 6 shows the results. Interestingly, the sign of the coefficient of the interaction variable *CAP*MUTUAL* is the same for *CAP*, although neither of them is statistically significant.

Based on these results, we estimate the benefits – in terms of reducing the probability of default – of raising initial capital by a fixed amount. As shown in equation [3] we compute the estimated effect associated with progressive increases of initial capital by 1

[31] As further controls, we replaced *HHI* with an alternative definition of the concentration index (based on loans rather than on branches) and we used credit/GDP ratios at province level as an alternative proxy for local financial development. The results (not shown) reveal that the coefficients of these variables are never statistically significant.

million (from M to $M+1$),[32] keeping all the other covariates at their sample average value X^*. Table 7 shows the results.

$$Benefit_{M+1} = \Pr(DEF\_5 = 1 \mid e^{CAP} = M+1, X^*) - \Pr(DEF\_5 = 1 \mid e^{CAP} = M, X^*) \quad [3]$$

The marginal benefit of increasing initial capital fades out progressively. When moving from € 2 to € 3 million it is 4.5 per cent, whereas for the highest starting level of capital it becomes negligible: the marginal benefit of increasing initial capital from € 13 to € 14 million is around 1 per cent.

As the threshold grows up, regulators' incentives in raising up the minimum level of initial capital progressively decreases, because of the decreasing effectiveness of capital in sheltering a bank from an early default. Thus, from a regulatory perspective, our results may provide a tool for setting minimum initial capital; the optimal level may be set by targeting the probability of default desired by regulators.

5. Counterfactual analysis

So far, using different approaches, we investigate the role played by initial capital endowment in a start-up bank. We find capital does play a role, even after controlling for a broad set of covariates: a higher initial capital endowment significantly affects the likelihood of failure for a start-up bank over a pre-defined time horizon.

The decision regarding its level can be driven by different factors. These include optimal liability structure and shareholders' value maximisation. But does banking regulation play a role in this process? Indeed one sensible question centres on the existence of a causal link between a *regulatory* minimum initial capital and the likelihood of default for a bank start-up.

[32] *CAP* is expressed in logarithmic terms; thus an increase in initial capital from M to $M+1$ corresponds to an increase in *CAP* from e^M to e^{M+1}.

In order to analyse further the link between regulatory initial capital and the probability of default, we can exploit a regulatory shift which affected only one category of banks in our sample. Indeed, the rules governing market access for Italian mutual banks were toughened during the late 1990s, when, in January 1999, the minimum capital required for their start-up was almost doubled, from ITL 2 billion (€ 1.03 million) to € 2 million. Notably, this decision affected only mutual banks; both commercial banks (*COM*) and popular banks (*COOP*) were unaffected, with a minimum capital threshold already set at € 6.3 million.

This regulatory shift allows us to turn to a quasi-experimental setting and run a counterfactual evaluation of the effect of a "treatment".[33] With respect to a target variable – in our setting, the five-year probability of default, *DEF_5* – we are able to compare the pre- and post-intervention performance, not only within the treated group (the mutual banks),[34] but also with respect to a control group of non-treated banks (popular and commercial ones).

Before engaging in the difference-in-difference (*D-I-D*) estimation, we first need to rule out the possibility that the change in the target variable is determined by a parallel change in the observed covariates, rather than by the treatment itself. The reliability of the estimated average effect of the treatment crucially depends on the absence of a significant pre-post shift in the controls other than the one undergoing the regulatory reform (*CAPITAL*).

Table 8 shows the average values of probability of default, capital and the other independent covariates, separately for the treated (Mutual) and the control group (Non-

[33] See Meyer (1995) and Wooldridge (2002).

[34] The regulatory change involved 32 mutual banks out of 69.

mutual), distinguishing the period before (1994-1998) and after (1999-2006) the policy shift (the "treatment"). The differences between the two groups (Mutual - Non-mutual) and between the two periods (pre - post 1999) are also reported; standard errors of the T-test for the equality of pre-post means are in parentheses.

The top panel of Table 8 illustrates the behaviour of the target variable (*DEF_5*): the pre-post difference for mutual banks is equal to -0.24, largely exceeding the contemporary reduction of -0.06 for non-mutual banks. Although we should refrain from reading this partial result as a causal effect, it seems that the regulatory decision on an initial capital increase is associated with a reduction in the probability of default on intervened banks.

Among the covariates of Table 8, capital is the one undergoing the regulatory reform. Related values measure the intensity of the treatment, as a consequence of the regulatory tightening of the minimum initial capital endowment from its previous level. For treated banks, average capital endowment increases by 86 per cent with respect to the pre-treatment period; the analogous time dynamic for non-treated institutions is less accentuated (+ 79 per cent).

As required for the reliability of the *D-I-D* analysis, the bottom panel of Table 8 shows that the differences for all the other observed covariates are mostly negligible; a t-test on pre-post difference fails to reject the null hypothesis of means equality for all the variables except *ORG_SCORE* and *EXPER*. As suggested by the results of the previous sections, the latter covariate plays a minor role. Conversely, *ORG_SCORE* is found to be significant in all the model specifications we use. Nevertheless, the sign of its pre-post differences can only reinforce the *D-I-D* results: mutual banks show, on average, a *worse* organisation (a higher score) after the treatment, while the control sample displays the opposite dynamic. Should we find a link between the policy shift and an improvement in the target variable, it would be despite the trend of *ORG_SCORE* (the same holds for *EXPER*).

Moving to the core counterfactual analysis, we estimate a linear probability model such as the following:

$$\begin{aligned} DEF\_5_b &= \alpha + \beta_1 MUTUAL_b + \beta_2 T1999_b + \beta_3 T1999 * MUTUAL + \\ &+ \sum_{i=1}^{n} \gamma_i (controls_b) + \varepsilon_b \end{aligned} \qquad [4]$$

As in the standard *D-I-D* approach, our variable of interest is *T1999*MUTUAL*. This interaction variable captures the casual effect on the five-year probability of default of the treatment on mutual banks (*MUTUAL*) established since January 1999 (*T1999*).[35]

Table 9 shows the set of results.

Column I reports the estimates of model [4] without control variables. The coefficient of the interaction variable *T1999*MUTUAL* is -0.18 and it is equivalent to the pre-post difference of *DEF_5* between Mutual and Non-Mutual, reported in Table 8. As in the previous sections, in column II we add the same set of control variables used in restricted versions of both model [1] and model [2]; the inclusion of management and market variables does not drastically affect the magnitude of the coefficient of the interaction term, which increases to -0.28. Moreover, adding a set of control variables does affect the precision of the estimate: the corresponding T-statistic is equal to -2.1.

The results obtained by this *D-I-D* linear probability regression suggest that the treatment – in the form of a regulatory increase affecting mutual banks since 1999 –

[35] In contrast to Section 4, we use a linear probability model rather then a logit model: indeed, the former allows a direct comparison of the average effects on probability of default for treated and non-treated banks.

determined, on average, a 28 basis point reduction in their probability of defaulting within the fifth year.

