Adverse Selection in Credit Markets: Evidence from a Policy Experiment

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Abstract: We find evidence for adverse selection in credit markets: risker borrowers are willing to pay higher interest rates than safer borrowers are. The data are from an Indian financial institution where interest rates are determined by competitive bidding. The government imposed an interest rate ceiling in 1993. We examine changes in default patterns before and after this policy change. We find adverse selection despite the use of collateral as a screening device. We are able to isolate adverse selection from moral hazard and to distinguish private information on riskiness from publicly observed riskiness.

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1 Introduction

Many economists believe that asymmetric information impedes the efficient functioning of markets, particularly financial markets. Even though there is substantial and influential theoretical research assuming asymmetric information, the empirical evidence is rather thin (Chiappori and Salanie (8)). The importance of asymmetric information seems to depend crucially on the specific market studied. In this paper we find evidence suggesting that asymmetric information may indeed be an impediment to credit markets. We test a well-known prediction of adverse selection models (Akerlof (1), Stiglitz and Weiss' (20)): riskier borrowers have a higher willingness to pay than safer borrowers.

There are three main empirical challenges with such a study. First, researchers are themselves uninformed about whether differences in riskiness are truly private information or not. Suppose the lender observes the riskiness of borrowers but the researcher does not. Then observationally riskier borrowers will be willing to pay higher interest rates. But a researcher who is uninformed about the lender's information set may conclude that there is adverse selection (on unobserved riskiness). Secondly, lender often have mechanisms in place such as collateral to prevent adverse selection (Bester). So even if adverse selection in an underlying friction in financial markets, it is entirely possible that the lender is able to mitigate adverse selection through carefully designed contracts. So tests of adverse selection must be carefully distinguished from tests for whether uninformed lenders are able to overcome adverse selection. Finally, researchers only have information on defaults. It is entirely possible that higher interest rates make borrowers default more for moral hazard reasons instead of adverse selection. It is difficult empirically to distinguish between the two.

We address all three empirical challenges in this paper. We distinguish unobserved riskiness from observed riskiness, allow for the effects of collateral, and disentangle adverse selection from moral hazard. We find robust evidence of adverse selection – privately informed risky borrowers are willing to pay higher interest rates than privately informed safer borrowers are despite collateral requirements in place to deter them.

Our paper contributes to an empirical literature on information asymmetries (Chiappori and Salanie (8)). More narrowly, our results are closely related to tests of specific information asymmetries in credit markets (Ausubel (4), Karlan and Zinman (11)). We use a natural experiment involving a South Indian financial institution; Ausubel (4) and Karlan and Zinman (11) use randomized field experiments. We are able to isolate adverse selection from moral hazard, while Ausubel (4) is not. We find adverse selection in a sample that is predominantly male; Karlan and Zinman (11) find adverse selection among women but not among men. Both Ausubel (4) and Karlan and Zinman (11) use data from uncollateralized lenders. These lenders screen by classifying borrowers into risk categories and offering different terms to borrowers in different categories. So they cannot test whether collateral overcomes the adverse selection problem but we can.

The financial arrangements we study are Roscas (or Rotating Savings and Credit Associations). These are popular in many developing countries. In a Rosca, a group of people get together regularly, each contributes a fixed amount, and at each meeting one of the participants receives the collected pot. In random Roscas, the pot is awarded in each round by lottery.[1] In bidding Roscas, the subject of our study, the pot is awarded to the highest bidder.[2] Once a participant has received a pot he is ineligible to bid for another. Such a participant may default (stop making contributions) after receiving the pot. Unlike textbook financial markets, every participant in a bidding Rosca effectively receives different financial terms (loan sizes, durations and interest rates all differ).

Bidding Roscas induce participants to self select. In each round, the person willing to pay the highest interest rate wins the pot. Early recipients owe more contributions that late recipients. So selection is *adverse* if riskier (or dishonest) participants select earlier pots while safer (or honest) participants choose later pots. There would be no adverse selection if early recipients were inherently just as risky as later recipients.

We test for adverse selection by exploiting an exogenous policy shock. In September

[1] Anderson and Baland (2) and Besley, Coate and Loury (5) examine the rationale for random Roscas.

[2] The literature on bidding Roscas includes Calomiris and Rajaraman (7), Eeckhout and Munshi (9), Klonner (12), (13), (14) and Kovsted and Lyk-Jensen (16).

1993, the government unexpectedly imposed a bid ceiling (of 30% of the total pot). This effectively transformed the early rounds of bidding Roscas into random Roscas since many participants bid up to the ceiling but only one of them chosen by lottery received the pot in each round.

Riskier participants can bid high, take early pots and then default in the absence of a ceiling. So later recipients are always much safer than early recipients. We show theoretically how the bid ceiling induces a *flattening* of riskiness. The ceiling effectively ties the hands of riskier participants – they can no longer just outbid the others and be assured of receiving early pots. The lottery induced in early rounds can push riskier participants to later pots. At an extreme, if the bid ceiling were set at 0%, the Rosca would be transformed from pure bidding to completely random. The average riskiness of all recipients whould be the same after the policy shock (implying a completely flat risk profile).

We use a difference-in-difference specification to capture this flattening. The difference between the riskiness of early and late recipients will be smaller after the policy shock compared with before if there is adverse selection but will be unaffected if there is none. We do not observe borrower riskiness, however – we only observe default rates. We find that defaults are indeed flatter as a result of the policy shock exactly as predicted by adverse selection. But default rates could be flatter for reasons completely unrelated to adverse selection including (a) differences in publicly observed riskiness (b) moral hazard. (c) (c) changes in the composition of Rosca participants and (d) transitory aggregate shocks.

We address all the alternative explanations for the observed default patterns. We need to isolate the privately observed from the publicly observed component of riskiness to address concern (a) since we are interested in selection on the former and not on the latter. We do so by controlling for the collateral required of borrowers (since borrowers who are known to be riskier will be asked for more collateral than those who are known to be safer). Since loan terms become more favorable to early recipients relative to late as a result of the ceiling, moral hazard could cause the flattening of observed defaults (concern (b)). So we control for loan terms while taking the double difference in defaults isolate adverse selection from moral hazard. We address concern (c) by showing how our difference-in-difference test remains consistent even if changes in the average riskiness of the pool of participants (and not adverse selection) were causing the flattening of default rates We exploit the overlaps in the Roscas started before and after the policy shock to address concern (d). In all four cases we show how our estimator based on the conditional double difference in observed defaults consistently captures the double difference in unobserved riskiness.

Our main finding is that borrowers willing to pay higher interest rates are riskier than those who are not. We reject the null of no adverse selection. This result is remarkably robust to alternative specifications of the difference-in-difference test and to controls for observed riskiness, moral hazard and aggregate shocks.

The Roscas we study are large-scale and impersonal. A Rosca organizer (a for-profit financier) takes on the default risk and receives a commission in exchange. It is the Rosca organizer who imposes a collateral requirement (asks for cosigners with regular salaries) on the winners of the pot based on how risky he perceives them to be. Social ties between participants play no role in screening bad risks or in enforcing contributions in these Roscas. This is contrast with the personalized informal financial arrangements that are so popular in village economies. In those schemes, participants collectively take on the default risk and so social ties play an important role in screening and effective enforcement (Anderson et al (3) and Karlan (10)).

We calculate that the 1993 bid ceiling reduced the total default costs of the Rosca organizer from 1.7 to 1.4 percent. While the bid ceiling did have this benefit of reducing adverse selection, it also also shrank the Rosca organizer's business considerably. The organizer told us that the bid ceiling was unwelcome for this latter reason. Presumably if the gains outweighed the costs of imposing a bid ceiling, the Rosca organizer would have imposed it voluntarily in the absence of any government intervention. In contrast with equilibrium credit rationing that can arise endogenously (Stiglitz and Weiss' (20)), credit rationing was exogenously imposed here.

Finally we briefly clarify what we do *not* do in this paper. Since we do not observe the costs of collecting defaults by the Rosca organizer, we cannot estimate the welfare effects of the bid ceiling. Just as Ausubel (4) and Karlan and Zinman (11), we use the term adverse selection throughout this paper to refer to the positive relationship between willingness to pay and riskiness, not to the welfare loss relative to a full information environment. Further, we do not claim either the presence or absence of moral hazard. At first glance it might seem that the randomness introduced by the bid ceiling could also isolate moral hazard from other confounding effects. But as we discuss in detail in the conclusion, however, the structure of Roscas and data limitations makes isolating moral hazard in this way impossible.

Paper Outline

We proceed as follows. In section 2, we motivate our empirical strategy. In section 3, we provide additional background on bidding Roscas in South India, and on the policy shock. In section 4, we construct a simple model of bidding Roscas when riskiness is unobserved. In section 5, we discuss our identification strategy. In section 6, we discuss our results from the 1993 policy shock. We conclude in section 7.

2 Motivating Examples

In this Section we introduce the rules of bidding Roscas before and after the 1993 policy shock. We describe our empirical strategy for distinguishing adverse selection (on unobserved risk) from selection on observed risk and from moral hazard. We do so through some stylized examples. This section is intended to be an intuitive non-comprehensive introduction to several issues that will taken up more carefully later in the paper.

Unrestricted and Restricted Roscas

Bidding Roscas match borrowers and savers. Each month participants contribute a fixed amount to a pot. They then bid to receive the pot in an oral ascending bid auction where previous winners are not eligible to bid. The highest bidder receives the pot of money less the winning bid and the winning bid is distributed among all the members as a dividend. Consequently, higher winning bids mean higher interest payouts to later recipients of the pot. Over time, the winning bid typically falls as the duration for which the loan is taken diminishes.

To illustrate these rules, consider a 3 person Rosca which meets once a month and each participant contributes \$10. Suppose the winning bid is \$12 in the first month. Each participant receives a dividend of \$4. The recipient of the first pot effectively has a net gain of \$12 (i.e. the pot less the bid plus the dividend less the contribution). Suppose that in the second month (when there are 2 eligible bidders) the winning bid is \$6. And in the final month, there is only one eligible bidder and so the winning bid is 0.

We shall refer to the first recipient as the *early* borrower in what follows. The second recipient saves \$6 , then borrows \$16 and finally repays \$10. Since we focus on the risk of defaulting contributions owed subsequent to winning the pot in this paper, we will refer to the second recipient as the *late borrower* in what follows. The net gains and contributions in this Rosca with unrestricted bidding are:

Example U

Month	*1*	*2*	*3*
Winning bid	*12*	*6*	*0*
Early Borrower	*12*	*-8*	*-10*
Late Borrower	*-6*	*16*	*-10*
Saver	*-6*	*-8*	*20*

The early recipient receives \$12 and repays \$8 and \$10 in subsequent months (which implies a 43% monthly interest rate). The final recipient is a pure saver: she saves \$6 for 2 months and \$8 for a month and receives \$20 (which implies a 25% monthly rate). Notice that an unrestricted bidding Rosca allows participants to self select – a participant with a higher willingness to pay can outbid another.

Next we discuss an example where bidding is restricted by the policy change. The bid ceiling is 30% of the pot (or $9). In the first month, suppose that two participants are willing to pay more than $9. Since bidding stops at the ceiling, one of the two is chosen by lottery to receive the pot. The early recipient contributes $10, receives the pot less the bid, $21, and also receives a dividend of $3, which provides a net gain of $14 in the first month. Each of the other participants contributes $10 and receives a dividend of $3 in the first month. So the payoffs in this Rosca with restricted bidding are:

Example R

Month	*1*	*2*	*3*
Winning bid	*9*	*6*	*0*
Early Borrower	*14*	*-8*	*-10*
Late Borrower	*-7*	*16*	*-10*
Saver	*-7*	*-8*	*20*

The ceiling substantially lowers the interest rate on borrowing for the early borrower (and also the interest rate on savings for the saver). Further, the ceiling makes it difficult for participants to self select.

Empirical Approach

In example *U*, the early borrower has demonstrated a higher willingness to pay for the first pot than the late borrower has *for the first pot*. If participants differ in their inherent riskiness, then a riskier participant will be be willing to pay more for the first pot than a safer participant. This an implication of adverse selection analogous to Stiglitz and Weiss' (20). So riskier participants will all else equal choose early pots and safer participants will choose later pots. This is adverse for the lender as the early borrower has more contribute subsequent to winning than the late borrowers does (in both examples, the early borrower must contribute $18 and the later borrower must contribute $10 subsequent to winning). It is possible that self selection in Roscas is based only on other characteristics (e.g. impatience or productivity) but participants are all equally risky. In such a case, the early borrower would be no riskier than the late borrower; there would be no adverse selection.

Riskiness is unobserved however. We do have data on default rates, i.e. on whether contributions have been made subsequent to receipt of the pot. The empirical challenge is to infer whether there are differences in unobserved riskiness driving self selection in Roscas when we just have data on differences in observed defaults.

Consider first a naive empirical approach: simply taking the difference between individual default rates of early and late borrowers in example U. Adverse selection would imply that early borrowers have higher default rates than late borrowers. But even if early and late borrowers were inherently equally risky, there could be several other explanations for this difference in defaults. For instance, the early borrower has to make contributions at dates 2 and 3 while the late borrower has contributions due only at date 3. Assume that the first contribution after receiving a pot is easier to make than the second. So for this purely mechanical reason the early borrower may have a higher default rate than the late borrower. Taking the difference in defaults between early and late borrowers does not isolate adverse selection.

The restricted Rosca (example R) provides an opportunity to investigate default patterns when self selection is hampered. Consider another empirical approach: difference in differences. If there were adverse selection, riskier participants should win early pots before the ceiling but may not after the ceiling. For instance, the early borrower in example U has clearly demonstrated his willingness to pay more for the first pot than the late borrower has. So the early borrower will all else equal be riskier than the late borrower in example U. But in example R, it is conceivable that the early borrower has a lower willingness to pay for the first pot than the late borrower's willingness to pay for the first pot. In such a situation, the late borrower may be riskier than the early borrower. In other words, the bid ceiling may impede self selection on riskiness by changing the order of receipt of the pot. We would then expect to see a flattening in the default profile as we move from Example U to Example R. In other words, the difference between default rates of early and late borrowers should be higher in Example U than in Example R. But even if early and late borrowers were inherently equally risky, there could be several alternative explanations for a flattening of the default profile. For instance, the early borrower has more favorable loan terms in the restricted Rosca relative to the early borrower in the unrestricted Rosca. So moral hazard too would imply a flattening of default rates as a consequence of the bid ceiling.

