Long Term Benefits of Membership in Microfinance Programs*

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Abstract

This paper studies the benefits of membership in microfinance programs, and examines whether membership in these programs is an effective instrument in smoothing inter-seasonal consumption. We hypothesize that the benefits to participation accrue differentially over time, as more experienced participants are better equipped on their own to minimize per capita consumption fluctuations. Using an Euler equation approach, we show that consumption differentials across seasons are inversely related to length of membership. Estimates from the gender-stratified model suggest that for a female participant, one year of membership reduces the percentage change in per capita consumption, caused by a unit shock, by 6%. We present simulation results confirming that as length of membership increases, the "certainty equivalent" of the participant decreases.

1 Introduction

In a primarily rural economy such as Bangladesh, a significant proportion of the labor force is subject to the seasonality of the crop cycle that results due to marked weather patterns. Seasonality coupled with the lack of access to formal insurance mechanisms implies that poor rural households experience strong fluctuations in their annual income flows. Absent sources of income which do not depend on weather outcomes, seasonal fluctuations in income flows are fully reflected in the household's annual consumption stream.

Participation in microcredit programs improves the ability of members to withstand aggregate shocks that cause seasonal consumption changes. Credit provided is earmarked for non-agricultural, self-employment activities, and this establishes a source of cash for the household that is unlikely to covary with agricultural shocks such as weather. Furthermore, since the majority of participants are female, credit provided to them succeeds in diversifying income across individuals within a household.

The benefits to participation change with length of membership. Reasons for this include: (1) Precautionary - by virtue of membership in these programs, more experienced members would have accumulated assets over time, which may be used in a precautionary role to smooth consumption. Estimates from the data used here suggest that the real average amount of total assets for a member household before participation is Taka 54034, and after participation is Taka 63338. Hence, the average total assets for a household increase by approximately 17.2% after participation. (2) Availability of collateral to obtain loans - accumulated assets and savings may also be used as collateral for loans from other sources. (3) Build up of a "reputation" - membership in credit programs leads to the formation of a "reputation" for more experienced clients. By demonstrating their ability to meet regular installment payments, more experienced members signal their ability to be good credit risks for other lenders in the market. (4) Since longer duration members possess the means to avail of other sources of borrowing, their demand for credit conditional on a particular source is more price elastic. This implies that they are charged a lower interest rate (Iqbal 1988).

Not many studies in the literature have attempted to evaluate the long-run benefits of participation, or sought to understand how the behavior of participants evolves over time. Of most relevance to this research is Pitt and Khandker (1998a), which examines credit effects by gender, and finds that credit given to female participants has strong beneficial effects on household consumption, children's schooling, nutrition, and male and female labor supply. However, the study does not consider whether positive benefits extend into the future. Other studies that use these same data include Morduch (1998), Pitt et. al (1999a), and Pitt et. al (2003). Using a difference-in-difference technique, Morduch (1998) concludes that membership does little to reduce poverty; it may, however, reduce vulnerability. Pitt et. al (1999a) evaluates the effect of participation on fertility and contraception, and finds no support for the hypothesis that participation reduces fertility and increases the use of contraceptives. Pitt et. al (2003) studies the impact of participation on anthropometric measures of child health. It is found that credit provided to female participants has strong, positive influences on health indicators of male and female children. Again, these studies do not consider whether effects persist in the long run.

Although we use similar estimation techniques and the same data as the studies noted above, this research differs in several important ways. First, this paper evaluates long term effects by estimating the consumption smoothing ability of participating households, as a function of their length of membership. The idea is that if more experienced members develop a stronger bargaining position through the accumulation of wealth, then they are better able to withstand seasonal shocks to household per capita consumption. This enhanced ability to buffer consumption against shocks indirectly captures the household's long run capacity to survive independently without aid. Second, we report "simulations" in support of our estimates to demonstrate how increases in membership length translate into lower income variability. Third, we address the issue of cohort effects and mis-targeting, and provide some evidence that our results are robust to these concerns. Fourth, by using the concept of "certainty equivalents", we provide a welfare interpretation of the estimates of our empirical models. Finally, this research includes an analysis of whether participants engage in smoothing consumption by delaying repayments during times of most hardship (the "hungry" or "lean" seasons).

We motivate our empirical analysis by incorporating borrowing constraints into Euler equations (Flavin, 1981; Zeldes, 1989; Foster, 1995). This is done through the formulation of a time varying, household specific interest rate. The deviation of this household specific time varying interest rate from the average village interest rate captures the higher cost of borrowing faced by a poor household. The household specific interest rate reflects the individual shadow price of intertemporal resource transfer and may be affected by social costs and monetary costs, as well as the household's bargaining position in the village. The household interest rate is not just the interest rate on loans to the household, it is a measure that proxies for all costs associated with smoothing the effects of seasonal shocks to household consumption. The household interest rate thus takes into account all other factors that affect the household's ability to borrow (for example, reputation in the village, previous experience with managing loans, and so on), over and above the actual interest rate on a loan. Hence, the household specific interest rate is the rate at which the household borrows, where the rate is a comprehensive measure of all implicit and explicit expenses associated with smoothing seasonal consumption shocks. The average village interest rate captures effects of price changes, seasonal shocks, and taste shifters that affect *all* households in the village. Our hypothesis is that participation in microfinance programs reduces the deviation in the household specific interest rate from the average village interest rate.

By using maximum likelihood techniques to estimate the relationship between changes in per capita consumption across seasons and several household and village level factors such as length of membership, changes in prices, preferences, and the cost of borrowing, we demonstrate that membership length is inversely related to the change in seasonal per capita food expenditure. Our estimation technique allows for self-selection into programs, non-random program placement, and for membership effects to differ by season. Our estimation results suggest that when members are stratified by gender, a one year increase in the length of membership of a female participant reduces the percentage change in per capita consumption caused by a unit shock (the concept of a "unit shock" is explained below) by 6%. This means that for an implicit interest rate change of one standard deviation around the mean, a household that has a female participant who has been a member for approximately four years (50^{th} percentile of length of membership) experiences a consumption change of 11.48% between two seasons. As predicted by our theoretical model, non-participants experience a higher consumption differential of 24.02%. A section of the paper is devoted to conducting various robustness checks for the results obtained.

The paper is organized as follows. Section 2 provides an outline of the model that is used for the empirical estimations. Section 3 provides a summary of the data, and section 4 discusses issues involved in the estimation. Section 5 reports the results and section 6 discusses the welfare implications of the results. Section 7 provides further support for results and section 8 concludes.

2 Model

We begin by incorporating credit constraints into the standard Euler equation of a dynamic utility maximization model, which relates consumption changes to changes in prices, the interest rate, and the discount factor.

Consider a household that maximizes the expected value of a time separable lifetime utility function. In each time period t, household i in village j chooses consumption C_{ijt} to solve:

$$Max\ E_t \sum_{k=0}^{T-1} \beta^k U(C_{ijt+k})$$

subject to an asset update:

$$A_{ijt+k} = (1 + r_{ijt})(A_{ijt+k-1}) + Y_{ijt+k} - P_{jt+k}C_{ijt+k} \quad \forall k$$

where β is the discount rate, E_t is the expectations operator conditional on information available as of time t, T is the end of the household's horizon, r_{ijt} is the interest rate faced by the household, A_{ijt} are household assets, Y_{ijt} is household income, and P_{jt} are prices in village j at time t. The first order conditions are of the following form:

$$E_t \frac{U'(C_{ijt+1})}{U'(C_{ijt})} = \frac{1}{\beta(1 + r_{ijt})} \frac{P_{jt+1}}{P_{jt}} \tag{1}$$

or

$$\frac{U'(C_{ijt+1})}{U'(C_{ijt})} \beta(1 + r_{ijt}) \frac{P_{jt}}{P_{jt+1}} = 1 + e'_{ijt+1} \tag{2}$$