To verify both the magnitude and the statistical significance of causal effect, we check the robustness of this result through a falsification test. More specifically, we modify model [4] by assuming for mutual banks a "false treatment" in the form of a fake capital increase taking place one year before (*T1998*) the authentic treatment year. We expect that a fake treatment – in the form of a relatively small variation in the real treatment starting date – would determine a dramatic change in both magnitude and significance in the relevant interaction coefficient. Columns III-IV of Table 8 present the results obtained by running model [4] under these assumptions. The results drawn from this test go in the expected direction: by moving away from the true treatment year, the coefficients of the relevant interaction variables substantially reduce both in magnitude (-0.06) and significance (t-statistic: -0.4).

6. Robustness

In this section, we conduct a few additional exercises to support previous findings. These robustness checks are related respectively to testing the effect of a narrowing/widening of the time window, to detect a default, to assess the proportionality assumption in the Cox model, to verify the presence of a split in the population between failed and non-failed banks and to check for any unconfoundedness issue.

6.1. Different cut-off date for the probability of default analysis

The choice of an arbitrary time span of five years for identifying the short-run probability of default represents a limitation of the analysis. To address this issue, the first robustness check is therefore to verify whether the results are confirmed when equation [2] is estimated by using a different definition for the dependent variable. In particular, we consider two alternatives: a default taking place after either four years (*DEF_4*) or six years (*DEF_6)* since the date on which the bank was established. Columns II-III of Table 6 present the estimated results of the logit model with the two alternative definitions of the binary dependent variable.

The results are robust to both these changes. The estimated coefficients for *CAP* range from -1.603 to -2.574; the estimated coefficients are statistically significant in all the specifications.

Among the other covariates, *ORG_SCORE* confirms its role.

6.2. Extended Cox model

Proportional hazard represents one of the main features of the Cox model. The model assumes that the ratio of the hazard functions for any two bank subgroups (i.e. two groups with different values of the covariate *X*) is constant over time. Interestingly, while the hazard *ratio* is assumed to be constant, the hazard can vary freely over time.

The violation of proportionality does not jeopardise the estimation of the model: partial likelihood accommodates for non-proportionality.[36] Nevertheless, proportionality is a strong assumption that needs to be tested.

To assess the proportionality assumption, we estimate an extended specification of the Cox model [1], by including a set of time-dependent variables,[37] constituted by the (time-independent) variable and the interaction of (log of) the survival time. Estimating the

[36] In the presence of a time-dependent variable, the estimated coefficients of its time-proportional hazard specification provide an estimation of the average effect (see Allison, 1995).

[37] In this setting, we define as time-dependent any variable whose values differ over time.

extended Cox model allows us to test for departure from proportionality at the overall model level[38] by running a Likelihood Ratio test[39] of the form

$$\text{LR} = (-2\log(L_{\text{PH model}})) - (-2\log(L_{\text{Extended Cox model}}))$$

Table 5, column II presents the estimates of the extended Cox model.

At global level, we strongly reject the null hypothesis of proportionality: the LR test values (Table 5, bottom row) range from 258 to 271, above the 1 per cent critical level.[40]

6.3. Split-population model

One limitation of the models used for survival analysis is that they were originally designed and implemented for analysing changes in the survival of individuals with a finite life-span. This assumption does not hold for firms in general and banks are no exception. This thing considered, Cole and Gunther (1995) and Maggiolini and Mistrulli (2005) suggest that both the parametric and Cox models we use for survival analysis may produce biased results because both models assume that eventually all individuals default. If the bank population is split into two groups (one of banks that eventually fail the other of surviving banks) then the timing of failure, as well as failure itself, may depend on different factors.

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[38] Moreover, the extended Cox model allows us to assess departure from proportionality at the individual variable level, by testing the statistical significance of each individual interaction variable. For the sake of brevity, these results are not shown.

[39] The LR test is distributed as a χ^2 under the null hypothesis of proportionality. If the test is significant, we can reject the null and then the extended Cox model is preferred.

[40] At individual variable level, it might be interesting to investigate the consequences for the estimated coefficient for both *CAP* and the respective time-dependent variable *CAP_T*. The estimated coefficient for the time-invariant variable *CAP* is 8.68; the coefficient for the time-dependent interaction variable *CAP_T* (not shown) is -3.96. To some extent, these results suggest the effect of capital on the timing of default is not constant over time.

We can verify the presence of this split in our sample by estimating a split-population survival model, using log-normal distribution for the survival part and a logit for the probability of default component.[41] The results shown in Table 10, column I do not appear to support the assumption of a split in the population and the related bias in both lognormal and Cox model's estimates.

In particular, most of the variables that are found to have a statistically significant effect on survival time in the previous regressions are similarly significant in the split population model, although the estimated coefficients are smaller.[42]

6.4. Propensity score matching

The results of the counterfactual analysis in Section 5 hold if the differences in observable characteristics between treated banks and the control group are negligible. Previous analysis seems to show that differences in observed covariates are not material. The analysis shown in Table 8 seems to rule out this possibility. Nevertheless, the possibility remains that the found effect is significantly related to differences between treated and non-treated banks. To eliminate the consequence of these differences, we can compare the effects on two samples of banks that are similar in all their observed characteristics. As a final step to address the unconfoundedness issue, we run a propensity score matching estimation (Rosembaum and Rubin, 1983).

As a first step, for each bank in our sample we estimate a propensity score, as in standard propensity score matching, by including all the covariates used in the previous section. Subsequently, based on the estimated score, we then estimate the average treatment effect

[41] Along with the timing of failure, split-population models allow the estimation of effects of variables on the likelihood of banking failure.

[42] In particular, among the control variables, organisation effectiveness confirms its role, although the magnitude of the coefficients declines by 30 per cent.

between treated and non-treated banks through kernel matching. The results are shown in Table 11.

The estimated average treatment effect is -0.134, with a standard error of 0.15. While not highly significant (T-statistic = -0.9), it is close in magnitude to the previous finding. This result is robust to a different matching algorithm. Switching to the nearest neighbour matching method does not affect results significantly: the average treatment effect is -0.188 (s.e.: 0.19).

7. Conclusions

A minimum amount of capital is one of the pre-conditions for a bank to start operating. We explore the relationship between the initial capital endowment of a bank start-up and its survival features, controlling for a set of market, managerial and financial variables.

Both survival and probability of default analyses provide evidence of the importance of the initial capital endowment to avoid early default. Furthermore, this can be read as a causal effect, according to the result of a counterfactual analysis: by exploiting a regulatory change for a subsample of Italian banks, an increase in regulatory minimum initial capital actually improved the resilience of the "treated" banks.

