Consumption Smoothing and the Structure of Risk and Time Preferences: Theory and Evidence from Village India [¤]

Takashi KUROSAKI [y]

August 2001

Abstract

This paper investigates the extent to which rural households in developing countries are able to smooth consumption, using a theoretical model of full risk sharing, in which participating households have di®erent risk and time preferences. A resulting rule of resource allocation is characterized in an intuitive way, clarifying the e®ects of diverse preferences. Empirical models are applied to a household panel data collected from rural India. Estimation results strongly support the heterogeneity in risk preferences. In contrast, little evidence is found in favor of the intertemporal resource allocation across households according to di®erences in time preferences.

Keywords: insurance, consumption smoothing, risk attitudes, discount rates.

JEL classi¯cation codes: O12, D12, Q12.

[¤] I would like to thank the ICRISAT for allowing the use of its household dataset and John Pender of the IFPRI for helping constructing empirical variables. Comments from seminar participants at Stanford University, the ICRISAT, and the 1999 autumn meeting of the Japanese Economic Association were very helpful, especially from Marcel Fafchamps, Kei Otsuka, Futoshi K. Yamauchi, Yasu Sawada, Shahe Emran, and Satheesh Aradhyula.

[y] Institute of Economic Research, Hitotsubashi University, 2-1 Naka, Kunitachi, Tokyo 186-8603 JAPAN. Phone: 81-42-580-8363; Fax: 81-42-580-8333; E-mail: kurosaki@ier.hit-u.ac.jp.

1 Introduction

This paper investigates implications of full risk sharing among low income households for the case with households having heterogeneous preferences. Following Townsend (1994), the extent of consumption smoothing among villagers in developing countries has been investigated for various regions and with various methods in the recent literature [Townsend (1995); Ravallion and Chaudhuri (1997); Udry (1994); Ligon (1998); Ligon et al. (1999); Kochar (1999)]. Although the underlying theoretical model and empirical models based on it mirror similar work for developed countries [Mace (1991); Cochrane (1991); Hayashi (1997); Crucini (1999)], testing full insurance implications as a benchmark is especially important for low income countries because risk is expected to a®ect people's welfare more in an economy where farming is the main activity and markets are underdeveloped over space.

A relatively unexplored issue in that investigation is the e®ect of heterogeneity in risk and time preferences among villagers on risk sharing arrangements. Most of the studies for developing countries mentioned above implicitly or explicitly assume homogeneous preference in their empirical tests. This is unsatisfactory considering the accumulation of theoretical work on rural institutions to cope with risk, where di®erence in risk attitudes plays an important role in allocating risk [Stiglitz (1988); Hayami and Otsuka (1993)]. Furthermore, considering the prevailing poverty and the paucity in risk mitigating arrangements in rural economies in developing countries, incorporating heterogeneous risk preferences is especially relevant from development perspective.

A related issue in economic development is discount rates. If the future is heavily discounted, households may behave in a myopic way, resulting in lower savings, lower investment, and less sustainable long-term cooperation. Following the usage by Pender (1996), who implemented a rare empirical study on discount rates in developing countries, \discount rate" in this paper refers to \a measure of the intertemporal rate of substitution, which may be a®ected by either diminishing marginal utility of consumption or pure time preference" (p.259). The e®ects of diversity in the latter is a key issue addressed in this paper | pure time preference might di®er among households, according to their differences in demographic structure, education level, etc. The diminishing marginal utility of consumption is closely related with the curvature of utility function, i.e., risk preference, which is another key element investigated in this paper, although we do not estimate the intertemporal rate of substitution directly.

In the following, the basic model of full-information intra-village risk sharing is extended to a case where participating households may have different risk and time preferences. Among the existing studies, Townsend (1994) partially examines the effects of heterogeneous risk attitudes in its empirical part, without considering the possibility of heterogeneous time preferences; Cochrane (1991) gives brief discussion on heterogeneous preferences with respect to both risk and time in its theoretical part, without deriving its full implications to empirical work. A distinctive feature of this paper is that, first, a rule of risk allocation is characterized explicitly when preferences are heterogeneous and its empirical implications are explored for testing full insurance and the structure of risk/time preferences. Another feature of this paper is application of the empirical model to a popular data set on this subject, i.e., the ICRISAT (International Crops Research Institute for the Semi-Arid Tropics) household panel data from rural India. This application not only generalizes Townsend's (1994) tests but also is expected to shed light on the relationship between households' time preference and their actual economic behavior, for which there are few empirical studies.[1]

The paper is organized as follows. In Section 2, a theoretical model is introduced to investigate the effect of heterogeneous preferences and its empirical implications are explored to derive testable hypotheses regarding the structure of risk/time preferences. Econometric results based on the ICRISAT household data are reported in Section 3. The final section concludes the paper, comparing our results with those from the recent literature.

