Household Division and Rural Economic Growth

Andrew D. Foster

Brown University

Mark R. Rosenzweig

University of Pennsylvania

October 2001

The research for this paper supported in part by grants NIH HD30907 and NSF, SBR93-08405. We are grateful to three referees and the Editor for useful comments and suggestions.

I. Introduction

The recent availability of longitudinal data from low-income countries makes possible for the first time the identification of the consequences of growth-augmenting innovations for household income change. However, it has become increasingly recognized that both the analysis and design of panel surveys is importantly affected by the break-up of households over time. For example, at the end of the sample period of the Indian ICRISAT Village Studies Survey (Walker and Ryan, 1984), which followed households in three villages for ten years, over 13% of the households had divided. Lifetime retrospective information from the same survey indicated that 40% of the households had been formed from extended households prior to the death of the head of the extended household. A newly-available longitudinal national survey of rural Indian households (Vashishtha, 1986), described below, indicates that over a twelve-year period, land and other assets were divided among male adult family members and separate residence was established in 30% of the original sample farm households. Both surveys dropped entire or sub-parts of households that had not remained intact as part of the survey design. Inferences about economic mobility from these data would thus be biased to the extent that household splits are nonrandom.

The neglect of the possibilities of household break-up in the longitudinal study designs stems in part from the relative absence of attention in the theoretical and empirical literatures to the determination of household structure. Household structure is pervasively treated as an exogenous or fixed characteristic. An improved understanding of the determinants of household division is thus useful not only for dealing with the potential selectivity of panel designs that drop dividing households, but in studying household behavior and income change generally. For example, in assessing the distributional effects of income-augmenting development strategies, to the extent that, as is common in many low income rural areas, co-production is tied to co-residence, household *cum* asset division can alter income levels and growth if there are important scale economies of production. In that case, knowledge of the effects of an intervention on incomes for a given productive asset configuration is insufficient to predict actual income growth in the absence of knowledge about how assets are divided via household break-ups. Moreover, the recent panel survey of Indian households indicated that over a twelve-year period, 80% of households either experienced no change or a decrease in landholdings. Of the 20% of those households that had experienced a decrease in landholdings, in 79% of the cases this was due to household partition. And, weighted by the size of the changes in landholdings, partition accounted for over 90% of the decline in average household landholdings. Clearly household division, rather than market transactions, plays a dominant role in the evolution of landholdings over time.

The effects of economic growth on income inequality has been a subject long investigated by economists. Most studies have examined data describing changes in various measures of income inequality or of income changes within income ranks as functions of measures of economic development (e.g. Chenery, et al. 1974, and more recently, Deininger and Squire, 1996). This methodology, however, cannot provide information on how particular income groups benefit or lose from growth, since the data used, repeated cross-sections, do not link individuals or households across time. Another approach uses computable general equilibrium models to examine through simulation methods the effects on economic mobility - changes in the relative incomes of pre-defined income classes - from growth-inducing policies (i.e., Adelman and Robinson 1978). In these models, however, the links between changes in the returns to individual endowments yielded by the structural model parameters and changes in the incomes of households are based on a set of fixed coefficients obtained from cross-sectional data that are used to assign individuals of particular characteristics to particular household configurations.[1] To the extent that growth has direct effects on the composition and numbers of households, as is likely, this method may give misleading results.

[1]The importance for measurement of well-being of the distribution of individuals with different characteristics across households is clearly articulated by Basu and James Foster (1996) in their axiomatic characterization of measures of literacy.

In this paper we formulate, test, and estimate a structural model of household division. Our model is based on the recent primarily theoretical literature on the collective household (Chiappori 1988; Browning and Chiappori 1996). This literature, and empirical studies that have attempted to test "unitary" models of the household (Schultz 1990, Thomas 1990), have focused on the determinants of the allocation of resources within intact households. An obvious failure of unitary household models, however, is that they cannot account for the break-up of households. Our model of the collective household assumes, as in the literature on the collective household, that individuals optimize subject to a set of pre-defined entitlement rules (e.g., inheritance laws) and intrahousehold allocations are efficient. Gains from co-residence arise from cost-sharing a household-specific public good and lower barriers to information sharing on best-practice farming techniques, but whether such gains are sufficient to make co-residence desirable depends on the existence of scale economies or dis-economies in production and on how household structure affects risk diversification and risk-sharing. The model yields implications for how household size and intra-household inequality interact with exogenous income growth to affect (i) the amount of the household public good that is consumed, (ii) which households divide, (iii) the exact divisions of the assets among the new households and (iv) the evolution of the incomes of the new configurations of households.[2]

The implications of the model are tested using panel data describing Indian farm households starting from the onset of the Indian "green revolution" in the late 1960's through 1982. We find for example that, as predicted by the model, within-household inequality in schooling, marriages, and riskiness increase the probability of household division. The estimated parameters of the structural model

[2]Although collective models of household behavior with a household-specific a public good of this form have been used in empirical applications to household formation through marriage (Foster 1996), this paper represents the first to our knowledge to focus on household division. Hayashi (1993) tests for intergenerational altruism based on comparisons of two and single-generation households in Japan and develops a model in which the demand for coresidence derives from the demand for a household public good. As he notes, his estimates are potentially subject to selectivity bias because his data do not provide information on both generations in split households.

are used to evaluate the consequences of technical change for household income growth by economic strata. The estimates indicate that inattention to the consequences of technical change for household division can lead to a substantial overestimate of the extent to which better-off households differentially benefit from technical change and demonstrate that the amount of within-household inequality can have important effects on the evolution of land assets over time and for the interhousehold distributional consequences of economic growth.

II. Theory

a. Household Public Goods and Household Surplus

A model of the decisions by individuals to consume and work together requires the specification both of resource allocation across individuals co-residing in a joint household and the resources available to the individuals if they reside in separate households. We assume that in each joint household j there are N individuals (claimants or heirs) who have property rights with respect to a divisible asset that produces a risky income stream. In particular, each claimant has the right to a share κ_{ij} of the household production asset A_j (e.g., land) in that the claimant may appropriate the proceeds from his share or sell it. The property rights may be assigned by law (e.g., partible inheritance, primogeniture) or by norms (e.g., include only men among heirs). In the Indian context, inheritance customs determine the entitlement shares of the claimants in a joint landowning household. Claimants are males. Among claimants whose father is resident in the households, current claims are based on the shares of the land owned by their father that they receive immediately after the death of their father. For brothers jointly farming land (i.e., sons who have decided to continue to farm and reside together after the death of the father), the shares of the total household land to which they have current entitlements are based on their inherited shares.

Households are considered to be joint households if there is more than one claimant coresiding, and are divided households if claimants separate from the joint household. Claimants who are adult sons of the head and who leave the joint household take with them a portion of their father's asset that is a function of their future inheritance. However, joint households that remain intact experience changes in household composition that can have important effects on the gains from jointness. Claimants in household j may marry and have children and these changes may affect the value of joint residence. In addition, grown sons and daughters who leave the household to set up their own households do not necessarily break all economic relations with the original household. Evidence on interhousehold transfers in India suggests that daughters, who have no inheritance claims and leave the household upon marriage, participate in risk-sharing arrangements with their father's household (Rosenzweig, 1988; Rosenzweig and Stark, 1989) Symmetrically, claimants may marry and married women who enter a household contribute to that household's risk reduction. However, although transfers may take place between departed sons and fathers, because the new household of the sons is typically proximate to that of their father, such transfers do not significantly contribute to consumption smoothing.

We assume that "autarchic" income, the income for an individual claimant farmed with his own nuclear family (wife and own children), is the sum of labor income from his family W_i, the product of the individual-specific productivity factor θ_i and the stock of assets to which the individual has a claim, an income shock e_i, and net transfers τ_i. Net transfers to household i depend on the number of departed daughters D_i, the income shock and the asset claim. Autarchic income is then $y_i = \theta_i \kappa_i A_j + W_i + e_i + \tau_i$. Analogously, joint household income is $y_j^N = \theta_j^N A_j + NW_i + e_j^N + \tau_j^N$. The expected benefits from joint residence/production depend on the expected value of the difference between y_j^N and Σy_i and, given risk aversion, on the difference in the riskiness associated with residence choice.[3] Gains from joint residence thus may arise from scale economies in production but also arise from savings associated with the financing of household public goods, jointly consumed by all claimants and their families. Offsetting these consumption gains from joint residence that we incorporate in the model or in the empirical implementation of the model are a) a direct preference for autarchic residence, b) possible diseconomies to joint production, c) the possibility that the marginal utility for the public good is actually negative for

[3]Coresidence and joint production appear to be isomorphic among Indian households. Among the reasons may be lower costs of information transmission and less scope for moral hazard among coworkers who eat together. We do not model structurally these returns to coresidence and joint production.

some claimants in the joint household, and d) reduced insurance associated with interhousehold transfers. Whether households remain intact thus depends on the production technology, risk, the taste for privacy, individual preference heterogeneity, and the household technology.

We assume that each household claimant's nuclear family, described by a vector $\mathbf{n_i}$, consumes a private good x_{ij} and a household public good z_j .The vector $\mathbf{n_i}$ contains the characteristics of the claimant and his wife and children, including non-resident sons. Preferences also depend directly on household structure, as indicated by r_{ij}. The individual utility of claimant i in household j is thus given by

$$u_{ij}=u(x_{ij},z_j,r_{ij};\boldsymbol{n_{ij}}) \qquad \textbf{(1)}$$

and the budget constraint for the joint household is

$$\sum_{i=1}^{N} x_{ij}+z_j=y_j^N \qquad \textbf{(2)}$$

Given (1) and (2), a potential source of conflict that may induce separation of household claimants is differing preferences for the desired level of the public good, which must be consumed at equal levels by all claimants when co-resident. In order to capture in relatively simple form the idea that conflict over the provision of the household public good is central to the decision about joint residence, that the degree of conflict is influenced by changes in income and by the amount of within-household inequality, and that joint residence may be importantly influenced by risk considerations we impose three restrictions on behavior in joint households: A. Decisions about joint residence in a given period must be made before the income shocks are realized but consumption allocations in a given period are made afterwards. B. Intra-household allocations conditional on residence and income realizations must be *ex ante* efficient[4] in the sense that it is not possible for a household to improve the expected utility of one

[4]Note that *ex ante* efficiency coupled with A implies *ex post* efficiency–for each income realization and residence status it is not possible to improve the utility of one claimant without decreasing the utility of some other claimant.

claimant (or sub-household) without reducing the expected utility of some other claimant. C. Each claimant must be provided an *ex ante* expected utility level that is at least equal to that which he could obtain if separately resident.

We assume that joint households will divide whenever it is not possible to satisfy condition C. That is, a household breaks up whenever there are no gains from joint residence. The gains for household j with N claimants may be defined as the difference between expected utility achieved by any heir i, say the first born i=1, and his reservation expected utility subject to the condition that each of the N-1 other claimants receives his respective expected reservation utility. That is, the gains from joint residence are:

$$\begin{gathered}\text{maximize w.r.t. } x_i(s), z(s) \forall s \\ E(u(x_1(s), z(s), r_1; \boldsymbol{n_1}) - Ev(y_1(s), r_1, \boldsymbol{n_1})\end{gathered} \tag{3}$$

such that

$$Eu(x_i(s), z(s), r_i; \boldsymbol{n_i}) \geq Ev(y_i(s), r_i, \boldsymbol{n_i}) \text{ for } i=2 \text{ to } N \tag{4}$$

and $\sum_{i=1}^{N} x_i(s) + z(s) = y^N(s)$ where expectations are taken with respect to states of the world, s and $Ev(y_i(s), r_i, \mathbf{n}_i)$ is the expected utility of claimant i if he resided and worked outside the joint household with his property $\kappa_{ij}A_j$. The household remains intact as long as the maximand in expression (3) is positive. Note that the decision to divide the household depends only on whether the *current* surplus is negative. Presumably forward-looking household members experiencing a negative surplus today might consider the possibility that surplus might be positive in the future and not split if costs of household reconstitution are high. However, as discussed below the processes that generate changes in household surplus, such as technical advances and fertility, make reversals of surplus unlikely.

b. Parameterization of Preferences

The model as specified is silent on how the gains from joint residence are distributed. To obtain specific implications for household division and a tractable estimation strategy it is necessary either to specify distribution rules or, alternatively, to choose a form for the utility function from which implications can be derived that are independent of the intrahousehold distribution of coresidence gains. We choose the latter strategy. In particular, we assume a parametric utility function for each claimant of the form

$$u(x,z;\boldsymbol{n}) = \ln((x-\beta'\boldsymbol{n}-\alpha z)(z+\gamma)) + \delta r_i \qquad \textbf{(5)}$$

where r_i is assumed to be one for autarchic households and zero otherwise and thus δ denotes the direct preference for autarchic residence (privacy). This characterization of claimant preferences has the property that the household demand for the public good does not depend on the household distribution of coresidence gains, making it unnecessary to take a stand on which claimant is the maximizer. It nonetheless captures the key elements of the decision about joint residence. This result follows from the fact that (5) exhibits transferable utility. Bergstrom (1997) provides a useful discussion of transferable utility in the context of models of household behavior. The transferable utility assumption also implies that the distribution of assets within the household does not affect public goods consumption. Evidence based on household data from Cote d-Ivoire (Duflo and Udry (2001))is consistent with this implication of transferable utility.

