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SWITCHING COSTS IN LOCAL CREDIT MARKETS

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Abstract

Switching costs are a key determinant of market performance. This paper tests their existence in the corporate loan market in which they are likely to play a central role because of the complexity of contracts and informational problems. Using very detailed data at bank-firm level on four Italian local credit markets we empirically show that firms tend to iterate their choice of the main bank over time. This inertia is not related to unobserved and time invariant preferences of firms across banks and can be attributed to the existence of switching costs. We also offer evidence that banks price discriminate between new and old borrowers by charging lower interest rates to the former in order to cover part of the switching costs. The discount is about 44 basis points, equal to 7 per cent of the average interest rate. These results prove robust to a number of other potential identification drawbacks.

JEL Classification: L13, G21.
Keywords: switching costs, local credit markets, price discrimination, lending relationships.

Contents

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* Bank of Italy.

1. Introduction[1]

A buyer faces switching costs if an investment specific to her current seller must be duplicated for a new seller. This creates economies of scope among repeated purchases from the same supplier (Farrell and Klemperer, 2007). Switching costs have far reaching consequences on the standard competitive market equilibrium because they modify entry conditions as well as incumbent pricing strategies. In the case of banking sector switching costs are also relevant from a macroeconomic point of view. They may reduce price elasticity in retail markets so that the transmission of policy rate changes to retail rate dynamics may exhibit some form of sluggishness because banks may not find it profitable to adjust their prices frequently (European Central Bank, 2009).

Several arguments suggest that switching costs might be relevant in credit markets. First, there are transaction costs of closing the accounts with the current lender and opening new ones with another bank. Second, there exist learning costs such as costs of switching to a new bank following specific rules and practices in its lending activity after learning different rules adopted by the old lender. Third and more importantly, switching costs are also related to the investment in setting up a close tie with a bank (Boot, 2000). Changing the lender may imply the loss of a number of relationship-based benefits such as intertemporal smoothing, increased credit availability, enhancement of borrower's project payoffs, and more efficient decisions in case of financial distress.

In this paper we study switching costs in business local credit markets, by focusing on a specific kind of switching behavior that is the change of a firms' main bank. We focus on the main bank because multiple bank financing is a widespread phenomenon, even in the case of rather small firms (Detragiache et al., 2000) and, in this case, it not obvious how to define a switching episode. However, as indicated by Petersen and Rajan (1994), Elsas and Krahnen (1998) and Elsas, Heinemann and Tyrell (2004) multiple banking often coexists with the presence of one bank with a pivotal role, whose presence will reduce

[1] We are grateful to Marcello Bofondi, Luigi Buzzacchi, Giorgio Gobbi, Alfonso Rosolia and anonymous referees for useful comments. We also thank participants at the F.I.R.S. Conference on "Banking, Corporate Finance and Intermediation" (Shangai, June 2006), at the Conference on "The Changing Geography of Banking" (Ancona, September 2006) and at seminars held at the Bank of Italy and at the University of Bologna. Usual disclaimers apply. The views expressed in this paper are those of the authors and do not necessarily reflect those of the Bank of Italy.
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coordination costs of the other arm's-length lenders while the latter help mitigating the hold up problem generated by the privileged position enjoyed by the main (and more informed) bank. As a consequence also multiple bank firms are likely to face switching costs when they change their main bank, because at least the relationship lending-based investment is to be duplicated.

We analyze switching costs with two empirical exercises. First, we investigate their existence with a test that follows directly from the definition of switching costs: if they characterize the demand side, then choosing a specific banking partner today reduces the utility from selecting a different main lender tomorrow. Through a standard revealed preferences argument it is possible to show that this is equivalent to say that firms' choices across lenders are persistent over time. However, persistence in lending relationships could also be generated by unobserved time invariant bank-firm matches (the so called spurious state dependence). To take account of this, we propose a mixed logit model through which it is possible to measure true persistence in lending relationships by simultaneously controlling for time invariant preferences of borrowers across lenders. Using very detailed data at bank-firm level on four Italian local credit markets we find that firms changing their main lender incur significant switching costs. As far as we know, the assessment of switching costs through a mixed logit model aimed at detecting true inertia in buyer-seller relationships is new in the context of credit markets.[2]

Second, we test whether banks price discriminate between old and new borrowers offering more favorable conditions to the latter. In fact this is a generally agreed prediction in the Industrial Organization literature that analyzes pricing strategies in industries with heterogeneous switching costs and customer recognition (Chen 1997, Taylor 2003). Our empirical findings, mainly based on an interest rate equation, show that banks actually lure borrowers attached to competing main lenders with attractive entry-level offers. In our preferred specification, switching premium amounts to 44 basis points. This "paying customer to switch" evidence is robust to a number of controls including those for selectivity and firm-level omitted variables. Moreover teasing interest rates are also found in the case of multiple bank firms switching to an already known new main bank.

The existing literature on switching costs is huge and an exhaustive survey may be found in Farrell and Klemperer (2007). However, there are still few empirical contributions explicitly referred to the analysis of switching costs in business lending markets. Kim, Klinger and Vale (2003) infer the existence of switching costs and assess their magnitude in Norwegian credit markets by analyzing aggregate market share and interest rate dynamics. Gopalan, Udell and Yerramilli (2007) investigate motivations for firm switching to a new bank by using micro data. They find that firms decide to change their previous banking partner mainly to obtain higher loan amounts and hence to overcome borrowing constraints at their existing bank.

Another recent line of research analyzes whether switchers are offered a discount or alternatively pay a premium on the interest rates offered. Within the theory of insider vs. outsider lending (Sharpe, 1990) and using data drawn from the 1998 Survey of Small Business Finance, Black (2006) finds that outsider rates tend to be higher than insider rates.[3] Ioannidou and Ongena (2010) reach an opposite result: in their data on loans extended by Bolivian banks a firm borrowing from an outside bank is charged an interest rate that is more than 50 basis points lower than that charged on a comparable loan from its current inside banks.[4]

We contribute to these streams of literature in several ways. First, disentangling switching costs from unobserved heterogeneity in explaining the correlation over time of bank-firm matches has important consequences on the understanding of credit market dynamics. Consider, for example, a bank that makes a *transitory* loan interest rate cut. If the true model of firm behavior is characterized by unobserved heterogeneity and switching costs are absent, the price cut will give rise to a *transitory* market share increase for that bank. In presence of switching costs, however, the same strategy will generate a *non-transitory* increase in the number of attached borrowers and this, in turn, modifies dynamic pricing strategies, as our evidence on teasing interest rates shows. Second, our findings shed also a new light on the nature of bank-firm relationships in the Italian credit

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[2] Earlier empirical applications mainly regarded the realm of marketing science, health economics, transport economics and mobile telecommunications market. See, for instance, Erdem (1996), Johannesson and Lundin (2000), Brownstone et al. (2000) and Lee et al. (2006) and Grzybowski (2008), respectively.

[3] See also Black (2008). In this paper the author shows that theoretical predictions are unclear: the interest rate for firms borrowing from the inside lender may be higher or lower than those for firms that borrow from an outside lender.