[1] Some of the existing studies infer risk and time preferences from experiments [Binswanger (1981); Pender (1996)], some estimate risk preferences from observed economic behavior in developing countries [Kurosaki (1998) and Kurosaki and Fafchamps (forthcoming)], and others estimate risk and time preferences using observation from developed countries [Lense (2000)], but very few have investigated time preference for developing countries based on observed economic choices.

2 Theoretical Model and Empirical Specifications

2.1 Theoretical Model of Full Risk Sharing

Basic settings of the theoretical model in this paper follow a model of full-information intra-village risk sharing, adopted in Mace (1991), Cochrane (1991), and Townsend (1994). We consider a rural economy of N infinitely-lived households. Household i is faced with uncertainty denoted by the state of nature s in period t that occurs with probability π_{st}. The household is endowed with stochastic income y_{ist} and consumes c_{ist} from which it obtains von Neumann-Morgenstern utility denoted by $u_i(c_{ist})$ with $u_i' > 0; u_i'' < 0$. With an assumption of separability between consumption and leisure, the Pareto optimal resource allocation is obtained by solving the social planner's problem:

$$\max_{\{c_{ist}\}} \sum_{i=1} \lambda_i \sum_{t=1} \rho_i^t \sum_{s} \pi_{st} u_i(c_{ist}) \tag{1}$$

subject to a feasibility constraint

$$\sum_{i=1} c_{ist} \leq \sum_{i=1} y_{ist}; \qquad \forall (s;t); \tag{2}$$

and a set of non-negativity constraints for c_{ist}, where λ_i is a Pareto-Negishi weight for household i and ρ_i is a subjective discount factor of household i corresponding to the pure time preference of each household.[2]

Assuming an interior solution, the Pareto optimal allocation requires that:

$$\lambda_i \rho_i^t u_i'(c_{ist}) = \mu_{st}; \qquad \forall i; \tag{3}$$

where μ_{st} is the Lagrange multiplier associated with the feasibility constraint (2) in period t with state s divided by its probability π_{st}. Equation (3) simply states that λ-weighted marginal utility is equalized among villagers. Its important implication is that idiosyncratic income shocks should not affect individual consumption. What matters is the

[2] In this specification, it is implicitly assumed that the social planner maximizes the sum, over households, of each household's intertemporal utility that is individually evaluated using each household's subjective discount factor. It is not assumed that the planner maximizes the sum, over periods using his own subjective discount factor, of each period's utility sum in the village. The former assumption is adopted because it allows a consistent mapping between the social planner's solution and a competitive equilibrium solution under complete markets, even when households have heterogeneous time preferences, while the latter does not.

aggregate income shock in t that is completely represented by μ_{st}. This implication constitutes the basis of the existing empirical studies. Although the condition (3) is derived from the social planner's optimization problem, the same rule of resource allocation can be derived as a result of competitive equilibrium within a decentralized framework as long as there is no private information and markets for state-contingent claims are complete.

In empirical tests, most of the existing studies assume in addition homogeneous preferences among households with respect to risk and time. This additional assumption results in an empirically testable hypothesis that the level of consumption change should be the same among villagers and it should not be affected by shocks idiosyncratic to individual income levels, when utility function u exhibits constant absolute risk aversion (CARA). When u exhibits constant relative risk aversion (CRRA), the log consumption growth should be the same among villagers and it should not be affected by the idiosyncratic shocks.

In this paper, the assumption of homogeneous risk/time preferences is relaxed. For simplicity, subscript s is dropped below, since the focus is on ex post, observable allocation of consumption.