Consider first the case of a claimant who lives apart so that the budget constraint is x+z=y, suppressing the state s for notational convenience. Assuming an interior solution,[5] autarchic consumption of the public and private goods, respectively, are

$$z = \frac{1}{2(1+\alpha)}(y-\beta'\boldsymbol{n}-\gamma(1+\alpha)) \text{ and } x = \frac{1}{2(1+\alpha)}((1+2\alpha)y+\beta'\boldsymbol{n}+\gamma(1+\alpha)) \qquad \textbf{(6)}$$

[5]It can be shown that at low levels of income the nonnegativity constraint for the pubic good binds and all income is allocated to the private good. This appears reasonable given that at low levels of income expenses will be dominated by food which would appear primarily to be a private good. It is also worth noting that the parameter γ has a simple interpretation when this constraint binds: it is the marginal utility of income.

and therefore autarchic utility is

$$v(y,\boldsymbol{n}) = \frac{1}{4(1+\alpha)}(y - \beta'\boldsymbol{n} + \gamma(1+\alpha))^2 + \delta \qquad \textbf{(7)}$$

Expressions (6) and (7) indicate that claimants with differing characteristics or incomes will exhibit different demands for public and private goods in autarchy. If, for example, one element of **n** is the number of children and the corresponding element in β is positive, a claimant with more children will wish to consume more of his income in the form of the private good, such as food for his children, and less in the form of the public good.

The model can also accommodate a special role for the household "head." As heads gain experience, for example, they may acquire greater preferences for household public goods as they identify with the collective household. This is analogous to the household head caring more about the private consumption of other co-resident family members. Thus an experienced head may have higher demand for the public good compared with other household claimants, in which case one element of **n** might be whether an individual is an experienced head. Under these circumstances, the death of a head may then reduce the coresidence gains for the remaining younger family members who had benefitted from the former head's support of the public good. Indeed, we show below that under certain conditions any factor that increases the demand for the household public good also increase the gains from household jointness.

Carrying out the maximization given by equations (3) and (4) for the parameterization of utility (5) and the expressions for autarchic utility (7) yields an analytic expression for the optimal level of the public good in the joint household

$$z = \frac{1}{2(1+N\alpha)}(y^{N} - \sum_{i=1}^{N} \beta'\boldsymbol{n} - \gamma(1+N\alpha)) \qquad \textbf{(8)}$$

While it is possible to derive a closed-form expression for coresidence gains and thus to derive comparative static results for an arbitrary distributions of income shocks, the expression is simplified with two assumptions. First, we assume that the production shocks take on two values, $\pm\Delta$, that occur with equal probability scaled by a multiplicative parameter ζ that depends on local climatic conditions. In the good (bad) state joint and autarchic incomes are given by $\bar{y}^N+\zeta\Delta^N-\tau^N$ $(\bar{y}^N-\zeta\Delta^N+\tau^N)$ and $\bar{y}_i+\zeta\Delta_i-\tau_i$ $(\bar{y}_i-\zeta\Delta_i+\tau_i)$, respectively, where $\bar{y}$ is the expected value of income and the τ's denote transfers in the two states which are equal and of opposite sign and may depend on the magnitude of Δ and whether one is married and/or has married daughters. Second, we define surplus in terms of a monotonic transformation exp(.) of expected utility. The results for when a split takes place hold for any monotonic transformation and, by choosing this transformation, the magnitude of the coresidence gains, not just whether they are positive, is independent of the particular identity of the maximizer and of the decision about how the gains from joint residence are distributed. Under these conditions the gains from joining together all of the claimants in one household are

$$\begin{aligned}\psi(y_i,\boldsymbol{n_i},y^N,i=1..N) &= \frac{1}{4(1+N\alpha)}([\bar{y}^N-\sum_{i=1}^{N}\beta'\boldsymbol{n_i}+\gamma(1+N\alpha)]^2-(\zeta\Delta^N-\tau^N)^2)-\sum_{i=1}^{N}Ev(y_i,r_i,\boldsymbol{n_i})\\ &= \frac{1}{4(1+N\alpha)}[(\bar{y}^N-\sum_{i=1}^{N}\beta'\boldsymbol{n_i}+\gamma(1+N\alpha)]^2-(\zeta\Delta^N-\tau^N)^2)\\ &\quad -\frac{e^{\delta}}{4(1+\alpha)}\sum_{i=1}^{N}([\bar{y}_i-\beta'\boldsymbol{n_i}+\gamma(1+\alpha)]^2-(\zeta\Delta_i-\tau_i)^2)\end{aligned} \qquad \textbf{(9)}$$

Equation (9) characterizing household surplus has implications for how household size, nuclear-family characteristics and productivity affect household break-up. However, the expression also indicates that intra-household claimant heterogeneity and income risk may play a role in the propensity of households to break up. To characterize more easily how the number of claimants, the characteristics of their nuclear families, and productivity change affects the probability of the household dividing net of heterogeneity and risk, we assume that all of the claimants are identical; i.e., $\mathbf{n_i}=\mathbf{n}$, $\kappa_{ij}=1/N$, $\theta_i=\theta_j^N=\theta$, $\delta=0$, and $W_i=W$ and that $\Delta_i=\Delta^N=0$ and $\tau_i=\tau^N=0$. Thus equation (9) becomes

$$\psi^*(\theta,N,A,W,\beta'\boldsymbol{n})=\frac{1}{4(1+N\alpha)}[\theta A+NW-N\beta'\boldsymbol{n}+\gamma(1+N\alpha)]^2 \\ -\frac{1}{4(1+\alpha)}N(\theta\frac{A}{N}+W-\beta'\boldsymbol{n}+\gamma(1+\alpha))^2 \qquad \textbf{(10)}$$

with, for α>0,

$$sign(\psi(\theta,N,A,W,\beta'\boldsymbol{n}))=sign((\theta A+NW-N\beta'\boldsymbol{n})^2-\gamma^2N(1+\alpha)(1+N\alpha)) \qquad \textbf{(11)}$$

Thus for this homogenous household, household division will take place when

$$\psi^*(\theta,N,A,W,\beta'\boldsymbol{n}))=((\theta A+NW-N\beta'\boldsymbol{n})^2-\gamma^2N(1+\alpha)(1+N\alpha))<0 \qquad \textbf{(12)}$$

Differentiating (12) with respect to N yields the effect of an increase in household size (number of claimants) for the same total household asset on the propensity to split the household:

$$\frac{\partial\psi^*}{\partial N}=2(\theta A+NW-N\beta'\boldsymbol{n})(W-\beta'\boldsymbol{n})-\gamma^2(1+\alpha)(1+2N\alpha) \qquad \textbf{(13)}$$

which is guaranteed to be negative when the demand for the private good is sufficiently high relative to the wage (W-$\boldsymbol{\beta}$**'**$\boldsymbol{n}$<0) and public good consumption is at an interior solution (θA+NW-Nβ'n>γ(1+Nα)>0).The intuition underlying this result is that an increase in the number of household members without an accompanying increase in household assets decreases per-capita income, which decreases the gains from coresidence given the assumed structure. This is offset somewhat, but not completely under the specified conditions, by the fact that an increase in the number of household members decreases the average cost of the pubic good under joint residence.

Similarly, the sign of the effect of any element of **n** on surplus and thus the probability of division depends, for the homogenous case, only on whether that element increases or decreases the level of consumption of the household public good. In particular, the effect of the kth element of **n** on household surplus is

$$\frac{\partial \psi^*}{\partial n_k} = -2AN(\theta A + NW - N\beta' \boldsymbol{n})\beta_k \tag{14}$$

and the effect on public good consumption is

$$\frac{\partial z}{\partial n_k} = -\frac{N}{2(1+N\alpha)}\beta_k \tag{15}$$

As these two expressions have the same sign, anything that increases public good consumption will also tend to increase surplus and thus decrease the probability of household division. Thus, for example, if it is found that the presence of young children in a joint household increases the demand for private relative to public goods, then the birth of a child may increase the likelihood of household division.

The effect of an increase in productivity is similar also to the effect of increasing the number of claimants. Differentiating ψ*(.) with respect to θ gives

$$\frac{\partial \psi^*}{\partial \theta} = 2A(\theta A + NW - N\beta' \boldsymbol{n}) \tag{16}$$

For a household that is at an interior solution in terms of the consumption of the public good (θA+NW-$\boldsymbol{N\beta' n}$>γ(1+Nα)>0), $\partial\psi^*/\partial\theta>0$ so that income growth tends to increase surplus and thus discourages division.[6] In this case the increased income results in an increase in consumption of the public good and thus there is an increase in the amount that can be saved by joint consumption compared with consumption under separate residence.

Less obvious implications of the model are that both heterogeneity within the household and

---

[6]Note that it is only in the neighborhood of ψ*=0 that the sign($\partial\psi^*/\partial\theta$)≈sign($\partial\psi/\partial\theta$); however, this is precisely the neighborhood in which a change in technology is likely to affect household division.

income risk also affect the likelihood of the household breaking up. With respect to intrahousehold heterogeneity, it is easily seen that, for given mean of y_i-$\boldsymbol{\beta}'\boldsymbol{n}_i$ and $y^N - \sum \beta' \boldsymbol{n}_i$ an increase in the intrahousehold (inter-claimant) variance of y_i-$\boldsymbol{\beta}'\boldsymbol{n}_i$ will tend to reduce ψ and thus joint residence, as (9) can be rewritten

$$\begin{aligned}\psi \;=\; & \frac{1}{4(1+N\alpha)}([\bar{y}^N - \sum_{i=1}^{N} \beta' \boldsymbol{n}_i + \gamma(1+N\alpha)]^2 - (\zeta\Delta^N - \tau^N)^2) \\ & - \frac{Ne^{\delta}}{4(1+\alpha)}[var(\bar{y}_i - \beta'\boldsymbol{n}_i) + mean(\bar{y}_i - \beta'\boldsymbol{n}_i + \gamma(1+\alpha))^2 - mean((\zeta\Delta_i - \tau_i)^2)]\end{aligned} \tag{17}$$

Intuitively, a higher intrahousehold variance of y_i-$\boldsymbol{\beta}'\boldsymbol{n}_i$ implies greater differences among claimants in their autarchic demands for the public good, whose consumption must be equal in the joint household.

Intertemporal income variability also can affect household break-up in the model. There are two mechanisms. First, the scale of operation may reduce income variance if yields are not perfectly correlated across plots, that is if $\Delta^{N^2} < \sum_i \Delta_i^2$. Second, household structure may affect the propensity of family members to engage in risk-sharing arrangements. We assume that the number of marriages associated with the joint household - the number of married daughters of household claimants and their wives - affect interhousehold transfers, so that $\tau_N = \tau_N(\zeta\Delta^N, D_1, .., D_N)$, where the D_i represent the number of married daughters and/or wives of each claimant, and the first partial is negative and all of the other partials are positive. Similarly, for the autarchic household $\tau_i = \tau_i(\zeta\Delta_i, D_i)$.

