Adverse Selection in Credit Markets: Evidence from a Policy Experiment

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Abstract: We test if riskier borrowers are willing to pay higher interest rates than safer borrowers are as predicted by Stiglitz and Weiss (1981). 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 and then relaxed the ceiling in 2002. Changes in default patterns are analyzed before and after each of these policy changes. We find evidence of adverse selection despite the use of collateral as a screening device. This study isolates adverse selection from moral hazard and controls for information on riskiness observed by the lender but not by the researcher.

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1 Introduction

Many economists believe that asymmetric information impedes the efficient functioning of markets. Even though there is substantial theoretical research assuming asymmetric information, the empirical evidence is rather mixed and largely confined to insurance markets (Chiappori and Salanie 2003). In this paper we look for evidence of adverse selection in credit markets. We test if riskier borrowers are willing to pay higher interest rates than safer borrowers are. This well-known prediction of Stiglitz and Weiss' (1981) adverse selection model has had considerable influence in finance, development and macroeconomics but has received little empirical scrutiny.

There are three main empirical challenges for such a study. First, the lender may have "soft" information on borrower riskiness which by definition is difficult to transmit or codify (Berger and Udell, 1995,Petersen and Rajan 1994, Uzzi and Lancaster, 2003). A researcher who is uninformed about the lender's soft information may erroneously conclude that there is adverse selection on unobserved riskiness. Secondly, lenders often require collateral. Even if adverse selection is an underlying friction in credit markets, the lender may be able to mitigate adverse selection through collateral requirements (Bester 1985) or using cosigners as collateral (Besanko and Thakor 1987). Finally, researchers only have information on defaults. It is entirely possible that higher interest rates increase default rates for moral hazard reasons instead of adverse selection. It is difficult empirically to distinguish between the two.

The contribution of this paper is to address all three empirical challenges. We distinguish unobserved riskiness from riskiness that is observed by the lender but not by the researcher, we allow for the effects of collateral, and we 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 adverse selection. This includes tests for adverse selection in used car markets (Genesove 1993), annuity markets (Finkelstein and

Poterba, 2004) and most specifically, in credit markets (Ausubel 1999; Karlan and Zinman, 2007). We use a natural experiment involving a South Indian financial institution; Ausubel (1993) and Karlan and Zinman (2007) use randomized field experiments. We are able to isolate adverse selection from moral hazard, while Ausubel (1999) is not able to do so. We find much stronger evidence of adverse selection than Karlan and Zinman (2007) do. A final difference is that both Ausubel (1999) and Karlan and Zinman (2007) use data from uncollateralized lenders. So unlike us they cannot test whether adverse selection persists in the presence of collateral.

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. In bidding Roscas, the subject of our study, the pot is awarded to the highest bidder.[1] 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 receives loans of different sizes and durations and at different interest rates.

We test if riskiness is positively related to the willingness to pay by exploiting an exogenous policy shock. In September 1993, the government unexpectedly imposed a bid ceiling (of 30% of the total pot). This ceiling typically binds in early rounds of bidding Roscas transforming them into random Roscas since many participants bid up to the ceiling but only one of them chosen by lottery receives the pot. To test for self selection based on unobserved riskiness, we compare the riskiness profile of borrowers before and after the policy shock. Riskier borrowers who have a higher willingness to pay can simply outbid the others and take early pots in the absence of the ceiling. But the ceiling effectively ties their hands. Even though safer borrowers have lower willingness to pay for early pots, they

[1] The rationale for random and bidding Roscas has been examined by Anderson and Baland (2002), Besley, Coate and Loury (1993), Calomiris and Rajaraman (1998), Klonner (2003, 2006) and Kovsted and Lyk-Jensen (1999), among others.

may still win early pots because of the lottery introduced by the ceiling. Riskier borrowers are hence pushed back to later pots. In other words. the bid ceiling causes a *flattening* of the riskiness profile of borrowers. This is easiest to see in an extreme case: suppose the bid ceiling were set at zero percent transforming the bidding Rosca into a purely random Rosca. The average riskiness of all borrowers would be the same implying a completely flat risk profile.

We use a difference-in-difference specification to capture this flattening. We predict that the difference between the riskiness of early and late borrowers is smaller after the policy shock compared with before if there is self selection based on riskiness but is unaffected if there is no such self selection. Borrower riskiness is unobserved, however; only default rates are observed. We find that defaults are indeed flatter as a result of the policy shock exactly as predicted by selection. But default rates could be flatter for reasons completely unrelated to selection including (a) differences in riskiness observed by the lender but not necessarily by the researcher (b) moral hazard (c) changes in the composition of Rosca participants and (d) transitory aggregate shocks.

We address all four alternative explanations for the flattening of default rates. 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 collateral requirements have been shown to be based on observed riskiness (Berger and Udell 1995; Jimenez, Salas and Saurina 2006, Liberti and Mian 2006). Since loan terms become more favorable to early borrowers 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 to 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 alone caused 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.

We find that borrowers willing to pay higher interest rates are riskier than those who are not despite the use of collateral requirements. The flattening of defaults caused by self selection on unobserved riskiness is substantial. Our estimates suggest that defaults of early borrowers fell by one-fifth and defaults of later borrowers rose by one-fifth as a consequence of the bid ceiling.. We reject the null hypothesis of no self selection on unobserved riskiness. 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.

We also examine the effect of a partial relaxation of the bid ceiling in 2002 on arrears (overdues at the end of the loan term). We find that the 2002 reversal had the opposite effect on arrears. Riskier participants could get their hands on early pots more easily after 2002 and so observed arrear profiles became steeper. In sum, exogenously capping interest rates in 1993 flattened the default profile while relaxing the cap in 2002 made the arrear profile steeper exactly as predicted.

Financial constraints on development are a major academic and policy concern (Armendariz and Morduch, 2005, Banerjee 2003). Our results provide evidence of a specific impediment to financial access in developing countries. The bidding Roscas we study are not small scale and personalized (as in Anderson, Baland and Moene 2004). Instead, a commercial organizer arranges financial intermediation between Rosca participants who have no social ties. The organizer 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. Cosigners also potentially serve as a deterrent for unobserved riskiness, but we find that they are ineffective in doing so.

Finally we briefly clarify what we do *not* do in this paper.

1. We test for a specific information asymmetry: Do borrowers have private information on their own riskiness (or equivalently, dishonesty) that the Rosca organizer? There

may of course be other unobserved characteristics on which borrowers differ such as impatience or productivity that have nothing to do with riskiness but affect a borrower's willingness to pay for an early pot. We do not test for asymmetric information on those dimensions.

2. We find evidence that the bid ceiling does reduce self selection on unobserved riskiness. We follow Stiglitz and Weiss (1981) in referring to such self selection on riskiness as "adverse" but we make no claims about whether such self selection (or the bid ceiling that prevents self selection) has positive or negative welfare effects in bidding Roscas.[2] Just as Ausubel (1999) and Karlan and Zinman (2007) therefore, we use the term adverse selection throughout this paper to refer to the positive relationship between willingness to pay and unobserved riskiness, not to the welfare loss relative to a full information environment.

3. We do not claim either the presence or absence of moral hazard. At first glance it might seem that the lottery introduced by the bid ceiling could be used to isolate moral hazard. However identifying moral hazard in this way is impossible since only the identity of the lottery winner but not of all lottery entrants is recorded.

Paper Outline

We motivate our empirical strategy in section 2. We provide additional background on bidding Roscas in South India, and on the policy shock in section 3. We construct a simple model of bidding Roscas when riskiness is unobserved in section 4. We discuss our identification strategy in section 5. We discuss our main results from the 1993 policy shock in section 6. We examine a reverse policy shock in 2002 as a robustness check in section

[2] In contrast with equilibrium credit rationing in Stiglitz and Weiss (1981), credit rationing was exogenously imposed here. While the bid ceiling constrained riskier borrowers, it also reduced participation in Roscas which was undesirable from the Rosca organizer's perspective. 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.

7. We conclude in section 8.

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 is an implication of adverse selection analogous to Stiglitz and Weiss (1981). So riskier participants will all else equal choose early pots and safer participants will choose later pots. 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 self selection on riskiness.

Riskiness is unobserved by the researcher 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 illustrate, 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. It is possible that the early borrower in example R has a collateral 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 collateral 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 collateral 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 is important for another reason. A key empirical challenge is that the researcher is uninformed about what the organizer observes about the borrowers. Collateral required can be used as a summary measure of observed borrower riskiness. The organizer will all else equal impose a stricter collateral requirement on borrowers who are known to be riskier than on borrowers known to be safer (as indicated by our field interviews). By conditioning the double difference on defaults on the collateral requirement we are able to ensure that it is differences in *unobserved* riskiness that we are capturing.

The collateral 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 screening high risks we should see no flattening of defaults when we compare Roscas U' and R (where both have the same collateral 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 2004; Radhakrishnan 1975). 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 2004). 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. $10,000$. 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 savers 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 2004 who use the same policy shock to study matching among borrowers and savers).

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 (2004), 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 a participant would receive the pot.

As we mentioned in the Introduction, this bid ceiling was relaxed in 2002. We postpone discussion of this reversal to Section 7.

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 in his place. The organizer receives two forms of payment for taking on the default risk. 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.

To enforce repayments, the organizer relies on a collateral (cosigner) requirement. Winners of a pot are asked to provide cosigners (typically salaried employees) after they have won the auction but before winnings are disbursed. Cosigners put their future incomes at risk by agreeing to guarantee loans for a borrower – and the Rosca organizer enforces payment from cosigners of defaulting borrowers through a lengthy legal process. The organizer

[3] We observed the lottery procedure 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.

also tries to discourage risky participants by requiring that all borrowers prove that they have a regular income before pots are awarded to them. The threat of legal action against both cosigners and borrowers provides incentives for participants to continue making contributions even after they have received the pot. 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.

Sample

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 35 Roscas between October 1992 and September 1994. The numbers of Roscas of different denominations in the sample are in Table 1. According to Table 1, there was a substantial decrease in the number of Roscas formed for most of the Rosca denominations, including a drop of 20.8 percent for the popular $(40, 250)$ denomination. Table 1 shows that 72% of the winning bids in the first half of Roscas of this denomination hit the ceiling of Rs. 3000 after the policy shock. By way of comparison, 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.

Since the Rosca organizer receives the first pot, and the last recipient owes no future contributions, we only include data on recipients in rounds $t \in \{2, ..., n-1\}$ for our analysis. Between 10 and 25% of participants are institutional investors (depending on the Rosca denomination). Since these investors never default, all pots allocated to them are also excluded from the empirical analysis. Our sample then consists of the remaining noninstitutional borrowers in Roscas that started after September 1992 but before September 1994. This gives a total of 46,269 observations, where each observation refers to the winner in a particular round t in Rosca i of the 11 denominations in our sample.

Default Rates

We calculate the 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] The average default rate in the sample is 1.23 percent (Table 2).

In practice, the Rosca organizer updates repayments continuously as funds are collected from borrowers and their cosigners. The legal process for collecting funds can take up to four or five years since the courts are so slow. So default rates decrease over time in the months and years after a Rosca has ended. We collected data on default rates for this sample in January 2002. This is about five years after the conclusion of Roscas of the popular (40, Rs 250) denomination that were started in 1994.

Loan Terms

There are three components of the loan terms for a round t winner of a Rosca: the cosigner requirement c_t, the winning bid b_t and the repayment burden denoted q_t. We discuss each in turn. We use early and late throughout to refer to winners in the first and second half of Roscas respectively.

We use two measures of the cosigner requirement. The first is cosigner incidence – one or more cosigners were required for 39.3 percent of the entire sample (Table 2). The second

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

is the number of cosigners, which on average was 0.71 but with considerable variation. The bid ceiling did little to change the difference in cosigners required for early and late recipients.

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 2), 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, $\mathbf{b}_t = (b_{t+1}, ..., b_n)$. 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 a 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

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

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 2 shows that the average repayment burden before the policy shock 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 analyze a model of bidding Roscas in which borrowers differ in their riskiness (or dishonesty) and are prone to unpredictable needs for cash. The model incorporates the collateral (or cosigner) requirement that the Rosca organizer imposes on borrowers. 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 that if the collateral requirement is not a sufficient deterrent riskier borrowers will on average be willing to bid more for a pot than safer borrowers. Adverse selection then will refer to this positive relationship between willingness to pay and riskiness. If there is adverse selection then 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 adverse selection in section 5.

