Income Transfers, Intra-Household Resource Allocation and Marriage Markets *

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Abstract: This paper develops a marriage market model which provides a basis for understanding both changes in intra-household resource allocation and marriage market turnover as a result of changes in the women's outside options. Contracts determining the resource sharing rule between partners are negotiated at the beginning of the contract, and hence conditions of the marriage market at the time of the contract matter. The sharing rule affects intra-household resource allocation in subsequent periods and equilibrium levels of marriage market turnover as new information of the union is revealed. Using experimental evidence from a conditional cash transfer program in rural Mexico, I present credible support of the model in terms of changes in intra-household resource allocation and marital turnover. Two years after the start of the program, expenditures shares in girls' clothing changed substantially in households who received the cash transfers relative to the comparison group. Second, marital union dissolution patterns among indigenous women who received cash transfers increased by 0.6 percentage points over the two-year period, a substantial increase in separation rates among eligible women. The extent of behavioral responses varies substantially depending on women's economic opportunities outside of the current marriage.

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1. Introduction

Unequal 'power' among household members and conflict over resources allocation is a common phenomenon in less developed countries. The idea that improving women's status in the household is beneficial to the household as a whole is often advanced as an argument in favor of social policies targeted towards women. Consistent with this argument, recent empirical evidence has shown that the level of unearned income received by a female in the household leads to an improvement in children's health (Thomas (1990), Duflo (2000)); evidence which is inconsistent with the unitary household model of resource allocation. Various theoretical alternatives to the unitary model have been proposed: collective household models, which treat the household as composed of various members with heterogeneous preferences reaching a Pareto-efficient allocation (Chiappori (1998, 1992), Browning and Chiappori (1998), Manser and Brown (1980), McElroy (1981)), and non-cooperative bargaining models, which predict that conflict over resource allocation among heterogeneous household members could result in inefficient allocations (Lundberg and Pollak (1993, 1994), Adam et al (2003)). However, there is no complete understanding of the mechanisms that may lead to these outcomes; for example, these studies significantly limit the extent to which contracts in the marriage market explicitly determine intrahousehold allocations.[1]

This paper addresses this gap in the literature by developing an application of a simple model of contracting in marriage markets that provides a basis for understanding both intra-household resource allocation and marriage market outcomes. Since there is limited mobility of individuals in the marriage market, contracts determining the resource sharing rule between partners are negotiated once at the beginning of the contract, and hence conditions of the marriage market at the time of the contract matter; the contracted sharing rule affects intra-household resource allocation in subsequent periods and equilibrium levels of marriage market turnover as new information of the union is revealed.

[1] One exception is Chiappori et al (2002)'s contribution in modeling the sex ratio as a distribution factor within the collective framework.

I analyze conditions under which information on changes in the unearned income of women, by improving women's outside option relative to the household's production within marriage, reduce the share of household income under control of the husband, and therefore shift intrahousehold resource allocation to goods which women may have preference for. Moreover, the model highlights the composition of individuals who respond to the change in unearned income of the women due to heterogeneity of the participants in the marriage market. Essentially, the contracted payments to men are reduced significantly more in households where women have better outside options if the degree of complementarity in household production between her unearned income and her attributes is reduced in a union.

As a result of the changes in their income share, men will evaluate the benefits of marriage according to the current contract and the expected sharing rule in a future contract, given the expected changes in the sharing rule in a future union. Marital dissolution is thus an optimal response to information on the gains to marriage in the present and future potential contracts. However, marital dissolution will crucially depend on the random component of household productivity at the time of contracting; men with high income shares relative to the expected income shares from a new contract will choose to remain in union, whereas men with relatively low income shares will choose to dissolve the union. These effects are augmented by the magnitude of the change in income shares due to women's outside options.

Subsequently, using experimental evidence from a conditional cash transfer program in rural Mexico, I present credible support of the model in terms of its predictions on intra-household resource allocation and marital turnover decisions. In 1997, the Mexican government initiated the Progresa Program, a large-scale human development conditional cash transfer program for marginalized households in rural areas. [2] It consists of cash transfers *provided to mothers* contingent on certain requirements in terms of children's school attendance and family-level visits to health services. 506

[2] Progresa was renamed OPORTUNIDADES under the Fox Administration. For consistency, I will refer to the program as Progresa in this text.

communities were selected to participate in an experimental evaluation of the program; the communities were randomly divided into two groups, the treatment group being phased into the program in March-April 1998 and the control group in November-December 1999. This randomized design allows for clear estimation of the effect of changes in unearned income of the women on intrahousehold allocation and marital decision-making responses.

The empirical study takes into account the large degree of heterogeneity in ethnic background among the population of eligible households, since forty (40) percent of the women under study report being indigenous.[3] In this setting, it will be important to take into account the substantial degree of heterogeneity in social norms regarding household decision-marking processes and marital formation and dissolution. A small body of ethnographic evidence suggests that variation in gender inequality and 'power' between mestizo and indigenous communities is evident. Whereas in mestizo households, the role of *machismo* enforces highly unequal gender disparities within the household, social scientists have found in traditional indigenous cultures a striking lack of it (Chiñas, 1992; Elmendorf, 1972; Wolf, 1959, 1964).[4] In addition, certain indigenous communities tend to have much more fluid marriage markets; for example, the existence of trial marriages and cohabitation is very prevalent. The existing evidence suggests that communities with lower degrees of social stigma and greater degrees of gender equality will lead to a higher incidence of movement in the marriage market as a result of the program.

Consistent with the theoretical and ethnographic evidence, this study finds that, two years after the start of the program, expenditures shares in girls' clothing changed substantially in households who received the cash transfers relative to the comparison group, evidence consistent with an increase in the share of household income spent in accordance with women's preferences (Thomas (1990),

[3] Women report speaking an indigenous language.
[4] A formal definition of *machismo* refers to "a sense of exaggerated masculinity or a cult of virility whose chief characteristics are extreme "aggressiveness and intransigence in male-to-male interpersonal relationships and arrogance and sexual aggression in male-to-female relationship" (Stevens, 1972, p.315)" (Elmendorf, 1977, p.141).

Duflo (2001), Rubalcava et al (2001)). Furthermore, I find evidence that the extent of the intra-household resource re-allocation process varies according to the women's characteristics; changes in budget shares are highest for women with young children, who completed primary schooling or higher, but who do not have sources of cash income; that is, women with the best outside options a priori.

Second, the study finds that marital union formation and dissolution patterns *among indigenous women* who received cash transfers changed substantially. Their marital dissolution rates increased by 0.6 percentage points over the two-year period, a substantial increase in separation rates among eligible women. The effect on union formation of indigenous women is analogous: new marital union formation rates increase by 5.6 percentage points after two years, a 70 percent increase in union formation. Moreover, these results vary significantly depending on the demographic characteristics and the socio-economic status of the households, which provide evidence of substantial variation in the separation and union formation incentives given the economic opportunities of the women.

This paper potentially contributes to two different literatures. First, the model re-incorporates the idea that the marriage market is an important factor in determining intrahousehold resource allocation (Becker, 1991), by providing an explicit *behavioral* mechanism of how marriage markets at the time of contracting and women's outside opportunities interact to affect intrahousehold resource allocation and marital turnover decisions. Evidence from the U.S. consistent with this view is provided by Grossbard-Schechtman (1993), Grossbard-Schechtman and Neideffer (1997), Angrist (2002), and Chiappori et al (2002), who show that labor force participation and labor supply of married women decreases as a result of increases in the sex ratio.

Second, the study provides credible empirical evidence of how changing women's outside options affect both intrahousehold resource allocation and marriage market turnover decisions within the same population; existing studies that test these theories may suffer from identification problems such as omitted variables bias and reverse causality (Thomas (1990), Dercon and Krishnan (2000), Duflo (2000)). A recent paper by Rubalcava, Teruel, and Thomas (2003), using a similar identification strategy as the one presented below present evidence that households in which women gain an increase in unearned income tend to spend a larger share of the budget on educational expenditures, child goods, and food.

In terms of policy implications, the model and the empirical evidence suggests that income transfers to women in union may improve their well-being in both married and divorced states. However, the effects of this targeting of government transfers may affect children upon divorce. Although it is not possible to identify the effects of divorce on children's well-being in this context, Reyes (2003) finds that marital dissolution in the U.S. as a result of changes in the divorce legislation of the 1970s increased suicide rates of teenagers in the 15-19 age group. In the case of poor rural Mexican households, the effects on human capital accumulation may be substantial.

This study is also related to another strand of the literature: theoretical and empirical studies of the causes of marital dissolution in developed countries. Standard economic models of the marriage market, which assume the existence of a unitary household making family structure decisions, suggest that unexpected changes in the perceived match quality (e.g., earnings opportunities) of one of the partners could lead to the dissolution of the marriage, since information available at the time of marriage should have been taken into account when making the marriage decision (Becker et al (1977), Weiss and Willis (1997)).[5,6] Therefore, the model predicts that the separation decision should be caused by an unanticipated change in the information of the match quality of the couple. Various empirical studies have attempted to identify the effects of unexpected changes in income in the likelihood of divorce; however, problems of identification due to omitted variables bias and reverse causality plague this literature (Becker et al (1977), Weiss and Willis (1997), Charles and Stephens

[5] To my knowledge, the theoretical literature does not explain why negative shocks to husbands increases the probability of dissolution, whereas a negative shock to the wife decreases this probability. This should be an important theoretical contribution to the topic.