[4] Moreover, these authors detect an interesting time pattern in the data showing that the initial switchers' advantage in terms of lending rates will be reduced and even reversed as they become attached or established borrowers.

market. The existence of a “paying customer to switch” strategy even in the case of firms selecting a new main lender with which they already had a lending relationship in the past points to the fact that the main lender plays a special role among the firm’s creditors. Notably, this holds true even in the case of the Italian credit markets where the fragmentation of credit supply is high and resorting to multiple lending is very common. Finally, our joint evidence on the true persistence in lending relationships and on teasing rates gives some clues on how to disentangle between alternative models of banking competition. While models based on Betrand competition can explain the existence of poaching strategies, they come to terms when they have to explain true persistence in bank-firm relationships. On the other hand, adverse selection models can easily explain borrowers’ lock-in but but are unable to justify the discounts offered to the firms switching to an already known bank. Models with heterogeneous switching costs and customer recognition can easily accommodate the two pieces of evidence.

The rest of the paper is organized as follows. In Section 2 we briefly recall theoretical contributions dealing with credit markets with switching costs. Section 3 describes the data. Our main results are presented in Section 4 and discussed in Section 5. Concluding remarks are presented in Section 6.

2. Theoretical background

In credit markets banks deliver their services directly to customers and hence they are able to know whether a given borrower is one of its current clients and price discriminate on the basis of this knowledge. Moreover, switching costs are likely to be heterogeneous across firms: for instance switching is expected to be costlier for a small and opaque single-bank firm with a well established relationship with a bank than for a large firm with multiple lenders and characterized by a large amount of hard information. In the Industrial Organization literature the models that best fit these two features are those analyzing markets with heterogeneous switching costs and customer recognition (Chen, 1997; Taylor, 2003). One general conclusion of this literature is that in equilibrium firms offer discounts to their competitors’ customers and that clients with “low” switching costs (below a certain threshold) change their supplier.

Beyond customer recognition and switching cost heterogeneity, credit markets exhibit additional peculiarities. First, borrowers may also differ in the quality of their investment projects and therefore in their ability to repay debt obligations. Moreover, there exist relevant asymmetries of information both between lenders and borrowers and, on the supply side, between informed and uninformed banks (Sharpe, 1990; von Thadden, 2004).

Gehrig and Stenbacka (2007) is one of the few attempts to adapt switching cost models with customer recognition to the case of business credit markets. The authors assume that there exist only short term loan contracts and that a firm resorts to credit in each period.[5] An unattached borrower can freely choose across alternative competing banks. Once it made its choice, the existence of switching costs affects the current choice in the sense that the firm receives a higher payoff if it chooses again the same lender it elicited in the past. Under this respect, switching costs produce the effect of establishing a causal link between past and current choices. As it will become clearer in Section 4, we will exploit this fact in order to identify the presence of switching costs in credit markets by using data on individual borrower credit histories.

The model has two periods. At the beginning of the first period all borrowers are unattached. At the end of the same time span, each borrowing firm chooses a specific banking partner (let us call it Inside bank or bank I). In the second period[6], the borrower wants to finance a new investment project requiring one unit of capital and returning $q > 1$ at the end of the period with probability π and 0 in the case of failure occurring with probability $1 - \pi$. Bank I will offer the unit of capital and require an interest rate equal to R_I on this sum; an outside bank (henceforth bank O) can also offer a loan contract to the attached borrower charging an interest rate equal to R_O. Banks will be paid back only in the case of a successful investment project, otherwise will receive nothing. Borrowers bear switching costs equal to s when moving from I to O. These costs are assumed to vary randomly across borrowers according to a uniform distribution defined on the $[0; \bar{s}]$ interval. It is also assumed that borrowers are not aware of their idiosyncratic switching costs until period 2. An attached borrower will compare the two loan contracts and decide to switch whenever $\pi(q - R_I) \leq \pi(q - R_O) - s$. This inequality implicitly defines a threshold level $s^* = \pi(R_I - R_O)$ such that bank I's customers with switching costs below it

[5] The absence of long term loan contracts might seem an extreme assumption about the working of credit markets. However, what it is crucial for the validity of switching cost models is that the two parties cannot write *complete* long term loan contracts.
[6] Consistently with our empirical design excluding unattached borrowers we skip over the analysis of the first period.

will switch to bank *O* while those with $s > s^*$ will stay with bank *I*. Within this set-up, the authors confirm the predictions on the existence of teasing rates: in equilibrium the outside bank will finance part of borrowers' switching costs by charging interest rates that are lower than those set by bank *I*. Namely, they will pay borrowers to switch. Moreover the discount offered to rivals' attached borrowers will increase with the intensity of switching costs measured by $\bar{s}$.

The authors also assume that there will be a proportion of borrowers whose investment projects will fail with certainty (i.e. with $\pi = 0$). In the second period, Bank *I* will be able to identify those borrowers with certainty and react by not renewing credit to them. The latter will switch to the bank *O* that in turn will be unable to identify those bad borrowers from the pool of switching firms (adverse selection). It can be shown that market equilibrium remains the same as that described above if the proportion of bad borrowers is under a given threshold.

This particular strategy aimed at introducing adverse selection enormously simplifies the model. However, a deeper integration between adverse selection and switching costs within a unified setting is a challenging task.[7] Rather than following that line of research, here we will investigate weather the evidence presented in Section 4 can be explained by resorting either to switching costs or to adverse selection as they were two separate theories on the working of credit markets.

3. Data

Our main data source is the quarterly Survey on lending rates carried out by the Bank of Italy since 2004 and including about 300 Italian banks. The sample is representative of credit markets at local (provincial) level. Information is available at firm-bank level and for each record matched, revocable and term loans and the interest rates charged on these operations are reported. Data also include several borrower characteristics like sector of economic activity, legal form and the municipality where the firm is located.

We merge this data set with additional information on bank characteristics taken from the Bank of Italy supervisory reports displaying branch locations and loans broken down by area and sector of economic activity. We restrict our analysis to business lending.

[7] For a credit market model combining switching costs and adverse selection within a different market environment, see Vesala (2007). This model, however, is not fully adequate to our aim because switching costs are not heterogeneous across borrowers.

Borrowers reporting bad loans are dropped from the sample as we want to exclude switching episodes that are due to a firm pathological condition. Data at our disposal refer to bank-firm relationships at two dates (March 2004 and March 2005, throughout the paper we refer to them as $t - 1$ and t, respectively). Firms that were not present in both dates were also excluded from the sample. The latter choice has two advantages. First, it enables us to address a threshold effect: the Survey on lending rates, in fact, only includes loans above 75.000 euros; accordingly, a specific borrower can enter or exit the sample due to reasons we could not control for. Second, the presence of a firm at the two dates is required because of the kind of switching event we are examining, i. e. one based on the possibility that the *same* firm might change a bank partner within that time span (see more on this below).

Our empirical strategy is also influenced by the need to keep the computational burden associated to the estimation of a mixed logit model (see equation 2 below) within reasonable limits. This also explains why the analysis is restricted to two dates. Besides, the time span between t and t - 1 corresponding to a one year period in our data seems to be appropriate to analyze the switching event. Computational reasons also induce us to restrict our sample to lending relationships between one of the top 15 banks operating in a province and the borrowers located in the same area. In fact, the mixed logit model requires a fairly limited number of alternatives in the choice set to be empirically manageable.[8]

At last, the analysis is focused on four local provincial markets: Turin, Bologna, Rome and Naples considered as separated entities. Again, computational reasons related to the mixed logit model estimation prevented us from increasing the number of bank-firm matches beyond a certain limit. But this aspect in our data is hardly a problem for the analysis as the number of observations is huge in each market. Furthermore, spatial segmentation is usually associated to credit markets in the light of the limited geographical scope of many lending relationships (see Petersen and Rajan, 2002, for the US and Degryse and Ongena, 2005, for the Belgian loan market).[9] The four selected provinces

[8] These banks are defined as those that are at the top in the ranking based on the number of customers they have in each local market. On average, these banks represent about 70 per cent of the total loans in each province. As a robustness check we further restrict the sample to the top 10 banks: (unreported) results are qualitatively unchanged.