The Model

The model has three agents and three dates indexed by $t \in \{1, 2, 3\}$. Agents have additively separable, risk-neutral intertemporal preferences. The per-period individual discount factor is denoted by $\delta < 1$. Agents are endowed with an income stream of \$1 per period. Two of the agents are borrowers and one is a pure saver. Following Besley, Coate and Loury (1993) 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. We shall index borrowers by θ, where $\theta \in \{L, H\}$. A borrower's type θ refers to inherent riskiness or equivalently to the probability of default. One of the borrowers is safer, type L, while the other is riskier, type H, where $1 \geq H \geq L \geq 0$.

Following Calomiris and Rajaraman (1998) and Klonner (2003), 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 of type θ receives an instantaneous payoff x_t^θ from funding her cash needs where $t \in \{1, 2\}$. Throughout this section we will refer to random variables by upper case letters, and their realizations by lower case letters. We assume that X_t^θ is independently and identically uniformly distributed with mean μ and range α, i.e.$X_t^\theta \sim U\left(\mu - \frac{\alpha}{2}, \mu + \frac{\alpha}{2}\right)$.

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 $1 - \theta$ or pays neither contribution with probability θ. Similarly the date 2 borrower pays the third contribution with probability $1 - \theta$ and defaults with probability θ at date 3. The Rosca organizer takes on the risk of default, and contributes \$1 if the winner of a pot does not continue to pay.[5] We shall assume for tractability that the riskiness of each borrower θ and the realization of the shock x_t^θ are observed by all Rosca participants. But the Rosca organizer does not observe riskiness θ and the shock x_t^θ – and so cannot distinguish the riskier type from the safer type.

The Rosca organizer imposes a collateral requirement for winners of pots. Even though the organizer cannot make this collateral requirement contingent on the winner's type, the cost of providing the same collateral may vary between risky and safe types. We assume that the collateral causes an immediate loss in expected utility of $\pi^\theta \geq 0$ per dollar owed in present value terms. For instance, just as in Bester (1985) or Besanko and Thakor (1987) collateral costs are likely to be higher for the risky borrower than for the safe borrower: $\pi^H > \pi^L$. In our context π^θ can be interpreted as the bribe that a borrower must pay a cosigner in order to offset the cosigner's expected loss from default. Since a riskier borrower has to pay a higher bribe than a safer borrower, this would imply that $\pi^H > \pi^L$. In what

[5] 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.

follows, we do not restrict $\pi^H > \pi^L$ however.

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^θ 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^θ 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. If $H > L$, we can think of either (a) the high type is more dishonest than the low type, (b) the high type has a riskier project than the low type does or (c) the high type is better at hiding his money from the Rosca organizer than the low type is. Even though we model cash needs as unpredictable, our results would be unaffected if cash needs were anticipated instead, i.e. both x_1^θ and x_2^θ were realized at the beginning of date 1.

Next we discuss the payoffs to the participants in the Rosca. Let b_t be the winning bid at the date t auction when bidding is unrestricted. (When bidding is restricted by ceiling $\overline{b}$ then denote the winning bids as b_t^c, where $b_t^c \leq \overline{b}$ for each t). 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 at date 1 receives $3 - \frac{2}{3}b_1$. She needs additional finance of $\frac{2}{3}b_1$ to undertake the project. As in Kovsted and Lyk-Jensen (1999), 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 $k > 1$.[6] To provide a rationale for the existence of Roscas, we assume that, on average, it is not profitable to

[6] 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.

finance the investment from only concurrent income and the costly external source of funds,

$$E(X) - 2k < 1$$

where $E(X) - 2k$ is the expected instantaneous utility of a borrower who invests (he needs external finance of 2 which equals the cost of 3 minus his current income of 1). The right hand side is the payoff to an autarkic individual who does not invest, which equals her unit income. Since $EX = \mu$, we can state this assumption as:

Assumption 1 *Investment is not profitable in the absence of Rosca funding,*

$$\mu < 1 + 2k. \tag{3}$$

Next define a borrower of type θ's willingness to pay for a date t pot as $g_t(x_t, \Delta^\theta)$ where $\Delta^\theta = \theta - \pi^\theta$. Explicit expressions for $g_t(x_t, \Delta^\theta)$ are derived in the appendix by equating a borrower's payoff between winning the pot at date t and waiting till date $t+1$. Since we observe that winning bids are always non-negative, we shall assume that willingness to pay in the second auction is always non-negative.

Assumption 2 *The winning bid in the second auction is always positive, $g_2(x, \Delta^\theta) \geq 0$ for all x and Δ^θ or, equivalently,*

$$\delta \leq \min_{\theta \in \{L,H\}} \frac{\mu - \frac{\alpha}{2}}{\mu - \Delta^\theta} \tag{4}$$

Condition 4 states that individuals discount future income sufficiently to always ensure a positive willingness to pay for the second pot. Notice that $\mu - \alpha/2$ is the lower bound of X and $\mu - \Delta^\theta$ is type θ's gain in utility from winning the third pot.

By definition, Stiglitz and Weiss (1981)'s model implies that riskier borrowers have a higher willingness to pay than safer borrowers if there is no collateral required. In this paper though we wish to investigate if collateral can successfully overcome such self selection based on riskiness. In our model, any differences in willingness to pay for a pot will depend not just on differences in riskiness but also differences in collateral costs π^θ. If collateral is a sufficient deterrent, then the risky type will not have a higher willingness to pay than the safe type. The following proposition thus characterizes when self selection based on riskiness persists despite the collateral requirement, and when self selection is eliminated or even reversed because of the collateral requirement. All proofs are relegated to the Appendix.

Proposition 1 ***(Riskiness, Collateral Costs and Willingness to Pay for the First Pot)***

Conditional on the utility shock x,

(i) The riskier borrower has a higher willingness to pay for the first pot than the safe borrower does, $g_1(x, \mathbf{\Delta}^H) > g_1(x, \mathbf{\Delta}^L)$, *if and only if*

$$H - \pi^H > L - \pi^L. \tag{5}$$

(ii) Both types have the same willingness to pay for the first pot, $g_1(x, \mathbf{\Delta}^H) = g_1(x, \mathbf{\Delta}^L)$ *if and only if*

$$H - \pi^H = L - \pi^L \tag{6}$$

(iii) The riskier borrower has a lower willingness to pay for the first pot than the safe borrower does, $g_1(x, \mathbf{\Delta}^H) < g_1(x, \mathbf{\Delta}^L)$ *if and only if*

$$H - \pi^H < L - \pi^L \tag{7}$$

Proposition 1 makes precise the condition (5) under which adverse selection cannot be eliminated by the collateral requirement. In this context a positive difference in borrower riskiness, i.e. $H > L$ is not sufficient for adverse selection. Instead this difference has to be sufficiently large relative to the difference in the costs of providing collateral. For such a case, figure 2 depicts the borrowers' willingness to pay functions. The g_1 functions are parallel with $g_1(x, \mathbf{\Delta}^H)$ being larger than $g_1(x, \mathbf{\Delta}^L)$ for any given x.

On the other hand, there may indeed be no adverse selection even though there are unobserved differences in riskiness provided collateral is a sufficient deterrent (6). Finally, the proposition gives a case where the riskier borrower may have a lower willingness to pay than the safe borrower for the first pot because the collateral requirement is an excessive deterrent.

Unrestricted Bidding

We first discuss how winning bids are determined in the absence of the bid ceiling. Unlike standard auctions, the losers in any Rosca auction receive a one-third share of the winning bid – and so have an incentive to bid up the winner to her willingness to pay. In the equilibrium of this sequential auction game, the borrower with the higher willingness to pay wins the first auction at a winning bid equal to her willingness to pay for the first pot, while the other borrower wins the second pot at a winning bid equal to his willingness to pay for the second pot. Formally, we have

$$\begin{aligned} b_1 &= \max_{\theta} g_1(x_1^{\theta}, \mathbf{\Delta}^{\theta}), \\ b_2 &= g_2(x_2^{\theta}, \Delta^{\arg\min_{\theta} g_1(x_1^{\theta}, \mathbf{\Delta}^{\theta})}). \end{aligned}$$

Since each borrower's willingness to pay for the first pot is higher than for the second pot, on average this implies a decreasing winning bid over the course of a Rosca, a feature well in accordance with auction outcomes in actual Roscas.

Since X_1^L and X_1^H are independently and identically distributed, the high type is more likely to win the pot if condition (5) holds. Nevertheless 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 even if (5) holds.

Restricted Bidding

We turn to the effect of the bid ceiling $\bar{b}$ on winning bids. When bidding reaches $\bar{b}$, there is a lottery with equal odds among all bidders interested in obtaining the pot at that price.

We focus on a specific case in which the effect of the ceiling on riskiness is most clearly demonstrated. We will assume that the ceiling never binds in the second round. In the first round we require that the ceiling is interior: each borrower's willingness to pay for the first pot may be larger or smaller than $\bar{b}$ with positive probability. To illustrate in Figure 2: we require that the ceiling is larger than the ordinate of the left-most point on the dashed line and smaller than the ordinate of the right-most point of the solid line.

Assumption 3

(i) In the second auction, the ceiling is no lower than the highest possible winning bid,

$$\bar{b} \geq \max_{\theta} g_2\left(\mu + \frac{\alpha}{2}, \Delta^{\theta}\right).$$

(ii) In the first auction, each borrower's willingness to pay may be smaller or larger than $\bar{b}$ *with positive probability,*

$$\max_{\theta} g_1\left(\mu - \frac{\alpha}{2}, \boldsymbol{\Delta}^{\theta}\right) < \bar{b} < \min_{\theta} g_1\left(\mu + \frac{\alpha}{2}, \boldsymbol{\Delta}^{\theta}\right).$$

Part (i) of this assumption ensures that the willingness to pay functions in the first auction remain unchanged which simplifies the analytics. The inequality $\bar{b} < \min_{\theta} g_1\left(\mu + \frac{\alpha}{2}, \boldsymbol{\Delta}^{\theta}\right)$ in part (ii) ensures that the ceiling does have an effect on the identity of the winner of the first auction. The inequality $\max_{\theta} g_1\left(\mu - \frac{\alpha}{2}, \boldsymbol{\Delta}^{\theta}\right) < \bar{b}$ ensures that the ceiling is not always reached as is the case in the data.

The winning bids in rounds one and two are now obtained as

$$\begin{aligned} b_1^c &= \min\left[\max_{\theta} g_1(x_1^{\theta}, \boldsymbol{\Delta}^{\theta}), \bar{b}\right], \\ b_2^c &= \begin{cases} g_2(x_2^{\theta}, \Delta^{\arg\min_{\theta} g_1(x_1^{\theta}, \boldsymbol{\Delta}^{\theta})}), \text{ if } \min_{\theta} g_1(x_1^{\theta}, \boldsymbol{\Delta}^{\theta}) < \bar{b} \\ \begin{array}{l} g_2(x_2^{L}, \Delta^{L}) \text{ with probability } \frac{1}{2} \\ g_2(x_2^{H}, \Delta^{H}) \text{ with probability } \frac{1}{2} \end{array}, \text{ if } \min_{\theta} g_1(x_1^{\theta}, \boldsymbol{\Delta}^{\theta}) \geq \bar{b} \end{cases}, \end{aligned}$$

where the superscript c indicates the presence of the bid ceiling. Notice that, in the first round, a lottery between the two borrowers only takes place when both borrowers' willingness to pay exceeds $\bar{b}$. If only one borrower is willing to pay more than $\bar{b}$, the other will not choose to join the lottery once bidding has reached $\bar{b}$. Similarly, the saver will never join a lottery though he will drive the winning bid up to $\bar{b}$ if at least one of the borrowers is willing to pay that amount.

Testable Implications

The bid ceiling makes it harder for riskier participants to get their hands on the first pot when adverse selection persists despite the collateral requirement (condition 5). Let us illustrate the argument with Figure 2. Let $\widehat{x}^{\theta}$ be the utility shock that makes type θ's willingness to pay for the first pot exactly equal to the bid ceiling, i.e.

$$g_1(\widehat{x}^{\theta}, \mathbf{\Delta}^{\theta}) = \bar{b},\ \theta \in \{L, H\}.$$

Provided that $x_1^H \leq \widehat{x}^H$ or $x_1^L \leq \widehat{x}^L$, the identity and hence the type of the winner will be the same as in an auction without a ceiling, but not otherwise. When $x_1^H > \widehat{x}^H$ and $x_1^L > \widehat{x}^L$, a lottery with equal odds determines which one of the borrower gets the first pot while, with no ceiling, the H type is more likely to win.