Differentiating with respect to climatic variability yields:

$$\frac{d\psi}{d\zeta} \;=\; -\frac{1}{2(1+N\alpha)}(\eta\Delta^N - \tau^N)\Delta^N(1 - \frac{d\tau^N}{d(\zeta\Delta^N)}) + \frac{e^{\delta}}{2(1+\alpha)}\sum_i (\eta\Delta_i - \tau_i)\Delta_i(1 - \frac{d\tau_i}{d(\zeta\Delta_i)}) \tag{18}$$

In the absence of transfers, as long as there is a positive taste for privacy (δ>0) and imperfect covariances across land plots, it is clear that surplus is increasing in climatic variability:

$$\frac{d\psi}{d\zeta} = -\frac{1}{2(1+N\alpha)}\zeta\Delta^{N^2} + \frac{e^{\delta}}{2(1+\alpha)}\sum_i \zeta\Delta_i^2 > -\frac{1}{2(1+\alpha)}\zeta\Delta^{N^2} + \frac{e^{\delta}}{2(1+\alpha)}\sum_i \zeta\Delta_i^2$$
$$> \frac{\zeta}{2(1+\alpha)}(-\Delta^{N^2} + \sum_i \Delta_i^2) > 0 \tag{19}$$

However, to the extent that ex post risk-sharing, which relies on marital ties to specific individuals, depends on whether those individuals are co-residing with other claimants, the existence of income risk may change propensities to break up upon the marriage of a son or daughter. Differentiating with respect to the number of married daughters, say, of claimant i yields

$$\frac{d\psi}{dD_i} = \frac{1}{2(1+N\alpha)}(\zeta\Delta^N - \tau^N)\frac{d\tau^N}{dD_i} - \frac{e^{\delta}}{2(1+\alpha)}(\zeta\Delta_i - \tau_i)\frac{d\tau_i}{dD_i} \tag{20}$$

Without additional assumptions, expression (20) cannot be signed. The reason is that the effect on the surplus from the joint household of an additional marriage associated with claimant i depends on the relative magnitudes of transfers that the married daughter would make to the joint household and to claimant i's split-off household, τ^N and τ_i, and the differential responsiveness of transfers to shocks across the two types of households. If, for example, daughters of the household head are more willing to provide insurance-transfers to the head when he is alone then when he is coresiding with his adult sons (and their wives),[7] then additional married daughters - or the marriage of a claimant - will reduce surplus. This can be seen clearly in the case in which transfers are provided by i's daughter's (wife's) family to i's household only if i's residence is not shared. In that case expression (20) becomes

$$\frac{d\psi}{dD_i} = -\frac{e^{\delta}}{2(1+\alpha)}(\zeta\Delta_i - \tau_i)\frac{d\tau_i}{dD_i} < 0 \tag{21}$$

---

[7] Lower insurance-based transfers under joint residence may arise, for example, from lower mutual altruism that increases the potential for moral hazard in the generation of income and/or commitment failures.

which is negative as long as transfers do not fully offset production shocks and an additional married daughter or daughter-in-law increases compensatory transfers. Expression (21) also shows that the departure of a daughter for marriage or the marriage of a claimant in a joint household, when marriage-based contributions to consumption-smoothing only occur in autarchic households, has a stronger negative effect on joint household surplus the greater is the magnitude of the income shocks, as given by the parameter ζ. Thus, "selfish" risk sharing creates an association between the marriage of claimants and household division.

c. Incomes, Technology and Information

The model implies that changes in the level and riskiness of joint household income y^N, changes in the level and dispersion of inter-claimant autarchic incomes, and marriage ties potentially affect the patterns of household break-up over time. However, to obtain a model that is estimable, we need to be more specific about the production technology governing autarchic and joint income production and how these evolve over time. We assume that permanent changes in the productivity factors θ_i and θ_j^N arise from two sources - changes in agricultural technology and farming expertise. In particular, we assume, as suggested by evidence in Rosenzweig (1995) and Foster and Rosenzweig (1996), that schooling augments the productivity of available new technologies through faster adoption and more efficient use via more rapid learning by doing. Thus we parameterize autarchic productivity at time t for individual i in area k as

$$\theta_{it} = \eta + \eta_{\phi}\phi_k + \eta_{\phi t}\phi_k \times t + \eta_S S_i + \eta_{St} S_i \times t + \eta_{S\phi t}\phi_k \times t \times S_i \quad \textbf{(22)}$$

where S_i is i's schooling, ϕ_k is local agricultural technical change, and the η's are parameters. Accounting for i's claim on assets and his full labor income and introducing parameters that capture static (η_A) and dynamic ($\eta_{A\phi t}$) scale economies in farm profitability total autarchic income is

$$y_{it} = (\eta + \eta_{\phi}\phi_k + \eta_{\phi t}\phi_k \times t + \eta_S S_i + \eta_{St} S_i \times t + \eta_{S\phi t}\phi_k \times t \times S_i)\exp((\eta_A + \eta_{A\phi t}\phi_k \times t)\kappa_{ij}A_j)\kappa_{ij}A_j + W_{it} \quad \textbf{(23)}$$

While the signs of the scale economies parameters are not obvious, we expect that the sign of $\eta_{S\phi t}$ is positive, given the productivity-augmenting role of schooling in an environment with technical change. Thus, in technical progress areas, any within-household differences in schooling levels will result in dispersion in autarchic incomes that will become more pronounced over time. This would reduce surplus, for given total income in the joint household, due to a decreases in consensus on the level of the public good to provide, and thus a greater likelihood of household splits in technical-change environments. However, as noted, an overall increase in household income tends to discourage household division if it increases the demand for the public good.

Finally, information among household members is likely to be readily shared within the household and thus operates as a public good in production. In particular, when individuals live together and jointly farm, the choice of best practice is likely to be determined collectively based on the co-resident individual with the best information. If the most informed individual is the one with the most schooling, the productivity of household j in area k at time t is

$$\theta_{jt}^{N} = \eta + \eta_{\phi}\phi_{k} + \eta_{\phi t}\phi_{k}\times t + \eta_{S}S_{j}^{\max} + \eta_{St}t\times S_{j}^{\max} + \eta_{S\phi t}\phi_{k}\times t\times S_{j}^{\max} \tag{24}$$

where S_j^{max} is the maximum schooling in the household.[8] Total joint household income is then

$$\begin{aligned} y_{jt}^{N} = {} & (\eta + \eta_{\phi}\phi_{k} + \eta_{\phi t}\phi_{k}\times t + \eta_{S}S_{j}^{\max} + \eta_{St}t\times S_{j}^{\max} \\ & + \eta_{S\phi t}\phi_{k}\times t\times S_{j}^{\max})\exp((\eta_{A} + \eta_{A\phi t}\phi_{k}\times t)A_{j})A_{j} + \sum_{1}^{N} W_{it} \end{aligned} \tag{25}$$

There are thus two potential sources of (expected) income gain arising from joint production - the costless sharing of information, which should always increase surplus, and static and dynamic technical

[8]Evidence in Foster and Rosenzweig (1996) suggests, for example, that whether anyone in the household has a primary education is a more powerful predictor of the adoption of high-yielding seed varieties in the early stages of the green revolution than whether the household head had completed primary schooling. Joliffe (1997) and Yang (1996) also present evidence of the importance of the maximum of the schooling of household members in determining household income. The relevance of within-household information sharing for the distributional implications of educational inequality is considered from an axiomatic perspective by J. Foster and Basu (1997).

scale economies, which increase surplus only if positive. Given (17), (23) and (25) within-household schooling inequality among claimants might therefore increase surplus in a high-growth setting and reduce household break-up by producing a large difference between total household income y_{jt}^N and the sum of the individual incomes y_{it}. The role of technical change in affecting household division is thus unclear.

d. Head Mortality and Household Division

An important feature of the data describing household division, discussed below, is that most splits appear to occur at the death of the household head, among households that have at least two heirs. The model is capable of delivering the result that the death of a household head reduces household surplus and thus is associated with household division, but it is necessary to specify a particular role for the household head. As we have shown in (13), a reduction in the number of claimants if claimants are homogeneous increases surplus because the increase in per-claimant income increases public good consumption. Thus, if the head is like any other claimant, his death will not likely lead to division. Moreover, if the head is different in a random way from other claimants, in preferences for example, his death increases claimant homogeneity, which we have shown also increases household surplus.

Two cases in which the head's death may lead to division are when the head has above average preferences for the public good, as discussed above, and when the head has superior knowledge about agricultural practices. As we show below, because of secular trends in schooling prior to the ARIS, household heads tended to have higher schooling attainment than younger cohorts. If the head's schooling is the highest in the household and schooling provides superior skill, then the head's death reduces the surplus from co-production. We will assess whether the level of the maximum schooling of a household affects division and test directly whether a head's characteristics matter for household public good consumption

III. The Data Sets and the Sample

The data used in this study are constructed from data files produced by the National Council of

Applied Economic Research (NCAER) from their Additional Rural Incomes Survey (ARIS) and from their follow-up panel survey, the Rural Economic and Demographic Survey (REDS). The 1971 ARIS provides information on 4,527 households for crop year 1970-71 meant to be representative of households residing in all rural areas of India in that year excluding Andaman and Nicobar and Lakshadwip Islands. The surveyed households are located in 259 villages in 100 districts. In 1982, NCAER canvassed all of the households in all of the original ARIS villages except those in the state of Assam in order to create a sample frame for a follow-up panel survey of the 1971 sample. Every household in the non-Assam villages in 1982 was traced back to the set of households existing in the ARIS villages in 1971, including those in the 1971 sample. On the basis of these complete listings, a new survey (REDS) was undertaken in 1982. The REDS resurvey included by design a subset of the 4363 non-Assam households in the 1971 round of the ARIS survey that met one of three pre-specified household division criteria: (i) households for whom the household head remained in the village between 1971 and 1982, (ii) households in which the head had died but the rest of the immediate male relatives of the head had remained together, and (iii) households in which the head remained alive but the household had divided, defined as at least one adult male relative of the head permanently leaving the household. REDS identifies which households in the original 1971 survey met each of the three criteria so that it is possible to identify which of all of the ARIS original households in 1971 split by 1982 and which experienced a death of the household head. For households in which the head was alive in 1982 but which had split, however, information is only available on the individuals in the original household who stayed with the original head. Moreover, if the head died between 1971 and 1982 and the household divided, no information is available for any part of the household in 1982. These exclusions are what make it necessary to estimate household division, and the identification of the sampling strata for the 1971 households from the 1982 canvassing makes this possible.

We examine the break-up of the 1971 households between 1971 and 1982 using information on all farm households in the 1971 ARIS survey, regardless of their status in 1982, in which the head of the household was aged 40 and over and in which there was complete information on all residing household members in the 1971 ARIS survey. This sample numbers 1,784 households, of which 1,387 households had one or more coresident claimants, e.g., a head and at least one son or brother, and thus the potential for break-up.[9] Table 1 describes in detail the sample criteria for the ARIS households and their classification based on the 1982 REDS.

We chose households with older heads for three principal reasons. First, older heads are more likely to have adult heirs, and thus can divide. Second, older household heads are more likely to be farming with their inherited landholdings, as their fathers are likely to have already died. Land entitlements for the household members of these older-head households can thus be more accurately specified, as we discuss below. Third, according to sources at NCAER, older household heads in 1971 who were still alive in 1982 were less likely to have moved out of the village or to have refused to participate in the resurvey. This is important because ARIS households that were classified in the1982 survey as "refusers" and outmigrants are not identified and are thus are included in the category of 1971 households not reinterviewed along with those households that split and for which the head had died. If these survey-eligible but excluded households are numerous, then the households categorized as divided at the death of the head would include many households that did not actually divide. Fortunately, only 4.8% of the *total* 1971 sample eligible for the resurvey refused to be interviewed or moved out of the village. Based on this overall figure, if none of the household refusers had divided by 1982 then at most 12% of the households classified as having a head death and a division by 1982 would not have actually divided, and at most 8.5% of all 1971 households classified as dividing by 1982 would have been misclassified. The misclassification figure for the sample of 40+ farm household heads is likely to be significantly lower.