In all our regressions, we control for a wide set of variables: in order to exploit all possible information available at the moment of the start-up, we include many variables used for the licensing phase by different supervisory frameworks. As expected, most of the variables influencing survival time also drive the probability of default within the first years of life. Out of these, the strength of a bank's organisation and a balanced growth path may help in preventing its exit from the market just a few years after the start-up; by contrast, board experience, market characteristics and business cycle shows either limited or no significance.

Survival analysis also confirms that our sample of start-up banks follows a life-cycle path: the hazard of default is high and significant in the first years of activity after which it tends to decrease, whereas a bank reaches a "maturity" phase. The fifth year of life seems to represent a watershed.

Assuming start-up banks' stability to be a relevant supervisory objective, our work provides banking regulators with a tool for measuring the marginal benefit of different regulatory thresholds for banks' initial capital. The optimal minimum capital level may be set by targeting the probability of default desired by the supervisor. Any higher threshold should be considered in the light of potential costs it may introduce and in terms of entry barriers. The evaluation of this issue is beyond the scope of this paper, but costs associated with onerous entry procedures and its negative spillovers (Djankov et al. 2002) should be taken into account properly. Finally, knowledge of those variables affecting both the timing and likelihood of a bank's default can help regulatory triage: a preliminary assessment of the likely development of a bank activity – based, for example, on its mock activity plan – can help supervisors target it for closer supervision.

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Appendix A

Accelerated failure time (AFT) models describe, in their general forms, the relationship between each pair of individuals. Given $S_i(t)$ the survivor cdf for individual *i*, the AFT model for any other individual *j* is

$$S_i(t) = S_j(\phi_{ijt}) \text{ for all } t$$

where ϕ_{ijt} is a pair-specific constant.

We estimate a specific case of AFT model: in a general form, it can be described as

$$\log T_i = \beta \mathbf{X} + \sigma\, \varepsilon_i$$

where $\log T_i$ is the logarithm of the predicted value of T (positive for all possible values of T), β is a vector of coefficient, **X** is a matrix of time-independent covariates, σ is an estimated parameter that allows the random disturbance term ε_i to have variance and mean constant over *i*, as well as independent across observations.

As in general AFT models, ε_i can assume different distributions, holding the assumption of constant of mean and variance and cross-sectional independence: we examine four different cases, relevant for our analysis.

Log-normal model assumes a normal distribution for ε_i: in turn, the fact that $\log T$ is normally distributed implies that T has a log-normal distribution.

Additionally, ε_i can assume a standard extreme-value distribution: its pdf is $f(\varepsilon_i)=\exp[\varepsilon_i - \exp(\varepsilon_i)]$. In this case, $\log T$ also assumes an extreme value distribution, conditional on the covariates and on a value of σ constrained equal to 1.

This implies T assumes an exponential distribution, corresponding to a constant hazard function of the form

$h(t) = \exp(\beta X)$ or equivalently $\log h(t) = \beta X$.

By relaxing the $\sigma = 1$ constraint, while keeping the standard extreme value distribution for ε, T assumes a Weibull distribution, conditional on the covariates. The log-survival time model and the log-hazard one can be expressed respectively as

$\log T = \beta X + \sigma\varepsilon$

and

$\log h(t) = \alpha \log T + \beta^* X$ with $\beta^* = \frac{-\beta}{\sigma}$.

The (generalised) gamma model is a distribution with three parameters: β, δ (the scale parameter) and k (the shape parameter). The pdf of this specification can be expressed as

$$f(t) = \frac{\beta}{\Gamma(k)\delta}\left(\frac{t}{\delta}\right)^{k\beta-1} e^{-\left(\frac{t}{\delta}\right)^{\beta}}$$

where Γ is the Gamma function. Note that all the models presented above (log-normal, exponential, Weibull) can be treated as special cases of the generalised gamma model. This property derives from its higher number of parameters.

Appendix B

The Cox (1972) regression model consists of a hazard model (called proportional hazard model) and an estimation methodology (partial likelihood).

In its basic setting, the proportional hazard model can be expressed as

$$h_i\ (t,\mathbf{X}) = \lambda_0(t)\ \exp\left\{\sum_{n=1}^{k} \beta_{in} X_n\right\}$$

where the first part $\lambda_0(t)$ is a baseline hazard function, which corresponds to the hazard for the individual i given all covariates β equal zero, and the second one is an exponential function of the covariates.[43] The proportional hazard model is time-independent: in fact, for individual i, hazard can be expressed as

$$\frac{h_i(t)}{h_j(t)} = \exp\left\{\sum_{n=1}^{N} \beta_n \left(x_{in} - x_{jn}\right)\right\}$$

which is independent of time. The proportional hazard assumption has to be checked (via a graphical analysis of covariate-wise Schoenfeld residuals, inclusion of time-dependent covariates and goodness-of-fit test).

If the proportional hazard assumption is violated, we can shift to an extended Cox model. This model allows for time-dependent variables: the extended version of the model can be expressed as:

$$h_i\ (t,\mathbf{X}(t)) = \lambda_0(t)\ \exp\left\{\sum_{n=1}^{k} \beta_{in} + \sum_{n=1}^{k} \delta_n X_n g_n(t)\right\}$$

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[43] By further specifying $\lambda_0(t)$ it is possible to obtain a different model, such as the exponential and the Weibull ones.

where X(t) denotes all predictors, X_n the nth time-independent covariate, $X_n(t)$ the n_{th} time-dependent covariate and $g_n(t)$ a function of time (generally, the logarithm of time).

We can run a Likelihood Ratio test for assessing the proportional hazard assumption in the form:

$$LR = -2\log(L_{\text{PH model}}) - (-2\log(L_{\text{extended Cox model}})) \approx \chi_n^2$$

under the null hypothesis *H0* of:

H0: $\delta_1 = \delta_2 = \cdots = \delta_n = 0$.

Partial likelihood can be expressed as

$$PL = \prod_{k=1}^{K} L_j$$

that is, as a product of the likelihood for all the k event. For a given event j, *Lj* is the ratio of the hazard of individual *i* (as expressed above) over the hazard of the risk set, that represents the set of all the individuals at risk at a given point of time.

Tables and figures

Table 1 – The sample

	N.	%	% of default	% of default within the 5th year
Total	119	100%	60%	37%
by category:				
MUTUAL	69	58%	77%	54%
COOP	23	19%	57%	26%
COM	27	23%	19%	4%
by area:				
NORTH	44	37%	50%	30%
CENTRE	24	20%	50%	33%
SOUTH	51	43%	73%	45%

The table presents the sample, that consists of banks chartered between 1993 and 2006; bank default observed within first semester of 2011.
MUTUAL refers to mutual banks; *COOP* to cooperative banks; *COM* to commercial banks.
NORTH, CENTRE and *SOUTH* indicate a bank headquartered in the North, Centre or South of Italy respectively.