[6] The theoretical literature does not distinguish between divorce, separation and annulments.

(2001)). The most convincing evidence of this pattern is for low-income households in the United States. Groeneveld et al (1980) show experimental evidence that government transfers to households from the SIME/DIME program increased the probability of separation of eligible whites and African-Americans by 72 and 56 percentage points, respectively; they find no effect on marital dissolution for chicano households. Future theoretical and empirical work in this area should explicitly take into account the idea of heterogeneous preferences within the household and conflict over allocation.

The paper is structured as follows: section 2 provides the theoretical framework and predictions of the collective model, followed by the main testable implications and the empirical strategy. Section 3 gives a brief overview of the Progresa conditional cash-transfer program, the data used in the analysis, and the social and economic context. Section 4 presents the empirical results, and finally, section 5 concludes.

2. Theoretical Framework and Derivation of the Empirical Strategy

2.1 Model of Household Production and Consumption Contracts and the Marriage Market[7]

Consider a number of M risk-averse male and N risk-neutral female individuals (M>N) making marital union decisions. Household production within marriage is represented as the production of a single good, and the individual's utility per period associated with the consumption of c units of the good is given by U(c). Both males and females have a discount factor equal to β, and the probability that any will die in any given period is 1 – μ. Each of the individuals is assumed to have access to a technology that requires one partner (e.g., reproduction). The quantity of output from this technology is given by $\Phi_t(x_t^F, x_t^M, y_t^F, y_t^M, \varepsilon_t)$, where Φ_t represents the state of household production within marriage at time *t*; x_t^F, x_t^M are the individual female and male's characteristics which affect household productivity; y_t^F, y_t^M are the unearned income of females and males,

[7] This model is an application and extension to the marriage market of the Beaudry and DiNardo (1991) model of implicit contracts in the labor market.

respectively, and ε_t is a random shock to household productivity at time *t*. Assume that the gains to marriage or household production in union are well approximated by the following process:

$$\Phi_t(x_t^F, x_t^M, y_t^F, y_t^M, \varepsilon_t) = \Phi^*(x_t^F, x_t^M, y_t^F, y_t^M) + \varepsilon_t, \qquad \varepsilon_t \text{ is i.i.d.} \quad (1)$$

where Φ^* is the long-run level of household production. Men and women have outside options $A_t^M = \Phi^*(0, x_t^M, 0, y_t^M)$ and $A_t^F = \Phi^*(x_t^F, 0, y_t^F, 0)$, respectively. Assume that men in this population have reservation 'wage' outside of marriage A_t^M with distribution function $F(A_t^M)$. When men do not have access to capital markets, women may have incentives to offer marital contracts that protect men against the risks associated with the productivity shocks in household production. Competition in the marriage market will force such contracts to offer zero expected profits to women.

$$\underset{\{\omega_{t=I}\}}{Max} \sum_{i=0}^{\infty} (\beta\mu^2)^i E_t[U(\omega_{t+i})] \qquad \text{(P1)}$$

$$\text{subject to} \qquad \sum_{i=0}^{\infty} (\beta\mu^2)^i E_t[\Phi_{t+i} - \omega_{t+i}] = \sum_{i=0}^{\infty} (\beta\mu^2)^i E_t[A_{t+i}^F]$$

The solution of P1 is well known: the optimal contract is a fixed wage contract conditional on the survival of both parties. The wage paid at time *t+j* in a contract negotiated at time *t* is

$$\omega_{t+j,t} = (1-\beta\mu^2)\Phi_t + \beta\mu^2(\Phi^* - A_{t+j}^F) \qquad \text{for all } t. \qquad (2)$$

The proportion of men who will decide to participate in a union at time t is determined by the equilibrium contract payment at time t and the outside 'wage' of each individual. The marriage market will adjust such that the marginal worker is indifferent between accepting the marital union contract and postponing until next period the decision to form a union. The equilibrium condition related to this indifference relationship is given by

$$U(A_t^{M*}) + (\beta\mu)E_t[V(\omega_{t+1,t+1}, \Phi_{t+1}) \mid \Phi_t] = V(\omega_{t,t}, \Phi_t). \qquad (3)$$

$V(\omega, \Phi_t)$ is a function representing the discounted expected utility associated with being in a union with contract payment ω when the state of the gains to marriage is Φ_t. Individuals with own $A_t^M > A_t^{M*}$ remain single, while individuals with $A_t^M \leq A_t^{M*}$ join a union.

I proceed by stating assumptions on the household production function under which changes in the unearned income of women, by improving women's outside option relative to the household's production within marriage, reduce the share of household income under control of the husband, and therefore shift intrahousehold resource allocation to goods which women may have preference for. Denoting $\partial f(x_1,...,x_n)/\partial x_i$ as $f_i(x_1,...,x_n)$, these are stated below:

Assumption 1 (Household Production Function):

$$\Phi^*_{k_t}(x_t^F, x_t^M, y_t^F, y_t^M) > 0, \qquad \Phi^*_{k_t k_t}(x_t^F, x_t^M, y_t^F, y_t^M) < 0 \qquad \forall k_t \in \{x_t^F, x_t^M, y_t^F, y_t^M\}$$

Assumption 2 (Household Production Function):

$$\beta\mu^2\Phi^*_1(x_t^M = 0) > \Phi^*_1(x_t^M = \underline{x}_t^M) \qquad \beta\mu^2\Phi^*_{31}(x_t^M = 0) > \Phi^*_{31}(x_t^M = \underline{x}_t^M)$$

The first part of Assumption 2 just states that the marginal productivity of a female attribute is decreasing in the male attribute to the extent that the discounted marginal productivity of the female attribute outside of marriage is greater than that in a union matched with a male of the lowest possible realization of the same attribute. The second part states that the degree of production complementarity between unearned income and the female attribute is decreasing in the male's attribute to the extent that the discounted degree of complementarity outside of a union is greater than the one in a union matched with a male of the lowest possible realization of the same attribute. Although these are strong (and may seem like arbitrary) assumptions of the household production function, they are likely to hold for households production functions in which male and female attributes are substitutes or, formally, when: $\Phi^*_{12}(x_t^F, x_t^M, y_t^F, y_t^M) < 0$ and $\Phi^*_{312}(x_t^F, x_t^M, y_t^F, y_t^M) < 0$, respectively.

The following propositions characterize the changes in intra-household resource allocation and marital dissolution rates as a result of an increase in unearned income of women in both single and married states. The first result states that an increase in unearned income of women leads to a reduction in the contracted "payment" to the husband due to the extent of decreasing marginal productivity of income in household production and competitiveness in the marriage market. This straightforward outcome results from the marital contract rule that household members will share income such that women will never be worse off than in the single state, on average. Since the increase in unearned income of women improves her outside option to a larger extent than it does within marriage (Assumption 2, Part A), the payment to the husband will necessarily fall. Moreover, the contracted payments to the males are reduced significantly more in households where women have better outside options if the degree of production complementarity between her unearned income and the female attribute is decreasing in the male's attribute according to Assumption 2. These results are formalized in the following Proposition:

Proposition 1: Suppose Assumptions 1 and 2 hold. Then

$$\partial \omega_{t+j,t} / \partial y^F < 0 \qquad \text{and} \qquad \partial^2 \omega_{t+j,t} \big/ \partial x_t^F \partial y^F < 0 .$$

Proof [see appendix].

As a result of the changes in their income share, men will evaluate the benefits of marriage according to the current contract and the expected income gain in a future contract, given the expected changes in the sharing rule in a future union. Marital dissolution is thus an optimal response to information on the gains to marriage in the present and future potential contracts. However, marital dissolution will crucially depend on the random component of household productivity at the time of contracting; men with high income shares relative to the expected income shares from a new contract will choose to remain in union, whereas men with relatively low income shares will choose to dissolve the union. This marital dissolution effect is augmented by the magnitude of the change in income shares due to women's outside options.

<u>Proposition 2:</u> Suppose Assumptions 1 and 2 hold. Then,

$$\partial \Pr[A_{t+p}^{M} > A_{t+p}^{M*}] \big/ \partial y_{t+p}^{F} = \begin{cases} > 0 & if \quad H_{t+p}^{1} > 0 \\ < 0 & if \quad H_{t+p}^{1} < 0 \end{cases},$$

where $H_{t+p}^{1} = \dfrac{\partial U(\omega_t)}{\partial \omega_t} - \beta\mu^2 \cdot \dfrac{\partial E_{t+p}\left[U(\omega_{t+p+1}) \mid \Phi_{t+p}\right]}{\partial \omega_{t+p+1}}$. In addition,

$$\partial^2 \Pr[A_{t+p}^{M} > A_{t+p}^{M*}] \big/ \partial (y_{t+p}^{F})^2 = \begin{cases} > 0 & if \quad H_{t+p}^{1} > 0 \quad and \quad H_{t+p}^{2} < 0 \\ < 0 & if \quad H_{t+p}^{1} < 0 \quad and \quad H_{t+p}^{2} > 0 \end{cases}$$

where $H_{t+p}^{2} = \dfrac{\partial^2 U(\omega_t)}{\partial {\omega_t}^2} - \beta\mu^2 \cdot \dfrac{\partial^2 E_{t+p}\left[U(\omega_{t+p+1}) \mid \Phi_{t+p}\right]}{\partial {\omega_{t+p+1}}^2}$.

Proof [see appendix].