Because information is available from the 1971 ARIS data describing the characteristics of the individuals in the households, it is possible to measure household inequality at the start of the 1971-82

[9]There are 1,111 households in which there is more than one claimant in addition to the head, so that if the head died household division would remain a possibility.

interval. Moreover, the availability of fertility histories for all married women in the household enables the characterization of the nuclear families of the claimants, the projection of changes in the size and composition of the nuclear families over time due to fertility and mortality, and the inclusion in the analysis of the roles of departed daughters and already-departed sons of the head in the initial year.[10] In particular, we can take into account that some the households observed in 1971 had divided. prior to the 1971 survey, which is important for accurately computing asset claims. For each household in 1971, we computed the number of the head's sons who had already left the household as the difference between the number of adult (10+) sons accounted for in the household roster and the number of the head's surviving sons born at least 10 years prior to the survey. This information is used in the structural model to obtain a measure of the asset claims of the sons while their fathers are still alive and to measure the initial intact (asset) conditions of all households, as described below. We also constructed a measure of departed daughters using the same method as for the sons.

Another important feature of the data is that the ARIS survey took place in the early stages of the Indian "green revolution," so that the experience of the households subsequent to 1971 was importantly influenced by unanticipated agricultural technical change and economic growth. Productivity growth occurred at very different rates across areas of India due in large part to differences in the suitability of the soil and climate for the new high-yielding seeds of the green revolution. Estimates of district-specific rates of technical change over the period 1971-82, obtained from profit function estimates based on these data (Foster and Rosenzweig, 1996), are available and can be used as an instrument to predict crop productivity growth. In particular, because the technical change estimates have error, we use them as instruments for a Laspeyres-weighted index of district-level crop productivity growth. Thus, it is possible to examine the impact of agricultural technical change on household division, on the evolution of household configurations and on income net and gross of household division.

For the set of the ARIS households for which there is complete information in 1971 on

[10]Almost no coresident adult sons of the heads had children over 15 years of age, so that departures of their children were unlikely.

household composition, there are also data on profits, gross cropped area, asset changes, financial transfers, and indicators of village-level adverse weather in each of three crop years starting in 1968-69. This three-year panel was used in Foster and Rosenzweig (1995) to estimate a profit function for the farm households. We used the same data and estimation procedure to estimate a measure of per-area profit variances based on per-hectare measures of profits, land assets (tubewells, fences, pumps), farm equipment (ploughs, carts, blades) and area devoted to high-yielding seed varieties in the last two crop years. Table A in the Appendix describes the procedure and the variables used and reports the fixed-effects IV estimates of the per-hectare profit function. Based on these estimates we computed a farm-specific, per-hectare variance measure of profit shocks as the square of the difference between the per-hectare profit residuals (actual less predicted profits) across the two years. Because this measure of riskiness contains measurement error, we regressed these farm-specific variance measures on a set of 16 state dummies and the average gross-cropped area of the farms, to control for the possibility that operational scale may affect farm-level profit variances, as discussed. We use the coefficients on the state dummies as measures of the profit risk faced by the households.

There are two caveats to this procedure. First, the estimates are obtained only for households that did not split. Although the fixed-effects procedure yields consistent estimates of the profit parameters even if, as our model implies, riskiness affects the probability of splits, if riskiness is associated with unobserved fixed characteristics of farms then our estimated risk measure obtained from intact households will be incorrect for divided households. We believe, however, that the largest component of exogenous risk is spatial or is orthogonal to the detailed measured farm-specific characteristics we have in the data. Second, our measure of risk includes any time-varying unobservables of households. If we have omitted important time-varying farm assets, for example, we will have obtained incorrect risk measures.

The first two columns of Table 2 provide information on the sample of the 1,387 landowning households in 1971 with a head aged 40 years and over that could divide after the initial survey, classified by whether or not the household had actually divided by 1982. The third column provides for comparison descriptive statistics on landowning households in which the head is the only land claimant. As can be seen from the table, more than a third of the households with at least one heir divided by 1982. And, division does appear to account for most of the change in average per-household landholdings - among the households that did not split there was an average decline of 1.5 acres (14.1%), while among those households with the same head of household at each survey round but that divided, average household landholdings dropped by 2.7 acres (21.8%). Because the households that divided began the period with larger landholdings it appears that household division equalized per-household landholdings. However, as we show below, the gross relationship between the initial size of household landholdings, even on a per-claimant basis, and subsequent division is non-linear.

Another important characteristic of the sample households in1971, given the perspective of the model, is the high proportion who have multiple claimants (the head and his sons or brothers) with unequal levels of schooling (56%). Within-household (and cross-household) schooling inequality evidently arises importantly from nonlinear and strong aggregate trends in schooling achievement in the Indian economy combined with high fertility. These schooling trends also are responsible for the fact that household heads, who tend to be the oldest claimant in most households, have higher levels of schooling than other household members. Comparison of the households that split with those that did not indicates, consistent with the model, that households that had more inequality initially were also more likely to divide over the 12-year period - among the households that divided between 1971 and 1982 with multiple heirs, almost 61% had unequal schooling at the start of the period. Among the group of non-dividing households, the comparable figure was just over 38%. The division of households thus contributed to lessening within-household inequality in schooling but to increasing differentials in between-household average schooling levels.

IV. Household Structure, Marital Ties and Risk Reduction

The model suggested that claimants' ties to other households that create the potential for risk-sharing arrangements can affect the household's propensity to break-up, particularly in high-risk areas. Table 2 also indicates that households that eventually divided and those that did not differed at the beginning of the sample with respect to the state-level risk measure and the number of external household ties associated with the claimants. In particular, households that eventually divided resided in slightly riskier areas and had on average a greater number of daughters of the household head who had left the household in the 1971, presumably for marriage, and a greater number of married claimants initially residing within the household compared with households that remained intact over the period. We will incorporate the effects of marital ties on household surplus through risk sharing in the structural model.

The three-year panel information on transfers and year-specific village-level indicators of adverse weather in the ARIS data enables a more direct examination of the extent to which marital ties affect the budget constraints of joint and autarchic households, in particular, the extent to which financial flows between households tied by kinship and marriage serve to reduce consumption variability. The panel data indicate that 13% of the farm households received a transfer over the three-year period. Based on the information provided on whether the villages experienced weather conditions that had an adverse effect on crops in each of the three crop years, 61% of the households also experienced crop shortfalls in at least one year Although the data do not indicate the source of interhousehold transfers, it is possible using the methodology in Rosenzweig (1988) and Rosenzweig and Stark (1989) and the household structure and weather data to indirectly test whether household marital ties are related to consumption-smoothing via interhousehold transfers.

Probit estimates of the determinants of the probability of a household receiving a financial transfer in a given year based on the panel provide four results pertinent to the study of household division.[11] First, the probability of receiving a transfer responds inversely to adverse weather, as would be expected if such transfers were manifestations of interhousehold insurance arrangements. The estimates indicate that while on average a household had a 2.5% probability of receiving a transfer in any

[11]The estimates are reported in Foster and Rosenzweig (2000).

year, that probability increases by 31% in a bad-weather year. Second, although on average the incidence of transfers does not depend on the number of departed married daughters or co-resident daughters-in-law, the probability of a household receiving a transfer in a bad-weather year is significantly increased the more married daughters-in-law that are residing in the household and the greater the number of daughters who have departed the household to be married. The point estimates imply that adding a departed daughter or a daughter-in-law symmetrically increases the probability of a transfer in a bad-weather year by 44%. Third, having departed sons does not add significantly to the household's probability of receiving a transfer in adverse weather. This latter result does not necessarily reflect the fact that sons are less tied to the origin household than are daughters, but likely reflects the fact that departed sons remain in the village and thus experience the same adversity as the origin household at the same time. Most of the married daughters reside outside the origin village and most of the daughters-in-law come from households residing outside the village.[12]

Finally, the probit estimates suggest that household jointness reduces the willingness of the households connected via marriage to participate as fully in risk-sharing arrangements with joint households as with separate households. The estimated effect of an additional departed daughter on the transfer probability in adverse weather was substantially less for joint households than for households in which the head was by himself, although the difference is not measured precisely. This evidence suggests, according to the model, that the positive effects of income variance on joint residence will be lower in households with more marital ties relative to those with fewer marital ties.

V. Determinants of Household Division: Probit Estimates

We first use the data on the 1971 farm households that have the potential to divide to directly estimate the determinants of the probability of household division making use of the insights of the model. Table3 reports probit estimates of the determinants of the probability that a 1971 household with multiple claimants split by 1982 based on the 1982 sample classifications. In the first column, estimates

[12]The reduction in interhousehold risk covariances associated with marital migration is the reason given in Rosenzweig and Stark (1989) for marital exogamy in India.

are presented from a specification which ignores the potential effects of household inequality, income growth, income riskiness and departed kin on household surplus, including only the number, marital status and *average* schooling and age of the claimants along with owned household landholdings, the number of children by sex. These estimates are conventional and unsurprising, suggesting that household size (the number of claimants and the number of claimant wives) and the age of the head (which strongly predicts head mortality) have significant effects on division - more crowded households with older heads in 1971 are more likely to have divided by 1982. Size of landholdings, however, appears to have no direct effect on the probability of household division, net of the effects of the other variables.

The second specification adds measures of household inequality that the model suggests are important determinants of household division. In particular, the specification adds the within-household distribution of schooling as measured by the intra-claimant variance and the maximum of schooling in the household. As expected based on equations (17), (23) and (25), increases in the intrahousehold variance of schooling (and thus autarchic income), given maximum schooling (and thus joint household income), increase household division, but increases in the maximum of schooling for given mean and variance result in decreases in household division. An additional interesting result, also seen in the first specification, is that division is more likely the higher the number of young male children in the household, but not female children. There are two possible interpretations of this. First, if male children are allocated more (private) resources than are female children, consistent with the sex-bias observed in India (Rosenzweig and Schultz, 1983; Sen, 1990), then the presence of male children reduces public good consumption and thus household surplus. However, having more boys also predicts additional claimants in the future, and additional claimants reduce household surplus. Additional (future) adult daughters, who depart, also affect coresidence surplus, but that effect depends on the differential strength of ties to parents and by parental co-residence.

The model also suggests that technical change affects household division to the extent that it influences heterogeneity in inter-claimant preferences for the public good, the difference between joint household income and the sum of autarchic incomes, and the overall income level. Although because of these multiple pathways by which technical change may influence the surplus associated with co-residence it is not possible to sign the overall effects of technical change and schooling on household break-up, there are implications for interactions between technical change and other household characteristics: technical change should reduce household division in a homogenous household (equation (16)) and should increase division probabilities differentially in households with high maximum schooling for given mean and variance of schooling (equations (17), (23) and (25)).

The third specification includes district-level crop productivity growth over the period 1971 through 1982, instrumented with the district-level technical change estimates of Foster and Rosenzweig, along with the state-specific measure of income variance per hectare. These results imply that, conditional on the income riskiness measure, on average higher income growth has a marginally significant negative effect on the probability of splitting. In the fourth specification, the instrumented productivity growth measure is interacted with the mean, variance and maximum of schooling for the household. The estimates of the interaction coefficients provide additional support for the model. First, at zero variance for schooling among claimants and at the population mean and maximum of household schooling, the coefficients imply that a 10% increase in yield per acre above the mean of 817 results in a 7.1% decrease in splits in multi-claimant households. Second, the coefficient on the yield growth × maximum schooling interaction is negative and significant at the 10% level - division is less likely in a high-growth environment the higher is the schooling of the claimant with the most schooling.

The coefficient on the income risk measure implies that surplus in joint households is on average higher in environments in which income risk is greater. The effect is small - a one-standard deviation increase in the variance of profits per hectare decreases the probability of household division by 6.8%. This implies that there are relatively small gains from land consolidation associated with imperfect correlations in local profit shocks. However, the model suggests, along with the results from the investigation of the determinants of transfers, that the effect of riskiness on joint household surplus is reduced the greater the number of marital ties in the household and, conversely, that adding marital ties reduce the surplus from coresidence in risky environments. In the final specification, we add the number of departed household members - sons and daughters of the head - and interactions between the income-risk measure and the two variables representing marital ties - the number of daughters-in-law and the number of departed daughters - that the transfer results suggest play important roles in determining risk-reducing interhousehold transfers. The addition of these variables does not substantially affect the estimates of the other coefficients in the specification. However, the added variables have significant effects on the probabilities of household division. The signs of the interaction terms, moreover, are consistent with the implications of the model and the transfer estimates. The point estimates suggest that a departed son increases the probability by 14% of the remaining household dividing. At the mean level of risk, adding one daughter-in-law to the household increases the probability of division by 66% and the marriage of a daughter of the head, at the mean, has little effect on the split probability. Where profit riskiness is one standard deviation above the mean, however, adding a daughter-in-law doubles the probability of splitting, and the marriage of a daughter increase the division probability by almost 10%.