[9] Kim, Klinger and Vale (2003)'s paper lacks this local dimension as they consider the Norwegian loan market at national level. To get around this problem, the authors run different estimations by splitting their

exhibit sharp differences in terms of per capita income levels, sectoral specialization, quality of the local institutions, size and concentration of the loan market (see Table 2). This huge heterogeneity serves as a check that the validity of our main findings extends across local environments showing different structural characteristics.[10]

The final sample includes about 79,000 bank-firm relationships and 50,000 borrowers. Table 1 contains a detailed description of the variables included in our sample while Table 3 shows summary statistics.

In the literature on switching costs it is usually assumed that a customer obtains its service or product from a single supplier. Consequently the switching event can be defined as the change in the identity of this unique supplier between the two periods. In credit markets however firms usually borrow from more than one bank thereby making the definition of the switching event more complex. For instance, a single-bank firm in $t-1$ could start getting credit from a new bank in t without breaking its pre-existing relationship. To address this problem in this paper a switching occurrence is defined as the change between $t-1$ and t of the firm's main bank, i.e. the bank granting the highest loan amount. This choice is motivated by the special role played by the main creditor. In our data set those lenders cover on average 87 per cent of total bank credit extended to each firm. Even considering exclusively firms borrowing from more than one bank, this percentage amounts to 67 per cent of a firm bank debt. Thus, given this strong concentration, it is likely that a relationship with the main bank will generate stronger benefits for the borrower and, consequently, increases its lock-in.[11]

Notably, our definition of the switching event encompasses both the switch toward a main bank in t who was a not a creditor of the firm in t - 1 and the case in which the new main lender is chosen among the set of those banks granting credit in t - 1. In the latter circumstance, the firm would not necessarily incur in the costs related to initiating a new bank-firm relationship (for instance think about the contractual costs generated by opening a new account and by the need to learn the new rules). Even in this case however, it is likely that there will be positive switching costs motivated by the special role played by the

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sample according to bank size (measured by the number of branches). This is hardly a solution as far as small banks are concerned because they are assumed to compete in the same national market.

[10] In this respect, it is worth noting that investigating how switching costs may vary according to differences in local credit markets is not a goal of our analysis.

[11] Elsas (2005) empirically shows that banks are more likely to be Hausbanks when their share of borrower debt financing is higher.

main creditor and hence by the need for the firm to further investing into the relationship to adapt to the changed identity of the new main bank. For instance, the substitution of the main lender could involve the need for the firm to increase the frequency of the contacts with the new main bank's loan officers to better know each other and to establish new formal and informal rules to follow in the future transactions and that fit better with the new role of that bank. We will come back to this issue in the empirical section.

4. Methodology and results

4.1. State dependence in bank-firm relationships

In order to assess the existence of switching costs we look at inertia in bank-firm relationships. A genuine causal effect between past and present choices made by firms when selecting their main lender would signal the presence of switching costs. However identifying such an effect is a challenging task as there exist two possible explanations for a positive correlation between repeated choices (Heckman, 1981). On the one hand, borrowing from a bank in the past alters current debtor preferences (so called "true state dependence") but, on the other hand, choices over time may be correlated solely because of temporally persistent unobservable factors influencing both the current and the past choice ("spurious state dependence"). In our setting distinguishing between these two explanations is crucial since only true state dependence would be conclusive on the existence of switching costs. In what follows we test for this causal linkage in firms' repeated choices by using, within a discrete choice framework, a mixed logit model that allows to rule out spurious state dependence by controlling for unobserved heterogeneity in time invariant firm-specific characteristics (Train, 2003).

We start by assuming that the net indirect benefit firm *i* obtains from choosing bank *j* as its main lender at date *t* is given by:

$$\Pi_{ijt} = \alpha_j + \lambda_j' Z_{it-1} + \gamma' X_{ijt} + \delta W_{ijt} + \varepsilon_{ijt} \tag{1}$$

where α_j, γ, λ_j and δ are parameters to be estimated and ε_{ijt} are random terms i.i.d. according to type I extreme value distribution. The deterministic part of the net benefit includes the following variables:

(i) α_j are bank fixed effects picking up (net) benefits originating from a specific lender and that are common to all potential borrowers;

(ii) $Z_{it\text{-}1}$ is a vector of firm characteristics including: borrower sector of economic activity (agriculture, industry, constructions and services), $LSIZE_{it\text{-}1}$, a proxy for firm size, $MONO_{it\text{-}1}$, a dummy variable for firms lending from a single bank. Note that the effect of each variable in $Z_{it\text{-}1}$ on Π_{ijt} varies with j;

(iii) X_{ijt} denotes a set of firm-bank variables including interest rates ($INTRATE_{ijt}$) and lender-borrower physical distance ($DIST_{ij}$). Both regressors are expected to have a negative effect on Π_{ijt}; $DIST_{ij}$ is included because traditional shoe-leather costs, as well as other relational specific investment expenditure will all increase with it.

(iv) $W_{ijt} = Y_{ijt\text{-}1} - 1$ where $Y_{ijt\text{-}1} = 1\{$firm i chooses bank j as its main lender in $t - 1\}$ and $1\{\cdot\}$ is the indicator function that is equal to 1 if the condition in the brackets is satisfied and zero otherwise. Hence W_{ijt} equals 0 if the previous choice of the main lender in confirmed and –1 otherwise so that δ measures the disutility from switching.

The estimation of parameters in (1) is based on the observed benefit-maximizing choices Y_{ijt} made by each firm. A standard maximum likelihood argument leads to a conditional logit model specification according to which the probability that firm i chooses bank j in period t as its main lender is given by (McFadden, 1974):

$$P\left(Y_{ijt}=1\right)=\frac{\exp\left(\alpha_j+\lambda_j' Z_{it-1}+\gamma' X_{ijt}+\delta W_{ijt}\right)}{\sum_k \exp\left(\alpha_j+\lambda_j' Z_{it-1}+\gamma' X_{ijt}+\delta W_{ijt}\right)}.$$

In this formulation W_{ijt} captures the correlation between repeated choices so that δ may pick up both state dependence and unobserved heterogeneity.[12] To overcome this difficulty and hence to identify true state dependence we assume that δ is randomly distributed across borrowers according to a parametric density function $g\ (\delta \mid \theta)$. Resulting choice probabilities are defined according to a mixed logit specification as follows:

$$P\left(Y_{ijt}=1\right)\equiv P_{ijt}=\int\left(\frac{\exp\left(\alpha_j+\lambda_j' Z_{it-1}+\gamma' X_{ijt}+\delta W_{ijt}\right)}{\sum_k \exp\left(\alpha_j+\lambda_j' Z_{it-1}+\gamma' X_{ijt}+\delta W_{ijt}\right)}\right)g\left(\delta \mid \theta\right)d\delta \qquad (2)$$

where α_j, γ, λ_j, and θ are parameters to be estimated.