The intuition for this result comes from the structure of the contract, in which the male receives an income share *in every period* equal to the difference in the total household production at the time of contracting (Φ_t^*), and the income equivalent to the household production of the matched woman outside of marriage (discounted by $\beta\mu^2$). Since this contracted payment depends on the realization of the random productivity shock at the time of contracting, males with a relatively high (low) realization of ε_t will choose to remain in the (leave the) current union, given their expected income share in the potential new union. This contingency implies that the rate of marital dissolution may increase or decrease in the population, depending on the unobserved random shocks to household productivity at the time of contracting and the extent of risk-aversion of males.

In addition, Proposition 2 shows that the magnitude of the effect on the probability of separation is largest for households where women have better outside options. This result relies on the assumption of decreasing returns to scale of female attributes in household production, since this implies that income shares to husbands are smaller in unions with females of more productive attributes, and on the risk-aversion of males. Again, the contingency of the contract on the productivity shock at the time of contracting implies that the separation decisions will be taken relative to the realization of this random shock.

2.2 Empirical Strategy

The empirical strategy to identify the effects of government transfers to mothers on both the extent of intra-household resource allocation and the likelihood of separation is based on the random assignment of households to treatment and control groups in the Progresa program and the amount of the income transfer received by mothers (see section 3 for a description).

According to Proposition 1, we should observe a proportional increase in the share of household income under control of females and a reduction in household income under the control of males. Since it is not possible to observe household income shares, I focus the analysis on expenditure shares on child expenditures, 'goods' for which females may have arguably greater preference for than males, as shown by previous empirical results in Brazil and South Africa (Thomas (1990), Duflo (2001)). To the extent that household members allocate resources optimally according to their individual preferences, subject to their personal income budget constraints, then total demand for children's goods is a function of female income and male income (see Chiappori (1992) for a discussion of optimal intra-household resource allocation). Additionally, assuming that women have higher income elasticities of child expenditures than men due to their differences in preferences, and taking the income shares results discussed above, we should observe an increase in child expenditures at the household level.[8] Given the latter assumption, I estimate the following equation:

[8] Formally, note that consumption demand for children's goods is a function of female income and male income, or $c^{CH} = c^{CH}(\omega^F, \omega^M)$. Therefore,

$$s_{ict}^{CH} = \alpha + \theta y_{ict}^{PR} + \beta_1 Y_{ict} + X_{ict}^{\prime}\beta_2 + P_t^{\prime}\delta + \varepsilon_{ict} \tag{4}$$

where s_{ict}^{CH} is the expenditures budget share spent on child expenditures, such as boys' and girls' clothing in household *i* in village *c* at time *t*; y_{ict}^{PR} is the amount of cash transfers received by the woman as a result of the program; Y_{ict} are total household expenditures; X_{ict} is a set of individual, partner, and household (including detailed demographic) controls; P_t is a vector of time controls; and ε_{ict} is a disturbance term, which is allowed to be correlated at the community level. In this specification, θ is the average effect of women's unearned income on the share of the household budget spent on children's clothing, conditional on the size of the household budget. Equation (4) can be easily expanded to account for time-varying effects of the intervention and heterogeneous average treatment effects for different sub-groups, which allows us to test for differences in the response of intra-household resource allocation depending on female attributes, as discussed in Proposition 1.

However, since the program take-up decision may not be uncorrelated with unobserved factors influencing intra-household resource allocation, and total consumption may be an endogenous variable in equation (4), I use the eligible amount of transfer for the household and a household-level rainfall shock indicator as instrumental variables for the former. Therefore, I estimate the following set of equations:

$$y_{ict}^{PR} = \pi_{02} + \pi_{12} EY_{ict}^{PR} + \pi_{22} R_{ict} + X_{ict}^{\prime}\pi_{32} + P_t^{\prime}\pi_{42} + \upsilon_{2,ict} \tag{5}$$

$$Y_{ict} = \pi_{03} + \pi_{13} EY_{ict}^{PR} + \pi_{23} R_{ict} + X_{ict}^{\prime}\pi_{33} + P_t^{\prime}\pi_{43} + \upsilon_{3,ict} \tag{6}$$

$$\frac{\partial c^{CH}}{\partial y_{t+p}^{PR}} = \frac{\partial c^{CH}}{\partial \omega^F} \cdot \frac{\partial \omega^F}{\partial y^F} + \frac{\partial c^{CH}}{\partial \omega^M} \cdot \frac{\partial \omega^M}{\partial y^F}$$

Since $\partial \omega^F / \partial y^F > 0$ and $\partial \omega^M / \partial y^F < 0$ according to Proposition 1, and $\partial c^{CH} / \partial \omega^F > \partial c^{CH} / \partial \omega^M$ by assumption of gender-based preferences, then $\partial c^{CH} / \partial y^F > 0$.

where EY_{ict}^{PR} is the eligible transfer amount according to the program specification (based on gender and demographic characteristics of the household), and R_{ict} is an indicator variable for the household suffering from a flood in the past six months. To the extent that rainfall variation, which affects farm productivity, affects total household income, but not the composition of expenditures, the latter instrumental variable will satisfy the exclusion restriction.

In addition, according to Proposition 2, households would potentially choose to separate as a function of the unearned income of females. Therefore, I estimate the following model:

$$D_{ict} = \alpha + \theta y_{ict}^{PR} + X_{ict}'\beta + P_t'\delta + \eta_{ict} \tag{7}$$

where D_{ict} is an indicator variable equal to 1 if the union dissolves and 0 otherwise; y_{ict}^{PR}, X_{ict}, and P_t are defined as above; and η_{ict} is a disturbance term, which is allowed to be correlated at the community level; in the empirical section, I estimate probit specifications of this model. Equation (7) can also be easily expanded to account for time-varying effects of the intervention and heterogeneous average treatment effects for different sub-groups, which allows us to test for differences in the marital dissolution effects conditional on female attributes, as discussed in Proposition 2. Due to the potential endogeneity of or selective take-up of the cash transfers, which may lead to biased estimates, I also present reduced-form estimates using T_{ct}^{PR}, the community-level treatment assignment indicator variable, as the explanatory variable of interest.

Note that in the empirical section, I will be estimating post-treatment equations (4) - (6) for the sub-sample of households who remain in union. Therefore, to the extent that unions which would experience the largest changes in income share allocations are more likely to dissolve, this would result in sample-selection problems which could bias θ in equation (4) downwards. In a future version of the study, I will take this potential source of bias into consideration by estimating Lee (2002) treatment effect bounds.

An issue in the empirical analysis is the extent of sample attrition. If being out-of-sample is correlated with the likelihood of receiving treatment, then this could lead to bias in the coefficient estimates. Sample attrition rates are approximately 10 percent and 21 percent for the samples of women in union and separated at baseline, respectively (Table A1, regressions 1 and 5). Although attrition rates are balanced across treatment groups, the likelihood of attrition is highly correlated with individuals' observable characteristics (regressions 2-4, 6-8). Therefore, to reduce the extent of potential attrition bias, we control for women, partner, and household characteristics in all specifications. Future versions will include Lee (2002) bounds on program impacts, in order to assess the robustness of the results to differential attrition between the treatment and control groups.

3. Progresa Program, Data, and Social Context

3.1 Overview of Progresa Program

In 1997, the Mexican government initiated a large-scale Education, Health, and Nutrition Program (Progresa) aimed at improving the human development among children in marginalized rural areas in Mexico. The program targets the poor in marginal rural communities, where 40 percent of the children from poor households left school after the primary level. The program provides cash transfers to the mothers of over 2.6 million children conditional on school attendance, at an annual cost of approximately one billion dollars, or 0.2 percent of Mexico's GDP.

The education component of Progresa consists of subsidies provided to mothers, contingent on their children's regular attendance to school. These cash transfers are available for each child attending school in grades three (3) to nine (9) of primary and lower secondary school, and range from $70 to $255 pesos, depending on the gender and grade level the child is attending (with a maximum of $625 per month per family in 1998). Overall, the program transfers are important, representing 22 percent of the income of beneficiary families.

A distinguishing characteristic of Progresa is that it included a program evaluation component from its inception. Progresa was implemented following an experimental design in a subset of 506 communities located across seven states: Guerrero, Hidalgo, Michoacán, Puebla, Querétaro, San Luis Potosí, and Veracruz. Among these communities, 320 were randomly assigned into a treatment group, with the remaining 186 communities serving as a control group, thus providing an opportunity to apply experimental design methods to measure its impact on various outcomes.

Within these selected communities, a poverty indicator was constructed using the household income data collected from the baseline survey in 1997. A discriminant analysis was then separately applied in each of the seven regions in order to identify the household characteristics that best classify poor and non-poor households. Eligible households were identified on the basis of this welfare index, the resulting score of a discriminant analysis (see Skoufias et al (2001) for a more detailed description of the targeting process). While household eligibility was determined within all communities, only households below a welfare threshold and within the treatment villages became PROGRESA beneficiaries (a diagram which is often used to explain the design of the program is depicted in Figure 1).

3.2 Data

Since the baseline census in 1997, the program has conducted extensive biannual interviews on the 24,000 households of the 506 communities during March 1998, October 1998, May/June 1999, and November 1999, at the time of the implementation of the experimental phase of the program. Each survey is a community-wide census containing detailed information on household demographics and household income, expenditures and consumption, schooling, health, and nutrition at the individual and household levels. The surveys in October 1997, October 1998, May/June 1999, and November 1999 (survey rounds 1, 3, 4, and 5, respectively) collected information on the marital status of each person in the household 8 years old and older. From this data, I construct a variable equal to one if the mother was separated or divorced at each survey round, and zero otherwise.