Suppose that there were many three person Roscas all with a \$10 contributions both before and after the ceiling was imposed. The participants in these Roscas may differ not just in their riskiness but in other characteristics (such as impatience, productivity shocks or urgency of consumption needs). The winning bid for the early borrower will vary across unrestricted Roscas based on these differences. We shall exploit this variation and use conditional difference in differences to isolate adverse selection from moral hazard. To ilustrate, imagine another three person Rosca with unrestricted bidding (example U' say). Suppose the winning bid in round 1 just happens to be \$9 (which is 30% of the pot). There is no lottery at date 1 and so the participant with the higher willingness to pay receives the first pot for certain. Suppose that the winning bid in the second round is \$6. Payoffs are given by:

Example U′

Month	*1*	*2*	*3*
Winning bid	*9*	*6*	*0*
Early borrower	*14*	*-8*	*-10*
Late borrower	*-7*	*16*	*-10*
Saver	*-7*	*-8*	*20*

Suppose we compare Example U' with Example R and do find that the default rates of early and late borrowers are flatter in example R than in example U'. Crucially, since loan terms are the same for both borrowers in both examples the moral hazard explanation is no longer valid. On the other hand, adverse selection would imply that the early borrower is riskier than the second in Example U' but not necessarily in Example R. So a flattening of defaults is consistent with adverse selection. In this way, by conditioning on winning bids (comparing Example R with Example U' instead of with example U) we will isolate adverse selection from moral hazard. In section 5 we show that such a comparison will not by itself raise any endogeneity concerns.

So far we have ignored the collateral requirement. Borrowers may be asked to provide cosigners before the funds are released. This is effectively like providing collateral. Now it is possible that the early borrower in example R has a cosigner requirement while the early borrower in example U' does not. So even though loan sizes and repayment amounts are the same, the early borrower in example U' may be more likely to default if the cosigner requirement prevents moral hazard. To ensure that the flattening is driven by adverse selection and not moral hazard we would need to ensure that the cosigner requirements were the same for early and late borrowers in examples U' and example R. In other words, we will control for *all* the loan terms while taking the double difference in defaults.

Controlling for the collateral/cosigner requirement is important for another reason. A key empirical challenge is that the researcher is uninformed about what the organizer observes about the borrowers. The cosigner requirement and/or the number of cosigners required can be used as a summary measure of observed borrower riskiness. Clearly the organizer will all else equal impose a stricter cosigner requirement on borrowers who are known to be risker than on borrowers known to be safer. By conditioning the double difference on defaults on the cosigner requirement we are able to ensure that it is differences in *unobserved* riskiness that we are capturing.

The cosigner requirement may even be effective in deterring riskier borrowers from taking early pots (when riskiness is privately observed). . In other words, if the cosigner requirement is effective in preventing adverse selection we should see no flattening of defaults when we compare Roscas U' and R (where both have the same cosigner requirements). If we find a flattening of defaults (as we do) for the examples, that implies that adverse selection is a friction *and* the organizer has been unable to eliminate this problem through collateral.

Finally we flag two other identification concerns using these examples. Default differences between early and late borrower in example R could be smaller than the same difference in example U' for reasons unrelated to adverse selection. For instance, suppose there were no differences in inherent riskiness between participants but (a) there was an aggregate shock in date 2 of Example U' but no aggregate shock in Example R, or (b) the participants in Rosca R were simply safer than those who participated in Rosca U'. In both these situations the flattening of defaults cannot be attributed to adverse selection. We postpone our discussion of how our empirical approach deals with both these additional concerns till section 5.

3 Data and Institutional Context

In this section we provide some background information on the bidding Roscas we study. We describe the different Rosca denominations in our sample and the effects of the policy shock on bidding. We also discuss how contributions are enforced, how we calculate default rates, and how default rates were affected by the 1993 policy shock.

Denominations

In South India, Roscas originated in villages and small communities where participants were informed about each other and could enforce contributions (Klonner (13), Radhakrishnan (19)). The bidding Roscas in this study are larger scale and anonymous. Participants typically do not know each other and the Rosca organizer (a commercial company) takes on the risk of default. Bidding Roscas are a significant source of finance in South India (where they are called chit funds). Deposits in regulated chit funds were 12.5% of bank credit in Tamil Nadu and 25% of bank credit in Kerala in the 1990s, and have been growing rapidly (Eeckhout and Munshi (9)). There is also a substantial unregulated chit fund sector.

The data we use is from an established Rosca organizer with headquarters in Chennai. The organizer (a non-bank financial company) began organizing Roscas in 1973 and has been expanding gradually since then. The most common Rosca denomination offered in the early 1990s meets for 40 months (with 40 members) and has a contribution of Rs. 250. This Rosca has a pot of Rs. $10,000$. At that time, every year over one thousand Roscas were organized of this denomination alone (see Table 1). The organizer also offers other Rosca denominations, some with shorter durations (e.g. 25 months), others with longer durations (e.g. 50 months), and some with higher contributions and some with lower contributions. In this way, the organizer can match borrowers and lenders into Roscas that vary based on investment size and horizon. In what follows, we will refer to a Rosca of duration n (in months) and contribution m as (n, m). Since all Roscas administered by the organizer meet once per month, n also equals the number of members. The available pot is nm.

Policy Shock

In September 1993 the Supreme Court of India enforced the 1982 Chit Fund Act, which stipulated (among other regulations) a 30% ceiling on bids for every Rosca denomination. The stated purpose of the intervention was to prevent usurious interest rates. According to the Rosca organizer, there was considerable uncertainty over which parts of the 1982 Chit Fund would be enforced and when. Since the Supreme Court ruling was a surprise, it can reasonably be interpreted as an unanticipated policy shock (see also Eeckhout and Munshi (9)).

The ceiling effectively converted bidding Roscas into partial random Roscas. If several participants bid up to the ceiling only one of them received the pot by lottery.[3] This rule applied to all Roscas started after September 1993. Roscas that were started before September 1993 continued to operate without restrictions on bidding. According to Eeckhout and Munshi (9), interest rates fell from $14-24\%$ before the policy shock to $9-17\%$ after the shock. This is in accordance with the stated objective of the policy which was to prevent usurious interest rates in Roscas. The ceiling also created uncertainty about when

[3] We observed the lottery procedures during a recent field visit. Plastic tags for all lottery entrants are placed inside a large tin, then the tin is closed and shaken vigorously, and finally one plastic tag is drawn at random from the dark depths of the inside of the tin. The process is transparent and fair.

a participant would receive the pot.

For our core sample, we use Roscas that were started after September 1992 and before October 1994 (the two year period around the implementation of the bid ceiling). This has the advantage of taking care of potential seasonality of membership in Roscas as well as providing a sufficiently large number of observations. The latter is especially crucial for our analysis since default is a low probability event in our sample.

Our sample comprises those eleven denominations that were most popular around the time of the policy shock. More precisely, we include all denominations in which the Rosca organizer started at least 40 Roscas between October 1992 and September 1994. The numbers of Roscas of different denominations in the sample are set out in Table 1. According to Table 1, there was a substantial decrease in the number of Roscas formed for several of the Rosca denominations, including a drop of 37.5 percent for the popular $(40, 250)$ denomination. Table 1 shows that 72.7% of the winning bids in the first half of Roscas of this denominations hit the ceiling of Rs. 3000 after the policy shock. For sake of comparison, it is useful to see that the 79.1% of the winning bids in the first half of unrestricted Roscas (i.e. before the policy shock) were at or higher than Rs. 3000. There is a substantial fraction of early rounds for which the ceiling binds particularly in longer duration Roscas.

The length of time for which the ceiling binds in different Rosca denominations is summarized in Table 2. For the popular $(40, 250)$ denomination, the ceiling binds till round 17 in half the Roscas, till round 19 in a third of the Roscas and till round 20 in a quarter of the Roscas started after September 1993. In denominations of shorter duration, such as the $(25, 400)$ denomination, the ceiling binds for a smaller fraction of the life of the Rosca because the winning bid reflects an interest payment for a loan of shorter duration. In denominations of longer duration, by contrast, such as the $(50, 1000)$ denomination, the ceiling binds till the middle round for half the Roscas in the sample.

Enforcement

Early borrowers clearly have an incentive to drop out and stop making contributions. The organizer of the Roscas offers protection to participants against such defaults. If a borrower fails to make a contribution, the organizer will contribute the funds. The organizer receives two forms of payment. He acts as a special member of the Rosca who is entitled to the entire first pot (i.e. the first pot at a zero bid). He also receives a commission (usually $5-6$ percent) of the pot in each round.

Rosca participants do not put up any traditional collateral. Instead, the organizer relies on outside cosigners and the threat of legal action against both cosigners and borrowers to provide incentives for participants to continue making contributions even after they have received the pot. In related work, we show that the cosigner requirement is effective in reducing defaults (Klonner and Rai (15)). The organizer tries to discourage risky participants by requiring that all borrowers prove that they have a regular income before pots are awarded to them. The organizer also told us that screening and enforcement policies were the same across branches and these were not changed in response to the policy shock.

Individual Defaults

We calculate the individual default rate of a member of Rosca i who receives the pot in round $t \in \{1, ..., n\}$ as

$$y_{ti} = \frac{\text{amount not repaid by round } t \text{ borrower}}{\text{amount owed by round } t \text{ borrower}}. \tag{1}$$

So $y_{ti} = 0$ if the borrower has made all contributions, and $y_{ti} = 1$ if the borrower has made no contributions after winning the pot. Partial defaults are observed in the data, which means that y_{ti} is often less than 1.[4]

Since the Rosca organizer receives the first pot, and savers receive the last pot (but owe nothing to the Rosca organizer) we only include data on recipients in rounds $t \in \{2, ..., n-1\}$

[4] When a participant stops making contributions before receiving a pot, she is excluded from the group and replaced by another individual.

for our analysis. All pots allocated to institutional investors are also excluded from the empirical analysis because these investors never default. Between 10 and 25% of participants are institutional investors (depending on the Rosca denomination). Our sample then consists of the remaining non-institutional borrowers in Roscas that started after September 1992 but before September 1994. The legal process for collecting from cosigners can take up to five years after the end of the Rosca (Klonner and Rai (15)). So observed individual default rates decrease over time in the months after a Rosca has ended. We collected data on default rates for this sample in January 2002 about five years after the conclusion of Roscas of the popular (40, Rs 250) denomination started in the year after the policy change.

We are interested in how individual default rates were affected by the bid ceiling. The average individual default rate remained virtually unchanged from 1.3% before to 1.2% after the policy shock.[5] A dramatic change occurred in the timing of individual defaults, however (see Table 3). Individual default rates in early rounds dropped from 2.04 percent to 1.64 percent but rose in later rounds from 0.59 percent to 0.74 percent. The double difference of 0.55 is significantly different from zero. This flattening of defaults is illustrated in Figure 1 for the Roscas of a particular denomination (with 30 rounds and Rs. 500 monthly contribution).

For the Rosca organizer, profitability depends on the *total* default rate not the individual default rates. The total default rate is

$$x = \frac{\text{total amount not repaid by members}}{\text{total amount owed}}$$

The total default rate gives the percentage of funds lent out in all Roscas that the organizer failed to collect from the Rosca's participants. It thus reflects the risk for the organizer associated with lending in Roscas. For example, consider two situations with the same average individual default rates. In the first, the individual default rates are higher for early borrowers relative to late. In the second, the individual default rates are the same

[5] This is calculated by averaging individual default rates in columns $1-2$ and then $3-4$ in the first row of Table 3.

across early and late borrowers. The total default rate will be higher in the first situation than in the second.

The total default rate dropped from 1.71 percent before the shock to 1.41 percent after the shock. The Rosca organizer has an extraordinary collection record among these non-institutional customers: the average total default rate is less than 2 percent. To put this number into perspective, Ausubel (4) reports an average default rate of 4.5 percent in his study of credit card lending in the US.

Loan Terms

There are two components to the loan terms for a round t winner of a Rosca: winning bid b_t and repayment burden denoted $q_{t.}$ We discuss each in turn.

Winning bids were 35.07% of the pot on average for winners of early pots before the bid ceiling of 30% was imposed. Winning bids for both early and late rounds dropped as a result of the bid ceiling (in Table 3), but the drop was nearly six times as large for early winners. Clearly a Rosca winner benefits from a lower winning bid since he has to pay less for the pot. In other words, the policy shock improved the loan terms for both early and late borrowers but more so for early borrowers than for late borrowers.

We define the repayment burden q_t as the sum of future (net) contributions that a Rosca winner in round t must pay. Let $\mathbf{b}_t$ be vector of winning bids in all rounds after date t of the Rosca, $\mathbf{b}_t = (b_{t+1}, ..., b_{n-1})$. Notice that b_n always equals zero since there is no auction in the last round. That vector determines the repayment burden of the round t borrower. In a Rosca with an individual contribution of m dollars per round the net payment that borrower has to make equals $m - b_j/n$ in each round after receiving the pot, i.e. for $j = t+1, ..., n-1$. As a consequence, the (aggregate) repayment burden faced by the round t borrower equals

$$q_t \equiv (n-t)m + \frac{1}{n}\sum_{j=t+1}^{n-1} b_j \tag{2}$$

To give an example, the repayment burden for the round 10 winner in a 40 round Rosca is the sum of contributions less dividends in rounds 11 to 40. If there were no dividends paid from rounds 11 to 40, then the repayment burden would be 75 percent of the pot. The higher the dividends in subsequent rounds, the lower the repayment burden and hence the more favorable the loan terms. Table 3 shows that the average repayment burden across denominations for the winners of early pots (first half of Roscas) was 64.33 percent of the pot. The average of total contributions due in subsequent rounds (i.e. the term $(n-t)m$ in equation (2)) is 75 percent of the pot. So the difference of 10.47 percent is the average dividends in the last three-quarters of the Roscas.