Under rational expectations, e'_{ijt+1} is the expectational error which is uncorrelated with information known at time t. Assuming a CRRA utility function of the form:

$$U(C_t) = \frac{C_t^{1-\alpha}}{1 - \alpha}$$

and substituting its first derivative into equation (2) yields,

$$\left(\frac{C_{ijt+1}}{C_{ijt}}\right)^{-\alpha} \beta(1 + r_{ijt}) \frac{P_{jt}}{P_{jt+1}} = 1 + e'_{ijt+1}$$

defining $R_{ijt} = (1 + r_{ijt})$ and substituting into the above implies:

$$ln\left(\frac{C_{ijt+1}}{C_{ijt}}\right) = \frac{1}{\alpha}\left[ln\ (\beta R_{ijt}) + ln\ \left(\frac{P_{jt}}{P_{jt+1}}\right) - ln\left(1 + e'_{ijt+1}\right)\right] \tag{3}$$

Following Zeldes (1989), define $(1 + e'_{ijt+1}) = (1 + e^a_{jt+1})(1 + e_{ijt+1})$, where e^a_{jt+1} is the aggregate component of the expectational error and e_{ijt+1} is the idiosyncratic component. We define

$$\epsilon_{ijt+1} = -\left(ln(1 + e_{ijt+1}) + \frac{1}{2}\sigma^2_{e_{ijt+1}}\right)$$

where the expectational error and its idiosyncratic component are related as in Zeldes (1989). Substituting ϵ_{ijt+1} into (3) we arrive at:

$$ln\left(\frac{C_{ijt+1}}{C_{ijt}}\right) = \frac{1}{\alpha}\left[ln\ (\beta R_{ijt}) + ln\ \left(\frac{P_{jt}}{P_{jt+1}}\right) + \epsilon_{ijt+1}\right] \tag{4}$$

In order to analyze the cost of borrowing faced by a household, the term that captures the difference between the household interest rate and the average interest rate in the village needs to be incorporated. We achieve this by following Foster (1995) where the term $\frac{1}{\alpha}ln\ (\beta R_{ijt})$ is approximated to its first order Taylor series expansion about the village average interest rate R_{jt}. Doing so allows us to explicitly incorporate the fact that households without collateral face higher costs of smoothing and are unable to borrow at the average village interest rate. This yields,

$$ln\left(\frac{C_{ijt+1}}{C_{ijt}}\right) = \frac{1}{\alpha}\left[ln\ (\beta R_{jt}) + (r_{ijt} - r_{jt})\frac{1}{R_{jt}} - ln\ \left(\frac{P_{jt+1}}{P_{jt}}\right) + \epsilon_{ijt+1}\right]$$

where r_{jt} is the average interest rate in village j at time t, and r_{ijt} is the interest rate faced by household i in village j at time t. As noted above, the household interest rate is a comprehensive measure that proxies for all implicit and explicit costs associated with smoothing the effects of seasonal shocks to consumption. The household interest rate thus takes into account all other factors that affect the household's ability to borrow (for example, reputation in the village), over and above the actual interest rate on a loan. The incorporation of the term depicting the difference between the average village interest rate and the interest rate faced by a household i in that village, allows us to analyze the household's cost of borrowing relative to the rest of the village. Rewriting the above, we obtain:

$$\Delta ln\ C_{ijt+1} = \gamma_{0t} + \gamma_{1t}(r_{ijt} - r_{jt}) + \gamma_{2t}\Delta ln\ P_{jt+1} + \nu_{ijt+1} \tag{5}$$

where $\gamma_{0t} = \frac{1}{\alpha}ln\ \beta R_{jt}$, $\gamma_{1t} = \frac{1}{\alpha R_{jt}}$, $\gamma_{2t} = \frac{1}{\alpha}$ (The coefficient of relative risk aversion is assumed to be non-negative), and $\nu_{ijt+1} = \frac{1}{\alpha}\epsilon_{ijt+1}$. Those households with easier access to credit (that is, with small $(r_{ijt} - r_{jt})$) are better able to smooth consumption. Thus for them, changes in per capita consumption are primarily governed by changes in prices, preferences, and the average interest rate that affects all households in the village.

We hypothesize that in the presence of season specific shocks, the deviation between the average village interest rate and the implicit interest rate faced by a poor household depends on the length of time the household has been a member of the credit program. Membership length reduces the cost of borrowing since those who have been participants for long periods of time accumulate the means to minimize consumption fluctuations. The $(r_{ijt} - r_{jt})$ term captures the deviation in the household interest rate from the average village interest rate, where the latter reflects seasonal shocks to consumption in the village. The higher cost of borrowing faced by (poor) non-participants is captured in the $(r_{ijt} - r_{jt})$ term. For participants, the differential they face is tempered by length of membership. In order to reflect this, we adopt the following specification:

$$\gamma_{0t} + \gamma_{1t}(r_{ijt} - r_{jt}) = \mathcal{F}(D_{ijt})\mu_{jt} \tag{6}$$

$\mathcal{F}(D_{ijt})$ denotes that the household's cost of borrowing is a function of D_{ijt}, where D_{ijt} is the duration of membership (of household i in village j at time t) in a credit program, and μ_{jt} is a village season dummy that predicts average interest rates.[1] Where

$$\mathcal{F}(D_{ijt})\mu_{jt} = \left(e^{\delta_t D_{ijt}}\right)\mu_{jt}$$

we obtain,

$$\gamma_{0t} + \gamma_{1t}(r_{ijt} - r_{jt}) = \left(e^{\delta_t D_{ijt}}\right)\mu_{jt} \tag{7}$$

Substituting equation (7) into equation (5):

$$\Delta ln\ C_{ijt+1} = \left(e^{\delta_{t+1} D_{ijt+1}}\right)\mu_{jt+1} + \gamma_{2t}\Delta ln\ P_{jt+1} + \nu_{ijt+1} \tag{8}$$

Equation (8) forms the basis of the subsequent estimations. Variables such as characteristics of the household head as well as the quantity of land owned by the household also play a role in reducing

[1] Note that μ_{jt} picks up the effect of all variables (including interest rates) that change by seasons across villages. Hence, these parameters do not capture interest rate variations exclusively. Given dearth of data on actual measures of average village interest rates, we are unable to control for this. Since interest rate variations at the village level are still captured, we do not significantly alter our interpretation of this parameter.

the cost of borrowing. These may be included in (8) in a similar manner to the inclusion of D_{ijt+1}.[2] Note that in equation (8), village level fixed effects that predict influences of seasonal shocks on average village interest rates (the μ_{jt+1})[3] are estimated simultaneously with the δ_{t+1} and γ_{2t} parameters.

The coefficient δ_{t+1} is the marginal effect of interest, and the hypothesis that experienced participants are better able to smooth consumption predicts that $\delta_{t+1} < 0$. If seasonal shocks are fully reflected in the μ_{jt+1} parameters that measure average interest rate changes at the village level, then it is evident from equation (7) that for a unit shock (as noted before, the μ_{jt+1} parameters pick up the effects of seasonal shocks. In order to interpret a negative delta coefficient, we assume that μ_{jt+1} equals one in equation (7)), a negative delta coefficient implies that one period of participation reduces the effect of a shock by a magnitude of $e^{-\delta_t}$. Hence, the longer the household has been a participant, the smaller is the effect of the anticipated shock on inter-seasonal consumption differentials. In the limit, changes in consumption are attributable to changes in prices and preferences, and seasonal shocks have an effect only through average interest rate that affect all households in the village.

3 Data

The data used in this analysis were collected from rural Bangladesh during 1991-1992. The sample is drawn from 29 randomly selected thanas (districts), 24 of which had one of the three microcredit programs (Bangladesh Rural Advancement Committee (BRAC), Bangladesh Rural Development Board's (BRDB) RD-12 program, and the Grameen Bank) in operation, and 5 of which had none. Grameen, BRAC, and BRDB's RD-12 program are similar in the types of services they provide, and

[2]We assume that household characteristics affect the cost of borrowing but do not directly shift tastes in the utility function. Given our focus on the cost of intertemporal resource transfer, we do not believe that this assumption is overly restrictive. Moreover, it seems unlikely that taste shifters explain most of the variation in the data for poor households. For example, there is little reason to believe that the poor have a steep drop in tastes during the lean season. Hence, any such result is likely to reflect high interest rates and prices.