Table 2 – Initial capital - statistics by categories and status at the 5th year of activity

	Total					Start- up banks defaulted within the 5th year					Start-up banks not defaulted within the 5th year				
	N	*mean*	*q1*	*q3*	*st. dev.*	*N*	*mean*	*q1*	*q3*	*st. dev.*	*N*	*mean*	*q1*	*q3*	*st. dev.*
Total	119	5.7	1.4	7.8	6.0	44	2.7	1.1	2.9	2.8	75	7.5	2.2	8.9	6.7
MUTUAL	69	1.9	1.2	2.4	1.0	37	1.7	1.1	1.8	1.0	32	2.2	1.4	3.1	0.9
POP	23	8.6	5.7	9.7	4.3	6	7.8	5.7	9.7	2.6	17	8.9	5.7	8.7	4.8
COM	27	12.9	7.8	15.0	6.8	1	11.0	11.0	11.0		26	13.0	7.8	15.0	7.0

The table presents the summary statistics for initial capital endowment for the banks of the sample. Values in million of euros.
MUTUAL refers to mutual banks; *COOP* to cooperative banks; *COM* to commercial banks.

Table 3 – Main statistics

Description	Variable	N	mean	q1	median	q3	min	max	std dev	source
Initial capital endowment (million euro)	*CAPITAL*	119	5.7	1.4	3.2	7.8	0.6	33.0	6.0	Supervisory reports
Log of initial capital endowment (million euro)	*CAP*	119	1.3	0.4	1.2	2.1	-0.5	3.5	1.0	Supervisory reports
Mutual bank - dummy	*MUTUAL*	119	0.6	0.0	1.0	1.0	0.0	1.0	0.5	Supervisory reports
Supervisory score for the organization effectiveness	*ORG_SCORE*	113	2.6	2.0	2.0	3.0	1.0	4.0	0.8	Off-site Supervisory rating system
Proportion of board members holding a previous experience in the banking industry		119	10%	0%	4%	12%	0%	100%	17%	Or.So.
Higher than average experienced board members - dummy	*EXPER*	119	0.4	0.0	0.0	1.0	0.0	1.0	0.5	Or.So.
Proportion of graduate-level educated board members		119	47%	29%	46%	67%	0%	100%	26%	Or.So.
Higher than average graduate-level educated board members - dummy	*EDUC*	119	0.5	0.0	0.0	1.0	0.0	1.0	0.5	Or.So.
Urbanization index at town level (1=low;2=medium;3=high)	*URB*	119	2.4	2.0	3.0	3.0	1.0	3.0	0.7	Istat
GDP average growth rate at province level in the 2 years preceding chartering	*GDP_GWT*	119	1.7%	1.4%	1.5%	2.6%	0.3%	2.8%	0.7%	Istat
Hirschman-Herfindahl index for bank branches at the province level	*HHI*	119	10.8	7.4	10.6	14.2	3.7	28.1	4.6	Supervisory register
Headquarter in South of Italy - dummy	*SOUTH*	119	0.4	0.0	0.0	1.0	0.0	1.0	0.5	Supervisory reports
Per-capita GDP at province level (thousand euro, year 2000 figures)	*GDP_PC*	119	17.7	12.5	17.9	21.4	10.5	28.0	5.2	Istat
Proportion of loans to households and SMEs	*BUSINESS_MIX*	118	65%	48%	68%	85%	8%	100%	25%	Supervisory reports
Absolute value distance of five-year asset growth rate from sample median value (percentage points)	*DEV_ASST_GWT*	119	27%	8%	18%	36%	0%	322%	36%	Supervisory reports

Table 4– Survival analysis – parametric models: AFT models under 4 different distributions

Distribution:	Gamma		Weibull		Exponential		Log-normal	
	I.a	I.b	II.a	II.b	III.a	III.b	IV.a	IV.b
INTERCEPT	2.900 **	2.768 **	5.269 **	5.104 **	7.889	7.173	4.190 **	4.193 **
	(0,908)	(1,131)	(2,023)	(2,191)	(5,24)	(5,225)	(1,839)	(1,978)
CAP	0.953 ***	0.984 ***	0.516 **	0.719 **	0.767	0.992 *	0.556 **	0.681 **
	(0,122)	(0,206)	(0,229)	(0,231)	(0,55)	(0,513)	(0,225)	(0,227)
MUTUAL	1.000 ***	0.804 **	0.169	0.330	0.167	0.321	0.233	0.304
	(0,228)	(0,334)	(0,37)	(0,385)	(0,904)	(0,88)	(0,361)	(0,384)
ORG_SCORE	-0.354 ***	-0.286 ***	-0.402 ***	-0.424 ***	-0.578 **	-0.584 **	-0.381 ***	-0.390 ***
	(0,06)	(0,079)	(0,072)	(0,076)	(0,183)	(0,179)	(0,079)	(0,085)
EXPER	-0.435 ***	-0.374 ***	0.043	0.042	0.171	0.201	-0.055	-0.015
	(0,08)	(0,095)	(0,153)	(0,159)	(0,357)	(0,352)	(0,134)	(0,144)
EDUC	0.066	0.150	-0.126	-0.076	-0.206	-0.145	-0.079	-0.048
	(0,094)	(0,113)	(0,139)	(0,152)	(0,33)	(0,328)	(0,13)	(0,139)
URB	-0.072	0.110	0.156	0.181	0.161	0.230	0.173	0.209 *
	(0,069)	(0,142)	(0,11)	(0,118)	(0,263)	(0,26)	(0,107)	(0,113)
GDP_GWT	-0.332	-0.738	-0.703	-1.085	-1.564	-1.752	-0.451	-0.709
	(0,504)	(0,462)	(1,247)	(1,353)	(3,232)	(3,234)	(1,081)	(1,151)
HHI	-0.012	0.002	0.014	0.011	0.021	0.017	0.009	0.008
	(0,009)	(0,015)	(0,013)	(0,013)	(0,034)	(0,033)	(0,015)	(0,016)
SOUTH	-0.311	-0.395 *	-0.545 **	-0.564 **	-0.746	-0.770	-0.450 *	-0.458 *
	(0,204)	(0,212)	(0,227)	(0,244)	(0,588)	(0,578)	(0,241)	(0,252)
GDP_PC	0.000 ***	0.000 **	0.000 **	0.000 **	0.000 *	0.000 *	0.000 **	0.000 **
	(0)	(0)	(0)	(0)	(0)	(0)	(0)	(0)
BUSINESS_MIX	-0.249		-0.510		-0.673		-0.131	
	(0,334)		(0,396)		(0,995)		(0,417)	
DEV_ASST_GWT	-0.359		-0.445 ***		-0.567 **		-0.483 ***	
	(0,272)		(0,097)		(0,26)		(0,14)	
DUMMY YEARS	YES	YES	YES	YES	YES	YES	YES	YES
Number of observations	119	119	119	119	119	119	119	119
-2 Log Likelihood	103.953	118.456	137.802	150.733	198.128	202.506	132.616	144.309
AIC	172.471	180.092	202.919	209.137	259.921	257.750	197.732	202.714
LR test			33.849	32.277	94.175	84.050	28.663	25.853

The table shows maximum likelihood estimates from a parametric Accelerated Faliure Time (AFT) model with different distribution assumptions. Dependent variable: (the log of) semester of default. Definition of variables can be found in Table 3. LR tests from the generalised Gamma model for parameter restrictions of the nested models.
Standard errors in parenthesis. * , ** and *** denote statistical signifcance at the 10%, 5% and 1% critical levels, respectively.