4 Theory

In this section we describe a model of bidding Roscas in which borrowers differ in their riskiness (or dishonesty) and are prone to unpredictable needs for cash. The Rosca organizer, who effectively acts as the lender, can impose a collateral requirement on the winner of a pot, effectively the borrower. We analyze bidding behavior first in the absence of a ceiling (unrestricted bidding) and then in the presence of a bid ceiling (restricted bidding). We show under what conditions riskier borrowers will on average be willing to bid more for a pot than safer borrowers and hence obtain earlier pots on average. We refer to this as adverse selection in Roscas, which is completely analogous to adverse selection in credit markets. When there is selection on riskiness, the bid ceiling makes it harder for riskier borrowers to take the early pots and then stop contributing subsequently. We establish how the bid ceiling makes early borrowers safer and later borrowers riskier. This flattening of the riskiness profile forms the basis for our empirical tests for observed selection and adverse selection in section 5.

The Model

The model has three agents and three dates indexed by t. Agents have additively separable, risk-neutral intertemporal preferences with no discounting. Agents are endowed with a income stream of \$1 per period. Two of the agents are borrowers and one is a pure saver. Following Besley, Coate and Loury (5) we shall allow for "lumpiness" in the use of Rosca winnings. Borrowers have a fixed investment or purchase of size 3 at dates 1 and 2, while the saver does not. One of the borrowers is of type θ^H while the other is of type θ^L, where θ^H and θ^L refer to inherent riskiness. We shall index borrowers by i, where $i \in \{L, H\}$. We will assume that $\theta^H \geq \theta^L$.

Following Calomiris and Rajaraman (7) and Klonner (12), we shall assume that borrowers have unpredictable needs for cash that drives participation in Roscas. Borrowers draw idiosyncratic productivity/utility shocks at dates 1 and 2. A borrower receives an instantaneous payoff x_t^i from funding her cash needs where $i \in \{L, H\}$ and $t \in \{1, 2\}$. We assume that X_t^i is independently and identically distributed with

$$F(x) = 1 - e^{-\frac{x}{\mu}}$$

Throughout this section we will refer to random variables by upper case letters, and their realizations by lower case letters.

Agents form a bidding Rosca that meets at dates 1, 2 and 3. There is an open ascending bid auction at dates 1 and 2. Let the winning bid at each date be denoted b_t. There is no auction at date 3 because there is only one eligible bidder and so the winning bid $b_3 = 0$. The model has been chosen so that the saver always wins the date 3 pot and the borrowers win at dates 1 and 2. Each agent contributes \$1 at date 1 but borrowers may default on contributions in subsequent rounds. The date 1 borrower either pays the date 2 and date 3 contribution with probability $1 - \theta^i$ or pays neither contribution with probability θ^i. Similarly the date 2 borrower pays the third contribution with probablity $1 - \theta^i$ and defaults with probability θ^i at date 3. The Rosca organizer takes on the risk of default, and contributes \$1 if a participant does not.[6] We shall assume for tractability that the riskiness of a borrower θ^i and the realization of the shock x_t^i are observed by all Rosca participants. But riskiness θ^i and the shock x_t^i are not observed by the Rosca organizer.

[6] To keep the model simple, we abstract from any commission that the Rosca organizer is paid for performing this enforcement role. None of our results would be affected by including such a commission.

For the winner of the first pot, the Rosca organizer stipulates a collateral requirement, which causes an immediate loss in expected utility of π^i. As types are unobserved, the lender cannot adjust the collateral to the borrower's risk type. We assume that the Rosca organizer does not require a collateral for the winner of the second pot. This is purely for analytical convenience and all results go through if this assumption is relaxed. The high type's perceived collateral cost, π_t^H, may or may not exceed π_t^L. For instance, if the private cost of furnishing collateral is greater for the riskier borrower, we still have $\pi_t^H > \pi_t^L$. This is in fact the assumption in Bester (6) who shows how collateral can overcome the adverse selection problem in a credit market. In what follows, we will discuss implications of both scenarios, $\pi_t^H > \pi_t^L$ and $\pi_t^H = \pi_t^L$.

We shall interpret this model fairly broadly. Our interviews with bidding Rosca participants in South India suggest that funds are used both for investment needs of small businesses and for consumption needs in households. These cash needs are sometimes unpredictable. Under an entrepreneurial finance interpretation, the project size of 3 refers to a fixed investment size and the x_t^i refers to an immediate income from a productivity shock. Under a consumption finance interpretation, the project size of 3 refers to the size of the consumption expenditure and the instantaneous utility of x_t^i is associated with satisfying an urgent consumption need. In either interpretation, a participant fails to make his or her contribution after winning the pot with probability θ^H or θ^L. If $\theta^H > \theta^L$, we can think of either (a) the high type is less honest than the low type, (b) the high type is less vulnerable to the Rosca organizer's punishment than the low type is or (c) the high type is better at hiding his money than the low type holding fixed the Rosca organizer's enforcement capability. Finally, our results would go through however if production/consumption needs were anticipated, i.e. both x_1^i and x_2^i were realized at the beginning of date 1.

The winning bidder receives the pot of \$3 and, according to the rules, pays two thirds of the winning bid to the other losing bidders. Thus the winner receives $3 - \frac{2}{3}b_1$, where b_1 denotes the winning bid of the first auction. She needs additional finance of $\frac{2}{3}b_1$ to undertake the project. As in Kovsted and Lyk-Jensen (16), we assume that each member has access to costly external funds to finance this difference between the amount received from the Rosca upon winning the auction and the cost of the project. Each dollar borrowed from this source causes an instantaneous disutility of $c > 1$.[7]

A type i borrower who wins the first pot has expected payoff of

$$-1 + x_1^i - \frac{2}{3}b_1 c - \pi_1^i + (1-\theta^i)\left(-2 + \frac{1}{3}E[B_2^j]\right) \tag{3}$$

In the first period, the borrower contributes 1 and gains an instantaneous payoff of x_1^i less the cost of external finance $\frac{2}{3}b_1 c$ and the collateral cost π_1^i. The payoff from dates 2 and 3 zero if the borrower defaults, which occurs with probability θ^i, and -2, the contributions owed in periods 2 and 3, plus a share of one third in the expected period two winning, which is paid by the other borrower in the Rosca, $E[B_2^j]$. is $\theta^i(2) + (1-\theta^i)(\frac{1}{3}b_2 - 2)$. With probability θ^i the winner of the first pot keeps keeps her income in dates 2 and 3, while with probability $(1-\theta^i)$, she pays contributions and receives a date 2 share of the winning bid.

Of the two borrowers, the one who did not win the auction at date 1 will win the pot at date 2 at winning bid b_2. So a type i borrower's payoff from winning at date 2 is

$$-1 + \frac{b_1}{3} + \left(-1 + x_2^i - \frac{2}{3}b_2 c\right) + (1-\theta^i)(-1) \tag{4}$$

where $(-1 + \frac{b_1}{3})$ is the net contribution at date 1, $(-1 + x_1^i - \frac{2}{3}b_2 c)$ is the instantaneous income at date 2 and $(1-\theta^i)(-1)$ is the expected contribution at date 3.

Finally the saver wins the entire pot at date 3 with a lifetime payoff of

$$\left(-1 + \frac{b_1}{3}\right) + \left(-1 + \frac{b_2}{3}\right) + (-1+3)$$

So the saver saves $1 - \frac{b_1}{3}$ at date 1 and $1 - \frac{b_2}{3}$ at date 2, and receives 2 as a return at date 3. Higher winning bids then imply higher interest rates.

[7] This assumption is purely for analytical convenience; our results would be unchanged if borrowers had no access to outside funds and project size was variable.

Willingness to Pay

Let $g_1(x, \boldsymbol{\theta}^i, \pi^i)$ and $g_2(x, \theta^i)$ denote the maximum willingness to pay for the first and second pot, respectively, of a borrower of type i who observes shock x, where $\boldsymbol{\theta}^i = (\theta^i, \theta^j)$. The willingness to pay for the second pot of a borrower of type i is obtained from equalizing his period two expected payoffs from winning and losing the second auction,

$$\left(-1 + x_2^i - \frac{2}{3}b_2 c\right) + (1 - \theta^i)(-1) = \left(-1 + \frac{b_2}{3}\right) + (-1 + 3).$$

This gives

$$g_2(x_2^i, \theta^i) = \frac{3}{1 + 2c}\left(x_2^i - EX + \theta^i\right). \tag{5}$$

Since a high type retains his income stream with a higher probability than a low type does after winning the pot, he will have a higher willingness to pay all else equal.

As by assumption it is one of the two borrowers who wins the first auction, the bidders in the second auction are one borrower and the saver. Since each loser of the auction receives a one-third shares of the winning bid, the saver has an incentive to bid up the remaining borrower to her maximum willingness to pay. The winning bid in the second auction, b_2, thus equals $g_2(x_2^i, \theta^i)$.

To derive the willingness to pay for the first pot, we equalize a borrower's first period expected payoffs from winning and losing the first pot,

$$-1+x_1^i-\frac{2}{3}b_1c-\pi_1^i+(1-\theta^i)\left(-2 + \frac{1}{3}E[B_2^j]\right) = -1+\frac{b_1}{3}+\left(-1 + x_2^i - \frac{2}{3}E[B_2^i]c\right)+(1-\theta^i)(-1).$$

Substituting $E[g_2(X, \theta^i)]$ and $E[g_2(X, \theta^j)]$ for $E[B_2^i]$ and $E[B_2^j]$, we obtain

$$g_1(x_1^i, \boldsymbol{\theta}^i, \pi^i) = \frac{3}{1 + 2c}\left[x - EX + \left(1 + \frac{2c - \theta^j}{1 + 2c}\right)\theta^i + \frac{\theta^j}{1 + 2c} - \pi_1^i\right]. \tag{6}$$

As for the second period willingness to pay, g_1 is increasing in own riskiness, θ^i. It is, moreover, increasing in the other borrower's riskiness because an increase in θ^j implies a higher expected winning bid in the second period if j loses the first auction. Consequently, the expected payoff from winning the auction becomes larger for i.

The fundamental issue for adverse selection is whether the riskier of the two borrowers has a higher willingness to pay for the first pot. Notice that $g_1(x_1^i, \boldsymbol{\theta}^i, \pi^i)$ is increasing in θ^i, all else equal, exactly as in the model of Stiglitz and Weiss ((20)). In general, however, π^i may vary with θ^i. The following proposition thus characterizes under what conditions the riskier borrower does in fact have a higher willingness to pay for the first pot.

Proposition 1 ***(Collateral Costs and Willingness to Pay for the First Pot)***
(i) Conditional on the utility shock x, the riskier borrower has a higher willingness to pay for the first pot, i.e.

$$g_1(x, \boldsymbol{\theta}^H, \pi^H) > g_1(x, \boldsymbol{\theta}^L, \pi^L), \tag{7}$$

if, and only if,

$$\frac{4c}{1+2c}(\theta^H - \theta^L) > \pi_1^H - \pi_1^L. \tag{8}$$

(ii) Conditional on the utility shock x, both borrowers have the same willingness to pay for the first pot, i.e. (7) holds with equality, if, and only if, (8) holds with equality.

Proposition 1 makes precise the condition (8) under which adverse selection cannot be eliminated by the collateral requirement. In this context that the unobserved differences in borrower riskiness, i.e. $\theta^H > \theta^L$, is not sufficient for adverse selection. Instead this difference has to be sufficiently large relative to the difference in the costs of providing collateral. There may indeed be no adverse selection even though there are unobserved differences in riskiness provided collateral is a sufficient deterrent (part *ii* of proposition 1).

Since the losers of the auction receive a one-third share of the winning bid, they have an incentive to bid up the winners at dates 1 and 2 to their maximum willingness to pay. So given shock realizations x_1^L and x_1^H and public information among the bidders, the date 1 winning bid will be

$$b_1 = \max\left[g_1(x_1^L, \boldsymbol{\theta}^L, \pi^L), g_1(x_1^H, \boldsymbol{\theta}^H, \pi^H)\right] \tag{9}$$

Since X_1^L and X_1^H are independently and identically distributed, condition (8) determines whether the high type is more likely to win the first pot or not. Even if (8) holds though, there will be situations where a low type has a shock sufficiently larger than the high type's shock and so will win the first pot.

Next we turn to the effect of the bid ceiling $\bar{b}$. The willingness to pay for the second pot of a bidder of riskiness θ^i with utility shock x_2^i is the same with or without the ceiling, $g_2^c(x_2^i, \theta^i) = g_2(x_2^i, \theta^i)$. The superscript c will refer to "ceiling" throughout. Even though the willingness to pay for the second pot is unaffected by the ceiling, the winning bid may be constrained by it. The winning bid for a borrower of type θ^i with utility shock x_2^i in the second auction is

$$b_2^c = \min\left[g_2^c(x_2^i, \theta^i), \bar{b}\right] = \min\left[g_2(x_2^i, \theta^i), \bar{b}\right].$$

In the first round, the willingness to pay for the pot is reduced relative to unrestricted bidding because of a lower expected winning bid in the second auction. Solving backwards, we obtain

$$g_1^c(x_1^i, \boldsymbol{\theta}^i, \pi^i) = \frac{3}{1+2c}\left[x - EX + \left(1 + \frac{2c - \theta^j}{1+2c}\kappa(\bar{b})\right)\theta^i + \frac{\theta^j}{1+2c}\kappa(\bar{b}) - \pi_1^i\right], \qquad (10)$$

where

$$\kappa(\bar{b}) = 1 - \exp\left\{-\frac{\bar{b}(1+2c)}{3\mu}\right\}.$$

As the ceiling $\bar{b}$ approaches infinity, clearly $g_1^c(x_1^i, \boldsymbol{\theta}^i, \pi^i)$ approaches $g_1(x_1^i, \boldsymbol{\theta}^i, \pi^i)$.