[3]The μ_{jt+1} parameters predict average village interest rate changes caused by seasonal shocks (among other things), and are estimated for two seasons of the year. These are thus village time parameters that predict effects of largely anticipatable seasonal shocks on average village interest rates which influence village consumption.

the eligibility criterion did not vary across them at the time these data were collected. Hence, all three programs are treated identically for purposes of the following estimation.[4]

Three villages in each of the 29 thanas were randomly selected, although in the 24 program thanas, villages were randomly selected on the basis of their having had a program in operation for three or more years. Households within the selected villages were then identified as "target" (those who were qualified to join the program) or "non-target", and then participants and non-participants among the target households were separately identified. Those households who were participants as well as target non-participants were oversampled. The data has information on 1,798 households of which 1,538 were target households. Of the target households, 905 were participants.

The data track each household for three separate rounds corresponding to the three major rice crop seasons in Bangladesh. The first round coincides with the post harvest time of the *Aman* rice crop (December/January 1991). The second round coincides with the post harvest time of the *Boro* crop (April/May 1992), and the third round coincides with the post harvest time of the *Aus* rice crop (August/September 1992). The *Aman* rice crop is the largest in the year and the *Aus* season is traditionally seen as the "lean" season, with rural consumption levels reaching their annual lows in the months just before the *Aman* harvest.

In this paper, round 1 refers to the *Aman* season (season 1), round 2 to the *Boro* season (season 2), and round 3 to the *Aus* season (season 3).

Table 1 (all tables are at the end of the paper) provides the weighted means and standard deviations of all the independent variables used in the analysis. Length of membership is measured by the number of months the participant has been a member of one of the three programs. Two

[4]The earliest estimations of this study did report separate program coefficients by gender. It was found that Grameen Bank credit (to female participants) had the strongest most significant effects; the effects of credit from other programs had the correct sign, but were either not as significant, or were insignificant. Since our aim here is to evaluate the consumption-smoothing impact of microcredit in general, all three programs are treated similarly. In future work, we hope to address more fully the question of which specific program has the strongest effects on the household's long run ability to smooth seasonal shocks to consumption.

households had members belonging to more than one program, thus for each household, the maximum value for length of program membership is used in the estimations. In order to account for the choice based sampling of the data, means of variables are adjusted by weights that correct for the difference between the actual distribution of households in the villages surveyed and the distribution of households in the sample. Only those households that appeared in all three rounds of the survey were included (this excluded 29 households). Non-target households owning more than 5 acres of land were excluded in order to maintain the validity of the assumption that the "landed" may be pooled with the "landless". This excluded an additional 43 households. The sample thus consists of 1,726 households.

The dependent variable consists of two sets of differences in per capita consumption (between seasons 2 and 1, and between seasons 3 and 2), which are estimated jointly. Per capita consumption is measured by per capita food expenditure in the week previous to the survey. In these data, expenditure on food constitutes almost 80% of total expenditure at the household level, and the variable is constructed from a single question in the survey. Note that under assumptions regarding the timing of information availability (in formulating the Euler equation), the difference between seasons 3 and 1 is approximately the sum of differences between seasons 2 and 1, and seasons 3 and 2. Round 1 data are excluded from the estimation (although round 1 level variables are still present) because the theoretical model implies that the difference in per capita consumption between seasons 2 and 1 is a function of round 2 data, and the difference in per capita consumption between seasons 3 and 2 is a function of round 3 data.

Table 2 reports the weighted mean and standard deviation of the dependent variables. The flowchart at the end of the paper (along with the figures) depicts the structure of the data.

4 Estimation

Identification of the effect of an endogenous variable such as membership duration is achieved by using the quasi-experimental nature of the data.[5] There are a couple of issues of importance. The main aim of microcredit programs is to alleviate poverty and to provide poor people with the means

[5]Pitt and Khandker (1998a) describes the nature of the quasi-experiment.

of a steady livelihood. If programs are deliberately placed in areas that are relatively poorer than others, then estimates of the effects of program participation are necessarily biased. The use of village level fixed effects that capture systematic differences in attributes across villages, aids in removing this bias. Yet, without any further variation in program availability, it is usually not possible to separately identify the length of membership effect from the village fixed effect. Identification of the length of membership effect becomes possible if the sample includes households in villages that have the program, but that are excluded from participating due to an exogenous rule. Microcredit programs only lend to those who own less than half an acre of cultivable land, or in the absence of land ownership, those who own assets whose value equals or is less than that of one acre of medium quality land in that area. This rule provides the random assignment necessary for identification.[6]

Individual heterogeneity also needs to be taken into account since those who choose to participate could be systematically different from non-participants. Once a credit program is established in a village, participants self-select into groups to become members of the program. If those who join are more able at managing self-employment activities, or have higher than average entrepreneurial skill levels as compared to non-participants, then the estimated participation effects are biased. Individual and household level heterogeneity of this type could confound results and incorrectly attribute to the program those effects that arise from differences in the nature of household unobservables. We correct for this in our estimation by controlling for the correlation between unobservables in the behavior equation and the length of membership equation.

Consider the following reduced form for duration of membership in one of the credit programs:

$$D_{ijt+1} = X^D_{ijt+1}\beta_D + \mu^D_{jt+1} + \varepsilon_{ijt+1} \texttt{ if choice = 1} \quad (9)$$

$$= 0 \texttt{ if choice = 0} \quad (10)$$

[6]Morduch (1998) presents some evidence that in the relative short run, ownership of land cannot be treated exogenously. Hence, the half acre rule is not exogenous and identification based on it leads to biased results. We address this concern in detail in section 7.1. When our empirical models were re-estimated using higher land threshold values (to account for the fact that participants often owned more than half an acre of land), our results increased in magnitude and significance. As noted in section 7.1, this is because mis-targeting led to a conservative bias. Thus, correcting for it improved the size and significance of our results.

Equation (9) is the reduced form for duration of membership for all those who have the choice to participate. If households do not have choice, then duration of membership is identically zero. The μ^D_{jt+1} are village level fixed effects that captures non-random program placement, and X^D_{ijt+1} are exogenous variables that affect duration of membership. The behavior equation is equation (8) which is reproduced below, and which now includes other exogenous covariates X^C_{ijt+1}:

$$\Delta ln\ C_{ijt+1} = \left(e^{X^C_{ijt+1}\beta_C + \delta_{t+1}D_{ijt+1}}\right)\mu_{jt+1} + \gamma_{2t}\Delta ln\ P_{jt+1} + \nu_{ijt+1} \qquad (11)$$

Since D_{ijt+1} is endogenous in (11), identification of its effect requires the use of instrumental variables. If a household is a target household (owns less than half an acre of land) in a village with a program, the household is considered to be eligible and to have choice (choice is a dummy variable). The interaction of choice and the exogenous variables are the identifying instruments. This is because exogenous variables can have an effect on length of membership only if the household is eligible and has choice. For households without choice, length of membership is identically zero. *It is important to note that the exogenous variables in of themselves are not the instruments, the interactions of the exogenous variables with choice constitute the instrument set.* As noted before, length of membership is endogenous for participants and identically zero for non-participants. The interactions of choice and the exogenous variables form the set of instruments, and these instruments affect length of membership discontinuously (the exogenous variables have an effect only if the household owns less than half an acre of land, and resides in a village with a program).

Assuming that the errors in (9) and (11) follow a jointly normal distribution, maximum likelihood estimation of equation (11) provides consistent and efficient estimates. As mentioned before, two sets of consumption differences are estimated jointly. The change in consumption between seasons 2 and 1 is a function of a village-season 2 dummy (which captures seasonal shock induced interest rate variations that affect all households in the village in season 2), and the change in consumption between seasons 3 and 2 is a function of a village-season 3 dummy (which captures seasonal shock induced interest rate variations that affect all households in the village in season 3). Thus the village-season dummy is allowed to have a different effect on the two sets of consumption changes that are estimated.

5 Results

5.1 Stratification by Gender

Microcredit programs such as BRAC, BRDB, and the Grameen Bank specifically court female members. The reason they target women is because in Bangladesh, women have little access to the wage labor market and are more credit constrained than men. Credit provided to female members has a stronger impact (as compared to credit given to men) on household outcomes such as children's schooling and per capita total expenditure.[7] Data used in this study reveal that on average at the household level, 67% of the participants are female. This relative dearth of male participants in the sample leads to the imprecise estimation of male effects, as is evident in table 3.