VI. Estimation of the Structural Model

While the estimates in Table 3 permit an assessment of the validity of the proposed model and show the effects of changing household size, adding marital ties, the distribution of schooling, and rates of technical change on household division probabilities, they cannot be used to draw inferences about the composition of split households – who splits from whom. Consequently, given the non-linearities associated with technological returns to scale and within-household human capital spillovers associated with information sharing, the probit and regression estimates can provide only limited insight into the implications of technical change for growth rates in household incomes and the differentials in growth by economic strata. Moreover, the finding that diversity and household break-up are correlated is consistent with alternative mechanisms than disagreement about the household public good, such as increased coordination or information costs. To examine these issues we adopt a more structural approach.

A key advantage of the structural approach is that with structural estimates from the model we can assess how technical change affects incomes inclusive of its effect on household division. We can compare these results to what would have been predicted ignoring household break-up, that is, as if all households stayed intact. We estimate the parametric structural model described in equations (9) and (25) based on the explicit asset claims of household members by fitting predictions of the model, for given parameter values, to the ARIS-REDS data on initial and final-period incomes and land holdings, and information on household division using a simulated method of moments procedure as discussed below. The estimated model parameters can then be used to assess the importance of endogenous household division in affecting the growth of household incomes induced by improvements in agricultural productivity.

a. Asset Claims and Family Assets

Estimation of the structure of the model, which involves the computation of household surplus at any given moment of time, so that we can estimate the division of landholdings depends importantly on the claim of each household member on household assets. Given that we do not explicitly observe the division of assets at the time of division, and could not in any case observe the claims on assets for household members that do not leave, it is necessary to explicitly specify asset claims in the model. The institutional rules governing claims on family assets in India are reasonably clear and can be used, in conjunction with the individual relationship information in the data, to impute asset claims for each household member. For each household in the 1971 sample, the set of individual asset (land) claimants was defined as the household head and any son or brother of the head of the household aged 10 years or older. Based on the practice of patrilocal partible inheritance that predominates in South Asia, we determined, based on the number and relationship of each of the claimants to the household head, the share of household land that each claimant was entitled to if he should separate from the household. In particular, under these rules upon the death of their father brothers are entitled to equal shares of the land.

The land for which the head has formal ownership rights does not correspond to the observed landholdings A_j of the head's household in cases in which any of the sons of the household head have already left the household and "taken" some of the land with them. We assume that sons splitting from the household before the death of the father have a claim to some but not all of their eventual inheritance. In particular we assume that sons are entitled to a fraction λ of the share of the father's total inherited land A_j^* to which, by the inheritance rules, they would be entitled upon the death of their father. Thus in a household with b co-resident brothers of the head (assuming brothers who have left the household have "taken" their land with them) and s sons of the head gs of whom have left the household with their asset shares, the amount of land over which the head has formal ownership rights, that is, pre-division assets A_j^*, is given by the formula $A_j^* = (A_j/(1 - ((gs)\lambda/(1+b)s)))$. This implies that the entitlement κ_j for a son in a household in which some of his brothers have already departed is $1/((1+b)s/\lambda - gs)$. Upon the death of the head, we assume that the departed sons take the remaining fraction of their inheritance from the household.

Because information on the number of sons that have previously left a household is critical for computing land entitlements among the remaining claimants in the household, such information is also necessary for predicting future household divisions. Indeed, the relationship between entitlements, observed household landholdings, and departed sons may account in part for why in the probit regression the number of departed sons, for given observed household landholdings, was correlated with the probability of a household's subsequent division. Information on the number of already-departed sons and knowledge of λ, the pre-inheritance entitlements of sons, are also necessary to backcast or reconstruct the pre-division landholdings of already divided households based on observed household landholdings in the current period. Such information thus can be used to more accurately predict household division prior to the sample survey. However, λ is not observable and can only be estimated as part of the structural model.

b. Fertility, Marriage, Preferences for Private Goods and Income Variance

Because changes in family size among claimants and the addition of marital connections can affect household surplus, to assess changes in surplus over time it is also necessary to project both fertility and marriages for each household claimant in the 1971 survey.[13] We do this by using the fertility and marriage histories and information on coresident daughters available in the demographic component of the 1982 survey to estimate fertility, nuptiality, and residence prediction equations that condition on claimant age and schooling, as discussed in Appendix A. Starting in 1971, births by sex and changes in marital status in 1972 were then drawn at random from the distributions implied by these estimated equations given the claimant's characteristics in 1971 (age, marital status, schooling, and composition of children by sex). Marital status and family size by sex were then updated and used as a basis for simulating demographic changes in 1973 given the simulated characteristics in 1972 and so on. Simulated marriages of daughters in that year were used to increment the stock of absent daughters for each claimant. Similarly, simulated household divisions, as discussed below, were used to increment the stock of absent sons. This process was repeated through 1982 yielding the required birth, marriage, absent children, and residence trajectories that were used to construct the indices $\boldsymbol{\beta}'\mathbf{n}$ determining the claimants' demand for private consumption in each year. The number of absent daughters, along with landholdings, an indicator for joint residence, and the marital status of the claimant, are used to compute the effective income variances, that is the income variances net of insurance-based transfers, for autarchic and joint households. In particular, the effective income variance for autarchic and joint households are:

$$(\zeta\Delta_i - \tau_i)^2 = \zeta^2 \times \exp(\xi_0 + \xi_1 D_i + \xi_2 mstat_i + \xi_3 \lambda\kappa A^N)$$

and

$$(\zeta\Delta^N - \tau^N)^2 = \zeta^2 \times \exp(\xi_0 + \xi_4 \sum_i (\xi_1 D_i + \xi_2 mstat_i) + \xi_3 A^N),$$

[13]Of course, from that fraction of the 1971 household that is reinterviewed in 1982 it is in principle possible to construct actual fertility and marriage histories. But no information is available for those individuals who split off and the fertility and marriage patterns in 1982 among intact households are obviously selective with respect to household division over the 1971-1982 period.

respectively, where D_i denotes the number of absent daughters of claimant i and mstat takes on the value of one if claimant i is married.

c. Estimation Method

For a given set of parameter values, the data on asset claims, schooling, technical change, demographic characteristics and expected incomes for each claimant and for the household as a whole are used to compute the degree of surplus ψ for any potential subset of coresident claimants based on (9) in each year. Computation of division configurations based on the surplus calculations proceeded as follows: The viability of a given household represented in the 1971 survey in 1972 was determined by examining whether the collection of 1971 claimants, given the predicted 1972 values of their individual and joint incomes and their individual demographic variables, generated positive surplus. If so, the household was assumed to remain intact in 1972. If not, the surplus was computed for the subhousehold that would remain if the i^{th} claimant were to depart from the household and take his share of the land $\lambda\kappa_{ij}A_j$.[14] This computation was made for each of the N-1 claimants other than the head for each set of parameter values. The 1972 joint household was then determined to be that subhousehold that provided the greatest surplus. This procedure is equivalent to choosing the split-off that maximizes the sum of the utilities of all claimants. Given that the functional form for individual utility (4) is consistent with transferable utility any division that yielded a lower sum of utilities would not be efficient.

Similarly, the determination of whether a household division would occur in 1973 was made by examining whether the resulting household from 1972 would generate positive surplus in 1973, and so forth for every year until 1982. In cases in which the death of a head was recorded over the 1971-1982 interval, it was assumed that the death took place in 1977, at which time headship was reassigned to the next oldest claimant, land that formed the remaining portion of the inheritance (1-λ) for sons that had previously left the household was allocated to them and the pre-inheritance parameter λ was set to one

[14]This algorithm does not allow for the possibility of splits by coalitions. We also attempted to estimate a model with the same structure but in which all possible combinations of household members were computed in each period. Although a variety of different initial values were tried for the parameters, this procedure failed to converge.

for the land shares of the head's sons. The household surplus algorithm was then applied to the new household, with the surviving joint household dividing only if it was determined that the computed surplus among the surviving claimants, given their new entitlements, was negative. The resulting predicted household structure in 1982 was used to compute 1982 household income as well as the land that would be held by the household given the divisions that are predicted to have taken place.

Parameter values were obtained using the method of simulated moments, with four moment conditions matching (i) predicted incomes in 1971 and (ii) an indicator of whether a split was predicted over the 1971-82 period to the actual data for all 1,387 households and (iii) predicted landholdings and (iv) incomes in 1982 to the values provided for the 1,022 intact and divided panel households in the 1982 survey. We chose a simulation-based method because of the need to use simulated rather than actual trajectories of births and marriages to determine surplus and thus split probabilities. The method of simulated moments was selected rather than simulated maximum likelihood because the former yields consistency without the requirement that the number of simulations grow with the sample size (Pakes and Pollard,1989). Details of the numerical procedure are given in the Appendix B.

VI. Structural Parameter Estimates

The top panel of Table 4 presents the estimates of the structural preference parameters of the model; the bottom two panels report the estimates of the income-generating η parameters from (25) that were jointly estimated with the structural parameters and the variance-effects parameters ξ, respectively. All of the parameters, except for that relating landholdings to income variance, are estimated with a reasonable degree of precision. The α parameter is positive, indicating a diminishing demand for the public good as incomes increase, as was assumed in the theoretical derivations of the effects of technical change and the number of claimants on division probabilities. The estimate of λ indicates that the head retains an important share of the heritable asset prior to his death - the share of the heritable asset that can be split off prior to the death of the head is evidently around 40 percent of that to which the claimant is entitled after the death of the head (father). δ and γ are also positive, indicating, respectively, that there is a positive demand for privacy and positive marginal utility of income for all positive levels of consumption.

The negative and statistically significant β coefficient for the household head indicates that the head has a special role in holding the household together.[15] In particular, household heads tend to value the private good less than do other household members. This suggests, net of the other reasons that the death of the head may affect household division embodied in the model, including those related to changes in the distribution of schooling, the number of claimants, and the claims on assets for sons of the head, why there is a high probability of division among households experiencing a death of the head. The positive β coefficients for the number of young daughters and sons indicate that the birth of a child increases the demand for private goods in the household, with private-good demand being higher by almost 21% for boy children than for children who are female. Note that this differential in the estimates of sex-specific intrahousehold child resource allocations, based solely on data describing household division and the total incomes of the households, is consistent with the observed higher rate of mortality for girls than for boys observed in India (Rosenzweig and Schultz (1983), Sen (1990)) which have been used to infer inequality in the intrahousehold allocation of resources to boys and girls. The marriage of a claimant, net of the effect of the birth of a child, decreases the demand for the private good.

The scale effects parameters associated with the production technology indicate modest static scale diseconomies - income gains evidently arise, for given schooling distributions, from farming with smaller landholdings. The point estimates indicate that for each acre increase in land, income is reduced by 3.8%.[16] It is possible that the evidence on scale diseconomies is biased due the existence of a negative correlation between average household land quality and size of landholdings, as found by Benjamin (1995) for Indonesia (using household structure as one identifying instrument). To assess the potential

[15]Note that by allowing head status to only affect tastes for the person who is household head at the beginning of the survey period we are assuming in effect that only an experienced head has different tastes for the public good and that the original head has served as head for some time. The absence of detailed data on the duration of headship precludes direct examination of this assumption.

[16]

importance of this we estimated using household fixed effects the relationship between farm profits per acre and size of landholdings, district-level productivity, and productivity interacted with landholdings based on our panel data. If land quality varies importantly across households - the source of land quality bias in the cross-sectional estimates of Benjamin - but land quality does not significantly vary across land plots within households then these within-household estimates are uncontaminated by the absence of a measure of land quality. The fixed-effects estimates (N=1231), which control for household-level land quality, also indicate the presence of both static and dynamic scale economies: profits per acre = -54.5*land - 68.0*land*productivity + 879.2*land productivity, with absolute values of t ratios of 1.43, 2.69 and 2.93 respectively.