Specification (2) allows to isolate the true state dependence by modeling the

[12] Econometrically, the inclusion of (a transformation of) the lagged endogenous variable among regressors may induce inconsistency in estimation if it is correlated with the current error term and this correlation is not modeled.

correlation between W_{ijt} and the error term. The expected value of δ will measure only true state dependence while unobserved heterogeneity will be picked up by the variance of δ. This can be easily shown considering that individual parameters for W_{ijt} can be expressed as $\delta_i = \delta_{mean} + \eta_i$ where δ_{mean} is the population mean and η_i is the individual stochastic deviation. The effect of the previous choice on the current benefit is now split in two additive terms: $\delta_{mean}W_{ijt}$ and $\eta_i W_{ijt}$. The random part $\eta_i W_{ijt}$ enters the stochastic portion of Π_{ijt} which now equals ($\eta_i W_{ijt} + \varepsilon_{ijt}$). This term is correlated over alternatives and time due to the common influence of η_i. With this specification the correlation between the lagged dependent variable and the current error term is explicitly modeled and δ_{mean} estimate is no longer affected by the endogeneity bias because, conditional on η_i, W_{ijt} is no longer correlated with the error term. At the same time a positive estimate of the variance of δ will signal that switching costs are heterogeneous and/or that some unobserved heterogeneity in time invariant preferences is at work.

For computational reasons the model in equation (2) is estimated separately for the four provincial markets (Turin, Bologna, Rome and Naples); $g(\delta \mid \theta)$ is specified as a Lognormal distribution since we expect δ to have a non-negative sign. Estimating parameters in (2) involves a missing data problem because $INTRATE_{ijt}$ is observed only for the bank-firm relationships which are in place (including those with lenders that are not a main bank). To tackle this problem, we impute lending rates for unselected alternatives with the fitted values of $INTRATE_{ijt}$ obtained by running a regression based on equation (3) (see below). This procedure introduces a generated regressor in the model and standard errors should be bootstrapped to correct for the variability of the first stage estimation. Unfortunately this is not a viable option in the context of the mixed logit specification given the constraints on computational resources. Hence estimated standard errors are to be considered as lower bounds and inference on the statistical significance of parameters has to be considered with that caveat in mind. Note however that the estimates we are interested in are so significantly different from zero that inferential conclusions seem to be valid even without bootstrapping standard errors (see below).

Estimates of the relevant parameters in choice probabilities (2) are reported in Table 4 (coefficients for bank fixed effects and for their interactions with firm characteristics are not shown). An overview of the main findings shows that these do not vary much across different provincial markets. Note first that the borrower-lender distance has a negative and significant effect on the probability of observing a specific bank-firm relationship. Borrowers whose locations are further off from those of a bank's branches are less likely to choose that bank as their main lender. Coefficients on the interest rate variable are always negative and significantly different from zero in all provincial markets, consistently with a downward sloping credit demand schedule: other things being equal, a bank charging higher interest rates with respect to its competitors will reduce the probability of being chosen as the firm's main source of credit. All in all, the observed matching factors used in our specification have a significant impact on the firm's choice of its main banking partner.

The estimated mean of ln (δ) is positive and significantly different from zero in all the provincial markets. Having controlled for time invariant and unobservable firm preferences across banks, this result signals the existence of a genuine causal link between lender-borrower matching over time or, equivalently, the existence of switching costs in credit relationships. All else being equal, a borrower changing its main lender will suffer a disutility that is significantly greater than zero. We also checked our findings assuming that δ is normally distributed and the results (available upon request) are qualitatively similar.

The standard deviation of ln (δ) is always highly significant in all provincial markets. As explained above this dispersion could reflect both heterogeneity of true state dependence across firms and the fact that some firms are better matched with a specific bank than other firms are (unobserved heterogeneity). Disentangling between the two components is beyond the scope of the present paper.[13]

4.2. Price discrimination between old and new borrowers

Having assessed the existence of switching costs we now turn our attention to the impact they may have on bank competition. More precisely, we investigate whether switching costs are associated with bank strategies aimed at offering better conditions to the switching firms (see Section 2 above). The basic interest rate regression we run is the following:

[13] However, it can be argued that the observed matching factors included in V_{ijt} should control for time-invariant matching factors in a credible manner so that the variability of δ could reflects at least in part genuine differences in switching costs.

$$INTRATE_{ijt} = const + \xi' CONTROLS + \beta DNEW_{ijt} + \mu_{ijt} \qquad (3)$$

where $INTRATE_{ijt}$, the loan interest rate charged by bank j to firm i in period t, is regressed on a set of controls including market, bank, loan contract and firm characteristics, and on a dummy $DNEW_{ijt}$ that equals one when bank j is firm i's main lender in t and it was not in $t - 1$ and zero otherwise. As in the previous section, we dropped those borrowers that were not present in the two dates. Our interest is focused on the parameter β capturing differential loan conditions when a firm turns to a new main bank. Other explanatory variables (*CONTROLS*) are given by:

(i) lender fixed effects controlling for any bank-specific factor such as marginal cost of funding and bank efficiency that might have an impact on lending rates; local market fixed effects capturing the influence of local market conditions;

(ii) bank j's local market power as measured by its market share in the local credit market;

(iii) firm-specific variables picking up borrowers' credit worthiness and their degree of informational opaqueness; they include firm size, a set of dummies indicating single-bank firms and limited liability enterprises, sectoral fixed effects (the adopted classification encompasses 187 industries). Our baseline equation (3) also includes the composition of a firm bank debt in terms of matched and term loans shares to control for the possibility that lending interest rates vary with contract loan characteristics (maturity, collateral requirements and other technical details).

To avoid simultaneity all time-varying regressors but $DNEW_{ijt}$ are taken with one-year lag. Regression results are shown in Table 5. The estimation is carried out on the pooled data referring to the four provinces; moreover, in our baseline regression the sample is restricted to those loans offered by main banks to maintain consistency with the framework adopted in the discrete choice model (see equations 1 and 2). Estimated parameters for provincial, sectoral and bank fixed effects are not reported. Notably, they are all jointly significantly different from zero, denoting that idiosyncratic factors featuring individual banks and provinces do affect interest rates.

In our baseline specification (column 1) the market share held by a bank within each provincial market has a positive effect on the cost of credit, consistently with the idea that local market power allows banks to charge higher interest rates. Large firms, limited liability companies and firms having relationships with many banks pay lower interest rates probably because less risky and less opaque firms have favorable credit conditions. Moreover our evidence shows that firms with higher shares of matched and term loans will be charged lower interest rates. This could be explained by a sort of a positive sorting effect according to which firms using long-term and more stable sources of credit are expected to be less risky than the others. But the most important result is related to the dummy $DNEW_{ijt}$. All else being equal, those firms that change their main bank are offered a switching premium of about 44 basis points. This finding is very similar to results documented in Ioannidou and Ongena (2010) who show that engaging a new bank decreases the rate paid on a new loan by more than 50 basis points. On the contrary, Black (2006) finds that interest rates on outside loans are around 40 basis points higher than those charged on loans from inside banks.

Our findings on new-old borrower price discrimination are robust to a number of checks aimed at addressing the potential bias arising from omitted variables or selectivity. In a second specification we add to the set of explanatory variables the lender-borrower geographical distance that proved to shape loan price conditions (Degryse and Ongena, 2005, Petersen and Rajan, 2002) and, at the same time, might be correlated with $DNEW_{ijt}$. Differently from other papers, it turns out that the estimated parameter for distance is never significantly different from zero (column 2). Thus, we do not find evidence of spatial price discrimination in local credit markets. All the other estimated parameters are left unchanged by this additional control.