This is made precise in Figure 3. Panel A depicts a situation with adverse selection and no bid ceiling. As each bidder's willingness to pay is distributed uniformly, the probability of the H type winning the first auction is

$$\frac{\text{the area of the polygon } ABCDF}{\text{the area of the square } ABCE}$$

Panel B depicts the analogous situation with a bid ceiling of $\bar{b}$ in place. Here a lottery determines the winner when the willingness to pay are in the rectangle $IGCH$, which corresponds to $x_1^H > \widehat{x}^H$ and $x_1^L > \widehat{x}^L$. As the H type wins the lottery with probability one half, we may depict the corresponding area by the triangle IGC. Thus the total probability of the H type winning the first pot is

$$\frac{\text{the area of the polygon } ABCIF}{\text{the area of the square } ABCE}$$

In this example, the ceiling reduces the probability of the high risk type winning the first auction by the triangle ICD relative to the square $ABCE$.

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. Let $E(\theta_1)$ and $E(\theta_1^c)$ be the expected riskiness of the winner of the first pot without and with a ceiling in place. Then the expected riskiness falls as a result of the bid ceiling, $E(\theta_1) > E(\theta_1^c)$.

In contrast consider a situation where collateral eliminates adverse selection (condition 6), as illustrated in Figures 4 and 5. If both borrowers receive a utility shock $x_1^H \geq x^*$ and $x_1^L \geq x^*$, then there is a lottery just as before in the presence of a bid ceiling. On average, the riskier borrower has no higher willingness to pay than the safe borrower as in Figure 4. The expected riskiness of the date 1 winner is unaffected by the bid ceiling. In Figure 5 this results in the points C and D of Figure 3 coinciding. So the probability of the high risk type winning the first auction without and with a ceiling in place corresponds to the area of the triangle ABC relative to the square $ABCE$ which is of course precisely one half. Hence, in this case expected riskiness of borrowers is equalized, $E(\theta_1) = E(\theta_1^c)$.

The main implication of our model is that the bid ceiling makes the average riskiness of participants more equal over time if there is adverse selection but not otherwise. 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 if adverse selection persists despite the collateral requirement. Recall that $E(\theta_t)$ and $E(\theta_t^c)$ refers to expected riskiness of date t borrowers before and after the ceiling is imposed, where $t = 1$ are early borrowers and $t = 2$ are late borrowers.

Proposition 2 ***(Testable Implication)***

(*i*) *If there is adverse selection, as in (5), 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

$$E(\theta_1 - \theta_1^c) - E(\theta_2 - \theta_2^c) > 0 \tag{8}$$

(*ii*) *If there is no adverse selection, as in (6), then there is no difference in 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

$$E(\theta_1 - \theta_1^c) - E(\theta_2 - \theta_2^c) = 0 \tag{9}$$

(iii) *If collateral is an excessive deterrent, as in (7), then early borrowers are safer before the ceiling compared with after, and late borrowers are riskier 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 negative.

$$E(\theta_1 - \theta_1^k) - E(\theta_2 - \theta_2^k) < 0 \tag{10}$$

In other words, when there is adverse selection despite the collateral requirement, 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 6. In contrast, when there is no adverse selection (or collateral is effective at equalizing the costs of default for risky and safe types), the risk profile is completely flat with unrestricted as well as with restricted bidding. The risk profile is unaffected by the policy shock as in Figure 7. So our empirical strategy (which we discuss in detail in section 5) will be to take the difference between the slopes of the riskiness profile of borrowers before and after the ceiling is imposed. That difference is positive only with adverse selection (as in Figure 6 and equation 8).

Our null hypothesis that willingness to pay is unrelated to riskiness will be (9). Note that we will conduct a two sided test of the null of no adverse selection. Effectively we are testing for either of two alternatives (a) riskier borrowers have a higher willingness to pay than safer borrowers because collateral is an insufficient deterrent (equation 8) or (b) riskier borrowers have a lower willingness to pay than safer borrowers because collateral is an excessive deterrent (equation 10). In the latter case, the Rosca organizer has simply set the collateral requirement such that there is a steepening of the risk profile as a consequence of the ceiling. This is certainly an empirical possibility – collateral may sometimes be "too effective" at deterring risk.

5 Identification

Our goal is to test for self selection based on riskiness in these Roscas. Such self selection implies that the policy shock will have a flattening effect on the riskiness profile of Rosca borrowers (Proposition 2). As 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 borrowers have a higher willingness to pay than safer borrowers, 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

$$y_{ti} = \alpha_t + \xi \; after_i + \beta \; after_i \; late_t + u_{ti}, \tag{11}$$

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. If risk and willingness to pay are positively related then the ceiling flattens the default profile and the double difference $\beta > 0$ as in equation (8) in Proposition 2. If unobserved risk and willingness to pay are negatively related then there is steepening of defaults and $\beta < 0$ as in equation (8). Finally, if unobserved risk and willingness to pay are unrelated then there is no flattening or steepening of the default profile. This is the null hypothesis in equation 9.

The critical question for identification 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^u and measured with error. Assume the benchmark default generating process is:

$$y_{ti} = \delta_t + \lambda_t \theta^u_{ti} + \upsilon_{it} \tag{12}$$

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

If the data is indeed generated by (12) then β is a consistent estimator of the double difference in riskiness:

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

where $E\theta_1^u - E\theta_1^{uc}$ is the expected difference in riskiness for early rounds and $E\theta_2^u - E\theta_2^{uc}$ is the expected difference in riskiness for late rounds (the superscript c refers to Roscas that are restricted by the bid ceiling). Recall that, according to Proposition 2, flattening through a ceiling on bids decreases the average riskiness of early borrowers, $E\theta_1^u > E\theta_1^{uc}$, and increases the average riskiness of late borrowers $E\theta_2^u < E\theta_2^{uc}$. So under the null, $\lim_{N\to\infty} \widehat{\beta} = 0$ but under the alternative $\lim_{N\to\infty} \widehat{\beta} > 0$. For this admittedly unrealistic benchmark case, the test for adverse selection is consistent.

What if the collateral requirement is enough to deter unobservably riskier people from taking early pots? If condition (6) holds, then there is no adverse selection – the null hypothesis. Recall in this connection that, theoretically, there could be no adverse selection either because markets (or collateral) work efficiently to overcome the adverse selection problem or because borrowers do not differ in their unobserved riskiness even in the absence of collateral.

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 (12), the estimated double difference of defaults in specification (11) 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 (11) may be inconsistent. In what follows we discuss how we augment the specification (11) so that we can identify adverse selection when the defaults are not generated by (12).

Observed Riskiness

In practice defaults are certainly not generated only by unobserved riskiness as we assumed in (12). In this subsection we shall first pose the problem – the unconditional double difference of defaults in specification (11) picks up both observed and unobserved riskiness. Then we shall discuss a solution altering our estimation strategy to account for differences in observed riskiness.

Assume that riskiness comprises both observed and unobserved components, $\theta_{ti} = \theta^u_{ti} + \theta^o_{ti}$. Suppose defaults depend only on riskiness

$$y_{ti} = \delta^u_t + \lambda^u_t \theta_{ti} + \upsilon_{it}, \tag{13}$$

then if we run the benchmark econometric specification (11), there will be an omitted variable bias. We will pick up both the double difference in observed and the double difference in unobserved riskiness,

$$\begin{aligned} \lim_{N\to\infty} \widehat{\beta} &= \lambda^u_1(E\theta_1 - E\theta^c_1) - \lambda^u_2(E\theta_2 - E\theta^c_2) \qquad (14)\\ \lim_{N\to\infty} \widehat{\beta} &= [\lambda^u_1(E\theta^u_1 - E\theta^{uc}_1) - \lambda^u_2(E\theta^u_2 - E\theta^{uc}_2)] + [\lambda^u_1(E\theta^o_1 - E\theta^{oc}_1) - \lambda^u_2(E\theta^o_2 - E\theta^{oc}_2)] \end{aligned}$$

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 (11). 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 and, according to the organizer, does depend on all observed factors relevant for expected defaults. We next show that the double difference of defaults conditional on collateral required can isolate unobserved riskiness.

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

$$c_{ti} = \delta^o_t + \lambda^o_t \theta^o_{ti} \tag{15}$$

Substituting (15) into (13) and solving gives a default generating process that depends on unobserved riskiness and the collateral requirement (eliminating the observed riskiness term),

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

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,

$$y_{ti} = \alpha_t + \xi \ after_i + \beta \ after_i \ late_t + \eta_t \ c_{ti} + u_{ti}. \tag{16}$$

Here β estimates the double difference in defaults conditional on the cosigner requirement,

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

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 2007). 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

$$y_{ti} = \phi_{t0} + \lambda_t^u \theta_{ti} + \phi_{t1} b_{ti} + \phi_{t2} q_{ti} + \phi_{t3} c_{ti} + \upsilon_{it}, \tag{17}$$

then a test that rejects the null hypothesis of no adverse selection on the basis of specification (16) 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 in 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 (15) to

$$c_{ti} = g_{t0} + \lambda_t^o \theta_{ti}^o + g_{t1} b_{ti} + g_{t2} q_{ti} \tag{18}$$

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 (18) into (17) and solving to give

$$y_{ti} = \phi'_{t0} + \lambda_t^u \theta_{ti}^u + \phi'_{t1} b_{ti} + \phi'_{t2} q_{ti} + \phi'_{t3} c_{ti} + \upsilon_{it} \tag{19}$$

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 (19), we can augment specification (16) 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

$$y_{ti} = \alpha_t + \xi \ after_i + \beta \ after_i \ late_t + \eta_t \ c_{ti} + \psi_t b_{ti} + \zeta_t q_{ti} + u_{ti}, \tag{20}$$

Next we show that $\lim_{N\to\infty} \widehat{\beta} = 0$ under the null of no adverse selection. 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

$$e'\left\{\lambda_1^u\left(var\begin{bmatrix}After\\B_1\\Q_1\\C_1\end{bmatrix}\right)^{-1}cov\left(\begin{bmatrix}After\\B_1\\Q_1\\C_1\end{bmatrix},\widetilde{\theta}_1^u\right)-\lambda_2^u\left(var\begin{bmatrix}After\\B_2\\Q_2\\C_2\end{bmatrix}\right)^{-1}cov\left(\begin{bmatrix}After\\B_2\\Q_2\\C_2\end{bmatrix},\widetilde{\theta}_2^u\right)\right\}, \tag{21}$$

where $e' = [1\ 0\ 0\ 0]$, $\widetilde{\theta}_t^u$ denotes the unobserved riskiness of a round t borrower as a random variable. Under the null hypothesis, $\widetilde{\theta}_t^u$ has no variation, as all borrowers are of identical unobserved riskiness. So the two covariance terms in (21) are vectors of zeros which implies $\lim_{N\to\infty}\widehat{\beta}=0$.

Rosca Composition

In our data, individuals are not randomly assigned into unrestricted and restricted groups. Instead, only unrestricted Roscas were started before September 30, 1993, and only 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.[7]

[7] Eeckhout and Munshi 2004 find that borrowers and savers sort themselves across Rosca denominations before and after the policy shock in predictable ways. Since participants have no default risk both in their

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 willingness to pay according to inherent riskiness between Rosca 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, however.

In terms of our 3 period model, the two borrowers have riskiness $\Delta^{1,H} = \theta^{1,H} - \pi^{1,H} = \theta^{1,L} - \pi^{1,L} = \Delta^{1,L}$ before, and $\theta^{2,H} - \pi^{2,H} = \theta^{2,L} - \pi^{2,L} = \Delta^2$ after the policy shock, where potentially $\theta^{1,H} \neq \theta^{2,H}$ and $\theta^{1,L} \neq \theta^{2,L}$. Notice that in this case we have $E\theta_1 = E\theta_2 = (\theta^{1,L} + \theta^{1,H})/2 \equiv \theta^1$ and $E\theta_1^c = E\theta_2^c = (\theta^{2,L} + \theta^{2,H})/2 \equiv \theta^2$. 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 (21), which applies if (as before) defaults are generated by (19).

Proposition 3 ***(Change in Rosca Composition)*** *If there is no adverse selection in Roscas with and without a bid ceiling, i.e.* $\Delta^{1,L} = \Delta^{1,H}$ *and* $\Delta^{2,L} = \Delta^{2,H}$*,*
(i) $\lim_{N\to\infty} \widehat{\beta} = (\lambda_1^u - \lambda_2^u)(\theta^1 - \theta^2)$;
(ii) $\left(\lim_{N\to\infty} \widehat{\xi}\right) \left(\lim_{N\to\infty} (\widehat{\xi} + \widehat{\beta})\right) = \lambda_1^u \lambda_2^u (\theta^1 - \theta^2)^2 > 0.$

Part (i) of Proposition 3 gives conditions for when our test will inconsistent when if is no adverse selection and only a change in levels of riskiness. Whether the limit of $\widehat{\beta}$ 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, as is the case in our theoretical model, where $\lambda_1^u = \lambda_2^u = 1$. If there are round-specific differences in how defaults depend on riskiness then our test will be inconsistent.