In order to conduct the analysis, the sample is selected as all eligible mothers (belonging to households classified as poor) between the ages of 16 and 55. The sample is then subdivided into two groups: (i) mothers in a union (marriage or cohabitation), and (ii) separated mothers (separated or divorced) as of October 1997, which results in a sample of 13,505 and 618 mothers, respectively.

Given the random assignment of villages to either treatment or control groups, the characteristics of the mothers between these groups should be balanced. Mean levels of the observable characteristics used in the empirical analysis by treatment group, and their difference, are reported; there are no statistically significant differences in the observed characteristics of these individuals, in general (see Table 1).[9]

3.3 Descriptive Statistics and Social Context

Individuals in this sample come from poor socio-economic status households, since Progresa is targeted to poor individuals in marginalized rural communities in Mexico (see Table 1). Mothers in a union at baseline have 3.6 children, and have completed 3.1 years of schooling, on average. Forty (40) percent of the women in the sample come from an indigenous background, and 23 report being in a cohabiting union (Table 1, Panel A). Mothers separated at baseline tend to be of a slightly higher socio-economic status: they are two years younger than women in a union (32.0 versus 33.8 years old), have 2.5 children and have completed 3.9 years of education, on average (Table 1, Panel B). Note that only 29 percent report speaking an indigenous language.

In the case of rural Mexico, it will be important to take into account the substantial degree of heterogeneity in social norms regarding household decision-marking processes and marital formation and dissolution. A body of ethnographic evidence suggests that variation in gender inequality and `power' between mestizo and indigenous communities is evident. Whereas in mestizo households, the role of *machismo* enforces highly unequal gender disparities within the household, social scientists have found in traditional indigenous cultures a striking lack of it (Chiñas, 1992; Elmendorf, 1972;

[9] Behrman and Todd (1999) conduct an exhaustive analysis of the degree of success of the random assignment of villages in the PROGRESA Program, and conclude that the randomization was successful.

Wolf, 1959, 1964).[10] In addition, certain indigenous communities tend to have much more fluid marriage markets; the existence of trial marriages or cohabitation is very prevalent, as made clear by the following reference:

> "A large number of women are forced to substitute for marriage non-legal living arrangements in the form of transient relationships of short or ephemeral duration, consensual unions of some durability, and polygamous concubinal unions with married men." (Davis, 1964, p.37, as quoted in Youssef, 1974) [11]

Chiñas (1992) also provides evidence that upon marital dissolution, Zapotec women in the Isthmus of Tehuantepec (Guerrero) keep custody over children, and are expected to go back to their parents' or brother's household; there is a norm of providing family support if a separation occurs. This is interpreted by ethnologists as evidence consistent with a lower degree of social stigma upon dissolution. The existing evidence suggests that communities with lower degrees of social stigma and greater degrees of gender equality will lead to a higher incidence of movement in the marriage market as a result of the program.[12]

However, indigenous women tend to come from a lower socio-economic background than non-indigenous women. Indigenous mothers in a union at baseline have 2.5 years of education, on average, and tend to be worse-off in terms of household characteristics and assets (except for the proportion owning agricultural land) (see Table 1, Panel A). Similar patterns hold for separated women at baseline (Table 1, Panel B). This could represent a potentially important difference in household wealth. Therefore, the magnitude of the cash transfer may be proportionally larger for the former group, leading to the proportionally larger effects observed for this group.

4. Results

4.1 Intrahousehold Allocation of Households In-Union

[10] A formal definition of *machismo* refers to "a sense of exaggerated masculinity or a cult of virility whose chief characteristics are extreme "aggressiveness and intransigence in male-to-male interpersonal relationships and arrogance and sexual aggression in male-to-female relationship" (Stevens, 1972, p.315)" (Elmendorf, 1977, p.141).

[11] I am indebted to Josefina Carpena-Méndez for bringing these marital patterns to my attention.

[12] Need to discuss lack of ethno-linguistic fractionalization in community.

In this section, I present two-stage least squares estimates of the impacts of the conditional cash transfer program on the extent of intra-household resource allocation in Period 4, almost two years after the start of the program (see Table 2). The coefficient(s) of interest represent the effect of a one peso increase in unearned income on girls' clothing budget share. The TOT point estimate for all indigenous households is 0.0060 percentage points [standard error 0.0018, not shown] (significant at 99 percent confidence), whereas it is only 0.0036 percentage points for mestizo households [standard error 0.0014, not shown] (statistically significant at 99 percent confidence). The difference in effects between indigenous and mestizo households is 0.0025 percentage points [standard error 0.013, not shown] (significant at 90 percent confidence), which suggests that the reallocation process is substantially different between these two types of households.

Intra-household resource reallocation appears largest for groups with potentially better economic opportunities outside of marriage. For instance, there are larger increases for indigenous women with primary schooling or above than for mestizo women with the same schooling level (Table 2, regression 3); the difference of 0.0040 percentage points [standard error 0.0021] is significant at 90 percent confidence. Similar differences exist between indigenous and mestizo women with young children (no children ages 10-14 years), for whom the percentage point difference of 0.0053 [standard error 0.0026] is significant at 95 percent confidence (regression 1). Similar patterns are found by comparing indigenous and mestizo women without outside labor market experience at baseline (i.e., not wage laborers or self-employed) (regressions 4 and 5). As a further check, I classify households according to the program eligibility welfare index (see Section 3.1), into top quantile (top half) and bottom quantile (bottom half); results are consistent with the previous classifications (regression 6). The point estimates suggest a similar pattern for young women (14-25 years old), but the coefficient is not statistically significant (regression 2). These results are consistent with the model to the extent that these sub-groups of women tend to have better opportunities outside of marriage.

For groups with *a priori* worse outside options (e.g., women ages 26-55 years old, with less than complete primary schooling, with children ages 10-14 years old) the point estimates suggest smaller resource re-allocations; however, differences between the latter and the omitted groups are always statistically insignificant.

4.2 Separation Decisions of Mothers In-Union (at Baseline)

In this section, I present preliminary estimates of the impacts of the conditional cash transfer program on the probability of separation of households (see Table 3). Linear probability model intent-to-treat effects estimates for all and indigenous women are 0.14 and 0.29 percentage points, respectively (statistically insignificant) (Table 3, regression 1, Panels A and B). In period 4 (November 1999), almost two years after the start of the program, there is a significant increase in the proportion of separated women; the reduced-form estimates are 0.35 (0.60) percentage points for all (indigenous) women (both significant at 95 percent confidence) (regression 2). These effects represent a 55 (90) percent increase in marital dissolution rates amongst eligible women. The impacts on non-indigenous women are small and insignificant: the estimate is 0.19 percentage points [standard error 0.21 percentage points, not shown]. To check robustness of the latter results, we estimated probit ITT effects estimates and also included state indicator variables, and found similar results, with estimated marginal effects of 0.32 (0.30) percentage points [standard error 0.13 (0.11) percentage points, not shown]. These estimated impacts seem to be realistic, since marital dissolution rates are quite low for this population and in Mexico in general. Moreover, assuming that marital dissolution follows a constant hazard model, these estimates would imply an increase in the proportion separated of 1.7 (2.9) percentage points for all (indigenous) women over a period of ten years, a noticeable effect.

We present IV estimates of the effect of the monthly cash transfer amount (in 100's of pesos) on the probability of separation (following equations (6) and (7)), (regressions 3 and 4). The two years post-treatment effect estimate implies that a monthly cash transfer of 100 pesos increases the probability of separation by 0.30 percentage points for all women, and by 0.51 percentage points for indigenous women (statistically significant at 95 percent confidence) (regression 4, Panels A and B). The instruments, indicator variables for the treatment group (pooled and in each period, depending on the specification), are very robust. The correlation between the treatment-period indicators and the amount of cash transfer received is significant at 99 percent confidence, as expected, since the randomized design should ensure that transfers are received only by the treatment group (regressions 5 and 6, Panels A and B).

Why are there differences between indigenous and non-indigenous women in their response to the program? Note that indigenous women receive similar amounts of cash transfers, on average; the coefficients on the "first-stage" regressions have similar magnitudes for all women and indigenous women. Therefore, it should be that, (i) either the transfers represent a larger share of the women's income for the former, or (ii) there are differences in the 'costs of dissolution' between the two groups, leading to differences in the outside options between these.

Marital dissolution rates appear largest for women whose outside economic options are best, or who expect a greater degree of income support. Reduced-form probit marginal effects estimates two years post-treatment are 0.57 percentage points for women with no children ages 10-14 years, who attend (primary) school regularly, therefore the conditionality constraint is not binding and take-up is practically absolute (Table 5, regression 2) ; 0.79 percentage points for young women (ages 16-25) (regression 3); 0.50 percentage points for women who completed primary school or higher (regression 4) (all statistically significant at 95 percent confidence).[13] Marginal effects estimates are quantitatively similar for indigenous women (Table 6), whereas the effects for non-indigenous women are substantially smaller and statistically insignificant (although they follow the same qualitative patterns, not shown). In summary, the heterogeneous reduced-form estimates suggest that the conditional cash transfers have quantitatively and qualitatively larger effects for women whose gains from marriage are smaller, and whose possibility for successful renegotiation of household resource

[13] See Schultz (2002) for a discussion of the effects of Progresa on take-up and school enrollment for children in primary and secondary school.

distribution is limited. This evidence is inconsistent with the collective approach to intra-household resource allocation, and possibly consistent with non-cooperative decision-making models.