Table 5 – Survival analysis: Cox model

	I.a	I.b	II.a	II.b
CAP	-1.630 **	-1.286 **	8.941 **	8.682 **
	(0.587)	(0.619)	(3.099)	(3.379)
MUTUAL	-0.629	-0.221	14.751 **	12.256 **
	(0.96)	(0.985)	(5.416)	(6.223)
ORG_SCORE	0.964 ***	1.003 ***	2.418 **	3.369 **
	(0.201)	(0.206)	(1.213)	(1.574)
EXPER	-0.126	-0.082	5.637 **	7.698 **
	(0.383)	(0.393)	(2.555)	(2.996)
EDUC	0.221	0.337	-0.484	-0.564
	(0.357)	(0.346)	(1.794)	(2.277)
URB	-0.429	-0.401	-1.298	-1.656
	(0.285)	(0.288)	(1.717)	(1.841)
GDP_GWT	3.159	2.668	2.035	1.152
	(3.316)	(3.312)	(3.968)	(4.238)
HHI	-0.026	-0.034	0.397 *	0.406
	(0.034)	(0.036)	(0.236)	(0.261)
SOUTH	1.067 *	1.181 *	12.503 ***	12.607 ***
	(0.611)	(0.624)	(2.773)	(3.27)
GDP_PC	0.000 **	0.000 **	0.001 ***	0.001 ***
	(0)	(0)	(0)	(0)
BUSINESS_MIX		0.871		5.078
		(1.038)		(6.02)
DEV_ASST_GWT		1.539 ***		-0.289
		(0.311)		(2.082)
DUMMY YEARS	YES	YES	YES	YES
Time-dependent covariates	NO	NO	YES	YES
Number of observations	119	119	119	119
-2 LOG L	456.582	439.006	185.403	180.391
AIC	498.582	485.006	247.403	250.391
Score	84.548 ***	121.751 ***	266.609 ***	297.425 ***
LR test			271.179	258.615

The table shows partial likelihood estimates from a semi-parametric Cox model. Dependent variable: (the log of) semester of default. Definition of variables can be found in Table 3. Time-dependent covariates are the interaction between the original covariates and the (log of) time at default (in semester). LR test of the joint significance of time-dependent covariates in the extended Cox model (Column II).
Standard errors in parenthesis. * , ** and *** denote statistical signifcance at the 10%, 5% and 1% critical levels, respectively.

Tab. 6 – Probability of default analysis: logit model

	I.a	I.b	I.c	II.a	II.b	III.a	III.b
dependent variable:	*DEF_5*	*DEF_5*	*DEF_5*	*DEF_4*	*DEF_4*	*DEF_6*	*DEF_6*
INTERCEPT	-6.814	-7.425	-11.657	-4.936	-7.396	-5.281	-5.879
	(5.625)	(4.543)	(7.105)	(5.625)	(4.864)	(4.795)	(4.311)
CAP	-3.422 *-0.297*	-2.118 *-0.242*	-1.558 *-0.103*	-2.392 - *0.208*	-1.603 *-0.179*	-2.574 *-0.272*	-2.280 *-0.272*
	(1.282) **	(0.876) **	(1.853)	(1.231) *	(0.804) **	(1.034) **	(0.857) **
MUTUAL	-1.326 *-0.110*	-0.879 *-0.098*	2.885 *0.256*	0.718 *0.061*	1.214 *0.135*	-0.958 *-0.097*	-0.896 *-0.104*
	(1.72)	(1.441)	(3.765)	(1.757)	(1.548)	(1.481)	(1.374)
CAP*MUTUAL			-2.671 *-0.219*				
			(2.256)				
ORG_SCORE	2.747 *0.228*	2.156 *0.242*	3.085 *0.187*	1.569 *0.133*	1.392 *0.154*	1.712 *0.173*	1.761 *0.203*
	(0.704) ***	(0.5) ***	(0.828) ***	(0.539) **	(0.416) ***	(0.513) ***	(0.462) ***
EXPER	1.652 *0.137*	1.256 *0.141*	1.662 *0.050*	0.354 *0.030*	0.134 *0.015*	0.550 *0.056*	0.545 *0.063*
	(0.931) *	(0.774)	(0.957) *	(0.836)	(0.724)	(0.768)	(0.722)
EDUC	0.437 *0.036*	0.011 *0.001*	0.551 *0.062*	0.258 *0.022*	0.127 *0.014*	0.587 *0.059*	0.202 *0.023*
	(0.786)	(0.652)	(0.821)	(0.832)	(0.697)	(0.723)	(0.657)
URB	-1.303 *-0.108*	-0.767 *-0.086*	-1.273 *-0.046*	-0.822 - *0.070*	-0.595 *-0.066*	-0.537 *-0.054*	-0.488 *-0.056*
	(0.734) *	(0.548)	(0.732) *	(0.675)	(0.533)	(0.569)	(0.519)
GDP_GWT	0.834 *0.069*	0.524 *0.059*	0.924 *0.046*	-0.091 - *0.008*	-0.231 *-0.026*	0.472 *0.048*	0.349 *0.040*
	(0.584)	(0.456)	(0.599)	(0.564)	(0.454)	(0.472)	(0.436)
HHI	-0.133 *-0.011*	-0.127 *-0.014*	-0.119 *-0.007*	-0.140 - *0.012*	-0.105 *-0.012*	-0.083 *-0.008*	-0.074 *-0.009*
	(0.118)	(0.097)	(0.122)	(0.113)	(0.097)	(0.091)	(0.082)
SOUTH	2.019 *0.167*	0.655 *0.073*	1.750 *0.122*	1.466 *0.124*	0.804 *0.089*	1.458 *0.148*	1.024 *0.118*
	(1.579)	(1.194)	(1.625)	(1.583)	(1.305)	(1.258)	(1.139)
GDP_PC	0.000 *0.000*	0.000 *0.000*	0.000 *0.000*	0.000 *0.000*	0.000 *0.000*	0.000 *0.000*	0.000 *0.000*
	(0)	(0)	(0)	(0)	(0)	(0)	(0)
BUSINESS_MIX	-6.190 *-0.513*		-6.465 *-0.200*	-3.966 - *0.336*		-1.829 *-0.185*	
	(3.362) *		(3.415) *	(3.082)		(2.415)	
DEV_ASST_GWT	5.461 *0.452*		5.885 *0.388*	4.885 *0.414*		3.841 *0.389*	
	(1.79) **		(1.907) **	(1.629) **		(1.554) **	
Number of observations	119	119	119	119	119	119	119
Max-rescaled R-Square	0.719	0.591	0.727 ***	0.651	0.523	0.664	0.608
Likelihood Ratio	82.017 ***	62.744 ***	83.374 ***	65.602 ***	49.701 ***	75.991 ***	67.751 ***
Score	57.366 ***	49.813 ***	58.702 *	49.824 ***	39.446 ***	56.589 ***	53.797 ***
AIC	87.469	103.566	88.112	88.511	101.009	101.683	106.975