Given the importance of gender stratification, equations (9), (10), and (11) are disaggregated to analyze female and male effects separately. The estimation thus includes an equation for credit borrowed by male members, a separate equation for credit borrowed by female members, and a joint behavior equation for the change in consumption at the household level. Equation (11) is modified in the following way:

$$\Delta ln\ C_{ijt+1} = \left(e^{X^C_{ijt+1}\beta_C+\delta_{ft+1}D_{fijt+1}+\delta_{mt+1}D_{mijt+1}}\right)\mu^C_{jt+1} + \gamma_{2t}\Delta ln\ P_{jt+1} + \nu_{ijt+1} \tag{12}$$

Separate identification of δ_{ft+1} (female coefficient) from δ_{mt+1} (male coefficient) is possible because in villages with only female groups, men do not have the choice to participate, and vice versa. In the data, there are 22 villages with only female groups, 10 villages with only male groups, and 40 villages with both male and female groups. Table 3 reports results from the estimation of equation (12). Although length of membership is measured in months (as shown in table 1), for purposes of estimation, the variable is scaled to represent its yearly equivalent.

As predicted, with disaggregation by gender, maximum years of female membership in season 2 is negative and significant. As before, we assume that seasonal shocks cause variations in the average village interest rate parameter μ^C_{jt+1}, and for purposes of interpreting the δ_{ft+1} and δ_{mt+1} coefficients, assume that μ^C_{jt+1} equals one (a unit shock). Estimates suggest that one year of membership

[7] For example, for every additional 100 Takas of credit borrowed, annual household consumption expenditure increases by 18 Takas for women but only 11 Takas for men (Pitt and Khandker, 1998a).

where the participant is a female reduces the percentage change in per capita consumption caused by a unit shock by 6%. Using the estimated μ^{C}_{jt+1} coefficients that measure average village interest rate variations, this means that for an interest rate change of one standard deviation around the mean, a non-participating household experiences a consumption change of 24.02%.[8] For the same magnitude of interest rate deviation, a household with a female participant who has been a member for 4.15 years (50^{th} percentile) experiences a lower consumption change of 11.48%. A household that has a female participant who has been a member for 7.58 years (upper 95^{th} percentile) experiences a consumption change of only 0.85%, for the same magnitude of interest rate change. As hypothesized, experienced participants face smaller variations in consumption in response to seasonal shocks.

Female membership in season 3 although negative, is insignificant. This is similar to the effects of the male membership variable in both seasons 2 and 3. The fact that male variables are often insignificant is evident from other research.[9] Although the insignificance of the female membership variable in season 3 might at first seem unexpected (since season 3 corresponds to the *Aus* season which is traditionally seen as the "lean" season in Bangladesh), this is just the incremental effect over and above the effect of length of membership in round 2. A test for the equality of female membership effects in rounds 2 and 3 could not be rejected (χ^2 value=2.38 and p-value=0.88). If we define the dependent variable as the change in per capita consumption between seasons 3 and 1 (since season 1 is the time of least hardship and season 3 is the time of most hardship, the magnitude of change in consumption would be the largest in this case), then female membership in season 3 becomes significant. In this model, estimates for female membership in round 3 suggest that one year of participation reduces the percentage change in per capita consumption caused by a unit shock by 6.88%, with a t-statistic of -2.02.

Table 3 shows that variables such as education of the household head, age of the household head, and the dummy for no adult male in the household have the expected signs, but are insignificant. It appears that after controlling for several household characteristics, it is mainly length of program

[8]Standard errors were calculated using the delta method. These are not presented since all estimates are significant.

[9]Pitt and Khandker (1998a). As noted above, the imprecise estimation of male effects arises from their inadequate representation in the data.

membership that has the mitigating effect on consumption changes. The difference in log price of coarse grain rice between seasons 3 and 2 has a strong positive effect on the dependent variable. As noted before, per capita consumption is measured by per capita food expenditure. The positive coefficient on this variable implies that demand is price inelastic[10], which is expected since rice is the staple grain in Bangladesh. The difference in log price of rice between seasons 2 and 1 is also positive but marginally significant. There is not much variation in the price of rice between seasons 2 and 1, which are only three months apart in these data. The remaining regression coefficients (excluding fixed effects), separated by gender of the participant, are also reported (in the table entitled "Table 3 continued"). These correspond to cases where either only females or only males have the choice to participate. As is evident, many of the estimates have the expected signs, but are measured imprecisely. In the interests of brevity, we focus on the estimates in table 3, where the main emphasis is on the impact of membership time on reducing inter-seasonal consumption variability.[11]

The ρ parameters in table 3 measure the correlation between the error in the behavior equation (equation 12) and the errors in the male and female duration of membership equations (the equivalent of equation 9 where duration of membership is disaggregated by gender of the participant). These correlations correct for self-selection into the programs, and are allowed to vary by season. Since the behavior estimated is in its first differenced form, much of the household specific heterogeneity is removed. Thus, as expected, none of the correlations in table 3 are significant.

The first differenced form of the specification may explain why several coefficients are measured imprecisely in table 3. There are three months between the first and second rounds of the data, and approximately two to three months between the second and third rounds of the data. It is unlikely that there is a significant change in variables across these short time spans.

We could not reject a joint test for the equivalence of male and female coefficients in table 3 (χ^2 value = 0.7 and a p-value = 0.29). Thus male and female effects were constrained to be the same,

[10]There is little reason for us not to believe that increases in prices across commodities are relatively proportional.

[11]The full set of parameter estimates (in table 3 and table 3 continued) is reported only for the gender-stratified regression of equation (12).

results for this pooled model are reported below.

5.2 Pooled Estimates

Table 4 presents the estimates from the pooled model in equation (11). As expected, when men and women are pooled together in the estimation, household years of membership in season 2 is negative and significant. As before, we use the concept of a "unit shock" to interpret the coefficient on length of membership. Results suggest that one year of membership reduces the percentage change in per capita consumption caused by a unit shock by 4%. Using the empirical distribution of the estimated μ_{j2}^{C} parameters, it is possible to interpret this in the following manner. For an interest rate change of one standard deviation around the mean, a non-participating household experiences a consumption differential of 23.56% between seasons 2 and 1.[12] For the same magnitude of interest rate deviation, a household that has a participant who has been a member for the average years of duration (4.36 years) experiences a lower consumption change of 16.24%, while a household with a participant who has been a member for 8.08 years (upper 95^{th} percentile) experiences an even lower consumption change of 9.99%. As predicted, large interest rate variations result in relatively small consumption changes for households with more experienced members.

Table 4 also shows that characteristics of the household head such as age, sex, and education have the expected signs, but are insignificant. The ρ parameters in table 4 measure the correlation between the behavior equation (equation 11) and the duration of membership equation (equation 9). As before, these correct for self-selection into the programs and are allowed to vary by season.

The effect of length of membership in round 3 is insignificant, but again, this is just the incremental effect. The null hypothesis for equality of membership effects by rounds could not be rejected (χ^2 value = 1.77 and p-value=0.82). If we define the dependent variable as the difference in per capita consumption between seasons 3 and 1, membership as of round 3 is significant. In this case, results indicate that one year of membership reduces the percentage change in per capita consumption caused by a unit shock by 4.7%, with an absolute t-statistic of 1.73. The difference in log price of rice between seasons 2 and 1 as well as between seasons 3 and 2 has a strong positive

[12] As before, standard errors were calculated but are not reported since estimates are all significant.

effect, signifying that the demand for rice is price inelastic.

5.3 Variance of Consumption Growth

Another interpretation of the estimates in Tables 3 and 4 may be provided by computing the variance of log consumption change (between season 2 and season 1) for participants of different membership lengths, in response to random draws from the empirical distribution of average village shocks (μ_{j2}^C).[13] Note that as before, we assume that seasonal shocks influence average village interest rate variations, and these variations are picked up by the μ_{j2}^C parameters.

The estimates in Tables 5 and 6 show that households with more experienced participants face smaller variations in consumption growth between season 2 and season 1[14]. Thus, experienced participants are better able to withstand seasonal fluctuations in consumption.