4.3 Union Formation Decisions of Separated Women (at Baseline)

Eligibility to the program in the village suggests an increase in union formation of separated women, depending on their access to transfers and their 'outside options'. In this section, I present preliminary estimates of the impacts of the conditional cash transfer program on the probability of union formation of women separated at baseline (see Table 6).

Linear probability model pooled intent-to-treat effects estimates for all and indigenous women are 2.0 and 4.9 percentage points, respectively, but statistically insignificant (Table 6, regression 1, Panels A and B). Reduced-form ITT estimates are insignificant at any time period post-treatment (regression 2, Panels A and B), even though the amount of monthly cash transfers received by the household is quite large (regressions 5 and 6, Panels A and B). Probit model estimates suggest the same lack of relationship (not shown). Moreover, IV-2SLS estimates of the cash transfer amount on the probability of union formation of women (for all women and indigenous women) are not significantly different from zero (regressions 3 and 4, Panels A and B). The analogous estimates for non-indigenous women are also insignificant (not shown).

However, average program impacts may hide important differences in the behavior of women according to their social and economic constraints. The likelihood of forming a new union varies significantly between indigenous and non-indigenous women, but also by the extent of income support eligibility and (Table 7). We present reduced-form probit marginal effects estimates for various sub-groups of women; effects are 12.3 percentage points for indigenous women with young children (no children ages 10-14 years) (significant at 99 percent confidence); 11.5 percentage points for indigenous women ages 16-25 (significant at 90 percent confidence); 10.1 percentage points for women who did not complete primary school (significant at 95 percent confidence). However, the effects for non-indigenous women are small and insignificant (not shown). In summary, this suggests that the marriage market for indigenous women ‘accelerates’. Future work will establish whether the variability in these effects is due to the proportional increases in income support relative to the outside option of the women, their ‘marketability’ in the marriage market, and the existence of social norms limiting remarriage.

5. Conclusion

This paper addresses a gap in the intra-household allocation and marriage market literatures by developing an application of a simple model of contracting in marriage markets that provides a basis for understanding both intra-household resource allocation and marriage market turnover. I analyze conditions under which information on changes in the unearned income of women, by improving women’s outside option relative to the household’s production within marriage, reduce the share of household income under control of the husband, and therefore shift intra-household resource allocation to goods which women may have preference for. Moreover, the model highlights the composition of individuals who respond to the change in unearned income of the women due to heterogeneity of the participants in the marriage market. Essentially, the contracted payments to men are reduced significantly more in households where women have better outside options if the degree of complementarity in household production between her unearned income and her attributes is reduced in a union. As a result of the changes in their income share, men will evaluate the benefits of marriage according to the current contract and the expected sharing rule in a future contract, given the expected changes in the sharing rule in a future union. Marital dissolution is thus an optimal response to information on the gains to marriage in the present and future potential contracts.

Subsequently, using experimental evidence from a conditional cash transfer program in rural Mexico, I present credible support of the model in terms of its predictions on intra-household resource allocation and marital turnover decisions. Interestingly, behavioral reactions to the change in women's unearned income are for women with the best economic opportunities outside of marriage. In summary, as seen in the empirical section, the model seems to fit the marital experience of couples in poor rural Mexico.

To the extent that women's outside options affect intra-household allocation through competition in the marriage market, the model provides a behavioral explanation for the existence of distribution parameters in collective models of household behavior. That said, the additional structure in the theoretical formulation could potentially enable researchers to examine testable restrictions of the model relative to less restrictive alternatives. Future work should also try to assess the extent to which this simple contract model can explain changes in household behavior as a result of divorce legislation changes in the United States; Rasul (2003) has contributed in the latter direction within a search-theoretic framework. These are important extensions to consider in future research.

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Table 1: Mothers, Partners, and Household Baseline Characteristics, October 1997

	Treatment	Control	Indigenous	Non-indigenous
Panel A: Women In Union Sample				
Mother's characteristics (N=14064)				
Age	33.9	33.7	34.5	33.5
Years of schooling	3.1	3.1	**2.5**	**3.5**
Indigenous	0.39	0.40	1.00	0.00
Wage laborer	0.04	0.04	0.04	0.04
Self-employed	0.04	0.05	0.04	0.03
Partner's characteristics (N = 13783)				
Age	38.2	37.7	**39.0**	**38.1**
Years of schooling	3.4	3.4	*3.2*	*3.4*
Indigenous	0.41	0.41	**0.96**	**0.04**
Wage laborer	**0.68**	**0.63**	**0.74**	**0.67**
Self-employed	0.13	0.15	0.13	0.12
Agricultural laborer	0.10	0.10	**0.05**	**0.12**
Non-wage laborer	**0.05**	**0.08**	0.05	0.04
Household characteristics (N = 14064)				
Number of children	3.6	3.6	*3.53*	*3.65*
Num. of children ages 10-14 years	0.9	0.9	*0.86*	*0.91*
Cohabiting union	**0.25**	**0.19**	*0.30*	*0.24*
Dirt floor	0.65	0.62	**0.82**	**0.55**
Own house	0.94	0.95	**0.97**	**0.91**
Toilet	0.57	0.58	0.60	0.56
Agricultural land	0.62	0.66	**0.71**	**0.55**
Animals	0.38	0.40	**0.32**	**0.41**
Panel B: Separated Women Sample				
Mother's characteristics (N=626)				
Age	32.3	31.6	32.9	32.5
Years of schooling	3.8	3.9	3.7	3.9
Indigenous	0.31	0.26	**1.0**	**0.0**
Wage laborer	0.34	0.41	0.31	0.31
Self-employed	0.12	0.12	0.11	0.13
Household characteristics (N = 626)				
Number of children	2.4	2.5	2.3	2.5
Num. of children ages 10-14 years	0.6	0.6	0.5	0.6
Dirt floor	0.66	0.65	**0.83**	**0.59**
Own house	0.94	0.96	0.95	0.93
Toilet	0.58	0.57	0.58	0.59
Agricultural land	0.54	0.56	**0.63**	**0.49**
Animals	0.34	0.39	**0.27**	**0.35**

Notes: Mean of characteristics by groups are presented. Figures in bold represent statistically significant differences at least at 95 percent confidence levels; robust standard errors are allowed to be correlated within villages.

Table 2: Heterogeneous Effects on Intrahousehold Allocation of Households, Period 4

	Dependent variable: Budget Share in Girls' Clothing					
	(1)	(2)	(3)	(4)	(5)	(6)
	2SLS	2SLS	2SLS	2SLS	2SLS	2SLS
Transfer Amount, Indigenous	0.0080*	0.0065	0.0069**	0.0060**	0.0060**	0.0071**
	(0.0033)	(0.0040)	(0.0021)	(0.0018)	(0.0018)	(0.0020)
Transfer Amount, Non-indigenous	0.0028	0.0032	0.0029+	0.0035*	0.0036**	0.0038*
	(0.0019)	(0.0024)	(0.0017)	(0.0014)	(0.0014)	(0.0019)
Transfer, Indigenous w/ children 10-14 years	0.0057**					
	(0.0015)					
Transfer, Non-Indigenous w/ children 10-14 years	0.0036**					
	(0.0012)					
Transfer, Indigenous & Mother's Age 26-55 years		0.0060**				
		(0.0016)				
Transfer, Non-Indigenous & Mother's Age 26-55 years		0.0036**				
		(0.0013)				
Transfer, Indigenous & Mother's Schooling < Primary			0.0060**			
			(0.0018)			
Transfer, Non-Indigenous & Mother's Schooling < Primary			0.0037**			
			(0.0014)			
Transfer, Indigenous & Wage Laborer				0.0047+		
				(0.0025)		
Transfer, Non-Indigenous & Wage Laborer				0.0052*		
				(0.0025)		
Transfer, Indigenous & Self-Employed					0.0060+	
					(0.0031)	
Transfer, Non-Indigenous & Self-Employed					0.0034	
					(0.0029)	
Transfer, Indigenous & Low Welfare Quantile						0.0054**
						(0.0016)
Transfer, Non-Indigenous & Low Welfare Quantile						0.0032**
						(0.0010)
Difference ($\text{Non-Indigenous}_{\text{OmittedGrp}}$) - ($\text{Indigenous}_{\text{OmittedGrp}}$)	-0.0053*	-0.0033	-0.0040+	-0.0026+	-0.0024+	-0.0033*
	(0.0026)	(0.0030)	(0.0021)	(0.0013)	(0.0013)	(0.0017)
Difference ($\text{Non-Indigenous}_{\text{SubGrp}}$) - ($\text{Indigenous}_{\text{SubGrp}}$)	-0.0021+	-0.0024+	-0.0022	0.0005	-0.0027	-0.0022
	(0.0012)	(0.0013)	(0.0014)	(0.0025)	(0.0030)	(0.0016)
Instruments F-Statistic (25,11688)	1797.0**	1818.5**	1524.3**	1797.5**	1572.8**	1918.2**
Controls	YES	YES	YES	YES	YES	YES
Observations	11713	11713	11713	11713	11713	11713

Notes: Coefficient estimates from 2SLS regressions are presented. Robust standard errors in parentheses; disturbance terms are allowed to be correlated within villages; significant at (+) 90 percent, (*) 95 percent, (**) 99 percent confidence levels. Controls include mother and partner's age, schooling, and mestizo ethnicity indicators, occupation group indicators, household assets, and number of children, and a cohabitation indicator. Instruments are flood indicator, eligible transfer amount and eligible amount interacted with non-indigenous indicator, sub-group indicator, and sub-group-mestizo interaction indicator, respectively.