Part (ii) of Proposition 3 states a property of the coefficients ξ and β in estimating equation (20) when there is no adverse selection and only a change in average riskiness.

theory and empirics, however, their results are not directly applicable to our study.

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. This is illustrated in Figure 8, Panel B. With adverse selection and no change in average riskiness, on the other hand, the default profiles will intersect (Figure 8, Panel A), which corresponds to the situation in Figure 6. 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 $\text{sign}(\xi) = \text{sign}(\xi + \beta)$ or $\xi(\xi + \beta) > 0$. If the signs are equal then there is no intersection of default profiles while if the signs are different, the null hypothesis of no adverse selection can be confidently rejected. Notice that this test is conservative as it will also fail to reject if there is adverse selection but, simultaneously, a sufficiently large change in the average riskiness of Rosca participants such that the two default profiles do not intersect.

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 (20) 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 (20) 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 (20) as

$$y_{ti} = \alpha_t + \xi \ after_i + \beta \ after_i \ late_t + \eta_t \ c_{ti} + \psi_t b_{ti} + \zeta_t q_{ti} + \gamma_j \ quarterj_{ti} + u_{ti} \qquad (22)$$

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

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 (22).

Notice that we are assuming that transitory aggregate shocks affect defaults identically in different rounds of the Rosca. If 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 motivate our regression analysis with evidence from simple means comparisons. We proceed to discuss how to partition Roscas into early and late rounds. We then describe our main results testing for adverse selection and examine their robustness.

Simple Differences in Differences

We are interested in how 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.[8] A dramatic change occurred in the timing of defaults, however (see Table 2). 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).

As we noted in section 5, however, this double difference does not ensure identification of adverse selection. For instance, the double difference in the winning bids is also positive and significant indicating better loan terms for early recipients post-ceiling. Testing carefully for adverse selection will involve conditioning on cosigner requirement, winning bid and repayment burden, and allowing for aggregate shocks as in specification (22). Before doing so we shall briefly discuss how we will implement specification (22) empirically.

Early vs. Late Partitions

We showed that the specification (22) ensures identification of adverse selection. In this section, we adapt specification (22) 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,

$$y_{kti} = \alpha_{kt} + \xi_k \, after_i + \beta \, after_i \, late^{\tau}_{kt} + \eta_{kt} \, c_{kti} + \gamma_j \, quarterj_{kti} + \psi_{kt} b_{kti} + \zeta_{kt} q_{kti} + u_{kti}. \quad (23)$$

Notice that, except for the aggregate shock terms, we incorporate denomination-specific and round-specific controls for collateral c_{kti}, bid b_{kti} and repayment burden q_{kti} throughout.[9]

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

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

[9] In principle, could estimate (23) 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 (23) 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.

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, $late_t^{\tau}$ 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 later two thirds of a Rosca. Formally, $late_{kt}^{1/3} = 0$ for $t \leq T_k/3$ and $late_{kt}^{1/3} = 1$ for $t > T_k/3$.

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, i.e. $late_{kt}^{1/2} = 0$ for $t \leq T_k/2$ and $late_{kt}^{1/2} = 1$ for $t > T_k/2$.

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 (23) the estimates α_{kt} capture the trend of the default profile pre-ceiling and the estimated $\alpha_{kt} + \beta \; after_{ki} \frac{t}{T_k}$ 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

We turn to our main results. Recall that we are testing if the bid ceiling flattens the riskiness profile. Given that we only observe defaults (and not riskiness directly), we estimate a conditional double difference of defaults (specification 23) and contrast it with an unconditional double difference (benchmark specification 11).

The OLS results for the double difference in defaults (estimate of β) with and without controls are in Table 3. The specification 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

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

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 Relative Rounds specifications, again first without and then with controls. We use two alternative measures for the collateral requirement – the incidence of a cosigner (whether any cosigner was required or not) and the number of cosigners (which ranges from 0 to 5). Each set of controls is jointly significant, except for the number of cosigners (for which the F test fails to reject in all specifications). However, only the winning bid has a significant average effect. This positive average effect implies that 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, else we would have found a positive average effect. The cosigner incidence, repayment burden and aggregate shock controls have different signs in different rounds and denominations (not shown in the table). So we reject on the F-test, but average effects are not significantly different from zero.

Our main result is that the double difference $\widehat{\beta}$ is positive and (mostly) significant for all three specifications of the adverse selection effect and is remarkably robust to the introduction of controls for cosigners, winning bids, repayment burden and aggregate shocks. As expected, for each of the three early-late partitions, the standard errors for the estimated $\widehat{\beta}$ substantially increase when the controls are added. The point estimates, however, remain unchanged in sign and similar in magnitude to those without the controls. This establishes that riskier participants are willing to bid higher than safer participants are, despite the cosigner requirement, as in part (i) of Proposition 1. The effect of such self selection on unobserved riskiness is substantial: the point estimate of the flattening effect $\widehat{\beta}$ (in column 2 of Table 3) is about two-fifth of the mean default rate (column 1; Table 2). This can be interpreted as a one-fifth fall in early defaults and a one-fifth increase in late defaults.

Changes in Rosca Composition

One concern raised in section 5 is that the flattening of the default profile as measured by β in specification (23) may be due to a change in the composition of Rosca participants in response to the policy shock even when there is no selection on unobserved riskiness. Recall, however, that no adverse selection together with a change in the average riskiness of borrowers implies that all borrowers, both early and late, become either more or less risky in response to the policy shock. As we have shown in section 5, this implies that ξ and $(\xi + \beta)$ in

$$y_{kti} = \alpha_{kt} + \xi \; after_i + \beta \; after_i \; late^{\tau}_{kt} + \eta_{kt} \; c_{kti} + \psi_{kt} b_{kti} + \zeta_{kt} q_{kti} + \gamma_j \; quarterj_{kti} + u_{kti} \quad (24)$$

have the same sign. We thus conduct non-linear Wald tests of the hypothesis $\xi(\xi + \beta) \geq 0$ for different versions of equation (24). The alternative, $\xi(\xi + \beta) < 0$ is, of course, one-sided. Notice that equations (24) and (23) are identical except for the parametrization of the $after$ dummy. While we allow for a denomination-specific effect in (23), there is a uniform effect in (24), which is needed to conduct the Wald test in a tractable way.

The results are in Table 4. For all three partitions, we find indeed that early defaults decreased while late defaults increased, although not significantly so. Moreover, for the Relative Round early/late partition, we reject the null hypothesis of (simultaneously) no adverse selection and a change in average riskiness in response to the policy shock at the 90% significance level. We fail to reject for the other two partitions at conventional significance levels, however.

Another informal way to jointly assess the three outcomes is to consider the sum of the three Wald test statistics. If the three tests were from independent trials (which of course they are not), the sum of -2.34 would correspond to a $p-$value of 8.8 percent.

To summarize, there is suggestive evidence against a change in the composition of Rosca participants and no adverse selection in all three specifications considered here, with one specification giving a rejection which is statistically significant. Given the conservative nature of this test (see the discussion in Section 5), we believe that our results do reflect self-selection on unobserved riskiness and not changes in composition.

7 A Policy Reversal in 2002

In this section we conduct a robustness check. We test if a relaxing of the bid ceiling in 2002 led to the opposite effect on default profiles. In other words, we test if relaxing the bid ceiling made it easier for riskier participants to take early pot. We do find evidence of an "un-flattening" of the riskiness profile. In this section we shall first describe the data from the 2002 policy reversal. Then we shall discuss the results.

Data

The Rosca organizer lifted the bid ceiling on Roscas that started after January 1, 2002, from 30 to 40 percent of the pot in response to an Amendment to the Chit Fund Act passed by the Indian government. This was part of a general move towards financial liberalization in India. We collected data from the Rosca organizer on November 30, 2005 nearly three years after the 2002 policy shock. The organizer records arrears at the time Roscas are completed as the fraction of payments still outstanding from a borrower. Over time these arrears are adjusted downwards as legal action and other pressure from the organizer results in the further collection of repayments. When these arrear rates stabilize several years after the completion of the Rosca, we refer to them as defaults. Sufficient time has not elapsed since the 2002 policy shock for us to use eventual default rates of Rosca borrowers. For this reason we will use arears while analyzing the lifting of the bid ceiling.[10] Arrears are of course highly correlated with eventual defaults.

We constructed the 2002 sample along the same lines as the sample of Roscas we used for our analysis of the 1993 policy shock. So we included Roscas that were:

1. Started between January 1, 2001 and December 31, 2002
2. Completed and in the organizer's database at the time of data collection
3. Of a denomination for which there were at least 35 Roscas started between January 1, 2001 and December 31, 2002.

This leaves us with 12 denominations for the 2002 sample (Table 5). The pot sizes are (not surprisingly) considerably larger than with the 1993 sample. Just as before, we eliminated institutional investors, leaving us with $60,312$ observations. Table 5 shows that there was an increase in Roscas of all denominations as a result of the 2002 policy shock. As expected, the 30% bid ceiling binds for a large fraction of early rounds but the 40% bid ceiling binds for a substantially smaller fraction of early rounds. For instance, for the popular (40 round, Rs. 250) contribution, 54.7 percent of winning bids were at the Rs. 3000 ceiling (30% of the pot) before the policy reversal in the first half of Roscas. Only 21.8 percent were at the Rs. 4000 ceiling in early rounds after the policy reversal. A substantial fraction of the winning bids (65.6 percent) were between Rs. 3000 and Rs. 4000 in the early rounds after the policy reversal.

Table 6 shows that the arrear rates in the 2002 sample were 13.1 percent on average, over ten times as high as the default rates in the 1993 sample (Table 2). The incidence of the cosigner requirement and the number of cosigners required were also higher in 2002 compared with 1993. This presumably reflects the increased pot sizes (and correspondingly higher loans).

[10] It would be natural to think of the imposition of the bid ceiling in 1993 and the relaxation in 2002 as one large natural experiment with three time periods. Since our dependent variable differs between the initial policy shock and the reversal, however, we cannot do so.

Results

The theory predicts that lifting the bid ceiling in 2002 will "steepen" the risk profile. In other words, variations in riskiness across Rosca participants will be reflected in an increase in the difference between arrear rates of early and late borrowers after the ceiling is lifted. The double difference in arrears of -1.72 is negative and significant (Table 6). Arrear rates were flatter before the 2002 policy shock compared with after.

As we discussed in Section 5, unconditional difference-in-differences of arrear rates do not by themselves indicate variations in inherent riskiness across borrowers. For instance, loan terms also became less favorable for early borrowers relative to late as a result of the bid ceiling (the average winning bids increased for early borrowers but dropped for late borrowers in Table 6). That could indicate an increase in moral hazard propensity for early borrowers relative to late as as a consequence of the policy reversal.

So in Table 7 we estimate adverse selection effects including controls for moral hazard-mechanical effects and for aggregate shocks. Our identification strategy is the same as for the 1993 policy shock. For specification (23), we would now expect negative estimates for β indicating that riskiness profile before the policy reversal was flatter than after the policy reversal. This is broadly what we find in Table 7 in contrast with Table 3 (where the estimates of β were positive from the imposition of the bid ceiling). We find a significant negative estimate of β for one of the three specifications with controls (column 4, Table 7) and a barely insignificant negative estimate of β (column 6). The relaxation of the bid ceiling was only partial (from 30% to 40% of the pot size) – and so the ceiling continued to bind in the earliest rounds of post 2002 Roscas. It is no surprise then that the partition "Rounds in Later Two Thirds" yields an insignificant estimate of β (in column 2). Since the relaxation of the bid ceiling was most effective later in the first half (but not in the first third) of Roscas, we would expect a smaller and less significant estimate of steepening in the later two thirds (column 2) compared with the second half (column 4).

The economic magnitude on steepening of arrears in 2002 is smaller than the economic magnitude of the flattening was (in 1993). Relative to average arrears the estimated double difference in columns 4 and 6 of Table 7 are 8.2 and 6.2 percent only. while these ratios were much higher in the 1993 default data. There are two possible reasons. First, the 2002 reversal was less dramatic than the imposition of the bid ceiling in 1993. Secondly, the organizer's collection of claims is a lengthy process continuing well past the time of data collection. Some arrears then are probably delays arising from short-term shocks to borrowers, rather than systematic differences in riskiness relevant for eventual defaults.