6 Welfare Implications

To gauge the welfare implications of the results obtained, consider the amount an individual (of a given length of membership) is willing to pay to smooth consumption (eliminate uncertainty) across her lifetime, that is, the "certainty equivalent". In its simplest form, the optimization problem for an individual who lives for three time periods where r_2 is the interest rate in the second period and r_3 is the interest rate in the third period is:

$$Max\, U(C_1) + \beta U(C_2) + \beta^2 U(C_3)$$

subject to

$$Y_1 + \frac{Y_2}{(1+r_2)} + \frac{Y_3}{(1+r_3)} = C_1 + \frac{C_2}{(1+r_2)} + \frac{C_3}{(1+r_3)}$$

Where Y_1, Y_2, and Y_3 denote incomes in the three periods, C_1, C_2, and C_3 denote consumption in the three periods, and β is the discount factor. From the Euler equation and the CRRA utility

[13]In order for us to provide an alternative interpretation of the results in Tables 3 and 4, we assume that the econometricians instrument set is not very different from that of the individual.

[14]The delta method was used to compute standard errors for these estimates. Since all estimates are significant, the standard errors are not reported

function, we arrive at the following expression for consumption in period one (C_1):

$$C_1 = \left(\frac{Y_1 + \frac{Y_2}{(1+r_2)} + \frac{Y_3}{(1+r_3)}}{1 + (1+r_2)^{\frac{1}{\alpha}-1} + (1+r_3)^{\frac{1}{\alpha}-1}} \right)$$

where $\beta = 1$, and α is the coefficient of relative risk aversion.[15]

If the individual faces no interest rate change, $r_2 = r_3$ above and she decides C_1^*, C_2^*, and C_3^*, to smooth consumption over her lifetime. If we use C_1^* as the benchmark case for subsequent comparisons, the question is how much extra income in the first year does an individual of a certain membership length require in order to be indifferent (as compared to the case of no variations in the interest rate), given interest rate variability. Thus, given that $r_2 < r_3$ above, we solve for Θ in the following:

$$C_1^* = \left(\frac{\Theta + Y_1 + \frac{Y_2}{(1+r_2)} + \frac{Y_3}{(1+r_3)}}{1 + (1+r_2)^{\frac{1}{\alpha}-1} + (1+r_3)^{\frac{1}{\alpha}-1}} \right)$$

Θ is a measure of the extra income required for indifference. If more experienced individuals face lower costs of borrowing, they will require lower "compensations" (smaller Θs).

Given the relation in (7) and estimates from Tables 3 and 4, we calculate the cost of borrowing faced by individuals of different membership lengths. Income (Taka/year) for participants in the data in round 1 is Taka 34,881, round 2 is Taka 33,192, and round 3 is Taka 24,463. The average formal interest rate per year is 16.87%. Taking the average over rounds for income and constraining Y_1, Y_2, and Y_3 to be equal to this average income amount, we report the estimates in Tables 7 and 8. Both Tables 7 and 8 show that more experienced participants require significantly lower amounts of compensation in their first year. Figure 1 (all figures are at the end of the paper) graphs the estimates from Tables 7 and 8 (In Figure 1 "Delta" denotes Θ). It is evident that as length of membership increases, declining proportions of average annual income are required in order to compensate for higher interest rate variability, ceteris paribus.

[15] We compute α from the coefficient on the change in the price of rice obtained in the above estimations.

7 Further Support for Results

7.1 Mis-targeting

Although the de-jure rule for membership in these programs is that only those households who own less than half an acre of land may participate, members may own more land (when they join) than this cut-off value. Hence, the identification rule may not be exogenous (Murdoch, 1998).

In order to address this, both models were re-estimated using higher land threshold values. The result was that the length of membership effect in both models became more pronounced and significant. This is explained by the nature of the bias that results from mis-targeting. The bias arose from the fact that those households that actually had the choice to participate were treated as not having choice. The appropriate way to correct for this is to give such non-participating households the choice to participate. For example, if the exogenous rule is that households with less than 1 acre of land are eligible to participate, then those with land more than half an acre but less than 1 acre who were formerly treated as not having choice, should now be treated as having choice. Table 2 suggests that non-participants in program areas have higher per capita consumption levels on average, as compared to participants. Incorrectly treating those who actually have choice as not having choice means that with the half acre cut-off rule, effects of participation are *underestimated*. Correcting for mis-targeting should thus *increase* the size of the participation effect.

Table 9 (only variables for the male and female years of membership by round are shown) reports that when households with less than 1 acre of land are treated as having choice, the effect of length of membership between seasons 2 and 1 becomes larger and more significant. Estimates suggest that one year of membership for a female participant reduces the percentage change in per capita consumption caused by a unit shock (using the same interpretation as before) by 6.3% (with an absolute t-statistic of 2.01).

The pooled model was also re-estimated to test for mis-targeting, with all households owning less than 1.66 acres of land having the choice to participate. Table 10 reports that when households with less than 1.66 acres of land are treated as having choice, the effect of length of membership becomes stronger. Results suggest that one year of membership reduces the percentage change in per capita consumption caused by a unit shock by approximately 5.2% (absolute value of the t-statistic is 1.77).

7.2 Seasonality in repayments

Lower consumption variability for experienced members could be directly caused by the fact that they are more likely to be in arrears. This implies that there would be greater seasonality in repayment for such members, since they do not repay their loans in periods with low income realizations, but do repay in favorable times. Using repayment data collected from a few Grameen branches (these data are different from the ones used in the estimation above), it is possible to test for this idea. If this was indeed the case, then amounts repaid by experienced members would be low in the months of October/November (the pre-Aman months which are considered to be the months of most hardship), and would peak in the post harvest months of the Aman crop in December/January. The counter-factual to the idea that only longer duration members demonstrate greater seasonality in repayment would be if short duration members exhibit a similar pattern. Creating a binary variable 'y' which is equal to 1 if a participant pays exactly what they owe and 0 otherwise, the difference between maximum and minimum of mean 'y' (across the Aman, Boro, pre-Aus, and Aus seasons) for less experienced members should be greater than or equal to the difference for more experienced members. Table 11 reports that the repayment data show exactly this pattern.[16]

Another way to rationalize greater seasonality in repayment for experienced members is to assume that they use temporary default to smooth annual consumption flows. Experienced members do not repay when income is low, but "overcompensate" (pay more than what they owe that week) during other months when income is high so that on average, they do not default on their loan. It is possible to analyze the degree of overcompensation for more and less experienced members in the Grameen data. If experienced members use temporary default, then they are likely to overcompensate during the post-harvest Aman season. Using the same definition of "old" as above, the data reveal that approximately 0.51% of old members pay more than what they owe during the Aman season, but a slightly larger 0.6% of new members also overcompensate during this time. Temporary

[16]Note that in order to have equal proportions of "old" (more experienced) and "new" (less experienced) members, an "old" member is defined as someone who joined before 1989 (50.65% of all members).

default may also imply that a larger proportion of experienced members pay less than what they owe during other times of the year such as the Boro and Aus seasons. The data reveal no evidence for this fact. Estimates suggest that 8.55% of experienced members pay less than their due during the Boro and Aus weeks, but a larger 9.52% of younger members do likewise.

Figures 2 and 3 plot the lowess smoothed values of log deviation (from individual specific mean) in amount repaid per week for old and new members (same definition for "old" and "new" as above). Data shown are for the 52 weeks of 1997 (1997 is the only year for which all Grameen branches have information), in order to facilitate comparisons across seasons in the amounts repaid per week. The post harvest time of the Aman crop coincides with weeks 94 to 100, the Boro crop coincides with weeks 65 to 72, and the Aus crop with weeks 77 to 84. It can be seen from the figures that both old and new members show similar patterns in repayments.[17]

7.3 Cohort effects

If the earlier cohort of participants was more able as compared to later cohorts, then easier capital access for more experienced participants could be driven by this characteristic rather than by lower costs of borrowing. Given the data in this study, conventional tests for cohort effects cannot be conducted (note that the estimation already controls for self-selection). But using the education of the household head as a proxy for ability, it is possible to study the relationship between ability and length of membership after conditioning on village-level effects. Plotting villages that have had the program for more than 8.67 years (the upper 90% of length of time a program has been present in a village) and those that have had the program for less than 3.5 years (the lower 10%), there would be evidence of ability bias (self-selection) if the distribution of participants who join first in either case was higher than those who join later. Figure 4 controls for village-level fixed effects and shows a lowess smoothed plot of log education of the household head (adjusted for differences in average schooling) and the lag between program availability and time of joining (separately for the two groups of villages mentioned above). The lag variable (plotted along the x-axis) is the gap between the time the program was set up in the village, and the time that a particular household in that

[17]In fact, as compared to new members, the amount repaid per week by old members during the "Aus" (lean) season is relatively higher.

village joined to become a member. Figure 5 is a magnified view of Figure 4, and shows the data for participants who joined within the first three and a half years (lag <=3.5) of the program's operation in the two groups of villages considered above. From figures 4 and 5, it is evident that there is no consistent pattern to support the claim that high ability people always join first. Thus easier capital access for more experienced members is not being driven by the fact that earlier cohorts were more able.