Table 3: Program Impacts on Separation of Women in Union

Dependent variables:	Separation indicator				Transfer Amount (100's Pesos)	
	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	OLS	2SLS	2SLS	OLS	OLS
Panel A: All Women						
Pooled Treatment Indicator	0.0014				1.09**	
	(0.0013)				(0.03)	
Treatment Indicator, Period 2		0.0011				0.82**
		(0.0011)				(0.03)
Treatment Indicator, Period 3		-0.0003				1.35**
		(0.0018)				(0.04)
Treatment Indicator, Period 4		0.0035*				1.09**
		(0.0016)				(0.04)
Transfer, Pooled (100's Pesos)			0.0012			
			(0.0011)			
Transfer, Period 2 (100's Pesos)				0.0013		
				(0.0014)		
Transfer, Period 3 (100's Pesos)				-0.0002		
				(0.0012)		
Transfer, Period 4 (100's Pesos)				0.0030*		
				(0.0014)		
Treatment Group (Baseline)	0.0003	0.0003	0.0003	0.0003	0.03**	0.03**
	(0.0002)	(0.0002)	(0.0002)	(0.0002)	(0.01)	(0.01)
Constant	-0.0045	-0.0045	-0.0042	-0.0042	-0.21**	-0.21**
	(0.0030)	(0.0030)	(0.0030)	(0.0030)	(0.05)	(0.05)
Controls	YES	YES	YES	YES	YES	YES
Observations	49998	49998	49998	49998	49998	49998
Mean of dependent variable	0.0050	0.0050	0.0050	0.0050	-	-
Panel B: Indigenous Women						
Pooled Treatment Indicator	0.0029				1.04**	
	(0.0018)				(0.05)	
Treatment Indicator, Period 2		0.0023+				0.68**
		(0.0014)				(0.05)
Treatment Indicator, Period 3		0.0005				1.36**
		(0.0027)				(0.06)
Treatment Indicator, Period 4		0.0060*				1.09**
		(0.0026)				(0.06)
Transfer, Pooled (100's Pesos)			0.0026			
			(0.0016)			
Transfer, Period 2 (100's Pesos)				0.0034+		
				(0.0020)		
Transfer, Period 3 (100's Pesos)				0.0003		
				(0.0019)		
Transfer, Period 4 (100's Pesos)				0.0051*		
				(0.0022)		
Treatment Group (Baseline)	0.0005	0.0005	0.0004	0.0005	0.03*	0.03*
	(0.0004)	(0.0004)	(0.0004)	(0.0004)	(0.01)	(0.01)
Constant	-0.0071*	-0.0071*	-0.0070*	-0.0071*	-0.06	-0.06
	(0.0030)	(0.0030)	(0.0030)	(0.0030)	(0.07)	(0.07)
Controls	YES	YES	YES	YES	YES	YES
Observations	19754	19754	19754	19754	19754	19754
Mean of dependent variable	0.0043	0.0043	0.0043	0.0043	-	-

Notes: Coefficient estimates from OLS and 2SLS regressions are presented. Robust standard errors in parentheses; disturbance terms are allowed to be correlated within villages; significant at (+) 90 percent, (*) 95 percent, (**) 99 percent confidence levels. Controls include mother and partner's age, schooling, and mestizo ethnicity indicators, occupation group indicators, household assets, and number of children, and a cohabitation indicator. Instruments in 2SLS regressions are assignment to treatment indicators, pooled (column 3) and by period (column 4).

Table 4: Heterogeneous Effects on Separation of All Women, Period 4

	Dependent variable: Separation indicator							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Probit	Probit	Probit	Probit	Probit	Probit	Probit	Probit
Treatment Indicator, Period 4	0.0032*	0.0057**	0.0079*	0.0050*	0.0032*	0.0032*	0.0038*	0.0015
	(0.0013)	(0.0020)	(0.0033)	(0.0022)	(0.0014)	(0.0014)	(0.0018)	(0.0021)
Treatment * Children ages 10-14		-0.0043*						
		(0.0021)						
Treatment * Woman's Age 26-35			-0.0051*					
			(0.0023)					
Treatment * Woman's Age 36-45			-0.0047*					
			(0.0021)					
Treatment * Woman's Age 46-55			-0.0036					
			(0.0022)					
Treatment * Woman Sch. <Primary				-0.0027				
				(0.0027)				
Treatment * Woman Wage Laborer					0.0001			
					(0.0052)			
Treatment * Woman Self-employed						0.0002		
						(0.0067)		
Treatment * Upper Quart. W-Index							-0.0014	
							(0.0024)	
Treatment * Partner's Sch. <Primary								0.0028
								(0.0027)
Controls	YES	YES	YES	YES	YES	YES	YES	YES
Observations	11798	11798	11798	11798	11798	11798	11798	11798
Mean of dependent variable	0.0067	0.0067	0.0067	0.0067	0.0067	0.0067	0.0067	0.0067

Notes: Marginal probit estimates (evaluated at mean covariate values) are presented. Robust standard errors in parentheses; disturbance terms are allowed to be correlated within villages; statistically significant at (+) 90 percent, (*) 95 percent, (**) 99 percent confidence levels. Controls include woman and partner's age, ethnicity, and schooling indicators; occupation-type indicators, cohabitation, number of children indicators, and household dwelling characteristics at baseline. In the regressions with interaction terms (2-8), we include the relevant explanatory variable, as well as the interaction with the time controls, but do not report these in the table.

Table 5: Heterogeneous Effects on Separation of Indigenous Women, Period 4

	Dependent variable: Separation indicator							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Probit	Probit	Probit	Probit	Probit	Probit	Probit	Probit
Treatment Indicator, Period 4	0.0030**	0.0055*	0.0045+	0.0040+	0.0032**	0.0033**	0.0059*	0.0021
	(0.0011)	(0.0023)	(0.0026)	(0.0024)	(0.0012)	(0.0011)	(0.0025)	(0.0016)
Treatment * Children ages 10-14		-0.0038+						
		(0.0020)						
Treatment * Woman's Age 26-35			-0.0028+					
			(0.0015)					
Treatment * Woman's Age 36-45			-0.0002					
			(0.0032)					
Treatment * Woman's Age 46-55			-0.0017					
			(0.0016)					
Treatment * Woman Sch. <Primary				-0.0015				
				(0.0028)				
Treatment * Woman Wage Laborer					-0.0016			
					(0.0015)			
Treatment * Woman Self-employed						-0.0021**		
						(0.0007)		
Treatment * Upper Quart. W-Index							-0.0047+	
							(0.0027)	
Treatment * Partner's Sch. <Primary								0.0015
								(0.0027)
Controls	YES	YES	YES	YES	YES	YES	YES	YES
Observations	4386	4386	4386	4386	4386	4386	4386	4386
Mean of dependent variable	0.0064	0.0064	0.0064	0.0064	0.0064	0.0064	0.0064	0.0064

Notes: Marginal probit estimates (evaluated at mean covariate values) are presented. Robust standard errors in parentheses; disturbance terms are allowed to be correlated within villages; statistically significant at (+) 90 percent, (*) 95 percent, (**) 99 percent confidence levels. Controls include woman and partner's age, ethnicity, and schooling indicators; occupation-type indicators, cohabitation, number of children indicators, and household dwelling characteristics at baseline. In the regressions with interaction terms (2-8), we include the relevant explanatory variable, as well as the interaction with the time controls, but do not report these in the table.

Table 6: Program Impacts on Union Formation of Separated Women

Dependent variables:	Separation indicator				Transfer Amount (100's Pesos)	
	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	OLS	2SLS	2SLS	OLS	OLS
Panel A: All Women						
Pooled Treatment Indicator	0.020				0.80**	
	(0.021)				(0.05)	
Treatment Indicator, Period 2		0.021				0.61**
		(0.024)				(0.05)
Treatment Indicator, Period 3		0.027				0.93**
		(0.025)				(0.06)
Treatment Indicator, Period 4		0.012				0.85**
		(0.024)				(0.06)
Transfer, Pooled (100's Pesos)			0.025			
			(0.026)			
Transfer, Period 2 (100's Pesos)				0.034		
				(0.041)		
Transfer, Period 3 (100's Pesos)				0.030		
				(0.027)		
Transfer, Period 4 (100's Pesos)				0.014		
				(0.029)		
Treatment Group (Baseline)	0.002	0.002	0.001	0.002	0.04*	0.04*
	(0.003)	(0.003)	(0.003)	(0.003)	(0.02)	(0.02)
Constant	0.045	0.045	0.045	0.044	0.02	0.02
	(0.031)	(0.031)	(0.031)	(0.031)	(0.10)	(0.10)
Controls	YES	YES	YES	YES	YES	YES
Observations	2472	2472	2472	2472	2472	2472
Mean of dependent variable	0.071	0.071	0.071	0.071	-	-
Panel B: Indigenous Women						
Pooled Treatment Indicator	0.049				0.78**	
	(0.036)				(0.09)	
Treatment Indicator, Period 2		0.032				0.54**
		(0.042)				(0.10)
Treatment Indicator, Period 3		0.061				1.04**
		(0.045)				(0.14)
Treatment Indicator, Period 4		0.053				0.95**
		(0.045)				(0.12)
Transfer, Pooled (100's Pesos)			0.063			
			(0.049)			
Transfer, Period 2 (100's Pesos)				-0.020		
				(0.157)		
Transfer, Period 3 (100's Pesos)				-0.025		
				(0.082)		
Transfer, Period 4 (100's Pesos)				0.015		
				(0.081)		
Treatment Group (Baseline)	0.003	0.003	0.003	-0.001	0.00	-0.02
	(0.008)	(0.008)	(0.009)	(0.018)	(0.04)	(0.04)
Constant	-0.069	-0.069	-0.092	0.98**	0.37+	0.50*
	(0.051)	(0.051)	(0.057)	(0.094)	(0.22)	(0.23)
Controls	YES	YES	YES	YES	YES	YES
Observations	800	800	800	658	800	658
Mean of dependent variable	0.080	0.080	0.080	0.080	-	-