In contrast to unrestricted Roscas in which the borrower with the higher willingness to pay wins the first pot, here the winner of the first may have a lower willingness to pay than the loser of the auction. The willingness to pay for the high and the low type in restricted Roscas is shown in Figure 2. If both are willing to pay more than the ceiling, $g_1^c(x_t^i, \boldsymbol{\theta}^i, \pi^i) \geq \bar{b}$ for all i, there is a lottery at date 1 and with probability one half, the winner is the bidder with a lower willingness to pay for the pot. If, on the other hand, at least one of the borrowers has a willingness to pay of less than $\bar{b}$, the outcome is qualitatively as in the absence of a ceiling. In contrast to equation (9), the winning bid in the first auction in restricted Roscas is given by

$$b_1^c = \min\{\max g_1^c(x_t^L, \boldsymbol{\theta}^L, \pi^L), g_1^c(x_t^H, \boldsymbol{\theta}^H, \pi^H)], \bar{b}\}$$

Testable Implications

Provided there is adverse selection (i.e. (8) holds), the bid ceiling makes it harder for riskier participants to get their hands on the early pot. Let us illustrate the argument with Figure 2. Let x^* be the utility shock that makes the low type's willingness to pay for the first pot exactly equal to the bid ceiling, i.e.

$$g_1^c(x_t^L, \boldsymbol{\theta}^L, \pi^L) = \bar{b}$$

Provided that $x_1^H \leq x^*$ and $x_1^L < x^*$, the high type still has a higher chance of winning the pot, but not otherwise. In other words, the difference in riskiness between high and low types plays less of a role in determining the winner in round 1 of restricted Roscas compared with unrestricted Roscas. As the bid ceiling $\bar{b}$ approaches 0 for instance, the restricted Rosca approaches a purely random Rosca, and the probabilities of winning for the two types converge to one half. So the expected riskiness falls as a result of the ceiling. Denoting by $E(\theta_1)$ and $E(\theta_1^c)$ the expected riskiness of the winner of the first pot without and with a ceiling in place, respectively, we obtain that $E(\theta_1) > E(\theta_1^c)$.

The bid ceiling also lowers the winning bid in the inital round. There is of course the obvious reason (the ceiling caps bids). But there is an additional, not-so-obvious, reason. In unrestricted Roscas, a borrower of either type does not gain any surplus from waiting (must pay his maximum willingness to pay at date 2). In restricted Roscas, though, a borrower of either type expects to earn a surplus (if the ceiling binds at date 2) and so the willingness to pay is lower at date 1 that for unrestricted Roscas. This is captured by the term $\kappa(\bar{b})$ in (10). Since borrowers of both types have a lower willingness to pay for date 1 pots after the ceiling than before,

$$g_1^c(x, \boldsymbol{\theta}^i, \pi^i) < g_1(x, \boldsymbol{\theta}^i, \pi^i) \text{ for } i = L, H.$$

which lowers the winning bid when the ceiling is in place.[8]

We collect our observations about the date 1 borrowers of the pot as:

[8] A borrower's willingness to pay in the first round of a restricted Rosca will depend on the probability that she will receive the pot in the second round. The ceiling is more likely to bind for a high type than

Proposition 2 ***(Early borrowers)*** *For early borrowers (date 1 auction winners):*
(i) The expected riskiness falls as a result of the ceiling $E(\theta_1 - \theta_1^c) > 0$ *if, and only if there is adverse selection, i.e. (8) holds,*
(ii) Expected winning bid falls as a result of the ceiling $E(B_1 - B_1^c) > 0$

The main implication of the theory is that the ceiling makes the average riskiness of participants more equal over time. This *flattening* of types – the riskier type is pushed to a later pot while the safer type is more likely receive the earlier pots after the ceiling – is what we will use to identify adverse selection.

Proposition 3 ***(Testable Implication)***
(i) If there is adverse selection, i.e. (8) holds, then early borrowers are riskier before the ceiling compared with after, and late borrowers are safer before the ceiling compared with after.

$$\begin{aligned} E(\theta_1 - \theta_1^c) &> 0 \\ E(\theta_2 - \theta_2^c) &< 0 \end{aligned}$$

So the difference in difference in expected riskiness is positive

$$E(\theta_1 - \theta_1^c) - E(\theta_2 - \theta_2^c) > 0 \tag{11}$$

(ii) If there is no adverse selection, i.e. (8) holds with equality, then there is no variation either early or late riskiness

$$E(\theta_1 - \theta_1^c) = E(\theta_2 - \theta_2^c) = 0$$

and so the difference in difference in expected riskiness is zero

$$E(\theta_1 - \theta_1^c) - E(\theta_2 - \theta_2^c) = 0 \tag{12}$$

for a low type in the second round and so this anticipated effect reduces the willingness to pay of the high type by more than it reduces the willingness to pay of the low type relative to before the ceiling. This is an additional reason why early recipients are less risky after the policy shock.

In other words, when there is adverse selection, restricted bidding flattens the risk profile. With restricted bidding, the winner of the first pot is on average less risky while the winner of the second pot is more risky than with unrestricted bidding as in Figure 3. When, on the other hand, there is no adverse selection, the risk profile is completely flat with unrestricted as well as with restricted bidding, i.e. the risk profile is unaffected by the policy shock as in Figure 4. So our empirical strategy (which we discuss in detail in section 5) will be to take the difference in the riskiness profile of borrowers before and after the ceiling is imposed. Our null hypothesis that willingness to pay is unrelated to riskiness will be (12). The alternative that willingness to pay is increasing in riskiness will be (11).

5 Identification

Our goal is to test for adverse selection, i.e. to test if riskier participants have a higher willingness to pay that safer participants in these Roscas. Adverse selection implies that the policy shock will have a flattening effect on the riskiness profile of Rosca borrowers (Proposition 3). As we have discussed earlier the policy shock was unanticipated. So we are not confronted with any selection bias arising from deliberate choices by prospective Rosca members about whether to join unrestricted Roscas that started before September 1993 or to wait to join restricted Roscas that started after that date.

We do not observe the riskiness of Rosca participants however; we only observe their default rates. In section 3, we have described how the policy shock did lead to flattening of defaults. But one Rosca participant may have a higher default rate than another for a variety of reasons that are completely unrelated to inherent riskiness. We shall explain our identification strategy by first assuming as a benchmark that defaults only depend on riskiness, and then adding on other potential determinants of defaults (loan terms, compositional changes, aggregate shocks). Our aim is to make precise how the flattening effect on riskiness implied by the theory can be captured by empirical specifications that test for flattening in default rates.

In all that follows in this section, we shall consider the special case of Roscas with three participants and three dates. Only the first two recipients are at risk of default in such Roscas. As before, the term early borrower will refer to the date 1 recipient and the term late borrower will refer to the date 2 recipient. This will then allow us to consider identification problems in light of the theory developed in section 4.

Benchmark

To test if riskier participants have a higher willingness to pay than safer participants, we will take the difference in differences in default rates between early and late borrowers before and after the policy shock. The benchmark econometric specification

$$y_{ti} = \alpha_t + \xi \ after_i + \beta \ after_i \ late_t + u_{ti}, \tag{13}$$

where t denotes the round of receipt of the pot, and i indexes Roscas of a particular denomination. The unit of observation y_{ti} is the individual default rate of the borrower in round t of Rosca i. The intercept term α_t is round specific. The dummy variable $after_i$ equals one if Rosca i started after the policy shock and zero otherwise. The dummy variable $late_t$ is an indicator for whether the borrower in round t was a late (as opposed to an early) borrower. The interaction term $after_i \ late_t$ interacts the indicator for before/after policy shock with the indicator for early/late receipt of the pot.

The least-squares estimate of β is the difference between (i) the difference in the average default rate of borrowers of early and late pots with unrestricted bidding and (ii) the difference in the average default rate of borrowers of early and late pots with restricted bidding. When all members are of equal riskiness, the riskiness of early and late borrowers are identical before and after the policy shock (as in equation 12). This implies that the difference in defaults between late and early borrowers is the same in Roscas with restricted and unrestricted bidding. Thus the null hypothesis of equal inherent riskiness (or no adverse selection) can be tested through the statistical hypothesis $\beta = 0$. The decision rule for our double difference test is: reject the null hypothesis in favor of adverse selection if the point estimate of β is positive and significant but do not reject otherwise. Loosely speaking we shall refer to β as the *adverse selection effect* of the policy shock.

The critical question for indentification is whether such a test is valid. In other words, does the double difference of observed default rates capture the double difference of unobserved riskiness? For expositional purposes, we shall first consider a hypothetical case in which it does. Assume that defaults are only generated by unobserved riskiness θ_i^u and measured with error. Assume the benchmark default generating process is:

$$y_{ti} = \delta_t + \lambda_t \theta_i^u + \upsilon_{it} \tag{14}$$

The parameter $\lambda_t > 0$ represents the effect of unobserved riskiness on defaults while the parameter δ_t is a round specific intercept.

In the theoretical model in section 4, $\delta_t = 0$ and $\lambda_t = 1$ provided (8) holds. If the data is indeed generated by (14) then β is a consistent estimator of the double difference in riskiness:

$$\lim_{N\to\infty} \widehat{\beta} = \lambda_1(E\theta_1 - E\theta_1^c) - \lambda_2(E\theta_2 - E\theta_2^c) \tag{15}$$

where $E\theta_1 - E\theta_1^c$ is the expected difference in riskiness for early rounds and $E\theta_2 - E\theta_2^c$ is the expected difference in riskiness for late rounds. Recall that, according to Proposition 3, flattening through a ceiling on bids decreases the average riskiness of early borrowers, $E\theta_1 > E\theta_1^c$, and increases the average riskiness of late borrowers $E\theta_2 < E\theta_2^c$. So under the null, $\lim_{N\to\infty} \widehat{\beta} = 0$ but under the alternative $\lim_{N\to\infty} \widehat{\beta} > 0$. For this admitted unrealistic benchmark case, the test for adverse selection is consistent.

What if the cosigner penalty is enough to deterred unobserved riskier people from taking early pots? Emprically if condition (8) held at equality, then there would be no adverse selection – the null hypothesis. But there could be no adverse selection either because markets (or collateral) work efficiently to overcome the adverse selection or because borrowers do not differ in their unobserved riskiness.

In the data, defaults may be affected by observed riskiness, loan terms, composition of risks and by aggregate shocks in addition to unobserved riskiness. In such departures from (14), the estimated double difference of defaults in specification (13) may capture more or less than the double difference in unobserved riskiness – and so a test that rejects the null hypothesis of no adverse selection on the basis of specification (13) may be inconsistent. In what follows we discuss how we augment the specification (13) so that we can identify adverse selection when the defaults are not generated by (14).

Observed Riskiness

In practice defaults are certainly not generated only by unobserved riskiness as we assumed in (14). In this subsection we shall discuss how to alter our estimation strategy when defaults are generated by both unobserved and observed riskiness. We shall pose the problem – the unconditional double difference of defaults in specification (13) picks up both unobserved/observed riskiness — and then discuss a solution.

Assume that riskiness comprises both observed and unobserved components, $\theta_i = \theta_i^u + \theta_i^o$. Suppose defaults only depend only on riskiness

$$y_{ti} = \delta_t^u + \lambda_t^u \theta_i + \upsilon_{it} \tag{16}$$

then if we run the benchmark econometric specification (13), there will be an omitted variable bias. We will pick up both the double difference in observed and the double difference in unobserved riskiness

$$\begin{aligned} \lim_{N\to\infty} \widehat{\beta} &= \lambda_1^u(E\theta_1 - E\theta_1^c) - \lambda_2^u(E\theta_2 - E\theta_2^c) \qquad (17) \\ \lim_{N\to\infty} \widehat{\beta} &= [\lambda_1^u(E\theta_1^u - E\theta_1^{uc}) - \lambda_2^u(E\theta_2^u - E\theta_2^{uc})] + [\lambda_1^u(E\theta_1^o - E\theta_1^{oc}) - \lambda_2^u(E\theta_2^o - E\theta_2^{oc})] \end{aligned}$$

where it is just the first term that we are interested in. We would like to isolate the double difference in unobserved riskiness.

We need an appropriate measure of the Rosca organizer's information on borrower riskiness as a control in specification (13). As we discussed in section 4, the organizer imposes a collateral requirement on borrowers after they have won the pot. In practice, this collateral requirement is an attempt to prevent defaults so will depend on all observed factors that can help the organizer prevent defaults. We next show that the double difference of defaults conditional on collateral requirement can isolate unobserved riskiness.

Suppose first that the only such observed factor that determined whether collateral is required was riskiness:

$$c_{ti} = \delta_t^o + \lambda_t^o \theta_i \tag{18}$$

Substituting (18) into (16) and solving gives a default generating process that depends on unobserved riskiness and the collateral requirement (eliminating the observed riskiness term)

$$y_{ti} = \delta_t^{u'} + \lambda_t^u \theta_i^u + \lambda_t^{u'} c_{ti} + \upsilon_{it}$$

where

$$\delta_t^{u'} = \delta_t^u - \frac{\lambda_t^u}{\lambda_t^o}\delta_t^o \text{ and } \lambda_t^{u'} = \frac{\lambda_t^u}{\lambda_t^o}$$

So the following regression will capture the flattening in unobserved riskiness

$$y_{ti} = \alpha_t + \xi \ after_i + \beta \ after_i \ late_t + \eta_t \ c_{ti} + u_{ti}, \tag{19}$$

Here β is the double difference in defaults conditional on the cosigner requirement

The collateral requirement is based on observed factors in uncorrelated with observed risk, i.e. c_{ti} and θ_t^u are uncorrelated. And so, in contrast with (17), we have isolated the effect of unobserved riskiness in increasing the bids.

$$\lim_{N\to\infty} \widehat{\beta} = \lambda_1^u(E\theta_1^u - E\theta_1^{uc}) - \lambda_2^u(E\theta_2^u - E\theta_2^{uc})$$

Clearly isolating the effects of unobserved from observed riskiness by controlling for the collateral requirement depends on the very simple default generating process that we have assumed in (16).