The estimates thus indicate that there are two costs associated with coresidence/coproduction - reduced utility from joint residence and lower land productivity. On the other hand, there are dynamic scale economies - technical change enhances productivity per acre more for households with larger landholdings ($\eta_{\phi At}>0$), with the point estimate suggesting that after 11 years a one acre difference in landholdings is associated with an extra 2.8% increase in output per acre in low as compared with high technological change areas.[17] These effects together suggest that how households divide matters for the level of and changes in income per claimant and that technical change, net of income effects, increases the surplus from jointness. The estimates of the determinants of land productivity also indicate, however, that the effect of technical change on income growth is also greater the higher is the maximum level of schooling in the household ($\eta_{\phi St}>0$), with the point estimate indicating that a one-year increase in the maximum schooling of a household would after 11 years increase output per acre by 143 rupees more in high as compared with low-technical change areas. This result implies that inequality in schooling among claimants will be more strongly associated with the variance of autarchic income in high technical change areas than in those areas with stagnant technology.

[17]This is consistent with findings in Foster and Rosenzweig (1995) that larger farms optimally plant a greater amount of new-technology seeds in every period and thus, through learning-by-doing effects, experience more rapid productivity gains.

Finally, the estimated risk-based income variance parameters, again obtained based only on information on household splits and income change, suggest, consistent with the empirical findings on interhousehold transfers, that the marriage of a claimant who resides in a non-joint household reduces the net variability in household income (by 20% (ξ_2=-.2)) and that the contribution of the claimant's marital ties to variance reduction is 40% less if he resides in a joint household (ξ_4=.59). The absence of formal insurance markets thus enhances the prospects of household division upon the marriage of a claimant. However, one coefficient appears at odds with those obtained from the probit analysis of division and transfers, as the number of departed daughters (D_i) appears to be associated with increased net household income variability.

VII. Structural Model Validation

The structural model parameter estimates appear to predict reasonably well the patterns of household splits in the sample. To assess the relationship between initial inequality in household resources at the individual level and subsequent household division and income growth, we stratified the sample households, using 1971 sample weights, into four strata according to the amount of landholdings per claimant. The model mimics the observed inverted-u-shaped pattern of split probabilities across the four per-claimant landholding groups. The model tends, however, to underpredict household division in the bottom part of the distribution relative to that in the top quartile, possibly because households with smaller landholdings are the most likely to migrate and thus to be misclassified in the data as a divided household with a deceased head. Households with lower per-claimant landholdings also have small total holdings of land.

Although the estimated parameters appear to predict well actual income change for those households included in the panel - the correlation between predicted and actual income change is over 0.4 - the structural model estimates account for only a small part of the relationship between the odds of a household split and the 1971 household characteristics. We added to the probit specification in column five of Table 3 the predicted probability of a split for each household over the 1971-82 period obtained from the structural estimates. The coefficient was .369, with an asymptotic standard error of .102 and the variable increased the pseudo R^2 of the equation by by 12%. Thus the non-linear transformation of the regressors that should predict splits if the model is correctly specified contributes significantly to explanatory power even after controlling for the linear effects of these regressors. The coefficient and its statistical significance was similar when the other regressors were excluded, but the pseudo R^2 of the predicted variable alone was only .006, and the addition of the regressors increased the pseudo R^2 by a factor of ten. Clearly, the simplifications incorporated in the structural model, in particular the specifications used to predict the fertility and marriage trajectories, omit important factors leading to household division that are picked up by the regressors.

The availability of data in the 1982 survey describing household consumption expenditures provides a second basis for assessing the ability of the structural model to track household behavior as well as to assess the linkage between household jointness and public good demand that is the core of the theoretical model. Consumption expenditures are disaggregated into 12 categories, among which is expenditures on fuel and electricity. Such expenditures contain an important public goods component that is likely to make up a large share of the overall public good in a household. In implementing the estimation of the structural model, predicted public good consumption for each sample household at each point in time is computed based on the estimated parameters. If the model is appropriate and household fuel and electricity expenditure represents an important household-specific public good then, even though expenditure data were not actually used in the process of estimation, one would expect the predicted public good expenditure shares from the model to predict actual fuel expenditures.[18] We regressed actual household fuel expenditure in 1982 on the predicted public good expenditure for 1982 from the model and the household's actual 1982 total consumption level, with the latter included in order to allow for the possibility that fuel expenditures are only a component of public good expenditures and that within the

[18]We did not use the information on fuel expenditures to estimate the parameters of the model precisely because fuel expenditures contain both private and public-good elements. However, we would expect that the measure is highly correlated with true public goods consumption.

class of all public goods fuel expenditure may be more or less income elastic. Figure 1 presents the leveraged residual plot based on the estimated regression equation (for the 1,387 households):

$$Fuel\ Share\ =\ -4.39+8.08(4.19)Pred.\ Public\ Good+.0108(4.68)Consumption\ Expenditure+e \qquad \textbf{(26)}$$

with t-ratios in parentheses. The structural parameter estimates, obtained using only information on household division and income, appear to successfully predict household public goods consumption. It appears that household public good consumption and household division are thus importantly related as implied by the model. Predicted household public good consumption predicts household division. The coefficient on the predicted household public good in a probit regression of household division on the predicted household public good and household consumption is -.188, with a standard error of .0535.

VIII. Assessing the Effects of Technological Change

In this section we use the structural parameter estimates and the predicted growth variable to assess how economic growth effects income mobility, taking into account the endogenous division of households that in turn affects income. In particular, we want to assess how households classified by per-claimant landholding fare relative to each other when there are advances in agricultural productivity. In order to do this, we keep track of the changing household affiliations of the initial claimants, as predicted by the structural estimates. As noted, it is not obvious *a priori* how technical change, for given within-household schooling inequality, affects the propensity for households to divide - although higher technical change results in greater propensities of household division due to the higher variance in autarchic incomes, the gains from co-residence and co-production are increased both by the increase in incomes and the higher return to joint production arising from the sharing of productivity benefits accruing to the highest-schooled claimant (S^{max}) under rapid technical change. Nor is it obvious based on the parameter estimates, for analogous reasons, how an increase in the within-household variance of schooling will affect household division or how this effect is influenced by technical change, at least to the extent that an increase in schooling variance, for given mean schooling, is associated with an increase in household maximum schooling.[19] An assessment of the effects of technical change and of changes in within-household schooling inequality on household division and income change thus requires simulations.

In order to examine the role of household division in the relationship between technical change and income growth, we used the structural estimates to predict the rate of income change for low and high levels of technical change under two scenarios. In the first we allow households to divide, as predicted by the model, and sum incomes over the predicted component parts of the original household. We refer to the income summed over all of the subcomponents of the original households as "dynastic" income. In the second scenario we predict household income conventionally; that is, based solely on the existing asset distributions and determinants of income change; i.e., under the assumption that households remain intact (but age and reproduce). The simulations were carried out using the 1971 ARIS sample of cultivating households that were at risk of dividing over the 1971-1982 period. In particular, the model parameter estimates were used to generate, for each household, predicted 1971 income levels (inclusive of full-time labor income), predicted 1982 incomes for the same household if it remained intact or of each split-off household from the original household if it divided as determined by the structural model, and predicted 1982 income under the counterfactual assumption that the 1971 household remained fully intact.

Figure 2 plots the difference in average predicted dynastic and intact-household income growth by the 1971 per-claimant land quartiles, corresponding to zero technical change (low) and average technical change (high). Both sets of predictions indicate that claimants residing in households with more land per claimant benefitted more from accelerated agricultural productivity growth. However, comparison of the two sets of columns indicate that both the amount of and inequality in income growth is overstated if household division is not taken into account. In particular, dynastic income increases

[19]It is possible, at least in principle, to experience an increase in the intrahousehold variance in schooling with the maximum schooling in the household fixed. Such an increase in variance would unambiguously increase splits, with larger effects in areas with rapid technical change. This prediction is substantiated by the probit estimates from Table 3, which control for maximum schooling.

14.5% more than intact-household income in the lowest quartile group but by 14% and 27% less in the two highest groups. Accounting for household division is evidently important in assessing the distributional impact of technical change, and household division evidently dampens the adverse distributional consequences of augmenting agricultural productivity..

An important factor accounting for the difference between the counterfactual or naive simulation and that incorporating household division is the substantially stronger negative effect of increased agricultural technical change on division probabilities for the farm households in the top quartiles of the per-claimant land distribution. The first row in Table 5 displays the change in the rates of division for the four per-claimant landholding groups due to accelerated technical change. Moving from a low to a high technical change environment evidently slightly increases division propensities for the bottom two quartiles but decreases substantially household division in the top two quartiles. The relatively higher rates of land division in the lower quartiles than in the upper quartiles would raise incomes in the bottom groups relative to the top groups due to the operation of scale diseconomies.

Why did agricultural productivity growth reduce division in the upper quartiles? One reason is that increases in agricultural productivity growth raise incomes and therefore the demand for the household public good more in the land-rich households. The last row of Table 6 displays the effects of growth on the demand for the household public good in the original 1971 households by quartile, which shows a monotonic positive, but decreasing effect as per-claimant landholdings rise. Another contributing factor is that average 1971 within-household inequality in claimant schooling was larger in the households in the top two quartiles. Although high rates of technical change exacerbate autarchic income differentials for given schooling heterogeneity, and thus enhance disagreement in preferences for the household public good, this effect was evidently outweighed by the gains from human capital externalities associated with returns to information sharing that are enhanced when rates of technical change are high. To quantify the role of intrahousehold inequality in mediating the effects of technical change-induced income growth on household division, we carried out a simulation in which we eliminated all of the intrahousehold variance in schooling while maintaining the same average schooling level in each household. The second row in Table 6 provide the resulting growth effects on division in this scenario and can be compared to those predicted on the basis of the actual amounts of within-household schooling inequality in the first row. As can be seen, the mean-preserving decreases in the intrahousehold schooling variances reduce dramatically the negative effects of technical change in the top two quartiles, by 35% and 61% respectively. This is consistent with the fact that upper-income households evidently would lose the benefits from co-producing with a highly skilled claimant by dividing household assets among claimants with lower levels of schooling under a high technical-change regime. Within-household schooling inequality evidently plays an important role in the division of households and in income mobility associated with technical change.

IX. Conclusion

Despite long-standing concerns about the extent to which the benefits of income growth resulting from technological progress accrue primarily to better-off households, little evidence has, to date, been brought to bear on this issue. Empirical evidence on this subject consists largely of studies examining changes in the distribution of income across households over time using repeated cross-sectional data and the relationship between these changes and income growth. These studies are incomplete in two important ways. First, they do not distinguish between income growth that may in part be influenced by the income distribution and the effects of exogenous technological progress on the distribution of income. Second, because the data sources used do not permit the tracking of individual households over time, these studies can only indirectly address the central question: to what extent is the effect of technical change on household income growth higher among those with greater wealth?

In this paper we use newly-available panel data to examine the consequences of economic growth propelled by agricultural technical change on economic mobility. Long-term panel data of the type that we study reveals that an important impediment to the study of income mobility and the evolution of the distribution of income over time is the changing composition of households. In the context of rural India this is principally due to household division, which also is evidently a more important determinant of household wealth changes than is household-specific accumulation. Thus, to understand the consequences of economic growth for changes in the income distribution and for economic mobility as well as to fill in missing data arising from the incomplete follow-up of split-off households typical of a long-term in panel survey we develop and estimate a model of household division.

The model, consistent with conclusions drawn from the recent literature on the microeconomics of collective household behavior, of necessity treats the household as a collection of individuals rather than a single economic unit. In particular, we show that a model in which the decision about joint residence is driven by the presence of household-specific public goods captures key features of household break-up and yields sufficient structure to allow the identification of underlying preference parameters even in the absence of direct observations on resource allocation within the household or expenditure patterns. In so doing our results both show the potential importance of and suggest a new direction for research on non-unitary models of household behavior.

Our results, based on estimates of the structural model, show that consideration of the process of household division can enrich our understanding of the effects of technical change on inequality. In particular, inattention to the implications of technical change for household division would lead one to overestimate the extent to which higher levels technical change differentially benefitted better-off households. Due to the importance of human-capital externalities in production, combined with greater within-household schooling inequality in richer households, and the presence of decreasing returns to scale in production, technical change that occurred during the first decade of the green-revolution in India tended to differentially reduce household division among households with more land resources per capita. Because of these reductions, the average effect of technical change on income growth for the members of these resource-rich households was considerably attenuated relative the effects in less wealthy households at the beginning of the period.