Moreover one might argue that our findings on $DNEW_{ijt}$ are driven by the omission of guarantees, as far as new borrowers are requested to pledge more collateral and, at the same time, the latter reduce the cost of credit. Unfortunately data on collaterals are unavailable to us and hence we rerun regression (3) on the sub-sample of firms using only short term contracts (matched and revocable loans) which are typically not pledged with collateral (Sapienza, 2002). As shown in column 3 the new-old borrower price differential continues to be negative and highly significant.

Another robustness check is concerned with selectivity bias. Interest rates are observed only for the main bank relationships that are actually in place. The probability of observing a given bank-firm matching could depend on unobserved factors that might be correlated with residuals in the interest rate equation. Ignoring this circumstance might produce inconsistency in the estimated results. To tackle this potential shortcoming, a Heckman correction is introduced. Exploiting previous results on distance (see column 2), in the selection equation the probability of observing a bank-firm matching is estimated as a function of borrower-lender distance and the set of bank and firm characteristics used in the main equation. Parameters are estimated using a maximum likelihood full information method and results are shown in column 4. Unreported evidence on the selection equation shows that distance has a significant and negative effect on the probability of observing a bank-firm match (see also Table 4). A rho test rejects the independence between the two equations, showing that a selectivity bias may be an issue in our estimates. More importantly, selectivity does not affect the existence of a discount for firms switching to a new main lender.

Up to now, we have excluded loans offered by lenders that are not a main bank. In column 5 we remove this constraint and assume that all customers starting a new lending relationship with *any* bank may be offered different interest rates with respect to the established borrowers. Accordingly $DNEW_{ijt}$ is replaced by $DNEW_{ijt}'$ that is set to one if in t - 1 firm i did not borrow from bank j, regardless the status of j (i.e. being j either a main or a non main bank, see Table 1). Results confirm our conclusions on the existence of a switching premium. Moreover, a comparison between the two estimated parameters in colums (1) and (6) clearly show that banks have to offer larger discounts if they want to substitute a past main lender than those that they should offer in the case they would aim at substituting a non main bank. Hence these findings clearly point to the importance and the privileged role played by the main lender.[14]

Finally, another criticism which can be raised is that the omission of firm-level variables correlated with the switching behavior could generate a bias in the parameter for $DNEW_{ijt}$. A straightforward solution would consist in adding borrower fixed effects to the specification. Unfortunately, our cross section does not allow us to perform such an exercise for the entire sample. However, it is possible to use firm fixed effects by restricting the sample to borrowers resorting to multiple lending and focusing only on $DNEW_{ijt}'$. As shown in column 6, a new lending relationship is associated with a lower

[14] It would have been interesting to estimate the effect of "new customer" status without a pre-existing credit history with the whole banking system. Unfortunately data at our disposal do not permit this exercise because

interest rate even after controlling for any unobserved time-invariant borrower characteristic.

4.3. Refinements on switching behavior and loan rates

In the previous subsection we showed that borrowers who changed their main lender pay lower interest rates on their current loans. A question arises: does this discount depend on the mobility episode? Or may it be the case that switchers differ systematically from stayers in certain permanent characteristics that influence their switching behavior as well as the interest rates they pay and that are not controlled for? We think that the former interpretation is more plausible because *(i)* the controls included in specification (3) are likely to capture most of the unobserved heterogeneity and *(ii)* switchers are charged more favorable lending rates also in the regression with borrower fixed effects (see Table 5 column 6). Nevertheless these arguments are not conclusive and some sort of confounding factor may be at work. To explore further this issue we adopt a counterfactual approach. Namely we control for unobserved borrower features by explicitly comparing interest rates paid by switchers with those the same borrowers would have paid if they would not be new client of their current bank. While the first set of prices is observed, counterfactual interest rates are not and have to be estimated. To this aim we follow Ioannidou and Ongena (2010) and adopt an exact matching estimation strategy whose steps are as follows:

(1) we split the set of firms into two subsamples: the switchers (group T) and the stayers (group C);

(2) for each firm $i \in T$ we construct a control group $C(i)$ made of all firms belonging to C that are *similar* to firm i. A firm is considered *similar* to i if, in the period t, borrows from the same main bank of firm i and matches the same characteristics of i (e.g. size, sector of economic activity, etc.). For continuous firm characteristics (say, size) a firm is included in $C(i)$ if its size belongs to a (-20%, +20%) window of firm i size while for qualitative firm characteristics the matching is obvious. The matching schedule varies according to the number of characteristics required. A firm $c \in C$ may belong to more than one control group, that is it may happen that $c \in C(i)$ and $c \in C(j)$ for $i \neq j$ and $i, j \in T$ (matching with replacement).

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a firm could be present in the second period and not in the first one only because of a threshold effect (see also Section 3).

(3) for all $i \in T$ we take the spreads between the interest rate charged to firm i and those paid by firms belonging to $C(i)$;

(4) we finally regress the spreads on a constant and cluster errors at switcher level.

The results of this procedure are reported in Table 6 where each column corresponds to a different matching scheme. For example, matching on bank, province, sector of economic activity and size (column 1) leaves 918,797 observations and 5,481 switchers, implying that for each switcher there are on average 168 *similar* stayers. Spreads vary between – 12** and – 58*** basis points and they are significantly less than zero regardless of the matching schedule. Notably, in the last column, in which matching variables are the same as those used in the regression (3) (except for a less fine sectoral breakdown), a switcher receives a discount equal to 42*** basis points that is very similar to the result shown in Table 5 column 1.

5. Discussion

The econometric evidence collected so far leads to two main conclusions. Switching to a new main lender is costly in terms of a one-shot reduction of the indirect utility associated to the lending relationship. Moreover, banks price discriminate between old and new borrowers by offering a discount to the switchers. These findings are consistent with the tenets of switching costs models with customer recognition illustrated in Section 2. However, one could argue that the same evidence can be accommodated by models that do not explicitly assume the existence of switching costs and whose predictions are observational equivalent to them.

For instance, with Bertrand competition banks may reduce interest rates to attract more customers and this kind of strategy is not motivated by attached customers' lock-in. In this perspective, Fudenberg and Tirole (2000; henceforth FT) present a model of behavior based price discrimination whose predictions are fully consistent with our empirical results on the insider-outsider interest rate difference. In their model of duopoly with horizontal differentiation, short-term contracts and time-invariance of consumers' brand preferences, FT show that firms offer second period discounts to customers attached to their opponent in the first period. Moreover, a share of customers will switch supplier.[15]

[15] In FT original model the authors consider a Hotelling model with two firms located at the extremes of the unit line and a population of consumers that are distributed along the line according to a cumulative distribution satisfying the monotone hazard rate condition. There are two periods and consumers' preferences

Though similar in terms of predictions on price discrimination, the two models radically differ in the mechanisms triggering these results. Specifically, in switching costs models lending history is relevant because through the presence of exogenous switching costs borrowers will be ex post locked into the past relationship. Competing banks may have an incentive to lower prices in order to tease these rival's attached borrowers. In FT, there are no switching costs but borrowers' preferences are differentiated across banks. Thus, lending history matters because it reveals the intensity of borrowers' preferences across banks and in doing so signals which customers may be the target for rival banks' poaching strategies.