The table shows the estimates from a logit model; dependent variables: *DEF_5* indicates a default within the 5-th year since chartering date; *DEF_4* and *DEF_6* a default within 4-th and 6-th year respectively. Definition of variables can be found in Table 3.
Average marginal effects (in italics) are computed as the average of individual marginal effect; standard errors in parenthesis. * , ** and *** denote statistical signifcance at the 10%, 5% and 1% critical levels, respectively.

Table 7 – Marginal benefit of raising initial capital

Capital increase (mln.)	Marginal benefit
2 - 3	-4.5%
3 - 4	-4.2%
4 - 5	-3.8%
5 - 6	-3.4%
6 - 7	-3.0%
7 - 8	-2.6%
8 - 9	-2.3%
9 - 10	-1.9%
10 - 11	-1.7%
11 - 12	-1.4%
12 - 13	-1.2%
13 - 14	-1.0%

The table shows the marginal benefit - as reduction of five-year probability of default - associated with a 1 million euro increase in initial capital at different initial level of capital.
Marginal benefits for total sample estimated by keeping all the covariates but CAP at their sample mean values as shown in Tab. 3.

Table 8 – Variables' average value: difference before and after the treatment (year 1999)

Variable	Bank type	Regulatory time span I 1994-1998	Regulatory time span II 1999-2006	Difference III Pre-post '99	
		Dependent variable			
DEF_5	*Mutual*	0.65	0.4	-0.24	(0.12)
	Non-mutual	0.17	0.11	-0.06	(0.10)
	Mutual-Non- m.	0.48	0.29	-0.18	(0.15)
		Covariates			
CAPITAL	*Mutual*	1.38	2.57	1.19	(0.18)
	Non-mutual	7.64	13.73	6.09	(1.53)
	Mutual-Non- m.			-4.9	(1.31)
ORG_SCORE	*Mutual*	3.65	3.97	0.31	(0.20)
	Non-mutual	3.5	3.19	-0.31	(0.18)
	Mutual-Non- m.			0.62	(0.28)
EXPER	*Mutual*	0.27	0.44	0.17	(0.12)
	Non-mutual	0.21	0.89	0.67	(0.11)
	Mutual-Non- m.			-0.50	(0.16)
EDUC	*Mutual*	0.24	0.38	0.13	(0.11)
	Non-mutual	0.7	0.67	0.03	(0.14)
	Mutual-Non- m.			0.16	(0.17)
URB	*Mutual*	2.16	2.38	0.21	(0.17)
	Non-mutual	2.52	2.7	0.18	(0.17)
	Mutual-Non- m.			0.03	(0.24)
GDP_GWT	*Mutual*	1.89	1.66	-0.23	(0.16)
	Non-mutual	1.78	1.56	-0.23	(0.21)
	Mutual-Non- m.			0.00	(0.26)
HHI	*Mutual*	12.59	11.19	-1.11	(1.12)
	Non-mutual	10.26	8.08	-2.13	(1.10)
	Mutual-Non- m.			1.02	(1.16)
SOUTH	*Mutual*	0.67	0.47	-0.21	(0.12)
	Non-mutual	0.3	0.15	-0.16	(0.12)
	Mutual-Non- m.			-0.05	(0.17)
GDP_PC	*Mutual*	15.75	15.6	-0.15	(1.00)
	Non-mutual	19.33	21.65	2.31	(1.48)
	Mutual-Non- m.			-2.46	(1.71)

The table shows the variables average values for the two period 1994-1998 and 1999-2006 (left panel, column I-II respectively) and the difference of the variables average values between the two periods (right panel, colum III).
For each variable, the first row refers to mutual banks; the second row to non-mutual banks; the third row to the difference between the pre-post difference for the two sub-samples. Definition of variables can be found in Table 3. Standard errors in parenthesis.

Table 9 – Effects of regulatory increase of minimum initial capital endowment: difference-in-difference linear regression model

	True treatment		False treatment: 1 year before	
	I.a	I.b	II.a	II.b
INTERCEPT	0.174 **	-0.940 **	0.158 *	-1.046 **
	(0.079)	(0.384)	(0.083)	(0.405)
T1999	-0.063	0.018		
	(0.099)	(0.099)		
MUTUAL	0.475	0.424 ***		
	(0.111)	(0.102)		
*T1999*MUTUAL*	-0.180	-0.289 **		
	(0.153)	(0.137)		
T1998			-0.029	0.013
			(0.103)	(0.109)
MUTUAL			0.435 ***	0.403 ***
			(0.126)	(0.120)
*T1998*MUTUAL*			-0.064	-0.187
			(0.159)	(0.140)
ORG_SCORE		0.331 ***		0.327
		(0.045)		(0.044)
EXPER		0.071		0.064
		(0.082)		(0.088)
EDUC		0.015		0.006
		(0.078)		(0.082)
URB		-0.122 *		-0.127 *
		(0.063)		(0.065)
GDP_GWT		0.063		0.061
		(0.045)		(0.046)
HHI		-0.012		-0.013
		(0.007)		(0.007)
SOUTH		-0.010		0.057
		(0.126)		(0.138)
GDP_PC		0.000		0.000
		(0.000)		(0.000)
Nr of observations	119	113	119	113
Adjusted R2	0.181	0.387	0.148	0.361

The table reports the estimates from a Linear Probability Model; dependent variable: *DEF_5* indicates a default within the 5-th year since chartering date; *T1999* is a binary variable that assumes the value of 1 if the bank chartering date is before January 1999, 0 otherwise; *T1998* is a binary variable that assumes the value of 1 if the bank chartering date is before January 1998, 0 otherwisethe; definition of variables can be found in Table 3.
Robust standard errors in parenthesis.
* , ** and *** denote statistical signifcance at the 10%, 5% and 1% critical levels, respectively.