First, consider a case where households have CARA preferences, i.e.,

$$u_i(c_i) = -\frac{1}{A_i}\exp[-A_i c_i]; \tag{4}$$

where A_i is an Arrow-Pratt coefficient of absolute risk aversion. An explicit solution to equations (3) and (2) is obtained as

$$c_{it} = -\frac{1}{A_i}\ln\mu_t + \frac{1}{A_i}\ln\lambda_i + \frac{1}{A_i}t\ln\rho_i = \alpha_i\bar{c}_t + \beta_i + \gamma_i t; \tag{5}$$

where

$$\alpha_i \equiv \frac{1}{A_i}\left[\frac{1}{N}\sum_j \frac{1}{A_j}\right]^{-1}; \tag{6}$$

$$\beta_i \equiv \frac{1}{A_i}\left[\ln\lambda_i - \frac{1}{N}\sum_j \alpha_j \ln\lambda_j\right]; \tag{7}$$

$$\gamma_i \equiv \frac{1}{A_i}\left[\ln\rho_i - \frac{1}{N}\sum_j \alpha_j \ln\rho_j\right]; \tag{8}$$

and $\bar{c}_t$ is the village mean of consumption levels. Equation (5) intuitively shows that the optimal consumption consists of a variable part proportional to the village mean consumption at the rate of α_i and a fixed part $\beta_i + \gamma_i t$.

As a special case of $\rho_i = \rho$ for all i, γ_i term disappears, resulting in an expression analogous to the standard notation in the sharecropping literature. Definition (6) implies that when a household is more risk averse than the village average in the sense that $\frac{1}{A_i} \leq \frac{1}{N}\sum_j \frac{1}{A_j}$, α_i becomes smaller than unity, i.e., the household's share in variable consumption is smaller than the village average. This implication is similar to the one derived for sharecropping arrangements. For example, when there is only one tenant and one landlord and when enforcement of labor or effort is perfect (i.e., without moral hazard), the tenant's crop share rate is larger (smaller) than the landlord's share when the tenant is more (less) risk averse than the landlord [Stiglitz (1974, p.231); Hayami and Otsuka (1993, p.47)].

Definition (7) implies that the village economy allocates consumption to households according to the size of λ_i. Although the weights can take any positive values under the social planner's optimization framework, there exists a mapping from the consumption allocation under a full-information competitive equilibrium to the consumption allocation under the social planner's problem with a specific vector of λ. Under such competitive equilibrium, wealthier households who can contribute more to the village income on average are likely to be assigned higher λ_i and hence higher consumption.

Regarding the effects of diversity in time preferences, when a household is more myopic, i.e., ρ_i is smaller, γ_i becomes more negative. The fixed consumption of such a household should be decreasing over time. This allocation is efficient since more myopic households evaluate consumption in the immediate period more highly than less myopic households do.

An intertemporal change of consumption associated with equation (5) is characterized as

$$c_{i,t+1} - c_{it} = -\frac{1}{A_i}[\ln \mu_{t+1} - \ln \mu_t - \ln \rho_i] = \alpha_i(\bar{c}_{t+1} - \bar{c}_t) + \gamma_i. \qquad (9)$$

An important implication of this expression to empirical works is that, even when the first difference of consumption is used as the dependent variable to test the full risk sharing hypothesis, it should vary among households in a systematic way. In other words, household specific effects remain as a slope effect on the village average consumption as well as an intercept effect.

When households have CRRA preferences, i.e.,

$$u_i(c_i) = \frac{1}{1 - R_i} c_i^{1 - R_i}; \tag{10}$$

where R_i is an Arrow-Pratt coefficient of relative risk aversion, similar results can be obtained. An optimal consumption is defined as

$$\ln c_{it} = -\frac{1}{R_i} \ln \mu_t + \frac{1}{R_i} \ln \lambda_i + \frac{1}{R_i} t \ln \rho_i = \alpha_i' \overline{\ln c}_t + \beta_i' + \gamma_i' t; \tag{11}$$

where α_i', β_i', and γ_i' are the same as in definitions (6), (7) and (8) except that A_i is replaced by R_i, and $\overline{\ln c}_t$ is the village mean of log consumption. As before, the log of the optimal consumption consists of a variable part proportional to the village mean and a fixed part. An intertemporal change associated with equation (11) is characterized as

$$\ln c_{i,t+1} - \ln c_{it} = -\frac{1}{R_i}[\ln \mu_{t+1} - \ln \mu_t - \ln \rho_i] = \alpha_i' (\overline{\ln c}_{t+1} - \overline{\ln c}_t) + \gamma_i'; \tag{12}$$

which implies that the log consumption growth should vary among villagers. This CRRA case is of special interest because definition (10) together with (1) implies that the intertemporal elasticity of substitution, which is one of the factors determining discount rates, is constant for each household but varies across households taking the value $1/R_i$.