An empirical test that can discriminate between the two alternative explanations is beyond the scope of the present paper. However it is possible to argue that FT's model can hardly substitute for an explanation of our empirical evidence based on switching costs. In fact, although FT's predictions are consistent with the results stemming from the interest rate equation, they do not come up with our findings on inertia in lending relationships. In FT model consumer choices are positively correlated across time because preferences are invariant from one period to another. Once their time invariant nature is controlled for as in our mixed logit specification, there should be no reason for the past decisions to influence the current ones. Summing up, switching costs are able to explain the additional evidence on inertia that instead can be hardly accommodated within a model of horizontal product differentiation like that proposed by FT.

Inertia and interest rate discrimination across borrowers can also be predicted within adverse selection models. Apart from switching costs, borrowers may also differ in terms of the quality of their investment projects (see Section 2). Specifically assume that there are two types of firms: those having investment projects with an high probability of success π_G and those holding investment projects with a lower probability of success π_B (obviously, $\pi_G > \pi_B$). Switching costs are assumed to be zero. Thanks to its past relationship with the borrower, bank *I* is assumed to observe a signal about firms quality whose knowledge is not available to the outsider. Accordingly, the insider will bid aggressively for borrowers that are considered of good quality according to its signal,

over the two alternatives are time invariant and there are no switching costs in moving across these alternatives. In the second period, firms are able to recognize the customers they served in the past and can price discriminate accordingly.

while it will be less aggressive in bidding for bad quality firms. As a consequence, the *O* bank will be jeopardized in its ability to compete for borrowers as it has to face the risk of lending mostly to bad quality firms. The threat of adverse selection will reverberate on outsider pricing strategies and may limit the intensity of price competition.

These are the basic elements of second period Sharpe (1990)'s equilibrium model as emended by von Thadden (2004). A clear cut prediction from it is that borrowers' lock-in will increase with the severity of adverse selection. But what are model insights as far as the interest rates charged by bank *I* and *O* are concerned? Black (2008) recently shows that interest rates charged to firms that borrow from bank *I* may be higher or lower than those charged to firms that borrow from bank *O* in Sharpe-von Thadden model, depending on the values taken on by model parameters.[16] In particular, it is shown that when the difference between π_G and π_B is large, expected interest rates paid by borrowers switching to the bank *O* will be higher than those charged by the bank *I* to its loyal borrowers. The opposite holds true when parameter space is such that the difference between π_G and π_B is relatively small. Moreover, Ogura (2006) emphasizes that predictions from adverse selection models in terms of interest rate difference between insider and outsider banks are very sensitive to the way with which the rules of the bidding game are designed. In particular, he shows that when the bidding game between bank *I* and *O* is an English auction, switchers will pay higher interest rates than those charged by the *I* bank to the stayers. This uncertainty about theoretical outcomes is also reflected by recent empirical findings on this topic.

The ambiguous predictions on price discrimination obtained from adverse selection models make very difficult to compare them with those deriving from switching cost literature. Here, we circumvent this difficulty by proposing two tests that can shed some light on that issue. First, following Black (2006), we restrict our interest rate regression to small firms. In fact, the variance of π*'s* should be maximum within this size category.[17] Thus, according to Sharpe-von Thadden model, it should be more likely that switchers pay higher interest rates than those charged to borrowers that do not change their banking partner. This exercise leads to an estimate of the parameter of $DNEW_{ijt}$ equal to -0.623*** (standard error = 0.054), thereby showing that new-old borrowers interest rate difference is

[16] Black (2008) also show bank *I* interest rates are higher than those charged by bank *O* for given borrower type.
[17] Small (large) firms are those whose *LSIZE* is below (greater than) the median value.

still negative for the small firms sample and is even larger than that found for the whole sample. Thus we found evidence in favor of poaching strategies followed by the *O* banks even in the circumstance in which it should be more likely to observe the opposite according to the tenets of the adverse selection models. Accordingly, we conclude that our evidence on interest rate discrimination fits better a model with heterogeneous switching costs and customer recognition. Interestingly, this result may also indicate that switching costs are higher in the case of small businesses which on average take higher advantage from relationship lending and are more likely to have a single bank partner (that implies higher transaction and learning costs).

Our second test is aimed at comparing interest rates set by *I* and *O* banks when the intensity of adverse selection is at a minimum. This enables us to observe price discrimination in a setting where the potential role of adverse selection is strongly limited. This situation may occur when a borrower switch to a new main lender with which the borrower already had a lending relationship in the previous period. In fact adverse selection is likely to manifest itself more intensively when a completely new relationship with a bank has started. At the same time, as already explained, we could expect that some kind of switching costs will be generated even when switching occurred toward an already known bank. Hence we rerun regression (3) on the subsample of multiple-bank firms that either do not switch or switch to an already known bank. "Paying customers to switch" strategies are again confirmed by this additional evidence: the parameter for $DNEW_{ijt}$ equals -0.171*** with a standard error of 0.035. Interestingly, the estimated switching premium is negative and can be consistent with our interpretation that a change in the main lender can generate switching costs even if that bank was among the firm's creditors in period t - 1. Moreover, the estimated discount is much lower than that observed for the sample including also the switching to a completely new main lender (see table 5, column 1). This is consistent with the fact that in this case switching costs also include the costs to start a completely new lending relationship.[18]

6. Concluding remarks

[18] Interestingly, according to unreported evidence estimated switching costs (according to the procedure illustrated in Subsection 4.1) are significantly greater than zero (but with a lower magnitude) also in the case of multiple-bank firms. Results are available upon request.

This paper investigates the issue of switching costs in lending markets where they are expected to be relevant because of the complexity of bank-firm contracts and the asymmetries of information between inside and outside banks. Using bank-firm matched data on Italian local credit markets we identify two basic facts that have important consequences for that environment. First, through a mixed logit model we show that firms tend to iterate their choice of the main bank over time. Since this finding is not related to unobserved and time invariant firms' preferences across banks it signals the existence of switching costs: turning to a new bank is costly in terms of a one-shot reduction of the indirect benefit a firm receives from its lending relationship. Second, it turns out that banks offer lower interest rates to their new customers to cover part of these costs, consistently with the tenets of literature on switching costs with customer recognition. The magnitude of that discount is non-negligible: on average it amounts to about 44 basis points and is equal to 7 percent of the average interest rate.

In general, our results put emphasis on the relevance of switching costs for the analysis of bank-firm relationships and competition in credit markets. Moreover, they call for a stronger integration between the traditional topics of the banking literature like adverse selection, moral hazard and asymmetric information and those typical of theoretical and empirical contributions dealing with switching costs in the Industrial Organization literature.