Moral Hazard

Observed defaults may not only be a function of a borrower's inherent riskiness, but also of the terms at which the loan is obtained. The winning bid and the repayment burden (defined at the end of section 3) are two components of the loan terms. A third component is the collateral requirement. Recall that.the flattening of the winning bid as a consequence of the policy shock improved the loan terms for early borrowers more than it did for late borrowers. For moral hazard reasons, then one would expect a flattening of default rates after the policy shock. Further for completely mechanical reasons as well, favorable loan terms can lead to reductions in defaults and confound tests of asymmetric information (Karlan and Zinman (11)). We cannot separate out the moral hazard and mechanical effects of the policy shock, but we can try to isolate adverse selection from both these effects. In this section we shall explain in detail how we augment our empirical specification to allow for changes in winning bids, repayment burdens and the collateral requirement.

Suppose defaults depended on inherent riskiness and the terms at which a participant obtained the pot:

$$y_{ti} = \phi_{t0} + \lambda_t^u \theta^i + \phi_{t1} b_{ti} + \phi_{t2} q_{ti} + \phi_{t3} c_{ti} + \upsilon_{it} \tag{20}$$

then a test that rejects the null hypothesis of no adverse selection on the basis of specification (19) is inconsistent because of the omitted variables (the loan terms). In such a test, the estimate $\widehat{\beta}$ would capture the double difference in riskiness *and* the change of loan terms on default.

The cosigner requirement in turn may depend on the loan terms (winning bid and repayment burden) in addition to depending on observed riskiness. So we can augment (18) to

$$c_{ti} = g_{t0} + \lambda_t^o \theta_i^o + g_{t1} b_{ti} + g_{t2} q_{ti} \tag{21}$$

We can rewrite the default generating process for y_{ti} so that it only depends on unobserved riskiness θ_i^u and not on observed riskiness. This involves substituting (21) into (20) and solving to give

$$y_{ti} = \phi'_{t0} + \lambda_t^u \theta^i + \phi'_{t1} b_{ti} + \phi'_{t2} q_{ti} + \phi'_{t3} c_{ti} + \upsilon_{it} \tag{22}$$

where

$$\begin{aligned} \phi'_{tj} &= \phi_{tj} - \frac{\lambda_t^u}{\lambda_t^o} g_{tj} \text{ for } j = 0, 1, 2 \\ \phi'_{t3} &= \phi_{t3} + \frac{\lambda_t^u}{\lambda_t^o} \end{aligned}$$

Based on (22), we can augment specification (19) to include loan terms: winning bid and repayment burden. We run the following regression where β now denotes the double difference in defaults conditional on loan terms

$$y_{ti} = \alpha_t + \xi \; after_i + \beta \; after_i \; late_t + \eta_t \; c_{ti} + \psi_t b_{ti} + \zeta_t q_{ti} + u_{ti}, \tag{23}$$

Finally show that the $\lim_{N\to\infty} \widehat{\beta} = 0$ under the null of no differences in unobserved riskiness. In other words, controlling for the loan terms leads to a consistent test for adverse selection. Note that $\lim_{N\to\infty} \widehat{\beta}$ equals

$$[1\;0\;0\;0] \left\{ \begin{array}{c} \lambda_1^u \left(var \begin{bmatrix} After \\ B_1 \\ Q_1 \\ C_1 \end{bmatrix} \right)^{-1} cov \left(\begin{bmatrix} After \\ B_1 \\ Q_1 \\ C_1 \end{bmatrix}, \widetilde{\theta}_1^u \right) \\ -\lambda_2^u \left(var \begin{bmatrix} After \\ B_2 \\ Q_2 \\ C_2 \end{bmatrix} \right)^{-2} cov \left(\begin{bmatrix} After \\ B_2 \\ Q_2 \\ C_2 \end{bmatrix}, \widetilde{\theta}_2^u \right) \end{array} \right\}, \tag{24}$$

where $\widetilde{\theta}_t^u$ denotes the unobserved riskiness of a round t as a random variable. Under the null hypothesis, $\widetilde{\theta}_t^u$ has no variation, as all borrowers are of identical riskiness. Hence the two covariance terms in (24) are vectors of zeros, and hence $\lim_{N\to\infty} \widehat{\beta} = 0$.

Rosca Composition

In our data, individuals are not randomly assigned into unrestricted and restricted groups. Instead, there were only unrestricted Roscas before September 30, 1993, and (mostly) restricted ones after that date. Ideally for the researcher, an identical set of individuals signed up Roscas of a particular denomination before and after September 1993. But there are plausible reasons why the characteristics of Rosca participants may be different before and after the ceiling for Roscas of the same denomination. First, individuals may choose to join a different denomination when confronted with restricted instead of unrestricted bidding. Second, an individual who chooses to sign up for a certain Rosca denomination when bidding is unrestricted may choose not to join a Rosca and seek other forms of finance instead when bidding is restricted. This latter argument could, at least in principle, also work conversely: an individual for whom other sources of finance dominate a Rosca membership with unrestricted bidding may decide to join a Rosca when bidding is restricted.

In this section, we do not take a theoretical stand on how participants of different riskiness may sort themselves across Rosca denominations as a result of the policy shock.[9] Nor do we try to theoretically predict whether safer or riskier borrowers would choose to join or drop out of the overall pool of Rosca participants. Instead, we analyze whether our empirical test for adverse selection remains consistent under alternative assumptions about non-random assignment into Roscas before and after the policy shock.In particular, suppose that the null hypothesis were true. There are no differences in inherent riskiness between the participants both before and after the policy shock (and therefore, no adverse selection). The pool of participants after the shock may be riskier or safer than those before, owever.

In terms of our 3 period model, the two borrowers have riskiness $\theta^H = \theta^L = \theta^1$ before, and $\theta^H = \theta^L = \theta^2$ after the policy shock, where $\theta^1 \neq \theta^2$. The question is whether such a shift in riskiness could lead to a nonzero estimate of β. To give an answer, we consider the probability limit of $\widehat{\beta}$ as given by expression (24), which applies if (as before) defaults are generated by (22).

Proposition 4 ***(Change in Rosca Composition)*** *If all borrowers in unrestricted Roscas are of unobserved riskiness θ^1 and all borrowers in restricted Roscas of unobserved riskiness θ^2,*

(i) $\lim_{N\to\infty} \widehat{\beta} = (\lambda_1^u - \lambda_2^u)(\theta^1 - \theta^2)$;

(ii) $\left(\lim_{N\to\infty} \widehat{\xi}\right)\left(\lim_{N\to\infty}(\widehat{\xi} + \widehat{\beta})\right) = \lambda_1^u \lambda_2^u (\theta^1 - \theta^2)^2 > 0.$

[9] Eeckhout and Munshi (9) find that participants sort themselves across Rosca denominations before and after the policy shock in predictable ways. Since participants have no default risk both in their theory and empirics, however, their results are not directly applicable to our study.

Proof:

(i) For ξ in estimating equation (23), we have that

$$\lim_{N\to\infty}\widehat{\xi}=\lambda_1 e'\left(var\begin{bmatrix}After\\X_1\end{bmatrix}\right)^{-1}cov\left(\begin{bmatrix}After\\X_1\end{bmatrix},\widetilde{\theta}_1^u\right), \tag{25}$$

where

$$e=\begin{bmatrix}1\\0\\0\\0\end{bmatrix} X_t=\begin{bmatrix}B_t\\Q_t\\C_t\end{bmatrix}$$

and the prime superscript denotes the transpose operator.

First notice that

$$var\begin{bmatrix}After\\X_1\end{bmatrix}=\begin{bmatrix}var(After) & cov(X_1,After)'\\cov(X_1,After) & var(X_1)\end{bmatrix}. \tag{26}$$

Second, when $\widetilde{\theta}_1^u$ equals the constant θ^1 for all pre policy shift Roscas and θ^2 for all post policy shift Roscas, we have that

$$\begin{aligned}cov\left(\begin{bmatrix}After\\X_1\end{bmatrix},\widetilde{\theta}_1^u\right) &= (\theta^2-\theta^1)\begin{bmatrix}var(After)\\cov(X_1,After)\end{bmatrix} \\ &= (\theta^2-\theta^1)\begin{bmatrix}var(After) & cov(X_1,After)'\\cov(X_1,After) & var(X_1)\end{bmatrix}e.\end{aligned} \tag{27}$$

Substituting identities (26) and (27) into (25) immediately establishes

$$\lim_{N\to\infty}\widehat{\xi}=\lambda_1^u(\theta^2-\theta^1). \tag{28}$$

For $\beta+\xi$ in estimating equation (23), we have that

$$\lim_{N\to\infty}(\widehat{\beta}+\widehat{\xi})=\lambda_2^u e'\left(var\begin{bmatrix}After\\X_2\end{bmatrix}\right)^{-2}cov\left(\begin{bmatrix}After\\X_2\end{bmatrix},\widetilde{\theta}_2^u\right),$$

and an analogous argument establishes that

$$\lim_{N\to\infty}(\widehat{\beta}+\widehat{\xi}) = \lambda_2^u(\theta^2-\theta^1). \tag{29}$$

Subtracting (28) from (29) immediately establishes part 1 of the proposition. Multiplication of (28) and (29), and the fact that λ_1^u and λ_2^u are both positive by definition establishes part 2 of the proposition.■

Part 1 of the proposition states that our test for adverse selection may in general be biased when there is in fact no adverse selection and only a change in the average riskiness of borrowers. Whether $\widehat{\beta}$'s limit is different from zero depends crucially on how riskiness translates into default rates in different rounds of a Rosca. If $\lambda_1^u = \lambda_2^u$, our test will be consistent. This is in fact the case in our theoretical model, where $\lambda_1^u = \lambda_2^u = 1$.

Part 2 of the proposition states a property of the coefficients ξ and β in estimating equation (23) when there is no adverse selection and only a change in average riskiness. Notice that ξ captures the change in defaults of early borrowers and $\beta+\xi$ the change in defaults of late borrowers. With a change in average riskiness but no adverse selection, the conditional default profiles before and after the policy change do not intersect and hence early and late defaults move in the same direction. With adverse selection and no change in average riskiness, on the other hand, the default profiles will intersect. So, provided we find that $\beta > 0$, a test that it is solely a change in composition (and not adverse selection) that is causing $\beta > 0$ is $sign(\xi) = sign(\xi+\beta)$ or $\xi(\xi+\beta) > 0$. If the signs are equal then there is no intersection of default profiles while if the signs are different, then the null hypothesis of no adverse selection can be confidently rejected.

Aggregate Shocks

We turn finally to another explanation for why the double difference in observed defaults may not correspond with the double difference in unobserved riskiness. Rosca participants may experience transitory aggregate shocks unobserved by the researcher. For instance, there may be no differences in inherent riskiness, yet it is changes in aggregate shocks that drive the flattening of defaults. Just as in previous subsections, there is an omitted variable bias in specification (23) because aggregate shocks are ignored.

We control for such aggregate shocks by exploiting the fact that unrestricted and restricted groups in our sample overlap. To account for transitory aggregate shocks, we augment (23) by quarterly time dummies. Indexing quarters by j, we introduce the dummy $quarterj_{ti}$, which equals one for all observations in which round t of rosca i is in quarter j. We rewrite (23) as

$$y_{ti} = \alpha_t + \xi \ after_i + \beta \ after_i \ late_t + \eta_t \ c_{ti} + \psi_t b_{ti} + \zeta_t q_{ti} + \gamma_j \ quarterj_{ti} + u_{ti} \qquad (30)$$

As a consequence, identification of adverse selection β is solely based on differences in defaults of contemporaneous borrowers of pots.

To illustrate, consider the Rosca denomination $(40, 250)$ and Roscas that started between October 1992 and September 1994. The overlap between unrestricted and restricted Roscas in such a sample ranges from 39 months for two Roscas that started in September and October 1993, respectively, and 17 months for two Roscas that started in October 1992 and September 1994, respectively. The period covered by such a sample is 63 months, from October 1992 to December 1997. With a partition of that time period into quarters, we arrive at 20 dummy variables (starting with the first quarter of 1993 and ending with the last quarter of 1997). It is these quarterly dummies that we use in the regression specification (30).

Notice that we are assuming that transitory aggregate shocks affect defaults identically in different rounds of the Rosca. If the aggregate shocks affected say the defaults of round 5 winners differently from defaults of round 35 winners, then interacting quarterly dummies with rounds could pick up some of the variation and allow for a flattening of the default profile. But this would substantially reduce the degrees of freedom.

6 Results from the 1993 Policy Shock

In what follows we shall first discuss how to partition Roscas into early and late rounds when Rosca denominations vary and length of time for which the bid ceiling binds depends on the denomination. We then describe our main results testing for adverse selection and examine their their robustness. Finally we decompose total default costs of the organizer into adverse selection and moral hazard-mechanical effects.

Early vs. Late

At the end of section 5, we showed that the specification (30) ensures identification. Here we adapt specification (30) in two ways. First, we allow for 11 denominations in our sample indexed by k. Second, we allow for different partitions of early and late rounds by indexing the late indicator by τ.

$$y_{kti} = \alpha_{kt} + \xi_k \, after_i + \beta \, after_i \, late^{\tau}_{kt} + \eta_t \, c_{ti} + \gamma_j \, quarterj_{kti} + \psi_{kt} b_{kti} + \zeta_{kt} q_{kti} + u_{kti}. \quad (31)$$

Notice that, except for the aggregate shock terms, we incorporate denomination-specific and round-specific controls for bid b_{kti} and repayment burden Q_{kti} throughout.[10] This implies then that we could use either absolute bids or relative bids (bids as a fraction of the pot) as independent variables. For the same reason, we could use either absolute repayment burden or relative repayment burdens as independent variables.