2.2 Empirical Model and Testable Hypotheses

Based on the theoretical model above, an empirical model is proposed to examine the sensitivity of consumption changes (or log consumption growth) with respect to aggregate and idiosyncratic shocks. A straightforward way of implementing this examination based on equation (9) is to estimate

$$\Delta c_{it} = b_i + a_i \Delta \bar{c}_t + \zeta_i X_{it} + u_{it}; \qquad i = 1, \ldots, N; \quad t = 1, \ldots, T; \tag{13}$$

where $\Delta c_t = c_t - c_{t-1}$, b_i, a_i, and ζ_i are parameters to be estimated, X_{it} denotes idiosyncratic income shocks to household i, and u_{it} is an error term with zero mean. Parameter

η is allowed to vary among households since functioning of risk sharing arrangements may differ from household to household. The consumption variable c_t should be replaced by $\ln c_t$ for a CRRA specification based on equation (12).

An important empirical implication from the previous subsection is that, even when a first difference is used as the dependent variable, household specific effects remain. Parameters a_i and b_i correspond to these effects due to heterogeneity in risk and time preferences respectively. In addition, when households' preferences and the economy's welfare weights change over time, for instance, due to changes in demographic composition [Townsend (1994); Cochrane (1991)], the parameters should reflect these changes also. However, a crucial point is that even when these changes are absent or controlled in different ways,[3] the heterogeneity in time-invarying preferences with respect to risk or with respect to time necessitates the use of panel methods.

Equation (13) can be estimated by a time series regression for each household when the time horizon of panel data is sufficiently long. Unlike Townsend (1994), who applies this approach without γ_i term in equation (5), our model explicitly includes a term b_i to allow heterogeneity in time preferences.

From this estimation, we can expect to obtain insightful inference on the structure of time and risk preferences among sample households and the nature of consumption smoothing. If parameter b_i is positive (negative), such a household has time preference with a higher (lower) discount factor ρ_i than the village average. By testing whether $b_i = b$ for all i, we can investigate whether households have the same time preference.

Similarly, if parameter a_i is greater (smaller) than one, such a household bears more (less) of the common shock than the village average. By testing whether $a_i = a$ for all i, we can investigate the hypothesis of homogeneous risk preferences. If it is rejected, we will proceed to the identification of those with higher risk attitudes.

Finally, if the null hypothesis of $\eta_i = 0$ is accepted for all i, the village economy achieves efficient risk allocation with respect to idiosyncratic shocks. If not, the magnitude of parameter η_i will tell us how sensitive a household's consumption is to unpredicted,

[3] For example, Townsend (1994) explicitly derives an expression for changes in age-sex composition and adds its term to his empirical model.

idiosyncratic events. A difference in magnitudes of γ_i would show which households are more vulnerable.

In the next stage, we investigate whether parameters b_i, a_i, and γ_i are related with households' social positions in a systematic way. By "social positions," we mean inherent characteristics of households that determine preferences toward consumption. However, one fundamental question is that most of the proxies for the social positions, such as wealth, education, and demographic structure are endogenous to household decisions in the long run. We partially reduce this problem by taking the initial values of these variables. Furthermore, we include in the empirical model a variable for caste ranking, which could be safely treated as exogenous. Nevertheless, the problem of endogeneity should be taken care of in interpreting the empirical results in the next section.

To undertake this investigation, we adopt two approaches. First, we estimate equation (13) for each household to obtain a set of parameter estimates $(\hat{b}_i, \hat{a}_i, \hat{\gamma}_i)$. Then we estimate correlation coefficients between them and households' initial characteristics Z_i. By testing the statistical significance of the coefficients, we can infer the structure of risk and time preferences.

If we find particular household characteristics to be related with the estimates $(\hat{b}_i, \hat{a}_i, \hat{\gamma}_i)$ in the first approach, we may be able to replace household dummies in (13) by a function of those characteristics. This is the second approach in which we estimate an empirical model

$$\Delta c_{it} = (b_0 + Z_i b_1) + (a_0 + Z_i a_1)\Delta \bar{c}_t + (\gamma_0 + Z_i \gamma_1) X_{it} + u_{it}; \qquad i = 1, \ldots, N; \quad t = 1, \ldots, T. \tag{14}$$

Again, we can examine how risk and time preferences vary among households by testing the statistical significance of b_1, a_1, and γ_1. Specification (14) has a much higher degree of freedom than equation (13), a great advantage considering the short time horizon of household panel data available from developing countries.