7.4 Robustness checks

Note that length of membership is not endogenous due to drop out rates. In these data, only 2 households are reported to have dropped out of the program. Additionally, in order to ensure that results are being driven by length of membership as opposed to participation, several tests were conducted. We introduced a dummy for participation into the estimation models. With control for participation, both length of membership and the dummy for participation are insignificant. When the models are estimated with only the participation dummy as the endogenous variable, its effect is insignificant. Hence, it is length of membership that is pivotal in reducing seasonal changes in per capita consumption.

8 Conclusion and Policy Implications

This paper studies the long run benefits of credit program participation by examining the relationship between length of membership in these programs and the ability to smooth consumption across seasons. Although several studies (Morduch 1998, Pitt and Khandker 1998b) have shown that participation in these programs is motivated by consumption smoothing concerns, few studies have analyzed the magnitude of smoothing benefits that result from this participation, or recognized the fact that these benefits accrue differentially across time. By using a structural model that relates changes in consumption across seasons to length of membership, the cost of borrowing faced by a household, and changes in prices and preferences, this paper reaches several interesting conclusions. Length of membership is found to reduce fluctuations in the household's cost of borrowing, which in turn implies an enhanced ability to smooth inter-seasonal consumption changes. Under the assumption that the household specific time varying interest rate captures the shadow price of intertemporal resource transfer, households with experienced participants face lower costs of reducing seasonal consumption differentials.

Using data from the Grameen Bank, BRAC, and BRDB credit programs, results show that when members are stratified by gender in the estimation, a one year increase in membership of a female participant reduces the percentage change in per capita consumption caused by a unit shock (concept of a "unit shock" is discussed above) by 6%. This means that for an interest rate change of one standard deviation around the mean, a household with a female participant who has been a member for the average years of membership (4.15 years) experiences a consumption differential of 11.48% between seasons 2 and 1. But for the same magnitude of the interest rate deviation, a household with a female participant who has been a member for 7.58 years (upper 95^{th} percentile) experiences a consumption change of only 0.85%. Results from the pooled model further substantiate our hypothesis that experienced members are better able to smooth seasonal consumption differentials.

Simulation exercises undertaken to demonstrate welfare implications provide striking results. Estimates from the gender-stratified model show that whereas those with zero years of membership require 6.51% of average annual income as a "certainty equivalent", households with female participants who have been members for 4.15 years (50^{th} percentile value) require a smaller 3.21% of average annual income. Welfare implications for estimates from the pooled model are similar, and provide further support for our hypothesis that as length of membership increases, declining proportions of average annual income are required in order to compensate for higher interest rate variability.

The results of this paper have important implications for program structuring. Estimates presented here demonstrate that although membership has beneficial impacts on a household's consumption smoothing ability, members may become less dependent on programs after a few years of participation. If experienced members face different incentives, then the lending and repayment terms for them might need to be different, as compared to those for less experienced members. Hence for example, perhaps programs could have a limit on the number of years participants in good-standing are automatically eligible for the next loan. Given the results of this paper, a natural cut-off appears to be approximately ten years. Beyond this limit, programs could require the participant to demonstrate that they still depend on microcredit loans, and are not yet in a position to survive independently. Related to this is the idea that eligibility to join (wealth status) needs to be re-evaluated at regular intervals, instead of just at the very beginning as is now the practice. In place of a limit based on length of time of participation, perhaps a cut-off based on the level of household savings and assets may also be appropriate. Once households reach a critical level of assets, perhaps they could be encouraged to "graduate" from microfinance programs. Finally, since weekly meetings and contributions to insurance schemes (for example, the "Emergency Fund") impose costs of time and other resources, perhaps more experienced groups of clients with adequate savings could move to schedules where they repay bi-weekly, and do not need to contribute to group funds.

Recognition of the fact that the nature of participants changes over time will help in making microcredit programs more cost-effective in the future. Anecdotal evidence exists to suggest that experienced members are more likely to default on loans; this may be rational in environments where participants are no longer dependent on loans and where costs of membership are not insubstantial. By examining the effect of participation on consumption smoothing benefits, and by highlighting possible repercussions of differing member incentives on program effectiveness, this paper contributes to the research in the area.

References

Chaudhury, Rafiqul H, "The Seasonality of Prices and Wages in Bangladesh" in *Seasonal Dimensions to Rural Poverty*, edited by R. Chambers, R. Longhurst, and A. Pacey. London: Frances Pinter, 1981.

The Economic Theory of Agrarian Institutions. ed. Pranab Bardhan. Oxford: Clarendon Press, 1989.

Flavin, Marjorie. 1981. "The Adjustment of Consumption to Changing Expectations About Future Income", *Journal of Political Economy*, 89(5): 974-1009.

Foster, Andrew D. 1995. "Prices, Credit Markets and Child Growth in Low-Income Rural Areas", *Economic Journal*, 105(430): 551-570.

Foster, Andrew D. and Mark R. Rosenzweig. 1996. "Technical Change and Human-Capital Returns and Investments: Evidence from the Green Revolution", *American Economic Review*, 86(4), September: 931-953.

The Grameen Bank: Poverty Relief in Bangladesh .ed. Abu Wahid. Boulder: Westview Press, 1993.

Iqbal, F. 1988. "The Determinants of Moneylender Interest Rates: Evidence from Rural India", *Journal of Development Studies*, 24(3): 364-378.

Khandker, Shahidur. 2001. "Does Microfinance Really Benefit the Poor? Evidence from Bangladesh", Manuscript, Asian Development Bank.

Khandker, Shahidur. 2003. "Microfinance and Poverty: Evidence Using Panel Data from Bangladesh", Policy Research Working Paper No. 2945, World Bank.

Latif, Muhammad A. 1994. "Program Impact on Current Contraception in Bangladesh", *Bangladesh Development Studies*, 22(1): 27-61.

McKernan, Signe-Mary. 2002. "The Impact of Microcredit Programs on Self-Employment Profits: Do Non-Credit Program Aspects Matter?", *Review of Economics and Statistics*, 84(1): 93-115.

Menon, Nidhiya. 2005. "Inter-dependencies in Microcredit Groups: Evidence from Repayment Data", *Journal of Developing Areas*, forthcoming.

Morduch, Jonathan. 1998. "Does Microfinance Really Help the Poor? New Evidence from Flagship Programs in Bangladesh", manuscript, Princeton University.

Nanda, Priya. 1999. "Women's Participation in Rural Credit Programmes in Bangladesh and Their Demand for Formal Health Care: Is There a Positive Impact?", *Health Economics*, 8: 415-428.

Pitt, Mark M. and Shahidur Khandker. 1998a. "The Impact of Group-Based Credit Programs on Poor Households in Bangladesh: Does the Gender of Participants Matter?", *Journal of Political Economy*, 106(5): 958-996.

Pitt, Mark M. and Shahidur Khandker. 1998b. "Credit Programs for the Poor and Seasonality in Rural Bangladesh", manuscript, Brown University.

Pitt, Mark M., Khandker, Shahidur, McKernan, Signe-Mary, and M. Abdul Latif. 1999a. "Credit Programs for the Poor and Reproductive Behavior in Low-Income Countries: Are the Reported Causal Relationships the Result of Heterogeneity Bias?", *Demography*, 36(1): 1-21.