Our intention is to partition Roscas into early and late depending on how long the ceiling binds. Here $\tau \in (0,1)$ determines a "cutoff round" that separates early from late borrowers, where, qualitatively, early borrowers correspond to the borrower of the first, and late borrowers to the borrower of the second pot in our theoretical model. To be precise,

[10] In principle, could estimate (31) by OLS and/or by a Tobit specification (since the default rate is censored at 0 and 1). For the present application the results from an OLS specification are of primary interest since they are based on changes in the conditional expected value of actual defaults as opposed to changes in the unobserved default propensity variable. More practically, given that (31) implies a total of $1,251$ right hand side variables and that $121,943$ observations are used, it was computationally infeasible to maximize the associated likelihood function.

$late_t^\tau$ equals one if $t > n\tau$. For identification, τ has to be chosen such that flattening of the risk profile occurs on or around round $n\tau$. It is sufficient for identification that a lottery occurs in round $n\tau$ in at least some of the Roscas with restricted bidding. Here we discuss alternative specifications of this cutoff τ. The concern is flattening may depend on the extent to which the bid ceiling binds in each denomination. For the sake of robustness, we use three alternative specifications, which address the concern:

Rounds in Later Two Thirds We first specify the cutoff $\tau = \frac{1}{3}$ for all denominations. So $late_{kt}^{1/3}$ is a dummy equal to one for all rounds in the second half of a Rosca, which may be denoted by $late_{kt}^{1/3}$ where $late_{kt}^{1/3} = 0$ for $t = 2, ..., T_k/2$ and $late_{kt}^{1/2} = 1$ for $t = T_k/2 + 1, ..., T_k - 1$.

Rounds in Second Half We next specify the cutoff $\tau = \frac{1}{2}$ for all denominations. So $late_{kt}^{1/2}$ is a dummy equal to one for all rounds in the second half of a Rosca, which may be denoted by $late_{kt}^{1/2}$ where $late_{kt}^{1/2} = 0$ for $t = 2, ..., T_k/2$ and $late_{kt}^{1/2} = 1$ for $t = T_k/2 + 1, ..., T_k - 1$.

Relative Round Finally, we take the statement "flattening of the default profile" literally and specify the adverse selection effect as

$$\beta \ after_{ki} \frac{t}{T_k}$$

where $\frac{t}{T_k}$ is the "relative round". In (31) the estimates α_{kt} denote the trend of the default profile pre-ceiling and the estimated $\alpha_{kt} + \beta \ after_{ki} \frac{t}{T_k}$ denote the trend of the default profile post ceiling. So a flatter (and thus less downward sloping) default profile post ceiling results in a positive β. If $\beta = 0$, then the trend post-ceiling coincides with pre-ceiling.

Main Results

Recall that the double difference in mean default rates reproduced in Table 3 indicates a flattening of default rates. As we noted in section 5, however, this double difference does not ensure identification of adverse selection, however. Instead, we have argued that identification is assured with specification (31) that conditions the double difference in defaults on aggregate shocks, winning bids and repayment burdens.

The OLS results for the double difference in defaults (estimate of β) with with and without controls are in Table 4. The specifications in column 1 refers to the benchmark without controls with $\tau = \frac{1}{3}$:

$$y_{kti} = \alpha_{kt} + \xi_k \; after_i + \beta \; after_i \; late_{kt}^{1/3} + u_{kti}.$$

while column 2 includes all the controls as in

$$y_{kti} = \alpha_{kt} + \xi_k \; after_i + \beta \; after_i \; late_{kt}^{1/3} + \eta_t \; c_{ti} + \gamma_j \; quarterj_{kti} + \psi_{kt} b_{kti} + \zeta_{kt} q_{kti} + u_{kti}$$

Columns 3 and 4 correspond to the Rounds in Second Half specifications (with $\tau = \frac{1}{2}$) without and with controls. And columns 5 and 6 correspond to the Rounds Without Ceiling specifications, again first without and then with controls.

Our main result is that the double difference $\widehat{\beta}$ is is positive and (mostly) significant for all three specifications of the adverse selection effect and is remarkably robust to the introduction of controls for aggregate shocks, winning bids and repayment burden. For each of the three early-late partitions, the standard errors for the estimated $\widehat{\beta}$ change when the controls are added but the estimates are qualitatively unchanged. This establishes then that riskier participants are willing to bid higher than safer participants are despite the cosigner requirement. We reject the null of no adverse selection. Recall that we have shown in section 5 that our test for rejecting the null of no adverse selection is consistent when all the controls are included.

We use two alternative measures for the collateral requirement – the incidence of cosigner (whether cosigners were required or not) and the number of cosigners (which ranges from 0 to 5). We also control for aggregate shocks $quarterj_{kti}$ and loan terms, b_{kti} and q_{kti} All the controls are significant coefficient estimates, except for the number of cosigners (for which the F test fails to be rejected in all specifications). Only the winning bid has a significant average effect. This positive average effect implies that on average a higher winning bid results in significantly higher defaults. The zero average effect for cosigner incidence suggests that the lender manages to fully compensate the effect of higher observed riskiness through the cosigner requirement on average, otherwise we would have a positive average effect. The cosigner incidence, repayment burden and aggregate shock controls must have has different signs in different rounds and denominations (since we reject on the F-test, but average effect not significantly different from zero).

Finally the reader might also be interested in whether there is a positive relationship between willingness to pay and observed riskiness. It is possible even likely that observationally riskier participants may be deterred from taking early pots because of the cosigners required if they did. If there was no selection on observed risk, then the observed riskiness patterns should be unaffected by the ceiling. We run the following regression, where the dependent variable is cosigner incidence (our measure of observed riskiness)

$$c_{kti} = \alpha_{kt} + \xi_k \ after_i + \beta \ after_i \ late^{\tau}_{kt} + \gamma_j \ quarterj_{kti} + \psi_{kt} b_{kti} + \zeta_{kt} q_{kti} + u_{kti}.$$

The cosigner incidence should be no flatter after the ceiling compared with before if cosigner requirement is sufficient in deterring known risks. This is exactly what we find in Table 5 for alternative partitions of early-late. There is no flattening of observed riskiness. While winning bids are significant predictors for cosigners, and the average effect is positive, it is not statistically significant.

7 Conclusion

We have used a natural experiment to test for adverse selection. This experiment involved imposing a bid ceiling on bidding Roscas in 1993 which effectively made the early rounds more like random Roscas. The experiment did not substantially change overall default rates. But the difference between early and late defaulters changed substantially. This is what we use for identification of asymmetric information. We also use the small sample variation in auction outcomes to control for moral hazard (and other effects of different loan terms). We are thus able to identify a significant adverse selection effect. Further, we find the opposite adverse selection effect when the bid ceiling was partially lifted in 2002.

In this paper we do not explicitly test for moral hazard. A colleague has suggested to us, however, that the randomness introduced by the 1993 bid ceiling could be used to disentangle moral hazard from adverse selection. Consider a a three period Rosca where the ceiling binds at date 1. The lottery entrants at date 1 all selected in willing to pay the same interest rate but by random chance one received the pot at date 1 and another participant received the pot at date 2. Comparing the default rates of the date 1 winner with the date 2 winner could potentially isolate moral hazard. There are at least two difficulties with this approach. First, we only have data on the eventual lottery winner, not on all the lottery entrants. So we do not observe if the date 2 winner of a pot did actually take part in the date 1 lottery or not. Secondly, even in the absence of moral hazard, the round 1 winner of a lottery may have a higher default rate than a round 2 winner of the lottery for purely mechanical reasons (such as the size effect and the duration effect that we discuss in the paper). We leave the investigation of moral hazard to future research.

Appendix

To improve accessability of the subsequent proofs, we first derive each bidder's willingness to pay in each of the two auctions. We start with the second auction. For a bidder of type θ^i the willingness to pay for the second pot, p_2^i, is obtained by equating her expected payoff from winning the second auction, which is given in (4), with her payoff from losing the second auction. Notice that, since the saver always obtains the third pot, winning the first pot and losing the second pot are equivalent. Thus her payoff from losing the second auction is given by (3). Equating (3) and (4) and solving for b_2 gives

$$p_2^i = \frac{3}{2c+1}\left(x_2^i - \mu + \theta^i\right).$$

Notice that type θ^i's willingness to pay is unaffected by a ceiling on bids.

To derive type θ^i's willingness to pay for the first pot without a ceiling, one equates the expected payoff from winning the first auction,

$$-3 + x_1^i - \frac{2}{3}cb_1 + \left(2 - \frac{1}{3}E[P_2^j]\right),\ j \neq i, \tag{32}$$

and the expected payoff from losing the first auction,

$$-3 + \mu + \frac{1}{3}b_1 - \frac{2}{3}cE[P_2^i] + \theta^i,\ j \neq i. \tag{33}$$

This gives

$$p_1^i = \frac{3}{2c+1}\left[x_1^i - \mu + \left(1 + \frac{2c - \theta^j}{2c+1}\right)\theta^i\right].$$

This implies that P_1^i has a two-parameter exponential distribution,

$$P_1^i \sim \exp\left(\frac{3}{2c+1}\left[-\mu + \left(1 + \frac{2c - \theta^j}{2c+1}\right)\theta^i\right], \widetilde{\mu}\right),$$

where $\widetilde{\mu} = 3\mu/(2c+1)$. With a ceiling on bids in place, $P_1^{c,i}$ is obtained analogously, albeit with

$$E[P_2^{c,j}] = \left(1 - \exp\left\{-\frac{\bar{b}}{\widetilde{\mu}}\right\}\right)\frac{3\theta^j}{2c+1}$$

and $E[P_2^{c,i}]$ subsituted for $E[P_2^j]$ and $E[P_2^i]$ in (32) and (33), respectively. The two-parameter exponential distribution of $P_1^{c,i}$ is as follows,

$$P_1^{c,i} \sim \exp\left(\frac{3}{2c+1}\left\{-\mu + \left[1 + \left(1 - \exp\left\{-\frac{\bar{b}}{\widetilde{\mu}}\right\}\right)\frac{2c - \theta^j}{2c+1}\right]\theta^i\right\}, \widetilde{\mu}\right). \tag{34}$$

We introduce some additional notation. We denote by $\widetilde{B}_1^c$ the maximum willingness to pay among the two borrowers in the first auction of a Rosca with a ceiling,

$$\widetilde{B}_1^c = \max\left(P_1^{c,L}, P_1^{c,H}\right).$$

In the subsequent proofs of the propositions 2 and 3, the following lemmas will be useful.

Lemma 1 *In the first auction of a Rosca with ceiling $\bar{b}$, the probability of the high risk type having a higher willingness to pay than the low risk type is increasing in $\bar{b}$,*

$$\frac{d\Pr(P_1^H > P_1^L)}{d\bar{b}} > 0.$$

Proof. *Denote $P_1^{c,i}$'s lower bound by $\gamma_1^{c,i}$. Integrating, one obtains*

$$\Pr(P_1^H > P_1^L) = 1 - \frac{1}{2}\exp\left\{\frac{\gamma_1^{c,H} - \gamma_1^{c,L}}{\widetilde{\mu}}\right\}.$$

Thus $sign\left[\frac{d\Pr(P_1^H > P_1^L)}{d\bar{b}}\right] = -\ sign\left[\frac{d(\gamma_1^{c,H} - \gamma_1^{c,L})}{d\bar{b}}\right]$. From (34) we have that

$$\frac{d(\gamma_1^{c,H} - \gamma_1^{c,L})}{d\bar{b}} = \frac{1}{\widetilde{\mu}}\exp\left\{-\frac{\bar{b}}{\widetilde{\mu}}\right\}\frac{3}{(2c+1)^2}\left[(2c - \theta^L)\theta^H - (2c - \theta^H)\theta^L\right],$$

which is clearly greater than zero. ■

Lemma 2 *B_1 first-order-stochastically dominates $\widetilde{B}_1^c$.*

Proof. *Consider P_1^i's distribution. It follows from (34) that $\frac{d\gamma_1^{c,i}}{d\bar{b}} > 0$. Notice that P_1^i has the distribution of $P_1^{c,i}$ with an infinite $\bar{b}$. Hence P_1^i first-order stochastically dominates $P_1^{c,i}$ for $i = \{L, H\}$, which implies that $B_1^c = \max\left(P_1^L, P_1^H\right)$ first-order stochastically dominates $\widetilde{B}_1^c = \max\left(P_1^{c,L}, P_1^{c,H}\right)$.* ■

Some intuition for the following lemma follows. The expected riskiness of the first borrower is θ^H given a winning bid at the lower bound of the high type's willingness to pay. The probability that the high type has a lower willingness to pay than the observed winning bid is zero at the lower bound of the winning bid, and increases monotonically as the observed winning bid increases. This monotonic feature depends on the distribution of X_{it}. This implies then that the conditional expected riskiness given the winning bid is highest at the lower bound and is strictly decreasing in the winning bid

Lemma 3 *The conditional expected value of riskiness given the winning bid $E(\theta_1 \mid b_1)$ is strictly decreasing in the winning bid b_1 before the ceiling if there is adverse selection, i.e. (8) holds. The same is true after the ceiling: $E(\theta_1^c \mid b_1^c)$ is strictly decreasing in b_1^c if there is adverse selection, i.e. (8) holds.*

Proof

When there is no ceiling, $E[\theta_1|b_1] = \theta^L + (\theta^H - \theta^L)\Pr(P_1^L < P_1^H|b_1)$. Hence $\frac{d\Pr(P_1^L<P_1^H|b_1)}{db_1} < 0$ for all b_1 implies the claim. Notice that

$$\begin{aligned}\Pr\left(P_1^L < P_1^H|b_1\right) &= \frac{\Pr(P_1^L < b_1 \cap P_1^H = b_1)}{\Pr(P_1^L < b_1 \cap P_1^H = b_1) + \Pr(P_1^H < b_1 \cap P_1^L = b_1)} \\ &= \left[1 + \frac{f_1^L(b_1)F_1^H(b_1)}{f_1^H(b_1)F_1^L(b_1)}\right]^{-1} = \left[1 + \frac{\exp\left\{\frac{b_1-\gamma_1^H}{\widetilde{\mu}}\right\} - 1}{\exp\left\{\frac{b_1-\gamma_1^L}{\widetilde{\mu}}\right\} - 1}\right]^{-1}.\end{aligned}$$

Hence $\operatorname{sign}\left(\frac{d\Pr(P_1^L<P_1^H|b_1)}{db_1}\right) = -\operatorname{sign}\left(\frac{d}{db_1}\left\{\left[\exp\left\{\frac{b_1-\gamma_1^H}{\widetilde{\mu}}\right\} - 1\right] / \left[\exp\left\{\frac{b_1-\gamma_1^L}{\widetilde{\mu}}\right\} - 1\right]\right\}\right)$. To ease the notational burden, define $\widetilde{h} = \frac{b_1-\gamma_1^H}{\widetilde{\mu}}$, $\widetilde{l} = \frac{b_1-\gamma_1^L}{\widetilde{\mu}}$ and notice that $\widetilde{h} < \widetilde{l} < \widetilde{h} + \widetilde{l}$. We have that

$$\frac{d}{db_1}\frac{\exp\left\{\frac{b_1-\gamma_1^H}{\widetilde{\mu}}\right\} - 1}{\exp\left\{\frac{b_1-\gamma_1^L}{\widetilde{\mu}}\right\} - 1} = \frac{1}{\widetilde{\mu}}\frac{\left(\exp\{\widetilde{h}+\widetilde{l}\} - \exp\{\widetilde{l}\}\right) - \left(\exp\{\widetilde{h}+\widetilde{l}\} - \exp\{\widetilde{h}\}\right)}{\left(\exp\left\{\widetilde{l}\right\} - 1\right)^2},$$

which is clearly smaller than zero.