Finally we discuss the estimates on the controls in Table 6. In contrast with Table 3, the number of cosigners is significantly positively associated with arrear rates (when the number of cosigners exceeds 1). This implies than an increase in the number of cosigners does not fully compensate for the additional arrear risk of observationally riskier borrowers. As before the winning bid has a positive and significant average effect on arrear rates.

8 Conclusion

There is a long theoretical literature starting with Akerlof (1970) on adverse selection as an impediment to efficient trades. In this paper we have used a natural experiment to investigate adverse selection in credit markets. 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. We are thus able to identify a significant and strong adverse selection effect. Further, we find the opposite effect when the bid ceiling was partially lifted in 2002.

This paper highlights three methodological issues.

1. Loan officers often have private information on borrowers such as estimates of their ex ante riskiness. The researcher does not have access to the lender's soft information about the borrowers. Such soft information can confound tests for information asymmetries in credit markets. We propose a novel solution. The loan officer reveals his private and soft information on ex ante riskiness of borrowers by his decision on how

much collateral to ask them for. So controlling for collateral required can account for information asymmetries between the researcher and the loan officer.

2. Stiglitz and Weiss (1981) predict a positive correlation between willingness to pay and unobserved riskiness in a credit market model of adverse selection . An analogous positive correlation property has also been the basis of numerous tests of asymmetric information in the insurance market (Chiappori et al, forthcoming). In this paper we show theoretically that positive correlation between willingness to pay high interest rates and riskiness in bidding Roscas can be eliminated or even reversed with collateral. Then we find empirically that the positive correlation persists despite the use of cosigners as collateral in the bidding Roscas we study.

3. Adverse selection is often difficult to disentangle from moral hazard and other mechanical effects. This problem has been discussed in the context of insurance markets (Chiappori and Salanie 2003) and credit markets (Karlan and Zinman, 2007). In this paper we isolate adverse selection by controlling for loan terms before and after the bid ceiling.

Finally, our research has specific relevance for credit constraints and microfinance in developing countries (Armendariz and Morduch, 2005). The organized Roscas we study are similar to the Grameen Bank model of microfinance. Both are forms of financing that use local information in the absence of traditional collateral. The organizer of Roscas makes cosigners liable for repayments of borrowers while the Grameen Bank gives group loans. We find that adverse selection persists despite cosigned loans; Ahlin and Townsend (2007) find that adverse selection persists despite group loans.

9 Appendix

Derivation of the Willingness to Pay Functions $g_1(x, \mathbf{\Delta}^\theta)$, $g_2(x, \Delta^\theta)$

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. In the second round, the expected utility of a borrower of type θ with a second period shock of x from winning at date 2 is

$$\frac{b_1}{3} + \delta\left(x - \frac{2}{3}kb_2 - \delta\pi^\theta\right) + \delta^2\theta. \tag{25}$$

$b_1/3$ is her share in the first period's winning bid, $x - \frac{2}{3}b_2k$ is her payoff from winning the second auction while $\delta\pi^\theta$ is the cost of furnishing a cosigner in period 2 who guarantees for a liability of \$1 one period later. Finally, in the third period, θ is that borrower's expected payoff from defaulting and consuming her unit income. If that borrower loses the second auction, on the other hand, her expected utility is

$$\frac{b_1}{3} + \delta\frac{1}{3}b_2 + \delta^2\mu. \tag{26}$$

Define $\Delta^\theta = \theta - \pi^\theta$. The willingness to pay for the second pot of a borrower of type θ, $g_2(x, \Delta^\theta)$, is obtained from equating expressions (25) and (26) and solving for b_2,

$$g_2(x, \Delta^\theta) = \frac{3}{1+2k}\left[x - \delta(\mu - \Delta^\theta)\right]. \tag{27}$$

As the losing bidder in this auction has an incentive to drive up the winning bid as far as possible, $g(x_2^\theta, \Delta^\theta)$ will be the winning bid if the other bidder is the saver. Notice that this willingness to pay is increasing in the difference $\theta - \pi^\theta$.

We denote by θ' the type complementary to type θ, i.e. $\theta' = H$ if $\theta = L$ and $\theta' = L$ if $\theta = H$. Working backwards, a borrower of type θ who observes a first-period shock of x and wins the first pot, has expected utility of

$$\left\{x - \frac{2k}{3}b_1 - \pi^\theta\left[\delta\left(1 - \frac{1}{3}E(g_2(X, \Delta^{\theta'})\right) + \delta^2\right]\right\} + \delta\left\{\theta + (1-\theta)\frac{1}{3}E(g_2(X, \Delta^{\theta'}))\right\} + \delta^2\theta, \tag{28}$$

where the expectations are taken with respect to X, which is distributed as $U\left(\mu - \frac{\alpha}{2}, \mu + \frac{\alpha}{2}\right)$. In the first period, the borrower gains an instantaneous payoff of x less the cost of external finance $\frac{2}{3}b_1k$ and the cost of collateral. As the expected liability of the winner of the first pot is $1 - \frac{1}{3}E_X(g_2(X, \Delta^{\theta'})$ in the second and 1 in the third period, the cost of collateral equals π^θ times the expression in brackets. In the second and third period, with probability θ, the winner of the first pot keeps her income, while with probability $(1-\theta)$ she pays contributions in both subsequent periods and, at date 2, receives a one third share of the winning bid.

If a borrower of type θ loses the first, but wins the second auction, her expected utility is

$$\frac{b_1}{3} + \delta \left[\mu - \frac{2k}{3} E(g_2(X, \Delta^\theta)) - \delta \pi^\theta \right] + \delta^2 \theta. \tag{29}$$

Define $\mathbf{\Delta}^\theta \equiv (\Delta^\theta, \Delta^{\theta'})$. Equating (28) and (29) gives the willingness to pay for the first pot for a borrower of type θ with utility shock x,

$$g_1(x, \mathbf{\Delta}^\theta) = \frac{3}{1+2k} \left\{ x - \delta \left[\mu - \Delta^\theta + \frac{1}{3}(1 - \Delta^\theta) E\left(g_2(X, \Delta^{\theta'}) \right) + \frac{2k}{3} E(g_2(X, \Delta^\theta)) \right] \right\} \tag{30}$$

Notice that assumption 1 implies that $g_1(x, \mathbf{\Delta}^\theta)$ is positive for all x and θ because $L \leq H \leq 1$, $\pi^\theta \geq 0$, and so $\Delta^\theta = \theta - \pi^\theta \leq 1$ for all θ.

Proof of Proposition 1

Using (30) and (27) to calculate $E(g_2(X, \Delta^H))$ and $E(g_2(X, \Delta^L))$, we have that

$$g_1(x, \mathbf{\Delta}^H) - g_1(x, \mathbf{\Delta}^L) = (\Delta^H - \Delta^L) \frac{3\delta}{(1+2k)^2} \left[2(1+\delta)k - (1-\delta)(\mu - 1) \right].$$

The fraction $3\delta/(1+2k)^2$ is clearly positive. To determine the sign of the term in brackets, we use the upper bound on μ from (3). This yields

$$2(1+\delta)k - (1-\delta)(\mu - 1) \geq 2(1+\delta)k - (1-\delta)2k,$$

which is clearly greater than zero. Thus $g_1(x, \mathbf{\Delta}^H) - g_1(x, \mathbf{\Delta}^L)$ is proportional to $(\Delta^H - \Delta^L)$ with a positive factor of proportionality. This immediately establishes all three parts of the proposition.

Proof of Proposition 2

Denote by $\Pr_t(\theta)$ the probability that the borrower of type θ wins the first auction in a Rosca with no ceiling, and by $\Pr_t^c(\theta)$ the corresponding probability in a Rosca with a bid ceiling. We can write

$$E(\theta_1 - \theta_1^c) = (\theta^H - \theta^L) \left[\Pr_1(H) - \Pr_1^c(H) \right]. \tag{31}$$

Define $\underline{x} = \mu - \alpha/2$, $\overline{x} = \mu + \alpha/2$ and assume for the moment that $\Delta^H \geq \Delta^L$. Together with assumption 3 this implies that $g_1(\underline{x}, \mathbf{\Delta}^L) \leq g_1(\underline{x}, \mathbf{\Delta}^H) < \bar{b} < g_1(\overline{x}, \mathbf{\Delta}^L) \leq g_1(\overline{x}, \mathbf{\Delta}^H)$. We may now write

$$\begin{aligned} \Pr_1(H) \quad = \quad & \frac{1}{2} \Pr\left[g_1(X, \mathbf{\Delta}^H) \leq g_1(\overline{x}, \mathbf{\Delta}^L)\right] \Pr\left[g_1(X, \mathbf{\Delta}^L) > g_1(\underline{x}, \mathbf{\Delta}^H)\right] \\ & + \Pr\left[g_1(X, \mathbf{\Delta}^L) \leq g_1(\underline{x}, \mathbf{\Delta}^H)\right] \Pr\left[g_1(X, \mathbf{\Delta}^H) \leq g_1(\overline{x}, \mathbf{\Delta}^L)\right] \\ & + \Pr\left[g_1(X, \mathbf{\Delta}^H) > g_1(\overline{x}, \mathbf{\Delta}^L)\right]. \end{aligned} \tag{32}$$

The first line of the term on the right hand side captures the intersection of the domains of the random variables $g_1(X_1^L, \mathbf{\Delta}^L)$ and $g_1(X_1^H, \mathbf{\Delta}^H)$. As, conditional on this event, each individual is equally likely to have a higher willingness to pay, H wins in exactly half of those cases. The second and third lines of the right hand side capture the event where L has a willingness to pay smaller than the lower bound of $g_1(X_1^H, \mathbf{\Delta}^H)$'s distribution. In all of those cases, H wins with probability one.

When there is a ceiling in place we have that

$$\begin{aligned} \Pr_1^c(H) \quad = \quad & \frac{1}{2} \Pr\left[g_1(X, \mathbf{\Delta}^H) \leq g_1(\overline{x}, \mathbf{\Delta}^L)\right] \Pr\left[g_1(X, \mathbf{\Delta}^L) > g_1(\underline{x}, \mathbf{\Delta}^H)\right] \\ & + \Pr\left[g_1(X, \mathbf{\Delta}^L) \leq g_1(\underline{x}, \mathbf{\Delta}^H)\right] \Pr\left[g_1(X, \mathbf{\Delta}^H) \leq g_1(\overline{x}, \mathbf{\Delta}^L)\right] \\ & + \Pr\left[g_1(X, \mathbf{\Delta}^H) > g_1(\overline{x}, \mathbf{\Delta}^L)\right] \left\{ \Pr\left[g_1(X, \mathbf{\Delta}^L) \leq \bar{b})\right] + \frac{1}{2} \Pr\left[g_1(X, \mathbf{\Delta}^L) > \bar{b})\right] \right\}. \end{aligned} \tag{33}$$

Regarding the first line of the right hand side of this equation, notice that in comparison to equation 32 the ceiling does not change the probability of one half, with which H wins the first pot. Regarding the third line of the right hand side, when H's willingness to pay exceeds $g_1(\overline{x}, \mathbf{\Delta}^L)$, she wins for sure only when L's willingness to pay is smaller than $\bar{b}$. Otherwise there will be a lottery, which H wins with probability one half.

Substituting $\frac{1}{2}(1+\Pr\left[g_1(X, \mathbf{\Delta}^L) \leq \bar{b})\right])$ for $\Pr\left[g_1(X, \mathbf{\Delta}^L) \leq \bar{b})\right] + \frac{1}{2}\Pr\left[g_1(X, \mathbf{\Delta}^L) > \bar{b})\right]$ in (33), we can take the difference

$$\Pr_1(H) - \Pr_1^c(H) = \frac{1}{2} \Pr\left[g_1(X, \mathbf{\Delta}^H) > g_1(\overline{x}, \mathbf{\Delta}^L)\right] \left\{1 - \Pr\left[g_1(X, \mathbf{\Delta}^L) \leq \bar{b})\right]\right\},$$

which is clearly strictly greater than zero if, and only if $\Delta^H > \Delta^L$. If $\Delta^H = \Delta^L$, on the other hand, $\Pr\left[g_1(X, \mathbf{\Delta}^H) > g_1(\overline{x}, \mathbf{\Delta}^L)\right] = 0$ and $\Pr_1(H) - \Pr_1^c(H) = 0$. Combining this with equation 31 and noticing that $E(\theta_2 - \theta_2^c) = -E(\theta_1 - \theta_1^c)$ establishes parts (i) and (ii) of the proposition.

By symmetry, if $\Delta^L > \Delta^H$, we obtain that $\Pr_1(L) - \Pr_1^c(L) > 0$, which implies $\Pr_1(H) - \Pr_1^c(H) < 0$. This establishes part (iii) of the proposition.