There are a number of simplifications that have been made in the construction of the model and its application that reflect the constraints of the data set, practical considerations, and the particular setting being examined. Some of these simplifications may be relaxed in future research. First, we do not explicitly model decisions about human capital acquisition, marriage, and fertility that take into account their potential implications for household break-up. This simplified the analysis by permitting division probabilities to depend only on current-period observables; allowing fertility and marriage to be influenced by expectations of household splits would have necessitated the development of a full-scale dynamic model with forward-looking agents and thus substantially increased the computational burden. Moreover, these simplifications have some justification in the context of this paper. Levels of education and the marriage and fertility of the claimants in the data were at least partially resolved before the differential rates of technical change that resulted from the introduction of new seed varieties in India became known. The presence of early marriage and high marital fertility also suggest that biological constraints are likely to be of primary importance in determining variation in these measures. Clearly, however, if a model is to be used to predict income mobility over several generations in the presence of sustained levels of technical change it would be necessary to jointly model the processes of educational acquisition, marriage, fertility, and household division.

Second, reflecting the fact that the data set covers only rural households, the analysis focuses exclusively on agricultural households. Given low levels of migration and urbanization over the study period, this does not appear to be a major shortcoming, although again it is clear that a long-run assessment of the process of income growth in India is likely to require attention to the non-farm sector in general, and in particular to the process of urbanization. Perhaps a more important limitation of the analysis is its focus on the division of assets among land owning households, and thus its neglect of the poorest sector of the rural population, landless laborers, who make up a quarter of the agricultural workforce. Analysis of this group would require a detailed model of the labor-market implications of technical change and is thus left for future research. The basic structure of the model in which disagreement over the consumption of the household public good and the role of individual-specific autarchic incomes, however, would appear relevant for an examination of household division and mobility in all sectors of the economy.

Third, the static model we employ does not allow for storage possibilities. Our model and the empirical results suggest, however, that risk sharing considerations are important for understanding the structure of households. This simplification thus may be important. The literature on savings behavior, on the other hand, has assumed that the cross-sectional and intertemporal variation in household structure is exogenous (e.g., Deaton and Paxson, 2000). An important agenda for future research is thus the theoretical integration of decisions concerning savings and the organization of the household.

Finally, we have not used the data to examine the potential implications of the model for intra-household allocation decisions and expenditure patterns, which has been the primary focus of the existing literature on the non-unitary household. The model we develop has testable implications, for example, for the share of the budget spent on private goods such as food and for the extent to which private goods are allocated towards certain members of the household. Indeed, most previous work on non-unitary models focuses on intra-household division among married couples given the possibility of divorce, despite the fact that that option is not often exercised among those who have been married for some length of time. The fact that household splits are common in this population, not only makes the basic notion of division as an important factor in household allocation more plausible, but also makes it much more likely in longitudinal data that one will observe consumption patterns before and after a split has taken place, thus permitting proper account to be taken of heterogeneity across households that affects both the distribution of resources and the process of household division.

Appendix A

Simulating Changes in Marriage and Children

The estimated marital status logistic regression for men aged 10-40 in 1982 relating the probability of ever being married (mstat=1) to age and schooling (ed) was:

$$\text{Prob}(mstat=1|age,ed)=1/(1+\exp(-(10.2+.703age-.00815age^2-.143ed))) \tag{27}$$

with each of these coefficients being significantly different from zero at at least the 1% level. Marital status for those ever married in 1971 was assumed to remain unchanged over the subsequent 12-year period. Marriage probabilities for those unmarried in 1971 for each of the next 12 years was computed using (29) by deriving an expression for the marriage hazard:

$$\text{Prob}(mstat[x+1]|age[x+1],ed,mstat[z]=0)=\frac{\text{Prob}(mstat[x+1]|age[x+1],ed)}{\text{Prob(mstat[x]|age[x],ed)}}-1 \tag{28}$$

where age[x] denotes age in year x and so forth.[20] Given the absence of first marriage after the age of 40 in this population, the marriage hazard was assumed to be zero after that age. These predicated hazards were then use to stochastically simulate ages at marriage for the never married (in 1971) heirs and heads in the sample.

The stock of male and female children of each household claimant for each year in the interval 1971-1982 were constructed by accounting for the observed stock of children in 1971 and estimating fertility (net of deaths in the first year of life) equations by sex as well as rates of household departure for daughters.[21] Using the sample of married women in 1982 with husbands age 10-60 we estimated the

[20]Equation (24) follows from the assumption that there is no unmeasured heterogeneity in the marriage hazard underlying equation (23).

[21]Given patrilocal residence and the early age at marriage for women, the departure of daughters importantly affects the number and sex composition of the children of heirs in joint households and thus needs to be accommodated in the model. The departure of sons of the head is captured explicitly by the model. The sons of other heirs rarely leave the household before the heirs themselves have left and thus, for the purpose of examining the decision to divide the household, the independent departure of male

following multinomial logistic equations predicting the probability of having a surviving son or daughter that is one year old as a function of the father's age, education, and number of coresident sons and daughters (*boys* and *girls*, respectively) aged two and above :

$$\ln\left[\frac{\text{Prob}(boy\,|\,age,ed,boys,girls)}{\text{Prob}(none\,|\,age,ed,boys,girls)}\right]=-7.747534-.0236832ed+.3994859age \\ -.0066129age^{2}-.2731266boys-.025989girls \tag{29}$$

$$\ln\left[\frac{\text{Prob}(girl\,|\,age,ed,sons,boys,girls)}{\text{Prob}(none\,|\,age,ed,boys,girls)}\right]=-5.087247-.0168241ed+.2186073age \\ -.0039797age^{2}-.1701104boys+.0534046girls \tag{30}$$

With the exception of schooling and the number of girls, each of the estimated coefficients was significantly different from zero at at least the 5% level. Note that fertility has the expected concave shape in age and that additional sons lead to a significant reduction in subsequent fertility but additional daughters do not.

Finally, the departures by daughters from the household over the sample period were simulated by estimating OLS equations relating the numbers of girls living at home and away to husband's schooling, current age, and age at marriage using the same sample of married women as was used to estimate the fertility parameters:

$$girls\ \ at\ \ home=1.07-.0011ed-.0582age+.00121age^{2}-.0207age\ \ at\ \ marriage \tag{31}$$

$$girls\ \ away=-3.341+.0057ed+.239age-.00278age^{2}-.0139age\ \ at\ \ marriage \tag{32}$$

With the exception of schooling, each of the estimated coefficients is different from zero at the 1% level or greater. The departure probability of girls currently in the household as a function of their father's

children can be disregarded.

characteristics was computed from these equations as the ratio of the predicted change over consecutive ages (i.e., age x to x+1) in the number of daughters away from the household divided by the predicted stock of daughters at home at that father's age. The resulting departure probability was then applied to the stock of daughters at home in each year for each claimant to simulate according to a binomial distribution the number of daughters departing in that year.

Appendix B

Let $\mathbf{q}_j$ denote the four-dimensional vector of outcomes for household j and let $\mathbf{f}(\mathbf{x}_j,\mathbf{B},\mathbf{e}_{js})$ denote the predicted value of these outcomes from the simulations given the observables $\mathbf{x}_j$, the vector of coefficients $\mathbf{B}$, and the vector of simulation draws $\mathbf{e}_{js}$ for simulation s. The $\mathbf{f}$ function was smoothed with respect to the parameter vector using the logistic transformation. In particular, in the simulation a split was assumed to take place with probability $1/(\exp(\psi_j/\omega)$ when the simulated surplus was computed to be ψ_j. This yields consistent estimates as long as ω^{-1} grows at the same rate as the sample size.

Two simulations per household were then carried out and the criterion function

$$\sum_j (\boldsymbol{q_j} - \boldsymbol{f_j}(\boldsymbol{x_j}, \mathrm{B}, \boldsymbol{e_{j0}}))' W (\boldsymbol{q_j} - \boldsymbol{f_j}(\boldsymbol{x_j}, \mathrm{B}, \boldsymbol{e_{j1}}))$$

was constructed using a fixed weighting matrix W. A diagonal weighting matrix was used with elements 10, 10^{-2}, 10^{-6}, 10^{-6} chosen for the split probability, 1982 landholding, 1971 income, and 1982 income, respectively. These weights normalize the prediction errors from the four fitted components of the model so that they are of the same order of magnitude. Minimization of this expression with respect to the vector $\mathbf{B}$ yields dimension(B) moment conditions of the form

$$\sum_j \sum_{s=0,1} \left[\frac{\partial \boldsymbol{f}(\boldsymbol{x_j}, \mathrm{B}, \boldsymbol{e_{js}})}{\partial \mathbf{B}'} \right]' W \; (\boldsymbol{q_j} - \boldsymbol{f_j}(\boldsymbol{x_j}, \mathrm{B}, \boldsymbol{e_{j(1-s)}})) = \mathbf{0}$$

which were iteratively solved for $\mathbf{B}$. Standard errors were computed using 75 bootstraps of the sample. The bootstrap does not account for the fact that the prediction equations for fertility and marriage were estimated. To do so would be quite complicated as these equations were estimated from the same data set but quite different samples. Thus the standard errors may be downwardly biased.

References

Adelman, Irma and Sherman Robinson (1978) *Income Distribution Policies in Developing Countries: A Case Study of Korea*, Stanford: Stanford University Press.

Altonji, Joseph; Hayashi, Fumio, and Kotlikoff, Laurence (1992) "Is the Extended Family Altruistically Linked: Direct Tests Using Micro Data," *American Economic Review* 82, 1177-98.

Basu, Kaushik and James E. Foster (1996) "On Measuring Literacy", CAE Working Paper 96-06, Ithaca: Cornell University Center for Analytic Economics.

Bergstrom, Theodore (1997), "A Survey of Theories of the Family," *Handbook of Population and Family Economics*, Mark Rosenzweig and Oded Stark, eds., Amsterdam: Elsevier Science Publishers.

Benjamin, Dwayne (1995), "Can Unobserved Land Quality Explain the Inverse Productivity Relationship?" *Journal of Development Economics* 46(1), pp. 51-84.

Browning, Martin and Pierre-Andre Chiappori (1996) "Efficient Intra-household Allocations: A General Characterization and Empirical Tests," manuscript.

Chenery, et. al., (1974) *Redistribution with Growth*, London: Oxford University Press.

Chiappori, Pierre-Andre, (1988) "Rational Household Labor Supply," *Econometrica* 59(4) 925-51.

Deaton, Angus and Christina Paxson, (2000) "Growth and Saving among Individuals and Households," *Review of Economics and Statistics*, 82, 212-25.

Deininger, K. and Lynn Squire (1996) "Measuring Income Inequality: A New Data-Base", Harvard Institute for International Development, Development Discussion Paper No. 537.

Duflo, Esther and Christopher Udry (2001) "Risk and Intrahousehold and Resource Allocation in Cote d'Ivoire," Economic Growth Center, Yale University, manuscript.

Foster, Andrew and Mark Rosenzweig (1995) "Learning by Doing and Learning from Others: Human Capital and Technical Change in Agriculture," *Journal of Political Economy*, 103 (6): 1176-1209.

Foster, Andrew and Mark Rosenzweig (1996) "Technical Change and Human Capital Returns and

Investments: Evidence from the Green Revolution," *American Economic Review,* 86(4), 931-953.

Foster, Andrew and Mark Rosenzweig (2000) "Household Division and Rural Economic Growth," University of Pennsylvania, manuscript.

Foster, Andrew (1993) "Household Partition in Rural Bangladesh," *Population Studies*, 47: 97-114

Foster, Andrew (1996) "Analysis of Household Behavior when Households Choose Their Members: Marriage-Market Selection and Human Capital Allocations in Rural Bangladesh", manuscript.

Hayashi, Fumio (1993) "Is the Japanese Family Altristically Linked: A Test Based on Engel Curves," Discusion paper No. 80, New York, Graduate School of Business, Columbia University, Center Japanese Economics and Business.

Jolliffe, Dean (1997) "Whose education matters in the determination of household income: Evidence from a developing country" IFPRI Food Consumption and Nutrition Division Disc. Paper No. 39.