3 Application to ICRISAT Households in India

3.1 Data

In this section, the empirical model above is applied to the ICRISAT household data from rural India. Characteristics of study villages and sample households are fully described by Walker and Ryan (1990). The data set used in this paper is composed of household information spanning the ten-year period from 1975 to 1984, collected from three villages in two states of Andhra Pradesh and Maharashtra. All of the three villages belong to the semi-arid regions of Peninsular India.[4] Forty households (ten each from farming categories of landless, small farms, medium farms, and large farms) were surveyed in each village, each year. Due to attrition and household division, the complete panel of ten years is composed of 104 households from the three villages.

This data set has been used extensively in investigation of consumption smoothing mechanisms [Townsend (1994); Ravallion and Chaudhuri (1997); Ligon (1998); Ligon et al. (1999); Kochar (1999); Jacoby and Skou¯as (1998); etc.]. Empirical results from these studies show that consumption of the sample households is insulated from °uctuations in individual income much better than initially expected but the hypothesis of e±cient risk sharing is rejected in many cases. This paper re-investigates this issue with an extended model that allows heterogeneous preferences with respect to risk and time.

De¯nition and statistics of empirical variables are shown in Table 1. The consumption variable c_{it} in equation (13) is de¯ned as the total household consumption expenditure[5] in real Indian Rupees (1983 Rs.), divided by the total adult equivalent units of household members.[6] It is called \per-capita consumption" for short below and used in estimating level (¯rst di®erence) regressions derived from a CARA speci¯cation. Its natural log (¯rst di®erence) is used in log regressions derived from a CRRA speci¯cation.

---

[4] Due to space limit, this paper presents results pooling the three villages. See Kurosaki (1999) for results for individual villages.

[5] The total consumption is de¯ned in a way similar to Townsend (1994), based on \observed transactions." Ravallion and Chaudhuri (1997) criticize this measure since its measurement errors are likely to be correlated with those of income measures, suggesting an alternative measure based on \°ow accounting." We leave for further study the sensitivity of our results to this alternative measure of consumption.

[6] Adult equivalent units used in this section are: 1.0 for adult male, 0.9 for adult female, and 0.52 for children up to 12 years old.

In estimating equation (13), the right hand side variable of village-wide average of consumption change (or village-wide average of log consumption change) is approximated by the average of all sample households except for the specific household under scrutiny (i.e., the average of thirty-nine neighbor households), to minimize the possibility of spurious correlation. Since an measurement error is introduced by this approximation due to the fact that the sample does not cover all the villagers, the estimation results below are valid only under the maintained hypothesis that the measurement error is small. X_{it} in equation (13), whose coefficient γ_i represents excess sensitivity, is defined as $X_{it} \equiv y_{it} - y_{i,t-1}$ for the CARA case and $X_{it} \equiv \ln y_{it} - \ln y_{i,t-1}$ for the CRRA case, where y_{it} denotes per-capita household income in real Rs.

The maximum estimation period is ten years from 1975 to 1984. The quality of some data for the first year and the last three years may not be as high as that for other years.[7] However, if the six-year panel from 1976 to 1981 is used, the degree of freedom becomes too low when β_i, α_i, and γ_i are specific to each household. Therefore, the longest panel available is used in estimating equation (13). Then the sensitivity of our results to the choice of sample period is examined through estimating equation (14), in which the problem of the degree of freedom is less acute.

3.2 Estimation Results

First, we estimate equation (13) for each household as a time series regression by the OLS. Summary results for β_i, α_i, and γ_i are shown in Table 2. Most of β_i are insignificant, a result inconsistent with intertemporal redistribution according to differences in time preference. Both the CARA specification (level-change regressions) and the CRRA specification (log-change regressions) reject the null hypothesis that $\beta_i = 0$ at 5% level only in two percent of the sample.

In contrast, the null hypothesis that $\alpha_i = 1$ and the null that $\gamma_i = 0$ are rejected more frequently. The former is rejected at 5% level in 16% (CARA specification) and in 11%

[7] Production input data were not collected as frequently in 1984 as in previous years, while consumption data were not collected as in detail in 1975 and 1982-84 as in other years [Walker and Ryan (1990, p.67)]. In this paper, consumption data for these years are adjusted proportionally using the village average ratio of non-covered items in the period 1976-81.

(CRRA) of the total households. The null hypothesis that $\gamma_i = 0$ is rejected at 5% level in 18% of the total households in both specifications.