Tables

Table 1 – Variables definition

Variable	Definition
	Firm varying
Sector of economic activity	187 sectors belonging to agriculture, industry, constructions and services and broadly corresponding to the three digits ISIC (International Standard Industrial Classification of all Economic Activities) classification
$LSIZE_{it}$	Natural logarithm of the sum of loans extended to firm *i* in period *t*. The sum is over all the bank-firm relationships recorded in the Survey on lending rates and regarding firm *i*
$DLTD_{it}$	Dummy variable equal to one if firm *i* is a limited liability company in period *t* and zero otherwise
$MONO_{it}$	Dummy variable equal to one if firm *i* is a single-bank borrower in period *t* and zero otherwise
SHM_{it}	Share of matched loans in firm *i*'s bank debt portfolio in period *t*
SHT_{it}	Share of term loans in firm *i*'s bank debt portfolio in period *t*
	Firm-bank varying
$INTRATE_{ijt}$	Loan interest rate charged by bank *j* to firm *i* in period *t*. It is computed as a weighted average of interest rates charged on matched, term and revocable loans
$MS_{jp(i)t}$	Bank *j*'s loan market share in period *t* in the province *p*(*i*) where firm *i* is located
Y_{ijt}	It is set equal to 1 if firm *i* chooses bank *j* as its main lender in *t* and zero otherwise. Formally Y_{ijt} = 1{firm *i* chooses bank *j* as its main lender in *t*} and 1{·} is the indicator function that is equal to 1 if the condition in the brackets is satisfied and zero otherwise
$DNEW_{ijt}$	Dummy variable equal to one when bank *j* is firm *i*'s main lender in *t* and it was not in *t* – 1 and zero otherwise. It holds that $DNEW_{ijt} = Y_{ijt}(1 - Y_{ijt-1})$
$DNEW_{ijt}'$	Dummy variable equal to one when bank *j* is one of firm *i*'s lenders in *t* and it was not in *t* – 1 and zero otherwise
$DIST_{ij}$	Physical distance between firm *i* and bank *j*. It has been computed as kilometers between the municipality where the firm is located and the municipality where the bank has the nearest branch to that firm. For some bank-firm relationship distance is zero because the bank has at least one branch in the municipality where the firm's headquarter is located. To circumvent this problem we substitute zeros with the ray of the circumference with the same area of that of the municipality. It is equivalent to approximate the municipality surface with a circumference and to assume that branches are located in the centre of the circumference while firms are uniformly distributed on the boundaries. This seems to be a reasonable assumption since branches are usually located where the population density is higher while firms are generally located far from cities centers. With such a substitution it may happen that the distance within a municipality is greater than some of the distance between municipalities. In this case we pick up the minimum of the two distances
W_{ijt}	It is equal to zero if in *t* – 1 bank *j* is firm *i*'s main lender and – 1 otherwise. It holds that $W_{ijt} = Y_{ijt-1} - 1$

Table 2 – Main local market characteristics

	Turin	Bologna	Rome	Naples
Per capita real GDP (000 euros) - 2002	19.7	22.2	20.1	10.6
Value added composition (percent.) - 2003				
Agriculture, hunting, forestry and fishing	0.8	1.5	0.5	1.2
Manufacturing	24.7	25.6	8.2	12.2
Construction	4.6	4.3	4.3	5.2
Services	69.9	68.7	86.9	81.4
Market size (loans extended to firms, millions euros) – December 2003	22,649	16,018	64,521	11,184
Herfindhal index on loans	0.0605	0.0582	0.0352	0.0595
Social capital (# bags of blood donated per million inhabitants in 1995)	38.9	75.7	17.2	9.0

Table 3 – Descriptive statistics of the sample

Variable	No. of Obs.	Mean	Std. Dev.	Min.	Max.
			Turin		
Interest rate [*INTRATE*] – only main bank	14562	5.993	2.117	0.000	13.744
Interest rate [*INTRATE*] – all bank-firm relationships	23964	6.105	2.176	0.000	19.127
Size [*LSIZE*]	14562	11.944	1.525	0.000	20.169
Share of matched loans [*SHM*]	14562	0.229	0.323	0.000	1.000
Share of term loans [*SHT*]	14562	0.466	0.417	0.000	1.000
Single-bank borrower [*MONO*]	14562	0.650	0.477	0.000	1.000
Limited liability borrower [*DLTD*]	14562	0.429	0.495	0.000	1.000
New main relationship [*DNEW*]	14562	0.122	0.328	0.000	1.000
New relationship [*DNEW′*]	23964	0.078	0.268	0.000	1.000
Lender-borrower distance [*DIST*] - main relationship	14562	5.154	3.245	0.706	65.567
Lender-borrower distance [*DIST*] - all relationships	218430	7.755	6.757	0.706	76.338
Market share [*MS*]	15	0.045	0.044	0.011	0.150
			Bologna		
Interest rate [*INTRATE*] – only main bank	10466	4.765	1.555	0.842	15.660
Interest rate [*INTRATE*] – all bank-firm relationships	17791	4.866	1.644	0.842	15.660
Size [*LSIZE*]	10466	12.036	1.578	0.000	19.658
Share of matched loans [*SHM*]	10466	0.269	0.339	0.000	1.000
Share of term loans [*SHT*]	10466	0.471	0.410	0.000	1.000
Single-bank borrower [*MONO*]	10466	0.607	0.488	0.000	1.000
Limited liability borrower [*DLTD*]	10466	0.472	0.499	0.000	1.000
New main relationship [*DNEW*]	10466	0.129	0.335	0.000	1.000
New relationship [*DNEW′*]	17791	0.071	0.257	0.000	1.000
Lender-borrower distance [*DIST*] - main relationship	10466	5.355	2.311	1.395	45.522
Lender-borrower distance [*DIST*] - all relationships	156990	8.281	6.191	1.395	51.610
Market share [*MS*]	15	0.043	0.044	0.011	0.168

Table 3 – Descriptive statistics of the sample (continued)

Variable	No. of Obs.	Mean	Std. Dev.	Min.	Max.
			Rome		
Interest rate [*INTRATE*] – only main bank	16014	6.683	2.494	0.002	19.520
Interest rate [*INTRATE*] – all bank-firm relationships	23351	6.945	2.526	0.000	19.520
Size [*LSIZE*]	16014	12.064	1.721	0.000	20.968
Share of matched loans [*SHM*]	16014	0.159	0.287	0.000	1.000
Share of term loans [*SHT*]	16014	0.458	0.442	0.000	1.000
Single-bank borrower [*MONO*]	16014	0.702	0.458	0.000	1.000
Limited liability borrower [*DLTD*]	16014	0.646	0.478	0.000	1.000
New main relationship [*DNEW*]	16014	0.090	0.286	0.000	1.000
New relationship [*DNEW′*]	23351	0.072	0.258	0.000	1.000
Lender-borrower distance [*DIST*] - main relationship	16014	3.987	7.227	0.003	60.497
Lender-borrower distance [*DIST*] - all relationships	240210	8.529	10.136	0.003	60.497
Market share [*MS*]	15	0.031	0.029	0.005	0.100
			Naples		
Interest rate [*INTRATE*] – only main bank	8289	6.802	2.473	0.861	16.601
Interest rate [*INTRATE*] – all bank-firm relationships	13792	7.041	2.532	0.743	18.226
Size [*LSIZE*]	8289	12.035	1.635	0.693	18.276
Share of matched loans [*SHM*]	8289	0.185	0.298	0.000	1.000
Share of term loans [*SHT*]	8289	0.416	0.425	0.000	1.000
Single-bank borrower [*MONO*]	8289	0.648	0.478	0.000	1.000
Limited liability borrower [*DLTD*]	8289	0.559	0.497	0.000	1.000
New main relationship [*DNEW*]	8289	0.126	0.332	0.000	1.000
New relationship [*DNEW′*]	13792	0.086	0.280	0.000	1.000
Lender-borrower distance [*DIST*] - main relationship	8289	15.333	52.910	0.718	392.437
Lender-borrower distance [*DIST*] - all relationships	124335	41.944	98.982	0.524	399.448
Market share [*MS*]	15	0.047	0.040	0.014	0.147
			Entire sample		
Interest rate [*INTRATE*] – only main bank	49331	6.093	2.336	0.000	19.520
Interest rate [*INTRATE*] – all bank-firm relationships	78898	6.238	2.397	0.000	19.520
Size [*LSIZE*]	49331	12.018	1.621	0.000	20.968
Share of matched loans [*SHM*]	49331	0.207	0.314	0.000	1.000
Share of term loans [*SHT*]	49331	0.456	0.426	0.000	1.000
Single-bank borrower [*MONO*]	49331	0.657	0.475	0.000	1.000
Limited liability borrower [*DLTD*]	49331	0.530	0.499	0.000	1.000
New main relationship [*DNEW*]	49331	0.114	0.318	0.000	1.000
New relationship [*DNEW′*]	78898	0.076	0.265	0.000	1.000
Lender-borrower distance [*DIST*] - main relationships	49331	6.528	22.528	0.003	392.437
Lender-borrower distance [*DIST*] - all relationship	739965	13.863	43.134	0.003	399.448