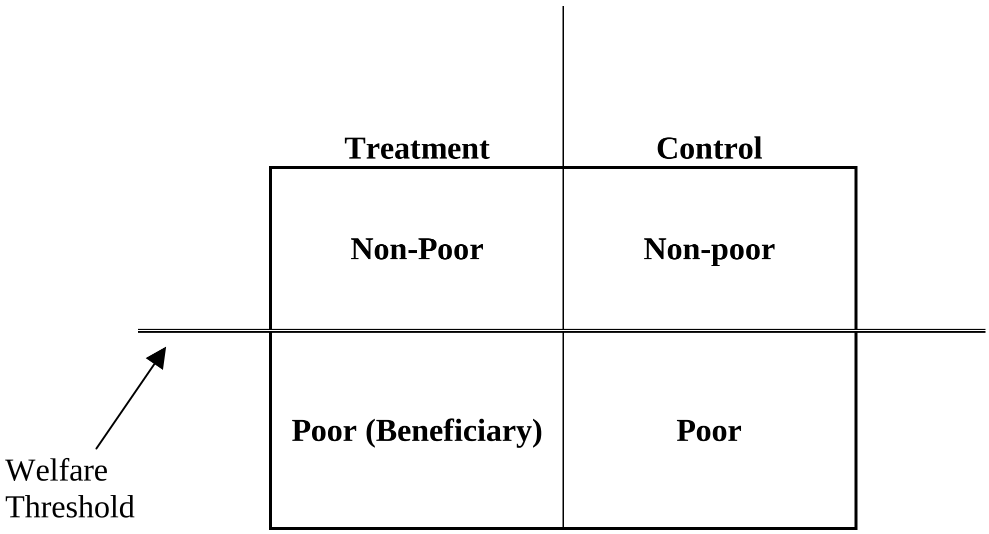
Notes: Coefficient estimates from OLS and 2SLS regressions are presented. Robust standard errors in parentheses; disturbance terms are allowed to be correlated within villages; significant at (+) 90 percent, (*) 95 percent, (**) 99 percent confidence levels. Controls include mother's age, schooling, and mestizo ethnicity indicators, occupation group indicators, household dwelling characteristics, and number of children, and a cohabitation indicator. Instruments in 2SLS regressions are assignment to treatment indicators, pooled (column 3) and by period (column 4).

Table 7: Heterogeneous Effects on Union Formation of Separated Women, Period 4

	Dependent variable: union formation indicator					
	(1)	(2)	(3)	(4)	(5)	(6)
	Probit	Probit	Probit	Probit	Probit	Probit
Panel A: All Women						
Treatment Indicator (Period 4)	0.010	0.037	0.019	0.015	0.019	0.053+
	(0.023)	(0.025)	(0.038)	(0.029)	(0.027)	(0.028)
Treatment * Children ages 10-14		-0.061*				
		(0.030)				
Treatment * Woman's Age 26-35			-0.024			
			(0.048)			
Treatment * Woman's Age 36-55			0.001			
			(0.048)			
Treatment * Woman Sch. ≥ Primary				-0.013		
				(0.039)		
Treatment * Woman Wage Laborer					-0.031	
					(0.036)	
Treatment * Upper Quart. W-Index						-0.073*
						(0.033)
Controls	YES	YES	YES	YES	YES	YES
Observations	618	618	618	618	618	618
Mean of dependent variable	0.07	0.07	0.07	0.07	0.07	0.07
Panel B: Indigenous Women						
Treatment Indicator (Period 4)	0.056	0.123**	0.115+	0.101*	0.074+	0.066
	(0.036)	(0.046)	(0.060)	(0.047)	(0.039)	(0.048)
Treatment * Children ages 10-14		-0.127**				
		(0.042)				
Treatment * Woman's Age 26-35			-0.077+			
			(0.046)			
Treatment * Woman's Age 36-55			-0.070*			
			(0.035)			
Treatment * Woman Sch. ≥ Primary				-0.070*		
				(0.033)		
Treatment * Woman Wage Laborer					-0.055	
					(0.042)	
Treatment * Upper Quart. W-Index						-0.021
						(0.059)
Controls	YES	YES	YES	YES	YES	YES
Observations	200	200	200	200	200	200
Mean of dependent variable	0.09	0.09	0.09	0.09	0.09	0.09

Notes: Marginal probit estimates (evaluated at mean covariate values) are presented. Robust standard errors in parentheses; disturbance terms are allowed to be correlated within villages; statistically significant at (+) 90 percent, (*) 95 percent, (***) 99 percent confidence levels. Controls include women's age group, primary school completion, non-indigenous, and occupation-type indicators; indicators for number of children ages 10-14 years, and dirt floor and own house indicator variables at baseline. In the regressions with interaction terms (2-6), we include the relevant explanatory variable, as well as the interaction with the time controls, but do not report these in the table.

Figure 1: Program Evaluation Design



Appendix A: Proofs

Proof of Proposition 1

AVERAGE PAYMENT EFFECT

Substituting equation (1) in equation (2) and rearranging terms gives

$$\omega_{t+j,t}(y^F) = \Phi^*(y^F) - \beta\mu^2 A^F_{t+j}(y^F) + (1-\beta\mu^2)\varepsilon_t \quad \text{for all t.} \qquad (4)$$

Taking derivatives with respect to y^F gives

$$\frac{\partial \omega_{t+j,t}}{\partial y^F} = \frac{\partial \Phi^*}{\partial y^F} - \beta\mu^2 \frac{\partial A^F_{t+j}}{\partial y^F} \qquad (5)$$

Note that $\partial\omega_t / \partial y^F < 0$ if $(\beta\mu^2)\partial A^F_t \big/ \partial y^F > \partial\Phi^* \big/ \partial y^F$ from equation (5).

PAYMENT EFFECT BY SUB-GROUP

The sign of the cross-derivative effect can be determined by differentiating both sides of equation (5) by x^F_t :

$$\frac{\partial^2 \omega_{t+j,t}}{\partial x^F_t \partial y^F} = \frac{\partial^2 \Phi^*(x^M_t, x^F_t)}{\partial x^F_t \partial y^F} - \beta\mu^2 \frac{\partial^2 \Phi^*(0, x^F_t)}{\partial x^F_t \partial y^F} \qquad (6)$$

Then

$$\frac{\partial^2 \omega_{t+j,t}}{\partial x^F_t \partial y^F} < 0 \qquad \text{if} \qquad \beta\mu^2 \frac{\partial^2 \Phi^*(0, x^F_t)}{\partial x^F_t \partial y^F} > \frac{\partial^2 \Phi^*(x^M_t, x^F_t)}{\partial x^F_t \partial y^F}$$

Since $\frac{\partial^2 \Phi^*(x^M_t, x^F_t)}{\partial x^F_t \partial x^F} < \frac{\partial^2 \Phi^*(0, x^F_t)}{\partial x^F_t \partial y^F}$ because $\frac{\partial^3 \Phi^*(x^M_t, x^F_t, y^F_t)}{\partial x^M_t \partial x^F_t \partial y^F} < 0$ (by assumption), then the condition is likely▪

Proof of Proposition 2

AVERAGE SEPARATION EFFECT

Using the marriage market equilibrium condition (3) at period $t+p$ (p periods after the start of the marital union contract):

$$U(A^{M*}_{t+p}) + (\beta\mu)E_t[V(\omega_{t+p+1,t+p+1}, \Phi_{t+p+1}) \mid \Phi_{t+p}]$$
$$= V(\omega_{t+p,t+p}, \Phi_{t+p}). \qquad (3')$$

Substituting for the value functions $V(\omega,\Phi)$ in equation (3') and eliminating terms results in:

$$U(A^{M*}_{t+p}) = U(\omega_t) + \frac{\beta\mu^2}{1-\beta\mu^2}\{U(\omega_t) - E_{t+p}[U(\omega_{t+p+1}) \mid \Phi_{t+p}] \qquad (3'')$$

Since $\partial\omega_{t+p,t} \big/ \partial y_{t+p}^F = \partial\omega_{t+p+1,t+p+1} \big/ \partial y_{t+p}^F$ (see equation (4)), differentiating both sides of equation (3'') with respect to Y_{t+p}^F and rearranging terms gives

$$\frac{\partial U(A_{t+p}^{M*})}{\partial A_{t+p}^{M*}} \cdot \frac{\partial A_{t+p}^{M*}}{\partial y^F} = \left(\frac{1}{1-\beta\mu^2}\right) \cdot \frac{\partial\omega_t}{\partial y^F} \cdot \left\{ \frac{\partial U(\omega_t)}{\partial\omega_t} - \beta\mu^2 \cdot \frac{\partial E_{t+p}\left[U(\omega_{t+p+1}) \,|\, \Phi_{t+p}\right]}{\partial\omega_{t+p+1}} \right\} \tag{7}$$