Tab 10 – Split-population model

	I.a	I.b
CAP	-2.129 ***	-1.159 ***
	(0.587)	(0.44)
MUTUAL	-1.436	-0.435
	(1.052)	(0.815)
ORG_SCORE	0.687 ***	0.763 ***
	(0.202)	(0.195)
EXPER	0.871 **	0.836 **
	(0.351)	(0.345)
EDUC	0.028	-0.126
	(0.308)	(0.311)
URB	-0.543 *	-0.498 *
	(0.286)	(0.268)
GDP_GWT	2.051 ***	-0.441 **
	(0.225)	(0.21)
HHI	0.011	-0.049
	(0.045)	(0.048)
SOUTH	0.099	0.701
	(0.618)	(0.722)
GDP_PC	0.000 *	0.000
	(0)	(0)
BUSINESS_MIX		-1.899 *
		(1.049)
DEV_ASST_GWT		1.078 ***
		(0.313)
DUMMY YEARS	YES	YES
Logit covariates	YES	YES
Number of observations	119	119

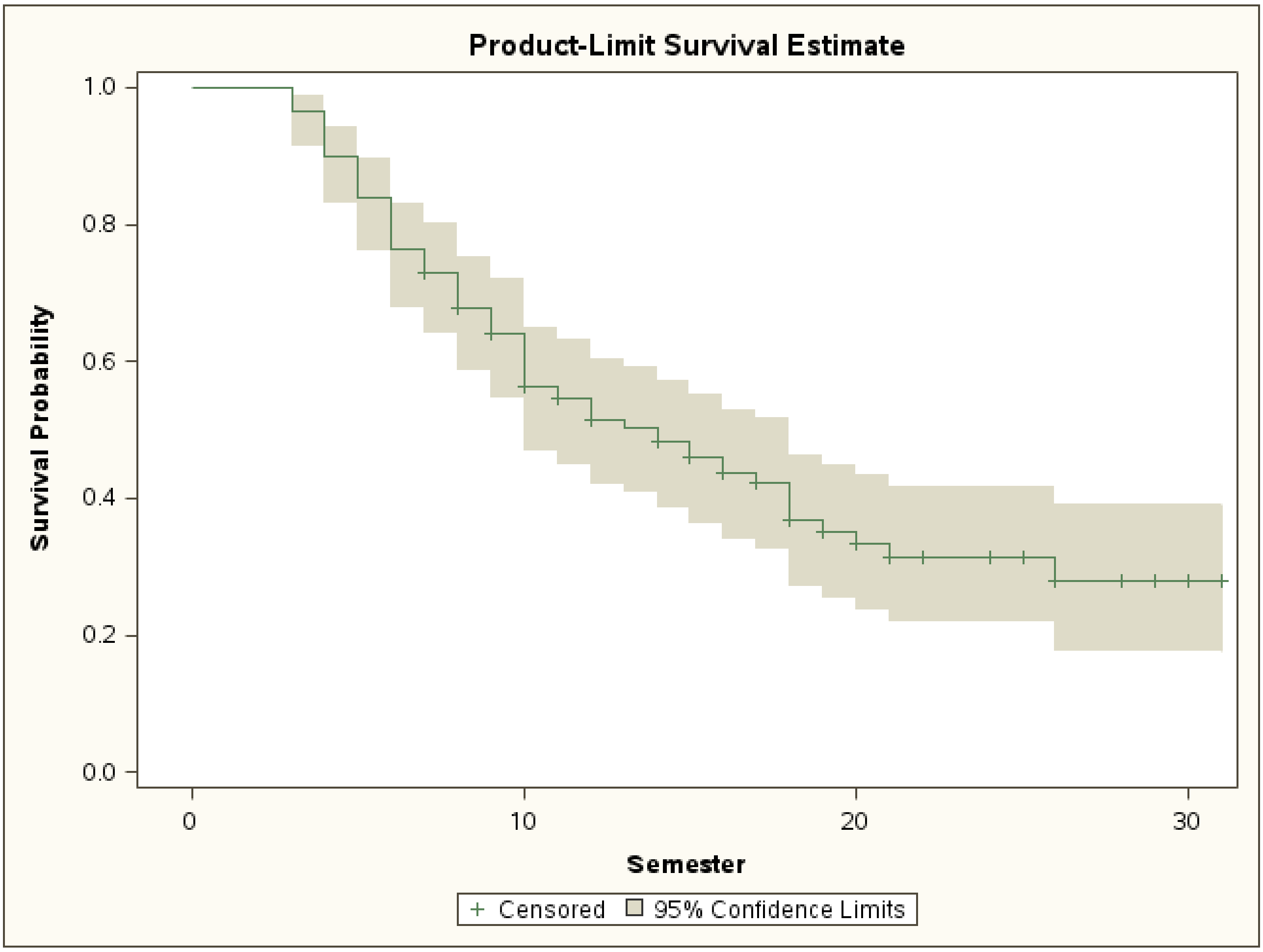
The table shows estimates from a split population model. Survival part: parametric log-normal AFT model; link part: logit model. Dependent variable: (the log of) semester of default. Definition of variables can be found in Table 3. Standard errors in parenthesis. * , ** and *** denote statistical signifcance at the 10%, 5% and 1% critical levels, respectively.

Tab. 11 – Propensity Score Matching

Matching algorithm	ATE
Kernel	-0.134
	(0.150)
Nearest neighbour	-0.188
	(0.190)