When there is a ceiling, $E[\theta_1^c|b_1] = \theta^L + (\theta^H - \theta^L)\Pr(H \text{ wins first auction}|b_1)$, where

$$\begin{aligned}&\Pr\left(H \text{ wins first auction}|b_1\right) \\ =& \begin{cases} \Pr(P_1^L < P_1^H|b_1), & \text{if } b_1 < \bar{b} \\ \frac{\Pr(P_1^{c,H} \geq \bar{b} \cap P_1^{c,L} < \bar{b}) + \frac{1}{2}\Pr(P_1^{c,H} \geq \bar{b} \cap \bar{b} \leq P_1^{c,L})}{\Pr(P_1^{c,H} \geq \bar{b} \cap P_1^{c,L} < \bar{b}) + \Pr(P_1^{c,H} \geq \bar{b} \cap P_1^{c,L} \geq \bar{b}) + \Pr(P_1^{c,H} < \bar{b} \cap P_1^{c,L} \geq \bar{b})}, & \text{if } b_1 = \bar{b} \end{cases}.\end{aligned} \tag{35}$$

Moreover, it follows from (34) that $\gamma_1^{c,H} > \gamma_1^{c,L}$ for all $\bar{b}$. Hence, the proof of

$$\frac{d\Pr(H \text{ wins first auction}|b_1)}{db_1} < 0,\ b_1 < \bar{b} \tag{36}$$

is analogous to the situation without a ceiling considered previously. For $b_1 = \bar{b}$, on the other hand, we have from (35) that

$$\Pr(H \text{ wins first auction}|\bar{b}) = \frac{1}{2}\left(\exp\{\widetilde{l}^c\} + \exp\{\widetilde{h}^c\} - 1\right)^{-1}, \tag{37}$$

where $\widetilde{l}^c = \frac{\bar{b}-\gamma_1^{c,L}}{\widetilde{\mu}}$, $\widetilde{h}^c = \frac{\bar{b}-\gamma_1^{c,H}}{\widetilde{\mu}}$. Notice that $0 < \widetilde{h}^c < \widetilde{l}^c < \widetilde{h}^c + \widetilde{l}^c$. On the other hand

$$\lim_{b_1 \nearrow \bar{b}} \Pr(H \text{ wins first auction}|b_1) = \left(1 + \frac{\exp\{\widetilde{h}^c\} - 1}{\exp\{\widetilde{l}^c\} - 1}\right)^{-1},$$

which is easily shown to be strictly larger than (37). In words, $\Pr(H \text{ wins first auction}|b_1)$ exhibits a downward jump at $\bar{b}$. This together with (36) establishes the claim for Roscas with a ceiling.

*

Proof of Proposition 2

(i) First notice that $E[\theta_1] = \theta^L + (\theta^H - \theta^L)\Pr(P_1^L < P_1^H)$, $E[\theta_1^c] = \theta^L + (\theta^H - \theta^L)\Pr(P_1^{c,L} < P_1^{c,H})$. Thus $\text{sign}(E[\theta_1 - \theta_1^c]) = \text{sign}(\Pr(P_1^L < P_1^H) - \Pr(P_1^{c,L} < P_1^{c,H}))$. Without a ceiling we have that

$$\Pr(H \text{ wins first pot}) = \Pr(P_1^L < P_1^H).$$

With a ceiling in place, we have that

$$\begin{aligned}
\Pr(H \text{ wins first pot}) &= E_{B_1^c}[\Pr(H \text{ wins first pot}|B_1^c)] \qquad (38)\\
&= E_{B_1^c}[\Pr(P_1^{c,L} < P_1^{c,H}|B_1^c < \bar{b})]\Pr(B_1^c < \bar{b})\\
+\Pr\left(H \text{ wins first pot}|B_1^c = \bar{b}\right)\Pr(B_1^c &= \bar{b})
\end{aligned}$$

Notice that the events $B_1^c < \bar{b}$ and $\widetilde{B}_1^c < \bar{b}$ are identical. Hence

$$\begin{aligned}
&\Pr\left(H \text{ wins first pot}|B_1^c = \bar{b}\right)\\
=\;&\Pr(H \text{ wins first pot}|\widetilde{B}_1^c \geq \bar{b})\\
=\;&E_{\widetilde{B}_1^c}\left[\frac{\Pr(P_1^{c,H} = \widetilde{B}_1^c \cap P_1^{c,L} < \bar{b}) + \frac{1}{2}\Pr(P_1^{c,H} = \widetilde{B}_1^c \cap \bar{b} \leq P_1^{c,L} < \widetilde{B}_1^c)}{\Pr(P_1^{c,H} = \widetilde{B}_1^c \cap P_1^{c,L} < \bar{b}) + \Pr(P_1^{c,L} = \widetilde{B}_1^c \cap P_1^{c,H} < \bar{b})}|\widetilde{B}_1^c \geq \bar{b}\right]\\
<\;&E_{\widetilde{B}_1^c}\left[\frac{\Pr(P_1^{c,H} = \widetilde{B}_1^c \cap P_1^{c,L} < \bar{b})}{\Pr(P_1^{c,H} = \widetilde{B}_1^c \cap P_1^{c,L} < \bar{b}) + \Pr(P_1^{c,L} = \widetilde{B}_1^c \cap P_1^{c,H} < \bar{b})}|\widetilde{B}_1^c \geq \bar{b}\right]\\
=\;&E_{\widetilde{B}_1^c}\left[\Pr(P_1^{c,L} < P_1^{c,H}|\widetilde{B}_1^c \geq \bar{b})|\widetilde{B}_1^c \geq \bar{b}\right].
\end{aligned}$$

Continuing (38), we have

$$
\begin{aligned}
& \Pr(H \text{ wins first pot}) \\
< \; & E_{B_1^c}[\Pr(P_1^{c,L} < P_1^{c,H} | \widetilde{B}_1^c < \bar{b})] \Pr(\widetilde{B}_1^c < \bar{b}) \\
+E_{\widetilde{B}_1^c}\left[\Pr(P_1^{c,L} < P_1^{c,H} | \widetilde{B}_1^c \geq \bar{b}) | \widetilde{B}_1^c \geq \bar{b}\right] \Pr(\widetilde{B}_1^c \geq \; & \bar{b}) \\
= \; & E_{B_1^c}[\Pr(P_1^{c,L} < P_1^{c,H} | \widetilde{B}_1^c)] = \Pr(P_1^{c,L} < P_1^{c,H}).
\end{aligned}
$$

A sufficient condition for the the claim to hold is thus $\Pr(P_1^L < P_1^H) - \Pr(P_1^{c,L} < P_1^{c,H}) \geq 0$, which in turn follows from Lemma 1. To see this notice that a Rosca without ceiling can be represented as the limit of a Rosca with ceiling where the ceiling tends to infinity. Hence,

$$\Pr(P_1^L < P_1^H) - \Pr(P_1^{c,L} < P_1^{c,H}) = \int_{\bar{b}}^{\infty} \frac{d \Pr(P_1^{c,L} < P_1^{c,H})}{d\bar{b}} d\bar{b}.$$

Lemma 1 establishes that the integrand is positive for all $\bar{b}$.

(ii) Lemma 2 establishes that B_1 first-order-stochasically dominates $\widetilde{B}_1^c$. As $B_1^c = \min(\widetilde{B}_1^c, \bar{b})$, B_1 also first-order-stochasically dominates B_1^c, which in turn implies the claim.

*

Proof of Proposition 3

(i) $E[\theta_1 - \theta_1^c] > 0$ follows from Proposition 2, part i. Moreover, $E[\theta_1 + \theta_2] = E[\theta_1^c + \theta_2^c] = \theta^L + \theta^H$, which implies that $E[\theta_2 - \theta_2^c] = -E[\theta_1 - \theta_1^c] < 0$.

(ii) If $\theta^L = \theta^H = \bar{\theta}$, then P_1^L and P_1^H, $P_1^{c,L}$ and $P_1^{c,H}$, P_2^L and P_2^H, and $P_2^{c,L}$ and $P_2^{c,H}$ have the same distribution. Hence $E[\theta_1] = E[\theta_1^c] = E[\theta_2] = E[\theta_2^c] = \bar{\theta}$, which proves the claim.

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Table 1. Descriptive Statistics for the 1993 Sample by Denomination

			Number of Roscas			Winning Bid ≥ 30% of Pot (%)			
						Before Shock		*After Shock*	
Duration	Contribution	Pot	*Before Shock*	*After Shock*	*Increase (%)*	*Early*	*Late*	*Early*	*Late*
			(1)	(2)	(3)	(4)	(5)	(6)	(7)
20	500	10,000	15	41	46.4	7.1	1.7	18.6	3.8
25	400	10,000	46	39	-8.2	24.6	0.0	44.0	0.3
25	1,000	25,000	26	28	3.7	28.1	0.0	42.4	0.4
30	500	15,000	100	117	7.8	40.5	0.0	57.2	0.6
40	250	10,000	423	331	-12.2	79.1	1.8	72.7	3.8
40	500	20,000	25	16	-22.0	83.2	3.8	87.9	7.2
40	625	25,000	101	54	-30.3	71.9	0.4	69.8	3.8
40	1,250	50,000	6	10	25.0	82.2	0.6	83.8	7.9
40	2,500	100,000	12	10	-9.1	71.4	0.4	67.2	5.2
50	1000	50,000	84	43	-32.3	91.6	4.8	91.9	14.2
60	1,250	75,000	15	7	-36.4	98.4	13.3	97.9	27.0

Notes:
Column 1: Roscas started between October 1992 and September 1993. No bid ceiling.
Column 2: Roscas started between October 1993 and September 1994. Bid ceiling 30% of pot.
Column 3: Percentage difference between values in columns 2 and 1.
Columns 4 to 7: Percentage of auctions in which the winning bid is no less than 30% of the pot.
Columns 4 and 6: First half of rounds in a Rosca.
Columns 5 and 7: Second half of rounds in a Rosca.

Table 2
Last Round in which Bidding Reaches the Ceiling after the Policy Shock in a Specified Fraction of all Sample Roscas by Denomination

		Fraction of Roscas		
Duration	Contribution	50%	33.3%	25%
20	500	.*	2	3
25	400	5	6	7
25	1000	5	6	7
30	500	8	10	10
40	250	17	19	20
40	500	19	21	22
40	625	15	17	18
40	1250	19	21	21
40	2500	16	18	18
50	1000	25	28	30
60	1250	37	39	40

Notes:
* The ceiling is never reached for half of all Roscas of the (20,500) denomination.
Illustration: For instance, for the (40,250) denomination, the last round for which the ceiling of Rs. 3000 is reached for half the Roscas of that denomination is round 19; the last round for which the ceiling is reached for a third (and a quarter) of the Roscas of that denomination is round 21.
Sample: all Roscas that were started between October 1993 and September 1994.

Table 3. Individual Default Rates, Winning Bids and Repayment Burdens
Mean Percentages

	Before Shock		After Shock		Single Difference		Double Difference
	Early	*Late*	*Early*	*Late*	*Early*	*Late*	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Default Rate (percent)	2.04	0.59	1.64	0.74	0.40	-0.15	0.55
	(89.46)	(4.64)	(8.13)	(5.17)	[0.06]	[0.05]	[0.08]
Winning Bid (percentage of pot)	35.07	11.51	27.17	10.22	7.90	1.29	5.61
	(11.09)	(6.76)	(5.49)	(6.06)	[0.07]	[0.06]	[0.09]
Repayment Burden (percentage of pot)	64.33	23.38	63.41	22.84	0.92	0.54	0.38
	(10.12)	(12.98)	(10.46)	(12.73)	[0.10]	[0.09]	[0.14]
Observations	13.034	13,977	8,744	10,514			

Notes: Standard deviations in parentheses for columns 1-4; Standard errors in brackets for columns 5-7.
Column 5 reports the difference between the values in columns 1 and 3, column 6 the difference between columns 2 and 4.
Column 7 reports the difference between the values in columns 5 and 6.