The expected change in the likelihood of separation of a household ($\Pr[A_t^M > A_t^{M*}]$) depends on the sign of $\frac{\partial\omega_t}{\partial y_t^F}$ and the following condition:

$$\frac{\partial U(\omega_t)}{\partial\omega_t} - \beta\mu^2 \cdot \frac{\partial E_{t+p}\left[U(\omega_{t+p+1}) \,|\, \Phi_{t+p}\right]}{\partial\omega_{t+p+1}} > or \leq 0\,. \tag{8}$$

SEPARATION EFFECTS BY SUB-GROUP

Differentiating both sides of equation (7) with respect to x^F gives:

$$\begin{aligned}
&\frac{\partial^2 U(A_{t+p}^{M*})}{\partial (A_{t+p}^{M*})^2} \cdot \frac{\partial A_{t+p}^{M*}}{\partial x^F} \cdot \frac{\partial A_{t+p}^{M*}}{\partial y^F} + \frac{\partial U(A_{t+p}^{M*})}{\partial A_{t+p}^{M*}} \cdot \frac{\partial^2 A_{t+p}^{M*}}{\partial x^F \partial y^F} = \\
&\quad = \left(\frac{1}{1-\beta\mu^2}\right) \cdot \left\{ \frac{\partial^2\omega_t}{\partial x^F \partial y^F} \cdot \left[\frac{\partial U(\omega_t)}{\partial\omega_t} - \beta\mu^2 \cdot \frac{\partial E_{t+p}\left[U(\omega_{t+p+1}) \,|\, \Phi_{t+p}\right]}{\partial\omega_{t+p+1}} \right] + \right. \\
&\quad \left. + \frac{\partial\omega_t}{\partial y^F} \cdot \left[\frac{\partial^2 U(\omega_t)}{\partial{\omega_t}^2} \cdot \frac{\partial\omega_t}{\partial x^F} - \beta\mu^2 \cdot \frac{\partial^2 E_{t+p}\left[U(\omega_{t+p+1}) \,|\, \Phi_{t+p}\right]}{\partial{\omega_{t+p+1}}^2} \cdot \frac{\partial\omega_{t+p+1}}{\partial x^F} \right] \right\}
\end{aligned} \tag{9}$$

Noting that $\frac{\partial\omega_t}{\partial x^F} = \frac{\partial\omega_{t+p+1}}{\partial x^F}$ and rearranging terms gives the cross-derivative effect on the equilibrium threshold level of A_t^{M*}. It is shown in equation (10) below:

$$\begin{aligned}
\frac{\partial^2 A_{t+p}^{M*}}{\partial x^F \partial y^F} &= \left[\frac{\partial U(A_{t+p}^{M*})}{\partial A_{t+p}^{M*}} \right]^{-1} \cdot \left(\frac{1}{1-\beta\mu^2}\right) \cdot \left\{ \frac{\partial^2\omega_t}{\partial x^F \partial y^F} \cdot \left[\frac{\partial U(\omega_t)}{\partial\omega_t} - \beta\mu^2 \cdot \frac{\partial E_{t+p}\left[U(\omega_{t+p+1}) \,|\, \Phi_{t+p}\right]}{\partial\omega_{t+p+1}} \right] + \right. \\
&\left. + \frac{\partial\omega_t}{\partial y^F} \frac{\partial\omega_t}{\partial x^F} \cdot \left[\frac{\partial^2 U(\omega_t)}{\partial{\omega_t}^2} - \beta\mu^2 \cdot \frac{\partial^2 E_{t+p}\left[U(\omega_{t+p+1}) \,|\, \Phi_{t+p}\right]}{\partial{\omega_{t+p+1}}^2} \right] \right\} - \left[\frac{\partial U(A_{t+p}^{M*})}{\partial A_{t+p}^{M*}} \right]^{-1} \cdot \frac{\partial^2 U(A_{t+p}^{M*})}{\partial (A_{t+p}^{M*})^2} \cdot \frac{\partial A_{t+p}^{M*}}{\partial x^F} \cdot \frac{\partial A_{t+p}^{M*}}{\partial y^F}
\end{aligned} \tag{10}$$

According to equation (10), the sign of $\frac{\partial^2 A_{t+p}^{M*}}{\partial x^F \partial y^F}$ depends on equations (6), (8), and condition (11) presented below:

$$\frac{\partial^2 U(\omega_t)}{\partial{\omega_t}^2} - \beta\mu^2 \cdot \frac{\partial^2 E_{t+p}\left[U(\omega_{t+p+1}) \,|\, \Phi_{t+p}\right]}{\partial{\omega_{t+p+1}}^2} > or < 0\,\cdot$$

Table A1: Relationship between attrition and characteristics of subjects at baseline

	Dependent variable: Attrition indicator							
	Panel A: Mothers In Union Sample				Panel B: Separated Mothers Sample			
	Treatment	Adjusted	Main effect	Interaction	Treatment	Adjusted	Main effect	Interaction
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	OLS	OLS	OLS	OLS	OLS	OLS	OLS	OLS
Treatment indicator	0.013	-	-0.053	-	0.004			0.127
	(0.010)		(0.062)		(0.028)			(0.166)
Mother's age		-0.001***	-0.002***	0.001		-0.005***	-0.006**	0.002
		(0.000)	(0.001)	(0.001)		(0.002)	(0.003)	(0.003)
Mother's years of schooling		-0.003**	-0.001	-0.002		-0.001	0.007	-0.014
		(0.001)	(0.002)	(0.003)		(0.005)	(0.008)	(0.011)
Mother indigenous		-0.006	-0.001	-0.007		0.036	0.047	-0.027
		(0.015)	(0.024)	(0.030)		(0.029)	(0.051)	(0.061)
Mother wage laborer		0.028*	0.072**	-0.068**		0.057**	0.028	0.051
		(0.015)	(0.030)	(0.034)		(0.028)	(0.046)	(0.057)
Mother self-employed		0.014	0.013	0.000		0.010	-0.012	0.031
		(0.013)	(0.023)	(0.027)		(0.037)	(0.057)	(0.074)
Partner's age		-0.001**	-0.001	0.000				
		0.000	(0.001)	(0.001)				
Partner's years of schooling		-0.002*	-0.002	-0.001				
		(0.001)	(0.002)	(0.002)				
Partner indigenous		-0.004	-0.011	0.010				
		(0.015)	(0.025)	(0.031)				
Partner wage laborer		-0.018	-0.062**	0.072**				
		(0.014)	(0.028)	(0.031)				
Partner self-employed		-0.014	-0.050*	0.056*				
		(0.015)	(0.026)	(0.032)				
Partner agricultural laborer		-0.021	-0.050*	0.047				
		(0.016)	(0.029)	(0.034)				
Partner non-wage laborer		0.023	-0.083***	0.149***				
		(0.033)	(0.029)	(0.049)				
Number of children		-0.009***	-0.005**	-0.007**		-0.014	-0.009	-0.010
		(0.001)	(0.002)	(0.003)		(0.012)	(0.019)	(0.024)
N. children 10-14 years old		-0.006**	-0.009**	0.005		-0.001	0.029	-0.045
		(0.003)	(0.004)	(0.006)		(0.019)	(0.035)	(0.041)
Cohabitation		0.007	0.013	-0.008				
		(0.007)	(0.012)	(0.015)				
Dirt floor		-0.010*	-0.008	-0.003		-0.006	0.041	-0.074
		(0.005)	(0.009)	(0.011)		(0.025)	(0.038)	(0.050)
Own house		-0.047***	-0.051***	0.007		-0.062	-0.035	-0.068
		(0.011)	(0.016)	(0.021)		(0.052)	(0.073)	(0.102)
Toilet		-0.024***	-0.019**	-0.008		0.034	0.006	0.036
		(0.006)	(0.009)	(0.013)		(0.026)	(0.044)	(0.055)
Agricultural Land		-0.008	-0.023*	0.024*		0.011	-0.019	0.064
		(0.007)	(0.013)	(0.015)		(0.027)	(0.041)	(0.053)
Animals		-0.010*	0.001	-0.017*		0.020	0.085	-0.110*
		(0.005)	(0.007)	(0.010)		(0.030)	(0.052)	(0.063)
Constant	0.101		0.364***	-	0.211***	0.424***	0.359***	
	(0.007)		(0.054)		(0.022)	(0.082)	(0.127)	
Observations	40515	40515	40515	-	1854	1854	1854	-
R-squared	0.00	0.02	0.02	-	0.00	0.03	0.04	-
F(Treatment Interactions)	-	-	-	1.76	-	-	-	1.00
P-value	-	-	-	0.02	-	-	-	0.45

Notes for Table A1: Robust standard errors in parentheses; disturbance terms are allowed to be correlated within villages, but not across villages. Significantly different than zero at 90% (*), 95% (**), 99% (***) confidence.. Columns 3 and 4 (and columns 7 and 8) presents results from one regression with main effects (col.3) (col.7) and all covariates interacted with treatment (col.4) (col.8).