Proof of Proposition 3

For ξ in estimating equation (20), we have that

$$\lim_{N\to\infty} \widehat{\xi} = \lambda_1 e' \left(var \begin{bmatrix} After \\ X_1 \end{bmatrix} \right)^{-1} cov \left(\begin{bmatrix} After \\ X_1 \end{bmatrix}, \widetilde{\theta}_1^u \right), \tag{34}$$

where

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

and the prime superscript denotes the transpose operator.

First notice that

$$var \begin{bmatrix} After \\ X_1 \end{bmatrix} = \begin{bmatrix} var(After) & cov(X_1, After)' \\ cov(X_1, After) & var(X_1) \end{bmatrix}. \tag{35}$$

Moreover, we have that

$$\begin{aligned} cov\left(After, \widetilde{\theta}_1^u\right) &= E(After\ \widetilde{\theta}_1^u) - E(After)E(\widetilde{\theta}_1^u) \\ &= E(After)E(\theta_1^{u,c}) - E(After)\left[(1 - E(After))E(\theta_1^u)\right] + E(After)E\left(\theta_1^{u,c}\right) \\ &= E(After)\left[1 - E(After)\right]\left[E(\theta_1^{u,c}) - E(\theta_1^u)\right] \\ &= V(After)\left(\theta^2 - \theta^1\right) \end{aligned}$$

An analogous argument establishes that

$$cov\left(X_1, \widetilde{\theta}_1^u\right) = cov(After, X_1)(\theta^2 - \theta^1).$$

Taken together, we have

$$cov\left(\begin{bmatrix} After \\ X_1 \end{bmatrix}, \widetilde{\theta}_1^u\right) = (\theta^2 - \theta^1)\begin{bmatrix} var(After) \\ cov(X_1, After) \end{bmatrix} \tag{36}$$

$$= (\theta^2 - \theta^1)\begin{bmatrix} var(After) & cov(X_1, After)' \\ cov(X_1, After) & var(X_1) \end{bmatrix} e.$$

$$cov\left(X_1, \widetilde{\theta}_1^u\right) = cov(After, X_1)(\theta^2 - \theta^1).$$

Substituting identities (35) and (36) into (34) immediately establishes

$$\lim_{N\to\infty} \widehat{\xi} = \lambda_1^u(\theta^2 - \theta^1). \tag{37}$$

For $\beta + \xi$ in estimating equation (20), 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

$$\lim_{N\to\infty} (\widehat{\beta} + \widehat{\xi}) = \lambda_2^u(\theta^2 - \theta^1). \tag{38}$$

Subtracting (37) from (38) immediately establishes part 1 of the proposition. Multiplication of (37) and (38), and the fact that λ_1^u and λ_2^u are both positive by definition establishes part 2 of the proposition.and the prime superscript denotes the transpose operator.

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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	32	84	162.5	3.5	0.0	19.5	0.3
25	400	10,000	92	79	-14.1	26.7	0.0	43.2	0.0
25	1,000	25,000	52	56	7.7	28.4	0.0	56.7	1.0
30	500	15,000	208	240	15.4	42.4	0.0	54.3	0.7
40	250	10,000	864	684	-20.8	79.1	2.2	72.0	3.7
40	500	20,000	52	32	-38.5	79.3	4.4	83.3	6.6
40	625	25,000	202	110	-45.5	70.9	0.6	76.1	5.1
40	1,250	50,000	12	23	91.7	80.6	0.0	84.8	11.9
40	2,500	100,000	24	20	-16.7	73.0	0.0	70.1	8.1
50	1000	50,000	170	86	-49.4	91.8	5.3	94.3	16.3
60	1,250	75,000	30	14	-53.3	98.7	12.9	100.0	31.4

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. Default Rates, Cosigners, Winning Bids and Repayment Burdens, 1993 Sample

Means

	Entire Sample	Before Shock		After Shock		Single Difference		Double Difference
		Early	*Late*	*Early*	*Late*	Early	Late	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Default Rate (percent)	1.23	2.04	0.59	1.64	0.74	0.40	-0.15	0.55
	(7.12)	(9.46)	(4.64)	(8.13)	(5.17)	[0.10]	[0.09]	[0.08]
Cosigner Incidence (percentage)	39.3	42.1	40.3	37.5	36.0	4.59	4.35	0.23
	(48.8)	(49.4)	(49.1)	(48.4)	(48.0)	[0.67]	[0.63]	[0.92]
Number of Cosigners	0.71	0.75	0.74	0.67	0.66	0.08	0.08	0.00
	(1.01)	(1.01)	(1.03)	(0.97)	(0.99)	[1.38]	[1.30]	[1.90]
Winning Bid (percentage of pot)	20.82	35.07	11.51	27.17	10.22	7.90	1.29	6.61
	(13.48)	(11.09)	(6.76)	(5.49)	(6.06)	[1.11]	[1.04]	[0.15]
Repayment Burden (percentage of pot)	42.36	64.33	23.38	63.41	22.84	0.92	0.54	0.38
	(23.50)	(10.12)	(12.98)	(10.46)	(12.73)	[0.16]	[0.15]	[0.22]
Observations	46,269	13,034	13,977	8,744	10,514			

Notes: Standard deviations in parentheses for columns 1-5; Standard errors in brackets for columns 6-8.
Column 6 reports the difference between the values in columns 2 and 4, column 7 the difference between columns 3 and 5.
Column 8 reports the difference between the values in columns 6 and 7.

Table 3. Selection on Unobserved Riskiness, 1993 Sample
OLS
Dependent Variable: 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)
<u>Controls</u>							
Cosigner Incidence		No	Yes	No	Yes	No	Yes
	F(388, 44293)		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, 44293)		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, 44293)		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, 44293)		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, 44293)		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. The total number of regressors in columns 2, 4 and 6 thus is 1,976.

Table 4. Wald Test for Change in Average Riskiness of Rosca Members, 1993 Sample

Specification of Late Rounds:	*Rounds in later two thirds*	*Rounds in Second Half*	*Relative Round*
	(1)	(2)	(3)
ξ	-0.0046*	-0.0027	-0.0061**
	(0.0024)	(0.0019)	(0.0031)
$\xi + \beta$	0.0007	0.0009	0.0040
	(0.0016)	(0.0018)	(0.0028)
W	-0.468	-0.57	-1.30*
	[0.320]	[0.284]	[0.097]

Notes:
* significant at 90%; ** significant at 95%; *** significant at 99%.
Each column presents part of the results from a single OLS regression of equation 25. Winning bid and repayment burden are in logarithms. There are 24 quarterly dummies (first quarter 1993 to 4th quarter 1998) in each specification. Number of observations in each specification: 46,269.
Standard errors in parentheses. *P*-values in brackets.
W is the test statistic for a non-linear Wald test of the null hypothesis $\xi(\xi + \beta) \geq 0$. The distribution of the test statistic is asymptotically standard normal. As the alternative is one-sided, the associated *p*-value equals the value of the cumulative distribution function of a standard normal random variable evaluated at W.

Table 5. Descriptive Statistics for the 2002 Sample by Denomination

			Number of Roscas			Winning Bid ≥ 30% of Pot				Winning Bid = 40% of Pot	
						Before Shock		*After Shock*		*After Shock*	
Duration	Contribution	Pot	*Before Shock*	*After Shock*	*Increase (%)*	*Early*	*Late*	*Early*	*Late*	*Early*	*Late*
			(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
25	400	10,000	299	306	2.34	44.0	0.0	30.9	0.1	5.7	0.0
25	1,000	25,000	204	187	-8.33	51.6	1.2	28.7	0.0	7.9	0.0
25	2,000	50,000	150	154	2.67	62.2	0.0	49.3	0.0	16.4	0.0
25	4,000	100,000	76	66	-13.16	69.0	0.5	53.0	0.0	16.2	0.0
30	500	15,000	286	244	-14.69	59.6	0.2	43.9	0.0	12.9	0.0
30	1,000	30,000	87	93	6.90	70.4	0.1	46.6	0.1	15.7	0.0
30	2,500	75,000	38	18	-52.63	73.9	0.8	64.1	0.0	23.9	0.0
30	10,000	300,000	28	26	-7.14	77.0	0.0	78.7	0.0	35.4	0.0
40	250	10,000	555	310	-44.14	54.7	0.5	65.6	0.4	21.8	0.0
40	625	25,000	155	112	-27.74	67.7	0.9	74.3	0.8	32.7	0.0
40	1,250	50,000	142	124	-12.68	78.0	2.2	86.9	1.8	44.7	0.1
40	2,500	100,000	130	138	6.15	86.8	4.9	89.7	2.4	57.6	0.0

Notes:
Column 1: Roscas started between January 1, 2001 and December 31, 2001. Bid ceiling 30% of pot.
Column 2: Roscas started between January 1, 2002 and December 31, 2002, and completed by December 2005. Bid ceiling 40% of pot.
Column 3: Percentage difference between values in columns 2 and 1.
Columns 4 - 7: Percentage of auctions in which the winning bid is at least 30% of the pot.
Columns 8 and 9: Percentage of auctions in which the winning bid equals 40% of the pot.
Columns 4, 6 and 8: First half of rounds in a Rosca. Columns 5, 7 and 9: Second half of rounds in a Rosca.

Table 6. Arrear Rates, Cosigners, Winning Bids and Repayment Burdens, 2002 Policy Shift

Means

	Entire Sample	*Before Shock*		*After Shock*		Single Difference		Double Difference
		Early	*Late*	*Early*	*Late*	*Early*	*Late*	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Arrear Rate (percent)	13.1	14.9	11.8	15.8	11.0	-0.92	0.79	-1.72
	(20.8)	(20.1)	(21.3)	(20.9)	(20.6)	[0.26]	[0.23]	[0.35]
Cosigner Incidence (percentage)	52.8	77.4	33.5	77.5	34.9	0.07	1.34	1.27
	(49.9)	(41.8)	(47.2)	(41.8)	(47.6)	[0.56]	[0.49]	[0.75]
Number of Cosigners	1.05	1.78	0.47	1.80	0.49	2.00	1.51	0.48
	(1.19)	(1.22)	(0.76)	(1.24)	(0.49)	[1.24]	[1.08]	[1.65]
Winning Bid (percentage of pot)	17.4	27.2	9.3	30.0	8.8	-2.84	0.57	-3.34
	(11.4)	(4.9)	(5.8)	(8.7)	(5.3)	[0.07]	[0.07]	[0.10]
Repayment Burden (percentage of pot)	40.7	65.8	21.5	65.9	21.7	-0.09	-0.26	0.17
	(24.9)	(10.8)	(12.8)	(10.2)	(12.8)	[0.15]	[0.13]	[0.20]
Observations	60,312	15,313	19,732	10,767	14,500			

Notes: Standard deviations in parentheses for columns 1-5; Standard errors in brackets for columns 6-8.
Column 6 reports the difference between the values in columns 2 and 4, column 7 the difference between columns 3 and 5.
Column 8 reports the difference between the values in columns 6 and 7.

Table 7. Selection on Unobserved Riskiness, 2002 Policy Shift
OLS
Dependent Variable: Arrear Rate

		(1)	(2)	(3)	(4)	(5)	(6)
After*Rounds in Later Two Thirds		-0.0103***	-0.0024				
		(0.0037)	(0.0059)				
After*Rounds in Second Half				-0.0170***	-0.0108**		
				(0.0034)	(0.0053)		
After*Relative Round						-0.0272***	-0.0162
						(0.0059)	(0.0106)
<u>Controls</u>							
Cosigner Incidence		No	Yes	No	Yes	No	Yes
	F(388, 58510)		2.14***		2.14***		2.14***
	Average Effect		-0.0172***		-0.0171***		-0.0171***
			(0.0039)		(0.0039)		(0.0039)
Number of Cosigners		No	Yes	No	Yes	No	Yes
	F(388, 58510)		1.20		1.20***		1.20***
	Average Effect		0.0159***		0.0159***		0.0159***
			(0.0016)		(0.0016)		(0.0016)
Winning Bid (logarithmic)		No	Yes	No	Yes	No	Yes
	F(388, 58510)		2.66***		2.66***		2.66***
	Average Effect		0.0601***		0.0594***		0.0594***
			(0.0050)		(0.0049)		(0.0049)
Repayment Burden (logarithmic)		No	Yes	No	Yes	No	Yes
	F(388, 58510)		0.92		0.92		0.92
	Average Effect		0.0260		0.0237		0.0243
			(0.0602)		(0.0602)		(0.0602)
Aggregate Shocks		No	Yes	No	Yes	No	Yes
	F(24,44265)		3.40***		3.40***		3.40***
R^2		0.0739	0.1113	0.0741	0.1113	0.0740	0.1113
Observations		60,312	60,312	60,312	60,312	60,312	60,312

Notes:
* significant at 90%; ** significant at 95%; *** significant at 99%.
Standard errors in parentheses in all specifications. All specifications include 354 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 12 denomination-specific dummies.
19 quarterly dummies (second quarter 2001 to 4th quarter 2005) 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 354 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 354 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 354 interactions) are added in specifications with controls for the winning bid. The logarithm of the amount owed interacted with denomination-specific round dummies (354 interactions) are added in specifications with controls for repayment burden. The total number of regressors in columns 2, 4 and 6 thus is 1,802.