Pitt, Mark M. 1999b. "Reply to Jonathan Morduch's 'Does Microfinance Really Help the Poor? New Evidence from Flagship Programs in Bangladesh' ", manuscript, Brown University.

Pitt, Mark M., Khandker, Shahidur, Chowdhury, Omar H., and Daniel Millimet. 2003. "Credit Programs for the Poor and the Health Status of Children in Rural Bangladesh", *International Economic Review*, 44(1): 87-118.

Srinivasan, T.N. 1989. "On Choice among Creditors and Bonded Labor Contracts", in *The Economic Theory of Agrarian Institutions*, edited by Pranab Bardhan. Oxford: Clarendon Press, 1989.

Zeldes, Stephen P. 1989. "Consumption and Liquidity Constraints: An Empirical Investigation", *Journal of Political Economy*, 97(2): 305-346.

Table 1: **Weighted Means and Standard Deviations of Independent Variables**

Independent Variable	**Mean**	**Standard Deviation**
Household cultivable land (in decimals**)	33.12	75.10
Highest educational level completed by household head (in years)	2.45	3.48
Sex of household head (male = 1)	.94	.24
Age of household head (in years)	40.87	12.73
Highest educational level completed by any female household member (in years)	1.57	2.80
Highest educational level completed by any male household member (in years)	3.03	3.78
No adult male in household	.04	.19
No adult female in household	.02	.13
No spouse present in household	.12	.33
Nontarget household	.30	.46
Difference in log price of coarse grain rice between season 2 and season 1	-.04	.10
Difference in log price of coarse grain rice between season 3 and season 2	-.10	.10
Household max. months member program*	52.31	24.15
Household max. female months member program*	49.77	21.95
Household max. male months member program*	54.71	27.03

Sample size is 87 villages and 1,726 households.
* Denotes endogenous variable.
** 1 Decimal = 1/100th of an acre.

Table 2: **Weighted Means and Standard Deviations of Dependent Variables**

	Difference in the log of per capita food expenditure last week between season 2 and season 1 (expenditure in taka per week)	**Difference in the log of per capita food expenditure last week between season 3 and season 2 (expenditure in taka per week)**
Participant (Program areas)	-.11 (.33) N = 1788	-.08 (.25) N = 1788
Non-participant (Program areas)	-.08 (.36) N = 1090	-.07 (.27) N = 1090
Total (Program areas)	-.09 (.35) N = 2878	-.07 (.26) N = 2878
Non program areas	.05 (.34) N = 574	-.10 (.26) N = 574
Aggregate	-.07 (.35) N = 3452	-.08 (.26) N = 3452

Standard errors in parenthesis.

Table 3: **Change in Log per Capita Food Expenditure - Gender Stratified Length of Membership Measured in Years**

Explanatory Variable	**Coefficient**	**t-statistic**
Maximum household years of female membership in season 2	-0.06	-1.82
Maximum household years of female membership in season 3	-0.01	-0.39
Maximum household years of male membership in season 2	-0.02	-0.45
Maximum household years of male membership in season 3	-0.01	-0.22
Log household cultivable land	0.03	0.29
Highest educational level completed by household head	-0.01	-0.38
Sex of household head	0.01	0.07
Age of household head	-0.01	-0.36
Highest educational level completed by adult female in household	0.03	0.28
Highest educational level completed by adult male in household	0.02	0.41
No adult male in household	0.04	0.22
No adult female in household	-0.07	-0.44
No spouse in household	-0.07	-0.35
Non target household	0.06	0.46
Difference in log price of coarse grain rice between season 2 and season 1	0.92	1.08
Difference in log price of coarse grain rice between season 3 and season 2	0.61	6.15
ρ (female, season 2)	-0.05	-1.07
ρ (female, season 3)	0.02	0.41
ρ (male, season 2)	-0.05	-1.24
ρ (male, season 3)	-0.00	-0.07

t-statistics are asymptotic t-ratios.
Fixed effect parameters are not reported.

Table 3 continued: **Change in Log per Capita Food Expenditure - Gender Stratified Length of Membership Measured in Years**

	Female		Male	
Explanatory Variable	**Coefficient**	***t*-statistic**	**Coefficient**	***t*-statistic**
Log household cultivable land	0.09	1.53	0.44	5.77
Highest educational level completed by household head	0.04	0.27	0.03	0.14
Sex of household head	-2.38	-2.84	5.92	2.24
Age of household head	0.06	3.85	-0.06	-2.6
Highest educational level completed by adult female in household	-0.04	-0.43	-0.14	-0.97
Highest educational level completed by adult male in household	-0.08	-0.62	0.30	1.55
No adult male in household	-0.37	-0.33		
No adult female in household			-1.39	-0.63
No spouse in household	-1.05	-1.55	-1.71	-1.43
σ (female)	4.63	33.86		
σ (male)			5.88	27.17
σ (outcome) (followed by t-statistic)	0.27(53.864)			

t-statistics are asymptotic t-ratios.
Fixed effect parameters are not reported.

Table 4: **Change in Log per Capita Food Expenditure - Pooled Duration of Membership Measured in Years**

Explanatory Variable	**Coefficient**	***t*-statistic**
Maximum household years of membership in season 2	-0.04	-1.64
Maximum household years of membership in season 3	-0.01	-0.38
Log household cultivable land	0.01	0.36
Highest educational level completed by household head	-0.01	-0.38
Sex of household head	-0.03	-0.41
Age of household head	-0.01	-0.34
Highest educational level completed by adult female in household	0.01	0.36
Highest educational level completed by adult male in household	0.02	0.40
No spouse in household	-0.07	-0.43
Non target household	0.05	0.45
Difference in log price of rice between seasons 2 and 1	1.30	2.58
Difference in log price of rice between seasons 3 and 2	0.61	6.55
ρ (season 2)	-0.05	-1.58
ρ (season 3)	0.02	0.52

t-statistics are asymptotic *t*-ratios.
Fixed effect parameters are not reported.

Table 5: **Estimates from the Gender Stratified Model Variance of Log Consumption Change Between Season 2 and Season 1**

Duration of Membership of Female Participant	**Variance of Log Consumption Change**
0 years	0.014
4.147 years (50^{th} percentile)	0.003
5.500 years (75^{th} percentile)	0.001
7.583 years (95^{th} percentile)	0.000

Table 6: **Estimates from the Pooled Model Variance of Log Consumption Change Between Season 2 and Season 1**

Duration of Membership of Participant	**Variance of Log Consumption Change**
0 years	0.014
4.359 years (50^{th} percentile)	0.007
5.667 years (75^{th} percentile)	0.005
8.083 years (95^{th} percentile)	0.002

Table 7: **Gender Stratified Estimates Percentage of Annual Income Required In Order To Be Indifferent**

Duration of Membership of Female Participant	**"Θ" As A Percentage of Average Annual Income**
0 years	6.51%
4.15 years (50^{th} percentile)	3.21%
5.50 years (75^{th} percentile)	1.80%
7.58 years (95^{th} percentile)	0.18%

Table 8: **Pooled Estimates Percentage of Annual Income Required In Order To Be Indifferent**

Duration of Membership of Participant	**"Θ" As A Percentage of Average Annual Income**
0 years	8.55%
4.36 years (50^{th} percentile)	6.04%
5.67 years (75^{th} percentile)	5.24%
8.08 years (95^{th} percentile)	3.74%

Table 9: **Gender Stratified Model**
Test for Mis-targeting: Households with < 1 Acre Have Choice
Dependent Variable: Change in Log per Capita Food Expenditure
Duration of Membership Measured in Years

Explanatory Variable	Coefficient	*t*-statistic
Maximum household years of female membership in season 2	-0.063	-2.012
Maximum household years of female membership in season 3	-0.013	-0.681
Maximum household years of male membership in season 2	-0.017	-0.300
Maximum household years of male membership in season 3	-0.003	-0.133

t-statistics are asymptotic *t*-ratios.
Fixed effect parameters are not reported.