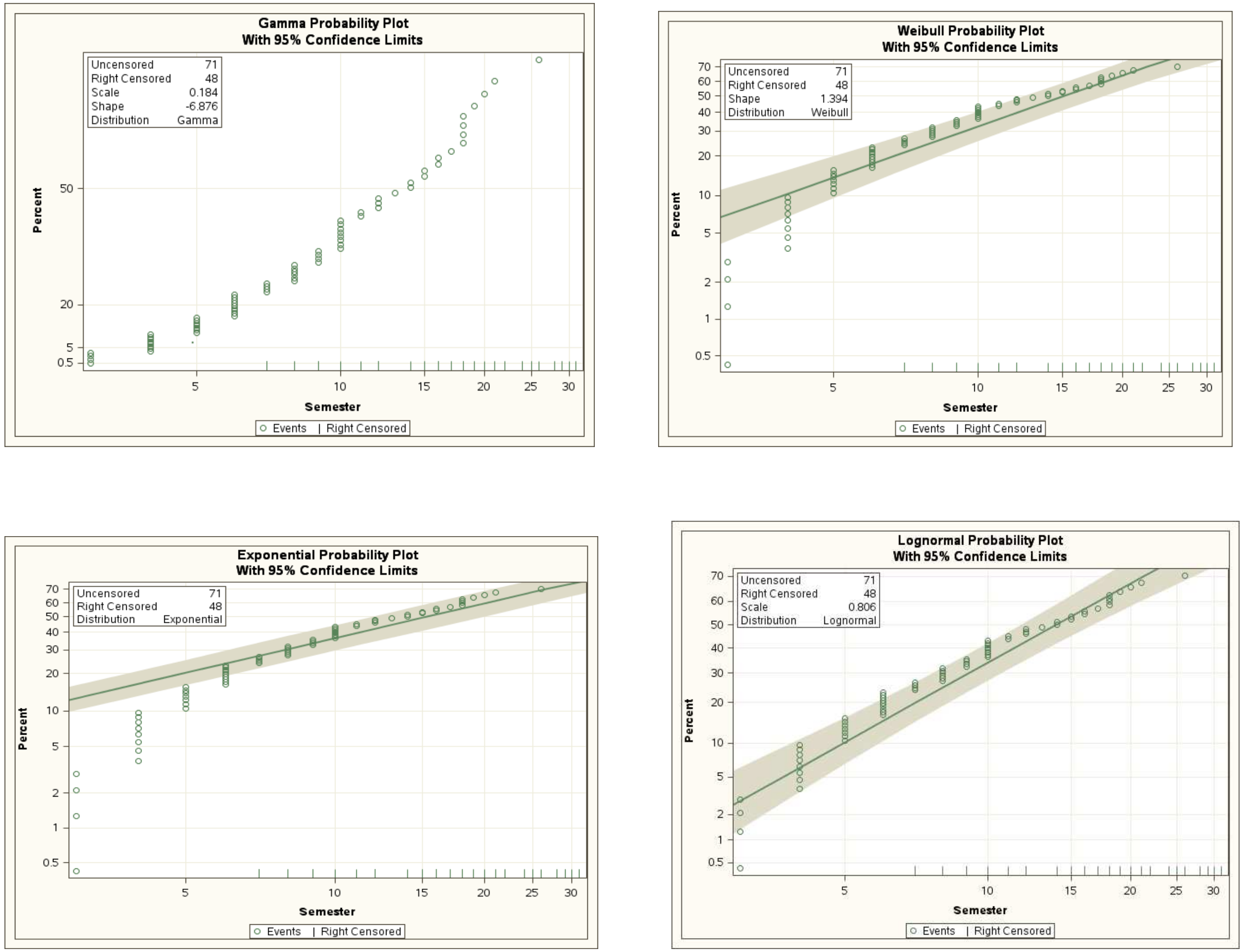
The table presents the average treatment effect (ATE) on the five-year probability of default estimated by propensity score matching. Standard errors in parenthesis.

Chart 1: Kaplan Mayer survival curve



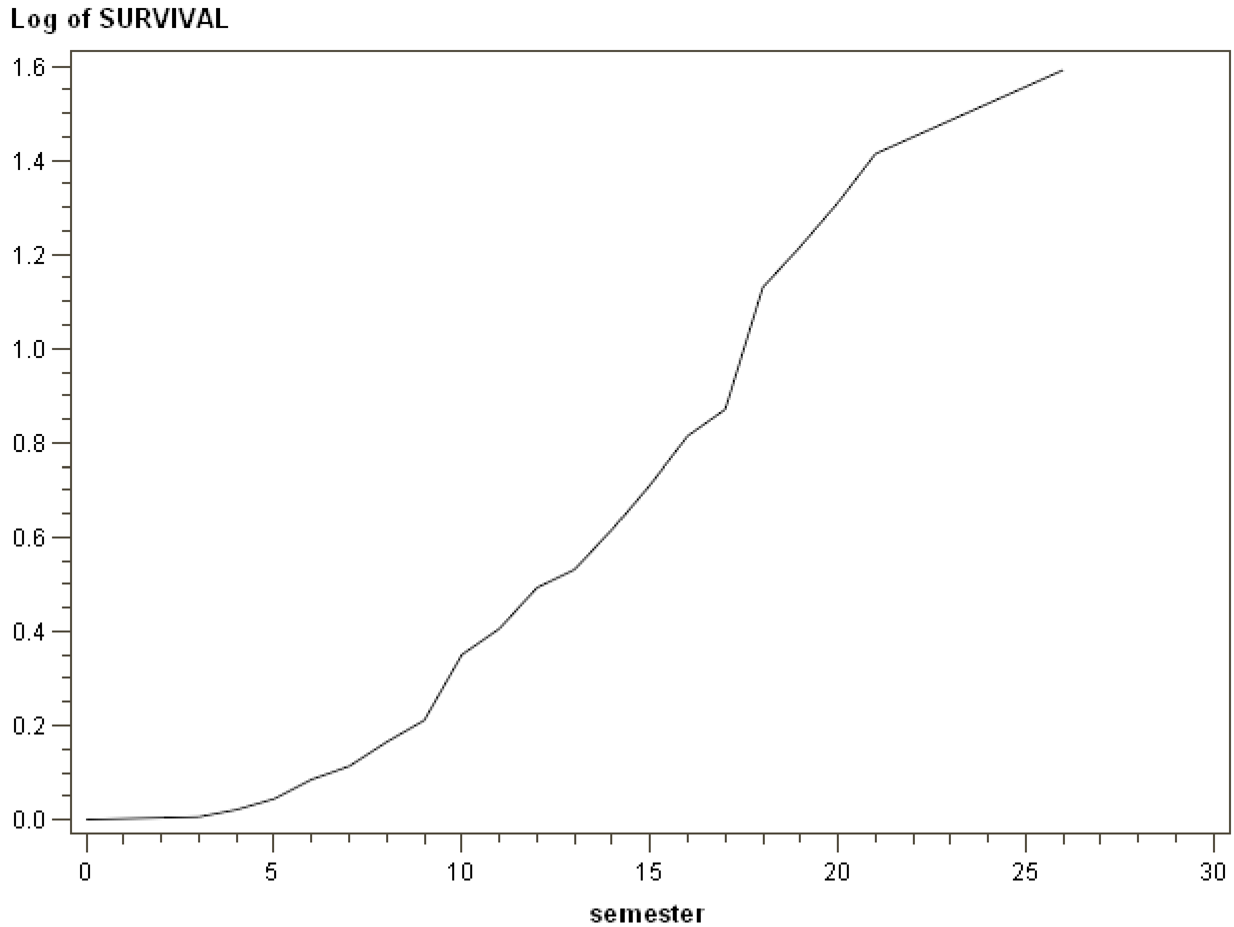
The chart presents the estimated survival function S(t) for the Kaplan-Mayer regression of time to default (in semester) without any predictor; the broken lines show a point-wise 95-percent confidence envelope around the survival function that represents the proportion of non-defaulted start-up banks.

Chart 2 – Probability plots of parametric models under four different distributions



The four charts show probability plots drawn using an inverse distribution scale, so that a cumulative distribution function (CDF) plots as a straight line. The straight line represents a nonparametric estimate of the CDF of the lifetime data.

Chart 3 – Probability plot of fitted Cox regression model



The chart shows the cumulative hazard (negative log-survivor) function .The slope of the solid line represents Cox model estimated hazard.

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