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.

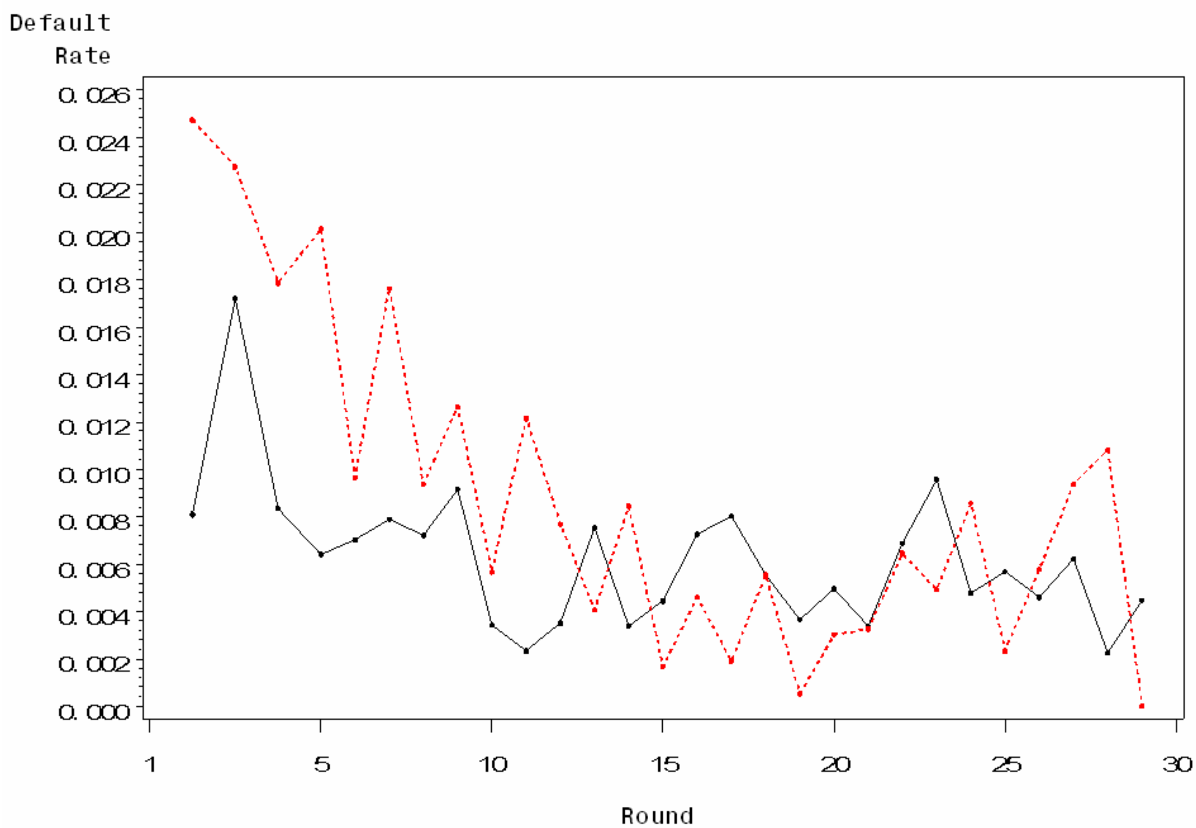


Figure 2. Willingness to Pay for the First Pot (Adverse Selection).
High Risk Type (dashed line). Low Risk Type (solid line). Bid Ceiling (dotted line).

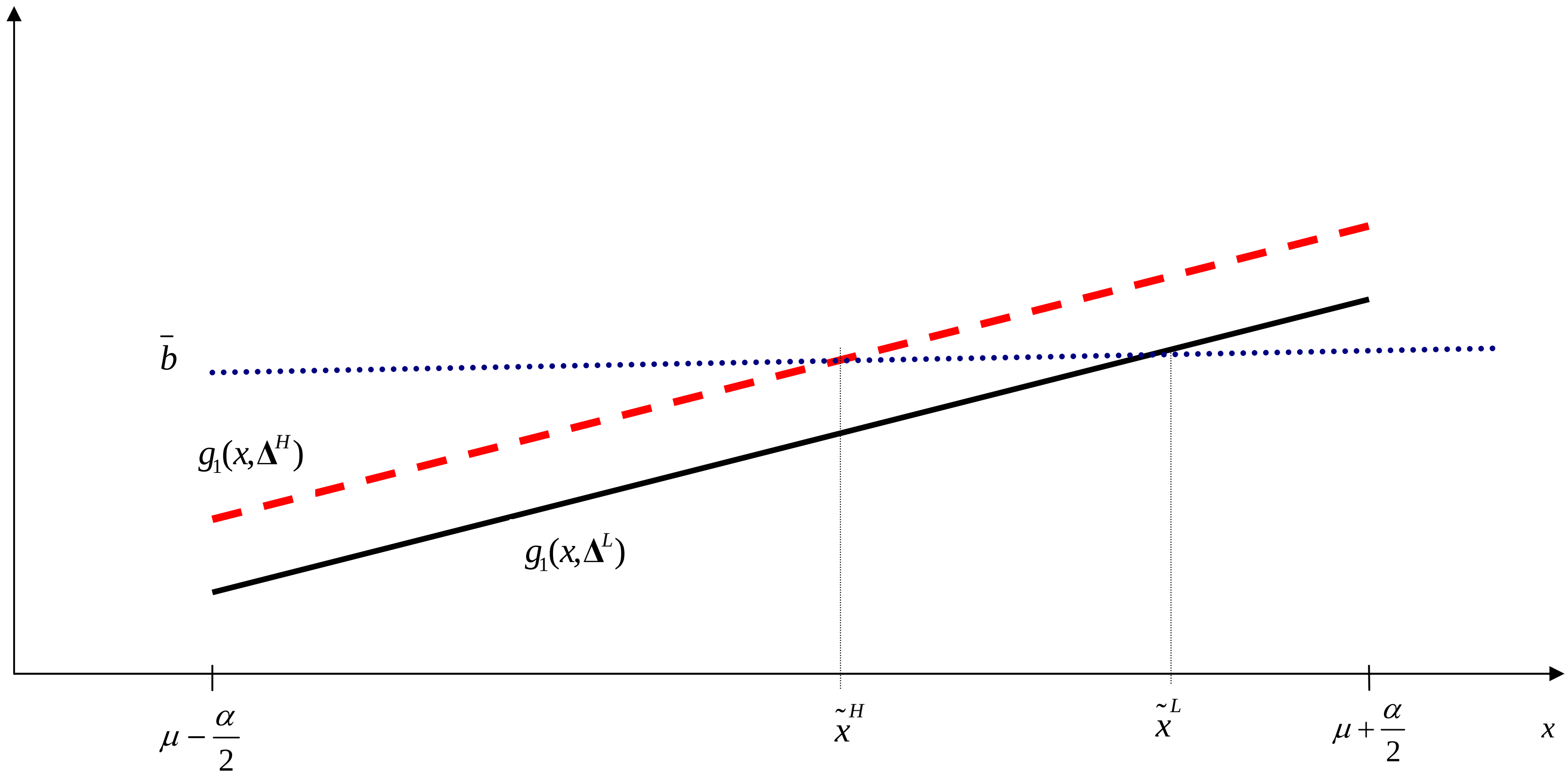
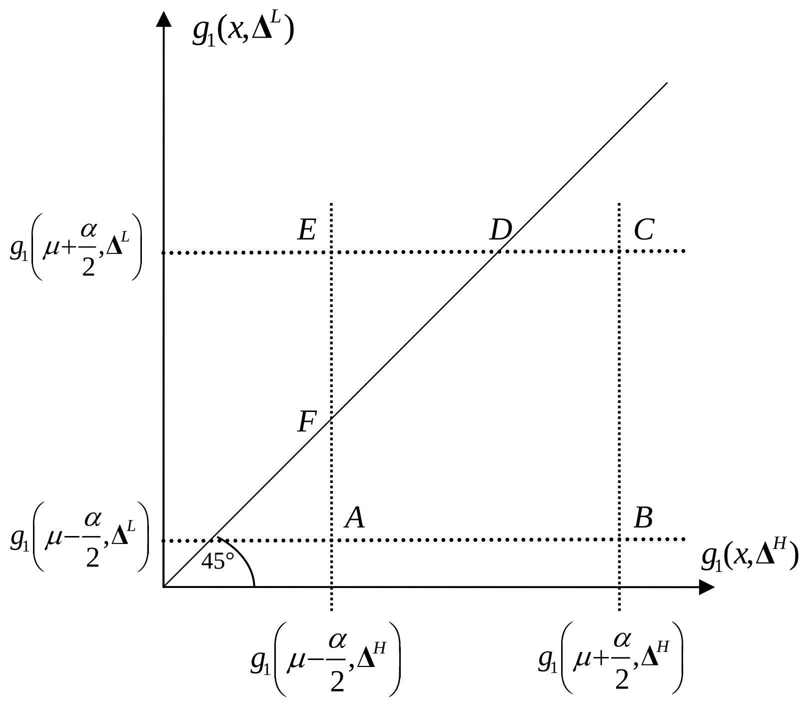


Figure 3. Probability of Winning the First Auction with Adverse Selection

Panel A. Unrestricted Bidding.



Panel B. Restricted Bidding.

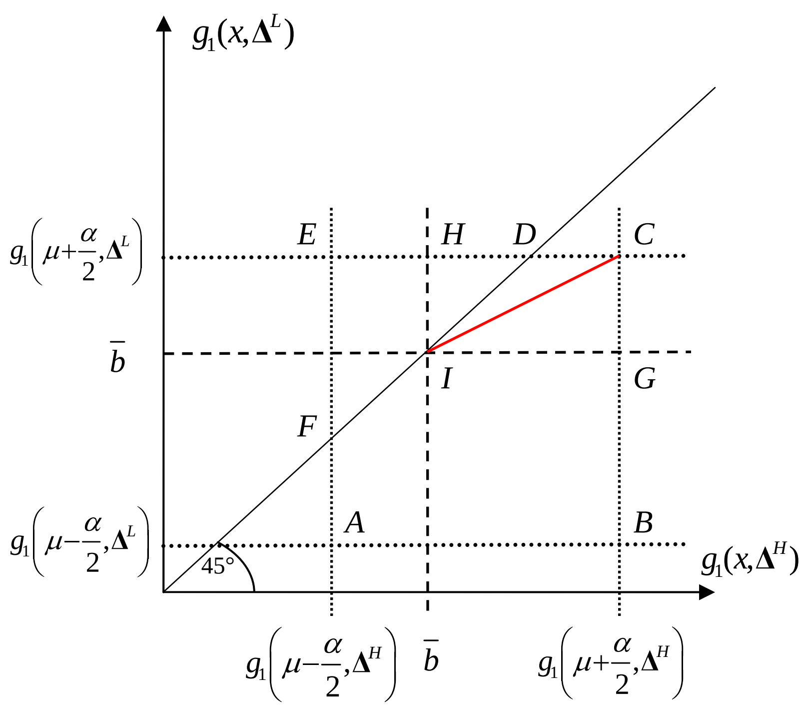


Figure 4. Willingness to Pay for the First Pot (No Adverse Selection).
High Risk Type (dashed line) Low Risk Type (solid line). Bid Ceiling (dotted line).