Table 10: **Pooled Model**
Test for Mis-targeting: Households with < 1.66 Acre Have Choice
Dependent Variable: Change in Log per Capita Food Expenditure
Duration of Membership Measured in Years

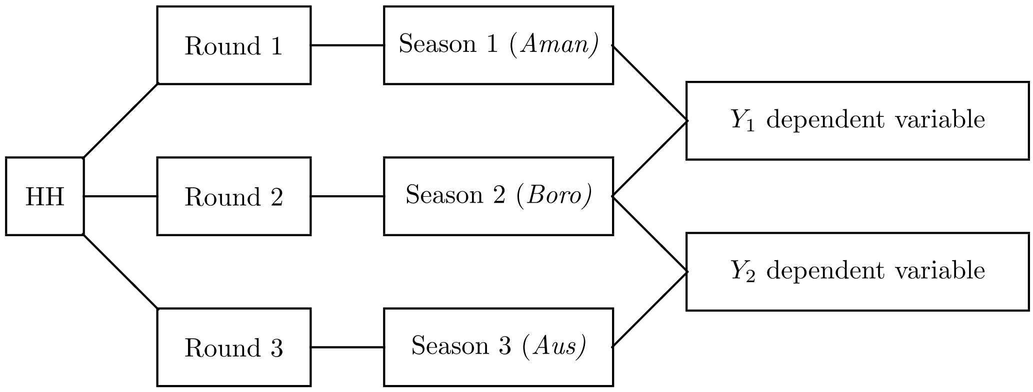
Explanatory Variable	Coefficient	*t*-statistic
Maximum household years of membership in season 2	-0.052	-1.765
Maximum household years of membership in season 3	-0.012	-0.479

t-statistics are asymptotic *t*-ratios.
Fixed effect parameters are not reported.

Table 11: **Average Repayment By Seasons in %**
"Old" = 1 If Date of Joining Was Before 1989
y=1 if Pay What You Owe That Week

Member Category	Pre-Aman	Aman	Boro	Aus	Difference (Max. - Min.)
Old	42.0	42.9	18.6	31.1	24.3
New	44.5	46.7	11.9	27.5	34.8

Structure of the Data



Y_1 is the difference of log per capita food expenditure last week between season 2 and season 1 (expenditure in taka per week).
Y_2 is the difference of log per capita food expenditure last week between season 3 and season 2 (expenditure in taka per week).

Figure 1: **Percentage of Annual Income Required for Indifference**

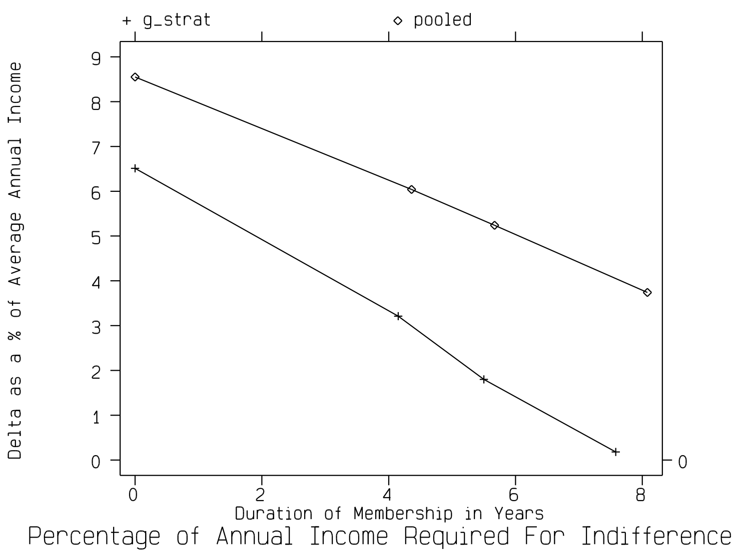


Figure 2: **Old Members - Log Deviation in Weekly Repayment**

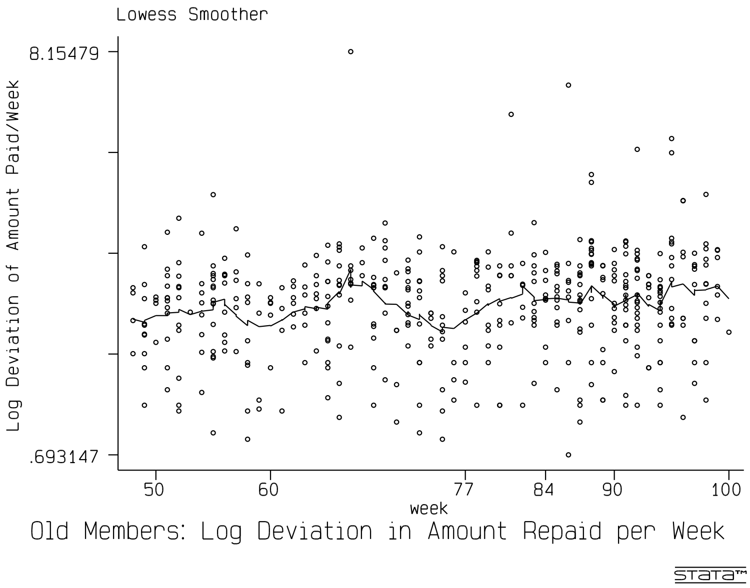


Figure 3: **New Members - Log Deviation in Weekly Repayment**

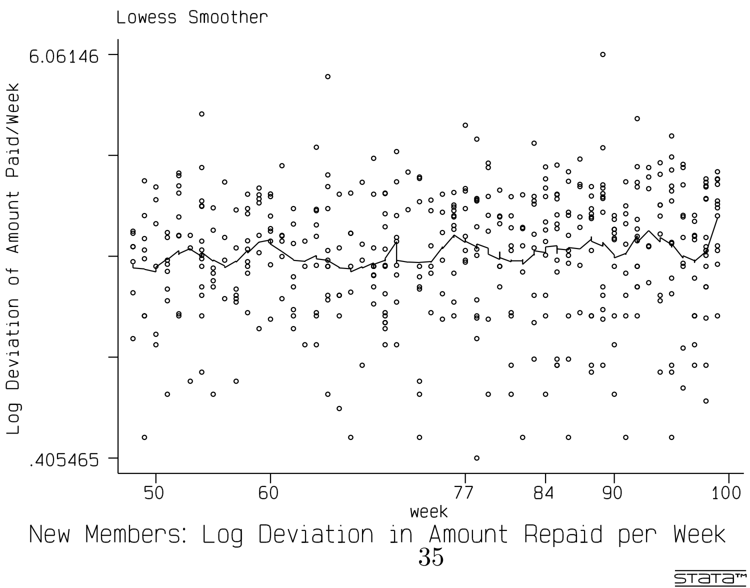


Figure 4: **Lag between Program Availability & Joining vs. Household Head's Schooling**

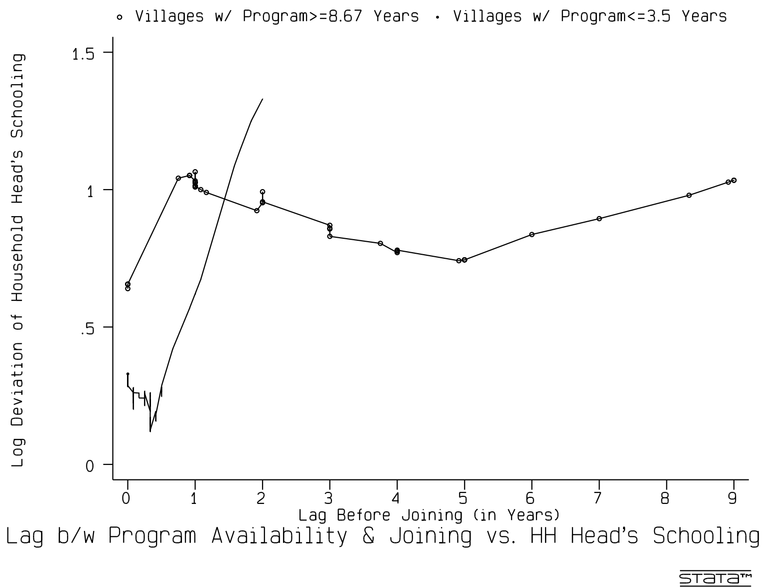


Figure 5: **Lag between Program Availability & Joining vs. Household Head's Schooling**