Table 4 – Mixed logit model

The table reports maximum likelihood estimates for the model defined in equation (2). The dependent variable is the probability that in period t (March 2005) firm i chooses bank j as its main lender. *INTRATE* is the interest rate charged by bank j to firm i in period t (March 2005). *DIST* is the physical distance between firm i and bank j. W is variable equal to zero if in t – 1 (March 2004) bank j is firm i's main lender and – 1 otherwise. All specifications include (unreported) bank-fixed effects interacted with firm characteristics including four dummies for the sector of economic activity (agriculture, industry, constructions and services), firm size proxied by the natural logarithm of the sum of loans extended by all banks to firm i in period t – 1 (March 2004) (*LSIZE*) and a dummy variable for single-bank firms (*MONO*). The symbol * indicates the coefficient is significantly different from zero at the 10 percent level; ** at 5 percent; *** at 1 percent.

	(1) Turin	(2) Bologna	(3) Rome	(4) Naples
Interest rate	- 0.667***	- 0.504***	- 0.510***	- 0.662***
[*INTRATE*]	(0.017)	(0.021)	(0.014)	(0.019)
Distance	- 0.078***	- 0.179***	- 0.038***	- 0.045***
[*DIST*]	(0.009)	(0.012)	(0.008)	(0.013)
W				
Mean of $ln(\delta)$	1.961***	1.928***	2.143***	2.018***
	(0.052)	(0.082)	(0.082)	(0.075)
Std. dev. of $ln(\delta)$	0.886***	0.892***	0.808***	0.918***
	(0.060)	(0.097)	(0.077)	(0.085)
Bank fixed effects	YES	YES	YES	YES
Log Likelihood	- 8,689	- 6,940	- 8,153	- 5,341
Likelihood ratio index	0.780	0.755	0.812	0.762
Observations	14,562	10,468	16,020	8,289
Number of cases	218,430	157,020	240,300	124,335

Table 5 - Interest rate regression

The table reports OLS estimates for the model defined in equation (3). The dependent variable *INTRATE* is the interest rate charged by bank *j* to firm *i* in period *t* (March 2005). *MS* is the loan market share of bank *j* in the province where firm *i* is located. *LSIZE* is the natural logarithm of the sum of loans extended by all banks to firm *i* in period t – 1 (March 2004). *SHM* is the share of matched loans in firm *i*'s bank debt portfolio in period t – 1 (March 2004). *SHT* is the share of term loans in firm *i*'s bank debt portfolio in t – 1. *MONO* is a dummy variable equal to one if firm *i* is a single-bank borrower in t – 1 and zero otherwise. *DLTD* is a dummy variable equal to one if firm *i* is a limited liability company in t - 1 and zero otherwise. *DNEW* is a dummy variable equal to one when bank *j* is firm *i*'s main lender in *t* and it was not in t – 1 and zero otherwise. *DIST* is the physical distance between firm *i* and bank *j*. $DNEW'$ is a dummy variable equal to one when bank *j* is one of firm *i*'s lenders in *t* and it was not in t – 1 and zero otherwise. Robust standard errors are reported in parentheses. The symbol * indicates the coefficient is significantly different from zero at the 10 percent level; ** at 5 percent; *** at 1 percent.

	(1) baseline	(2) distance	(3) short term	(4) heckman	(5) new all relat.	(6) firm FE
Provincial market share [*MS*]	2.371*** (0.489)	2.405*** (0.510)	3.361*** (1.034)	2.271*** (0.478)	1.514*** (0.383)	-0.210 (0.531)
Firm size [*LSIZE*]	-0.117*** (0.008)	-0.117*** (0.008)	-0.087*** (0.012)	-0.117*** (0.006)	-0.126*** (0.006)	
Share of matched loans [*SHM*]	-1.274*** (0.037)	-1.274*** (0.037)	-1.205*** (0.048)	-1.274*** (0.034)	-1.474*** (0.031)	
Share of term loans [*STM*]	-3.010*** (0.029)	-3.010*** (0.029)		-3.010*** (0.025)	-2.801*** (0.026)	
Single-bank firm [*MONO* = 1]	0.162*** (0.023)	0.162*** (0.023)	0.252*** (0.043)	0.164*** (0.022)	-0.202*** (0.019)	
Limited liability firm [*DLTD* = 1]	-0.232*** (0.019)	-0.232*** (0.019)	-0.391*** (0.038)	-0.232*** (0.019)	-0.237*** (0.016)	
New main relationship [*DNEW* = 1]	-0.438*** (0.031)	-0.438*** (0.031)	-0.363*** (0.057)	-0.339*** (0.131)		
Lender-borrower dist. [*DIST*]		0.001 (0.002)				
New relationship [$DNEW'$ = 1]					-0.287*** (0.031)	-0.384*** (0.044)
Rho				-0.023 (0.030)		
Prob (Rho = 0)				0.0450		
Province fixed effects	YES	YES	YES	YES	YES	YES
Sector fixed effects	YES	YES	YES	YES	YES	YES
Firm fixed effects	NO	NO	NO	NO	NO	YES
Bank fixed effects	YES	YES	YES	YES	YES	YES
Constant	9.235*** (0.117)	9.231*** (0.119)	9.344*** (0.198)	9.308*** (0.113)	9.691*** (0.095)	6.494*** (0.060)
Observations	49331	49331	15560	739965	78898	43947
R-squared	0.41	0.41	0.31		0.34	0.68
Log Likelihood				-147212.71		

Table 6 - Difference between interest rates charged on new borrowers and rates on attached firms

The table reports OLS estimates for the regression described in the subsection 4.3. The dependent variable is the spread between the interest rate charged to a switcher and the rates charged by the same bank to all stayers that are *similar* to the switcher. *Similarity* is based on the matching variables. Spreads are regressed on a constant. Robust standard errors, clustered at switcher level, are reported in parentheses. The symbol * indicates the coefficient is significantly different from zero at the 10 percent level; ** at 5 percent; *** at 1 percent.

Matching variables	(1)	(2)	(3)	(4)	(5)	(6)
Lender	YES	YES	YES	YES	YES	YES
Province	YES	YES	YES	YES	YES	YES
Sector of economic activity	YES	YES	YES	YES	YES	YES
Firm size	YES	YES	YES	YES	YES	YES
Single bank	NO	YES	YES	YES	YES	YES
Limited liability	NO	NO	YES	YES	YES	YES
Share of term loans	NO	NO	NO	YES	NO	YES
Share of matched loans	NO	NO	NO	NO	YES	YES
Spread	-0.121**	-0.314***	-0.292***	-0.578***	-0.244***	-0.416***
	(0.049)	(0.054)	(0.058)	(0.080)	(0.096)	(0.118)
Observations	918,797	363,560	213,912	49,737	68,376	22,600
Number of switchers	5,481	5,343	5,183	3,929	4,120	2,597
Average number of matches	168	68	41	13	17	9

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