Table 4. Selection on Unobserved Riskiness
OLS
Dependent Variable: Individual Default Rate

		(1)	(2)	(3)	(4)	(5)	(6)
After*Rounds in Later Two Thirds		0.0067***	0.0054**				
		(0.0015)	(0.0027)				
After*Rounds in Second Half				0.0037***	0.0039		
				(0.0013)	(0.0024)		
After*Relative Round						0.0082***	0.0113**
						(0.0024)	(0.0052)
Controls							
Cosigner Incidence		No	Yes	No	Yes	No	Yes
	F(388,44265)		3.97***		3.97***		3.97***
	Average Effect		-0.0002		-0.0002		-0.0003
			(0.0032)		(0.0032)		(0.0032)
Number of Cosigners		No	Yes	No	Yes	No	Yes
	F(388,44265)		1.04		1.04		1.04
	Average Effect		0.0004		0.0004		0.0004
			(0.0016)		(0.0016)		(0.0016)
Winning Bid (logarithmic)		No	Yes	No	Yes	No	Yes
	F(388,44265)		5.54***		5.54***		5.54***
	Average Effect		0.0172***		0.0177***		0.0182***
			(0.0049)		(0.0048)		(0.0048)
Repayment Burden (logarithmic)		No	Yes	No	Yes	No	Yes
	F(388,44265)		1.71***		1.71***		1.71***
	Average Effect		0.0202		0.0202		0.0202
			(0.1306)		(0.1306)		(0.1306)
Aggregate Shocks		No	Yes	No	Yes	No	Yes
	F(24,44265)		4.60***		4.60***		4.60***
R^2		0.0463	0.0917	0.046	0.0917	0.0461	0.0918
Observations		46,269	46,269	46,269	46,269	46,269	46,269

Notes:
* significant at 90%; ** significant at 95%; *** significant at 99%.

Standard errors in parentheses in all specifications. All specifications include 388 denomination-specific round dummies. There are $T_k - 2$ per denomination, where T_k is the number of rounds of a Rosca of denomination *k*. There are $T_k - 2$ instead of T_k interactions per denomination because there is no auction in the first and last round.
The *late* variable is interacted with 11 denomination-specific dummies.
24 quarterly dummies (first quarter 1993 to 4th quarter 1998) are added in specifications with controls for aggregate shocks. A dummy for the incidence of cosigners, which equals one if there is at least one cosigner attached to the loan and zero otherwise, interacted with denomination-specific round dummies (a total of 388 interactions) are added in specifications with controls for cosigner incidence. The number of cosigners attached to the loan interacted with denomination-specific round dummies (a total of 388 interactions) are added in specifications with controls for the number of cosigners. The logarithm of the winning bid interacted with denomination-specific round dummies (a total of 388 interactions) are added in specifications with controls for the winning bid. The logarithm of the amount owed interacted with denomination-specific round dummies (388 interactions) are added in specifications with controls for repayment burden.

Table 5. Selection on Observed Riskiness

OLS

Dependent Variable: Cosigner Incidence (= 0 if no cosigner is attached to the loan and one otherwise)

		(1)	(2)	(3)	(4)	(5)	(6)
After*Rounds in later Two Thirds		-0.0061	-0.0063				
		0.01	0.0184				
After*Rounds in Second Half				0.0002	0.01		
				0.0092	0.0162		
After*Relative Round						-0.0045	0.0126
						0.0161	0.0353
<u>Controls</u>							
Winning Bid (logarithmic)		No	Yes	No	Yes	No	Yes
	F(388,44265)		83.81 ***		83.81 ***		83.81 ***
	Average Effect		0.0149		0.0147		0.0153
			0.0332		0.0365		0.0309
Repayment Burden (logarithmic)		No	Yes	No	Yes	No	Yes
	F(388,44265)		1.01		1.01		1.01
	Average Effect		0.8985		0.8978		0.9001
			0.8952		0.8949		0.8951
Aggregate Shocks		No	Yes	No	Yes	No	Yes
	F(24,44265)		1.42		1.42		1.42 *
R^2		0.0573	0.0746	0.0573	0.0746	0.0573	0.0746
Observations		46,269	46,269	46,269	46,269	46,269	46,269

Notes:

* significant at 90%; ** significant at 95%; *** significant at 99%.

Standard errors in parentheses in all specifications. All specifications include 388 denomination-specific round dummies. There are $T_k - 2$ per denomination, where T_k is the number of rounds of a Rosca of denomination *k*. There are $T_k - 2$ instead of T_k interactions per denomination because there is no auction in the first and last round.

The *late* variable is interacted with 11 denomination-specific dummies.

24 quarterly dummies (first quarter 1993 to 4th quarter 1998) are added in specifications with controls for aggregate shocks. The logarithm of the winning bid interacted with denomination-specific round dummies (a total of 388 interactions) are added in specifications with controls for the winning bid. The logarithm of the amount owed interacted with denomination-specific round dummies (388 interactions) are added in specifications with controls for repayment burden.

Figure 1. Empirical Average Default Profiles with Unrestricted (dotted line) and Restricted Bidding (solid line) in Roscas of the <30 round, Rs 500 contribution> denomination

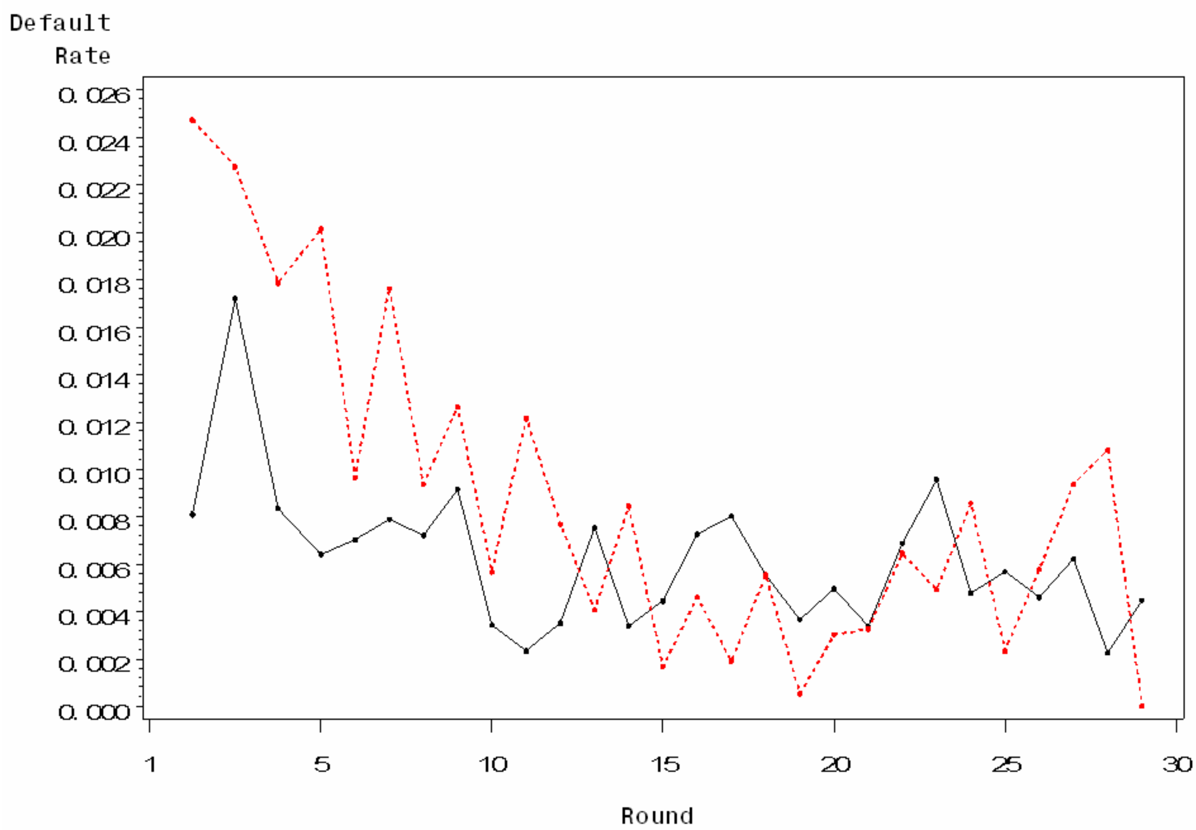

Figure 2. Willingness to Pay for the First Pot and Bid Ceiling.
High Risk Type (dashed line). Low Risk Type (solid line). Bid Ceiling (dotted line)

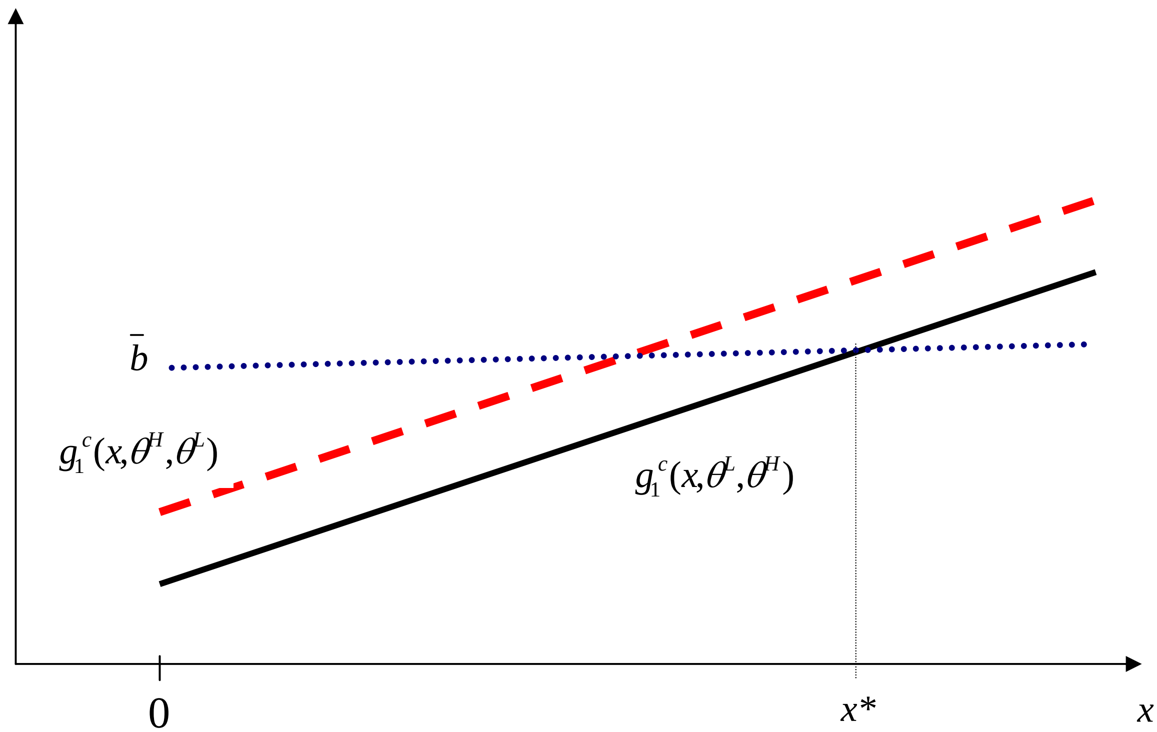

Figure 3. Riskiness Profiles in the Presence of Adverse Selection
Rosca with Unrestricted Bidding (dashed line) and Restricted Vidding (solid line).

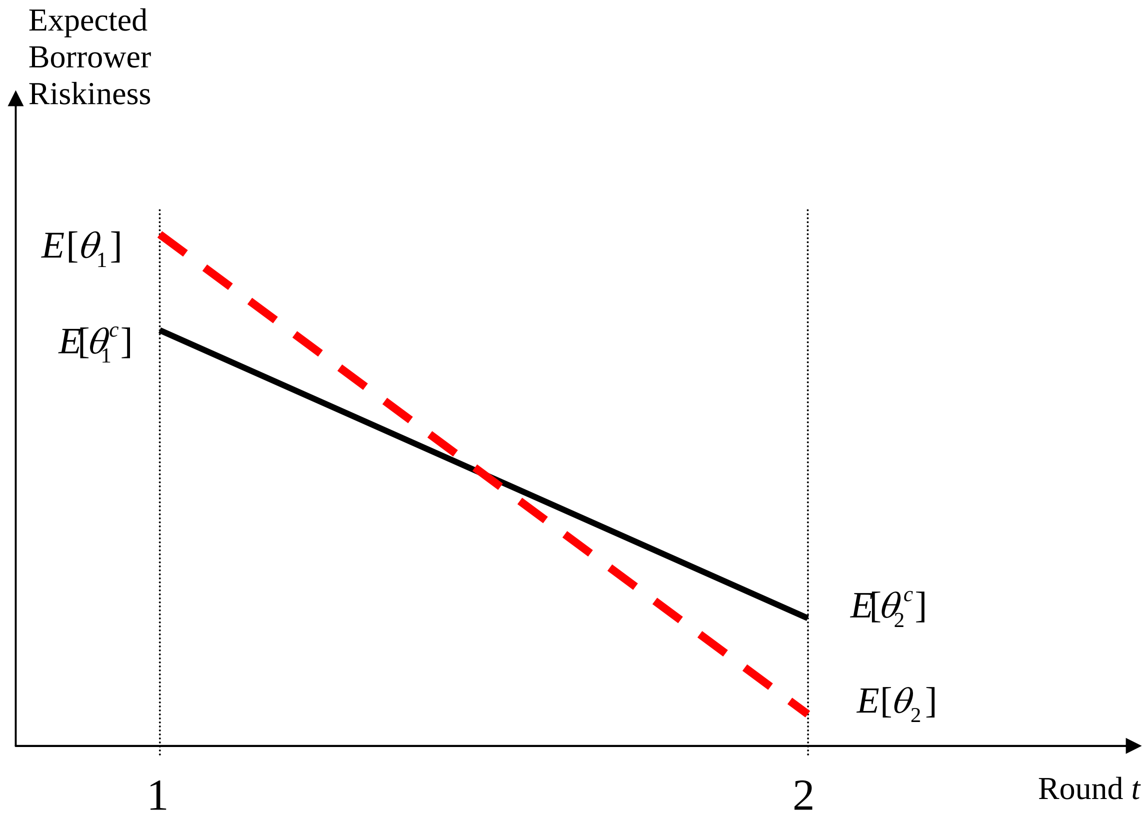

Figure 4. Riskiness Profile in a Rosca in the Absence of adverse Selection