Lam, David (1997) "Demographic Variables and Income Inequality," *Handbook of Population and Family Economics*, Mark Rosenzweig and Oded Stark, eds., Amsterdam: Elsevier Science.

Pakes, Ariel and David Pollard (1989) "Simulation and the asymptotics of optimization estimators", *Econometrica* 57:1027-57.

Rosenzweig, Mark and T. Paul Schultz (1983) "Market Opportunities, Genetic Endowments and Intrafamily Resource Distribution: Child Survival in Rural India," *American Economic Review*, 72: 803-15.

Rosenzweig, Mark R. and Kenneth J. Wolpin (1985) "Specific Experience, Household Structure, and Intergenerational Transfers: Farm Family Land and Labor Arrangements in Developing Countries," *Quarterly Journal of Economics* 100 (Supplement), 961-987.

Schultz, T. Paul (1990) "Testing the Neoclassical Model of Family Labor Supply and Fertility," *Journal of Human Resources* 25, 599-634.

Sen, Amartya K. (1990) "More Than 100 Million Women Are Missing," *New York Review of Books* 37:2 (20 December 1990), 61-66.

Thomas, Duncan (1990) "Intra-Household Resource Allocation: An Inferential Approach," *Journal of*

Human Resources 25, 635-664.

Walker, Thomas and James Ryan (1990) *Village and Household Economies in India's Semi-arid Tropics*, Baltimore: The Johns Hopkins University Press.

Table 1
Sample Sizes by 1971 and 1982 Sample Selection Criteria and 1971 Household Status in 1982

			Number of Households by 1971 Sample Criteria			
Status of 1971 Household in 1982	Household Interviewed in 1982	Household Divided by 1982	Household Resides Outside Assam (A)	(A) + Complete Information on HH Members (B)	(A) + (B) + Head ≥40 and Farm Household (C)	(A) + (B) + (C) + One or More Resident Heirs
Head alive, intact	Original household	No	2114	1942	843	636
Head died, intact	Original household	No	601	561	425	235
Head alive, split	Only part of original household with head	Yes	413	351	151	151
Head died, split	Not interviewed	Yes	1235	1116	365	365
Totals			4363	3970	1784	1387

Table 2
Characteristics of Landowning Households in 1971 with Head Aged 40 and Above, by Potential for Division and Division Status in 1982

Sample	Households with at least one heir in 1971		Households with no heirs in 1971
Status in 1982	Divided, with or without Head Death	Intact, with or without Head Death	Intact or Head Death
Percent head died by 1982	70.7	26.9	47.9
Mean age of head	55.5 (10.1)[a]	53.0 (9.44)	54.0 (16.0)
Mean landholdings in 1971 (acres)	12.4 (14.5)	10.6 (10.6)	8.14 (9.56)
Mean landholdings in 1982, if in 1982 resurvey[b]	9.66 (13.2)	9.16 (10.0)	7.67 (10.5)
Mean agricultural income (Rupees)	4829 (4588)	4172 (4087)	2117 (2163)
Mean per-hectare profit variance (Rupees squared)	54308 (5167)	54019 (4712)	52203 (5456)
Mean number of departed married daughters of the household head	.655 (.998)	.560 (.913)	.338 (.760)
Mean number of departed sons of the household head	.333 (.777)	.261 (.648)	.403 (.912)
Mean number of heirs in household	2.19 (1.18)	1.95 (1.17)	0
Mean age of heirs	35.0 (9.61)	33.6 (9.03)	-
Mean schooling of heirs (years)	4.68 (4.01)	5.02 (3.89)	-
Percent of heirs ever married	55.8	44.5	-
Percent of households with more than one heir	66.3	54.3	-
Percent of multi-heir households with unequal heir schooling	60.7	38.2	-
Number of households	516	871	397

[a]Standard deviation in parentheses.
[b]Includes 1982 intact households and divided households with the original 1971 household head.

Table 3
Maximum-Likelihood Probit Estimates: Determinants of the Probability of Household Division Between 1971 and 1982 for Households with at Least One Heir in 1971

	I	II	III	IV	V
Total claimants	.131	.200	.199	.168	.228
	(3.53)[a]	(4.61)	(3.24)	(2.66)	(5.07)
Number of wives of claimants	.367	.477	.506	.619	-2.90
	(2.64)	(3.34)	(2.87)	(3.46)	(2.00)
Head's age	-.0541	-.0538	-.0496	-.0463	-.0478
	(1.61)	(1.61)	(1.47)	(1.37)	(1.39)
Head's age squared ($x10^{-3}$)	.520	.504	.519	.539	.433
	(1.77)	(1.72)	(1.71)	(1.77)	(1.45)
Mean of claimants' schooling	.00874	.224	.222	-.592	-.427
	(0.88)	(3.90)	(3.85)	(1.78)	(1.27)
Variance of claimants' schooling	-	.0437	.0419	-.214	-.174
		(3.31)	(3.13)	(2.40)	(1.96)
Maximum of claimants' schooling	-	-.212	-.206	.436	.283
		(3.81)	(3.68)	(1.34)	(0.87)
Number of male children<15	.133	.175	.208	.210	.197
	(1.55)	(2.01)	(2.36)	(2.37)	(2.22)
Number of female children<15	.0263	.0356	.0311	.0339	.0223
	(0.35)	(0.47)	(0.40)	(0.43)	(0.28)
Land owned ($x10^{-2}$)	.195	.251	.428	.366	.00226
	(0.47)	(0.61)	(0.94)	(0.83)	(0.51)
Income growth, 1971-82 ($x10^{3}$)	-	-	-.378	-1.72	-1.79
			(1.70)	(3.79)	(3.89)
Interactions--Income growth[b]					
x Mean of claimants' schooling	-	-	-	.953	.768
				(2.48)	(1.99)
x Variance of claimants' schooling	-	-	-	.308	.259
				(2.97)	(2.50)
x Maximum of claimants' schooling	-	-	-	-.744	-.572
				(1.99)	(1.52)
Variance of profit shock ($x10^{-4}$)	-	-	-.151	-.306	-.828
			(3.52)	(3.44)	(4.11)
Interactions--Variance of shock ($x10^{-4}$)					
x Number of departed married daughters.	-	-	-	-	.186
					(2.01)
x Number of claimant wives	-	-	-	-	.658
					(2.37)
Number of departed married daughters	-	-	-	-	-1.01
					(2.06)
Number of departed sons	-	-	-	-	.138
					(2.26)
Constant	.181	-.0370	1.68	2.95	5.45
	(0.20)	(0.04)	(1.61)	(2.71)	(3.84)
χ^2 (df)	31.3 (8)	46.1 (10)	61.8 (14)	91.5 (15)	101.7 (19)

[a]Absolute values of asymptotic t-ratios in parentheses.

Table 4
Parameter Estimates from the Structural Model

Parameter	Coefficient	Standard Error
α	1.22	.0417
β:		
1971 Head	-1.30	.0755
Number of boys	2.45	.142
Number of girls	2.03	.0727
Married	-.223	.114
Constant	-1.42	.0864
δ	1.85	.0284
γ	1.51	.0350
λ	.430	.0380
Per-acre productivity:		
η	721.8	8.10
η_{ϕ}	-.341	.0192
$\eta_{\phi t}$	-.0274	.00292
η_{S}	25.0	1.73
η_{St}	-10.4	.174
$\eta_{\phi St}$	.0125	.000184
Scale Effects		
η_{A}	-.0377	.00281
$\eta_{\phi At}$	.00240	.000440
Net Income Variance Effects:		
ξ_3 (A^N)	-.0198	.135
ξ_2 (mstat)	-.203	.132
ξ_1 (D_i)	3.99	.343
ξ_0	0.913	.0860
ξ_4 (joint residence)	0.594	.0608

Table 5
Predicted Effects of Increasing Agricultural Technical Change
on Household Division and Household Public Good Consumption,
by Per-Capita Land Quartile Group, 1971-82

Quartile:	1	2	3	4
Household division				
Actual	.0129	.0163	-.1145	-.2190
No intra-household schooling inequality	-.0019	.0317	-.0744	-.0852
Public good consumption	.0909	.183	.229	.241

Table A

Fixed-Effects IV Estimates of Per-Hectare Profit Function, 1969-70 - 1970-71

Variable	Mean (S.D.)	FE-IV Estimates[a]	
Land assets (pumps, fences, pipes, canals) per hectare (Rupees)	403.8 (930)	.0319 (2.05)[b]	.0333 (2.06)
Farm equipment per hectare (Rupees)	113.2 (200)	.312 (1.80)	.336 (1.87)
Share of area devoted to HYV	.0303 (.106)	2840 (2.14)	-23.1 (0.99)
HYV share*whether maximum schooling in household is primary or above	-	-	8542 (2.72)
Number of farmers	2491	2491	2491

[a] Instruments include: 1970 assets, gross-cropped area, use of HYV seeds, village adverse weather in 1970-71, and investments in assets between 1968-69 and 1969-70. Mean profits/hectare=955.2.

[b] Absolute value of asymptotic t-ratio in parentheses.

Figure 1
Leveraged Residual Plot: Relationship Between Actual Household Fuel Expenditures in 1982
and Predicted Household Public Good Expenditure

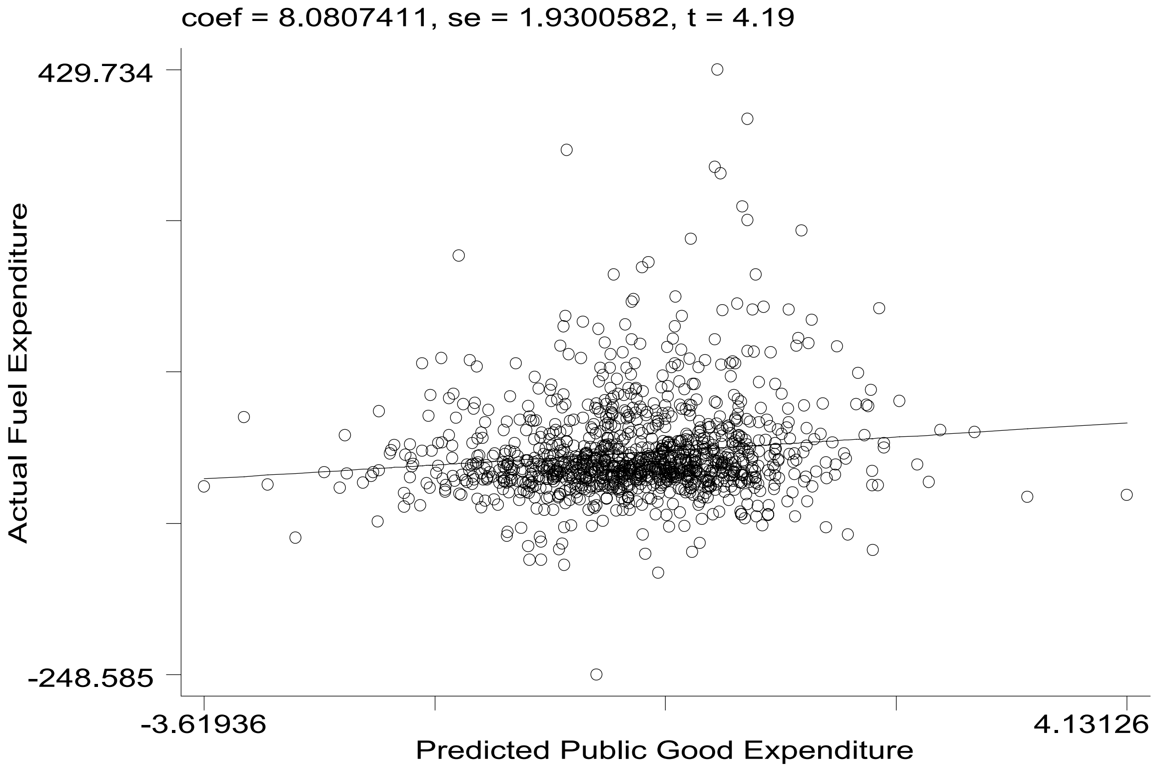

Figure 2
Predicted Effects of Increasing Agricultural Technical Change on Household Income Growth, by Per-Capita Land Quartile Group, 1971-82