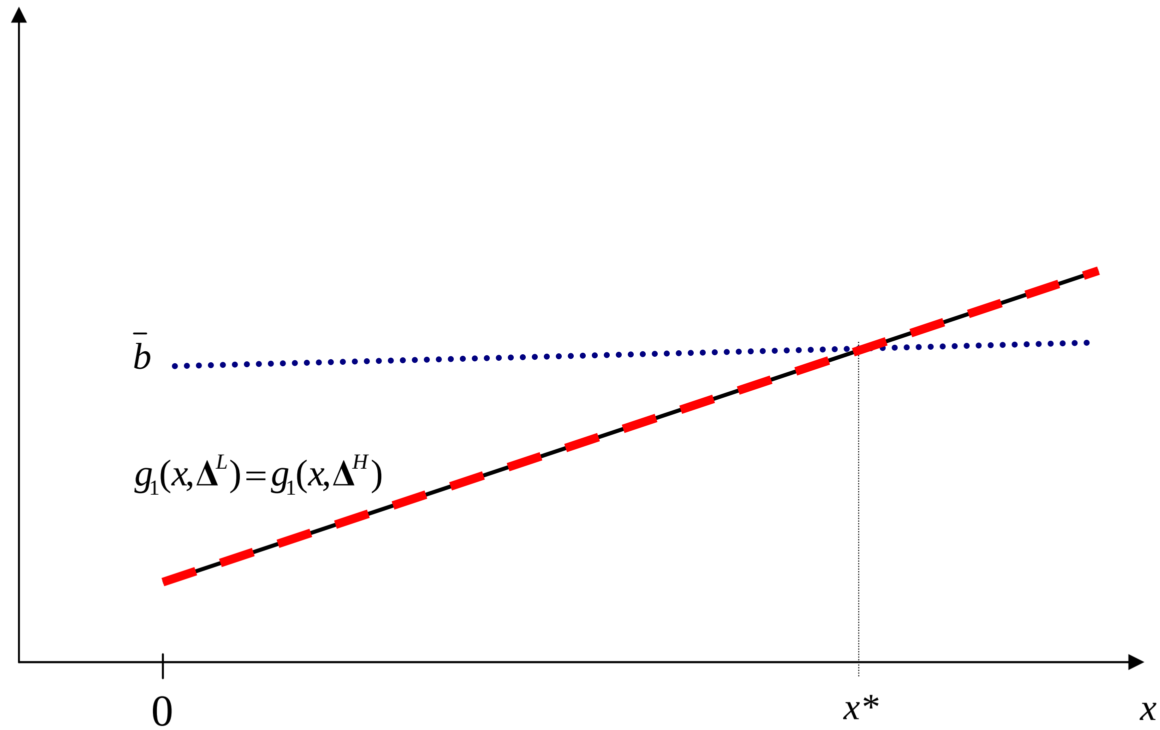


Figure 5. Probability of Winning the First Auction, No Adverse Selection.

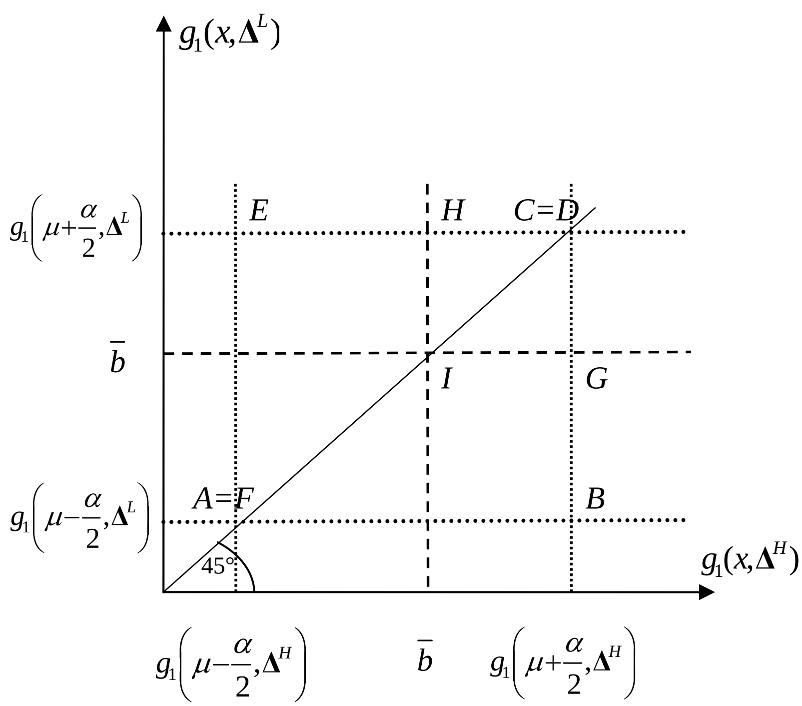


Figure 6. Riskiness Profiles with Adverse Selection.
Rosca with Unrestricted Bidding (dashed line) and Restricted Bidding (solid line).

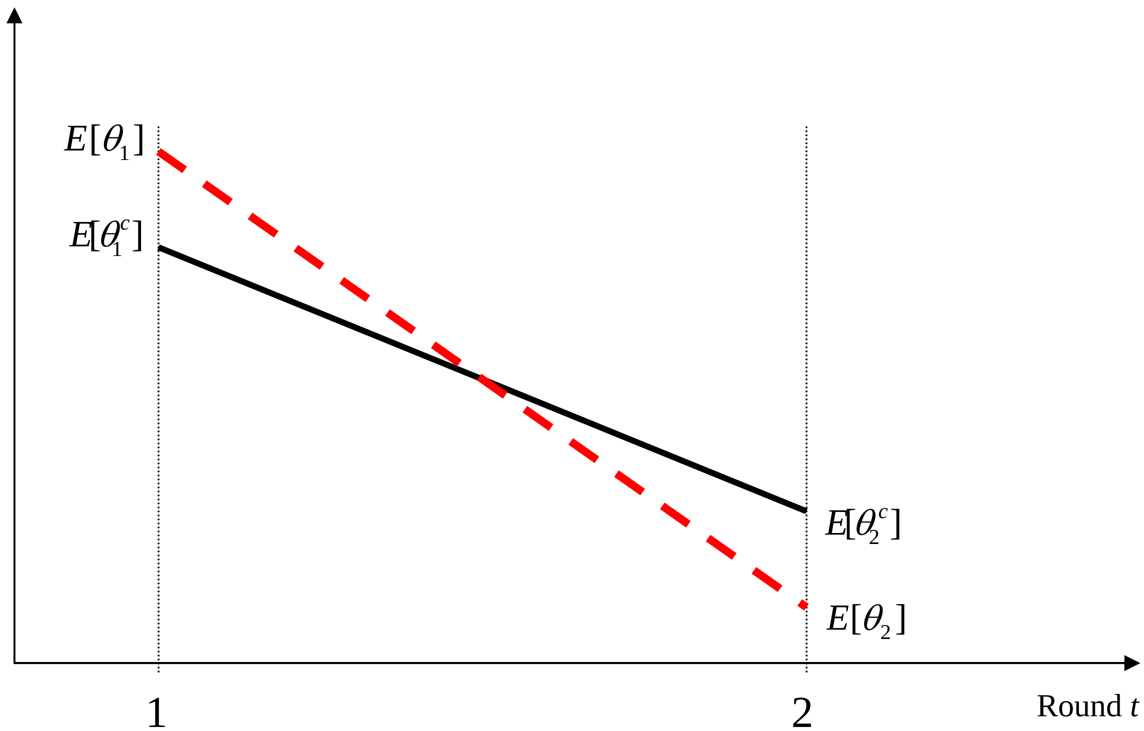


Figure 7. Riskiness Profiles with No Adverse Selection.

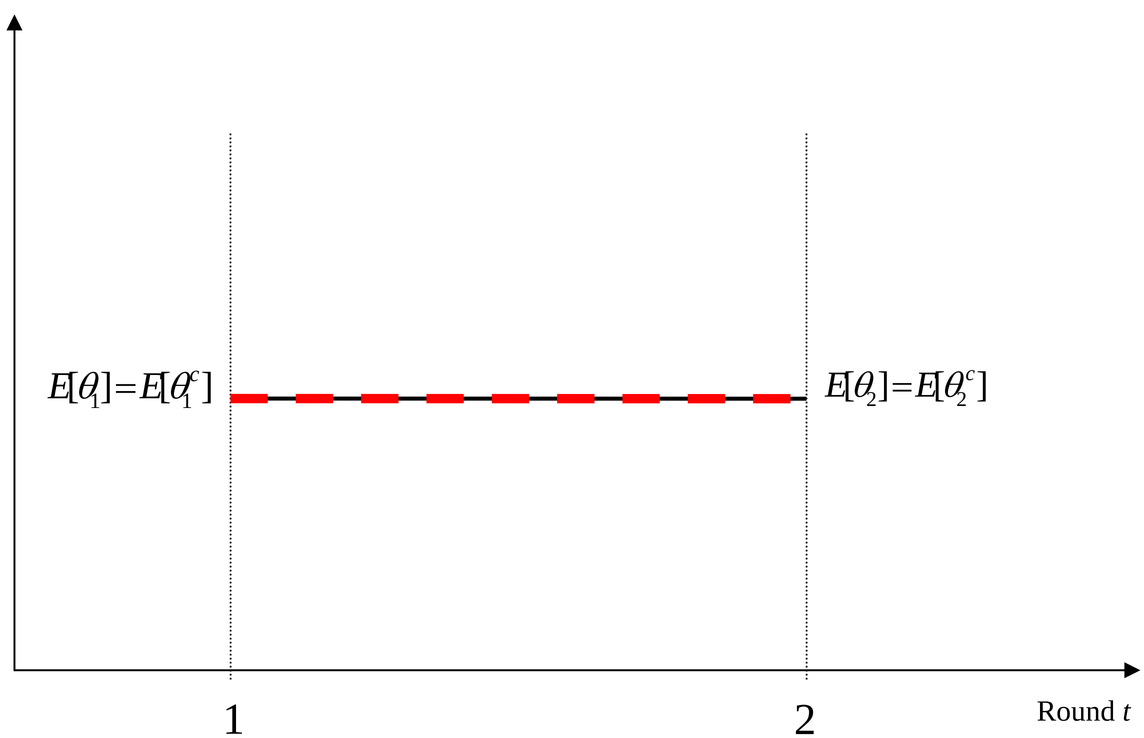
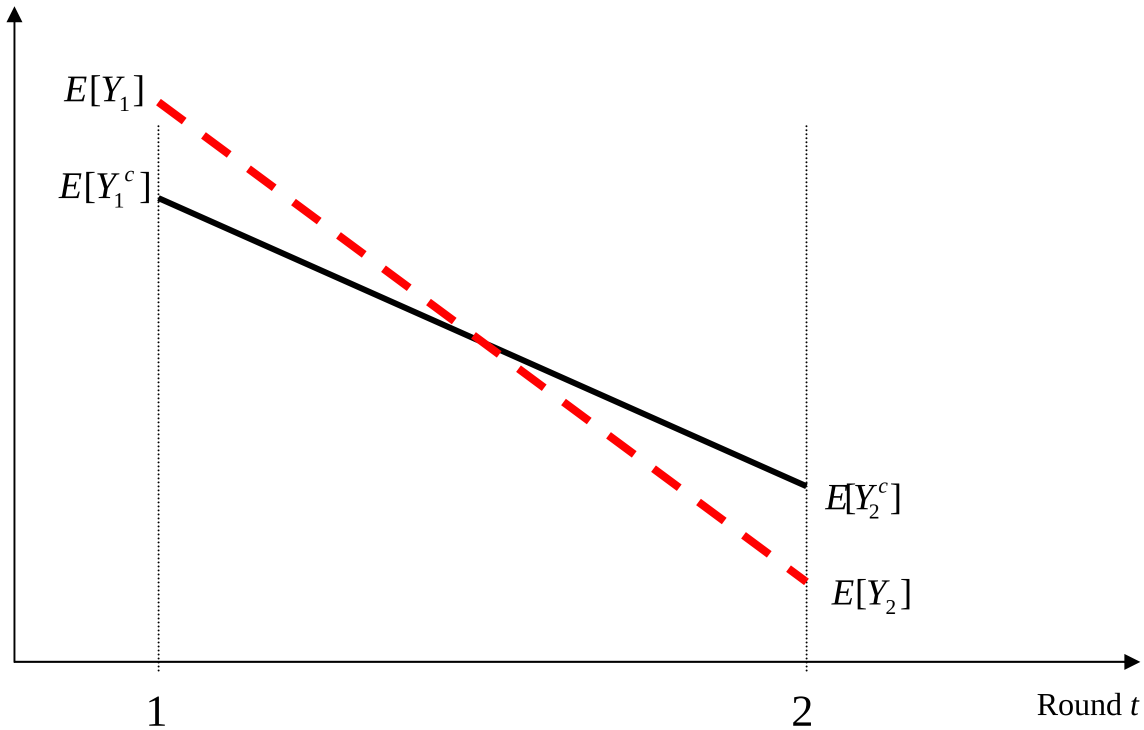


Figure 8. Default Profiles, Rosca with Unrestricted Bidding (dashed line) and Restricted Bidding (solid line).

Panel A. Adverse Selection, no Change in Rosca Composition.



Panel B. No Adverse Selection, Change in Rosca Composition.