To support the findings above statistically, the joint significance of heterogeneous b_i, a_i, and γ_i is tested, using panel estimation results. The fixed or random effects estimation corresponds to a restriction that $a_i = a$ and $\gamma_i = \gamma$ for all i. F tests for the joint significance of these restrictions show that the homogeneous assumption regarding a_i and γ_i is rejected in several cases [Kurosaki (1999, Appendix Table 1)]. Therefore, some evidence is found for the heterogeneity among households in their sensitivity to common and idiosyncratic shocks.

Second, we investigate whether parameter estimates for b_i, a_i, and γ_i from equation (13) are structurally related with households' social positions. Five variables that represent households' initial characteristics are used for Z_i: a dummy variable for ownership of agricultural land in 1975 (LANDD), its value per capita (LANDPC), education status of the household head (SCHOOL), demographic characteristics approximated by the share of children in household size (CHILDR), and caste rank (JGRRANK) (see Table 1). The marginal effect of land for owners is represented by LANDPC and its threshold effect for a landless to become a landed household is represented by LANDD. As discussed in the previous subsection, all these variables except JGRRANK are endogenous to household decisions in the long run. Therefore, we cannot interpret the relation as the one showing any causality.

Another practical issue is that the land variables (LANDD and LANDPC), SCHOOL, and CASTE are highly correlated.[8] In rural India, land ownership, education, and high caste ranking are a typical signal for a high social position. On the other hand, the demographic variable could represent other aspects that directly affect households' preferences. Therefore, although LANDD, LANDPC, SCHOOL, and CASTE may capture different aspects of household characteristics, we do not attempt multiple regressions but instead report bivariate correlation coefficients between each of the estimates $\hat{b}_i$, $\hat{a}_i$, and $\hat{\gamma}_i$, and one of the shifters in Z_i.

[8] From our data set, correlation coefficients between LANDPC and other four variables are: LANDD= 0.418 ***, CHILDR=0.0531, JGRRANK=-0.503 ***, and SCHOOL=0.588 *** (all three villages pooled), where *** shows that the coefficient is statistically significant at 1%.

Results in Table 3 show that $\hat{a}_i$ is significantly correlated with land ownership per capita (LANDPC). This is consistent with the risk sharing interpretation that more landed households tend to bear more of the common risk. The relation is almost nil for the time preference parameter $\hat{b}_i$. As is found by Townsend (1994), LANDPC is related with $\hat{\gamma}_i$ negatively but they are not significant when the three villages are pooled. Lower caste households (higher JGRRANK) tend to respond less to common risk (lower $\hat{a}_i$) but more to idiosyncratic risk (higher $\hat{\gamma}_i$), although their effects are statistically significant only when each village is investigated separately (Kurosaki [1999]). Demographic character does not seem to be related with these parameters in Table 3.

Third, we estimate equation (14) with structural shifters to investigate whether the relationship between these parameters and social status variables becomes significant with more degrees of freedom. This is because the significance levels of correlation coefficients are not high in Table 3 and show a wide difference across villages [Kurosaki (1999)].

Table 4 reports estimation results when the three villages are pooled. It is shown that, first, none of these shifters are significant in affecting b_i. Since we find only a weak and patchy evidence of demographic variables' relation with $\hat{b}_i$ in Table 3 and in Kurosaki (1999), we conclude that no strong evidence is found in support of the intertemporal redistribution according to differences in time preferences within each village. Second, land variables significantly increases a_i and significantly decreases γ_i in many cases. Therefore, Townsend's (1994) claim that landless households are more risk averse and more vulnerable to income shocks has been verified in our results also. He, however, shows this result by estimating sub-samples of landless and landed class separately, without formally testing the statistical significance of the difference. Our results are based on formal tests and show further that, among the landed class, more landholding (i.e., higher LANDPC) implies less risk aversion and more insulation from income shocks. Third, village-by-village estimation results reported in Kurosaki (1999) show a contrast among the three villages, regarding the non-significant variables among the four correlated shifters of LANDD, LANDPC, SCHOOL, and JGRRANK. The village difference is consistent with the contrast among the villages with respect to social and economic infrastructure reported in Walker and Ryan (1990). Although education, land variables, and caste are correlated, their effects are not the same at the village level.

Finally, we investigate the robustness of the results above in two ways. First, borrowing the idea of Ravallion and Chaudhuri (1997), we re-estimate equation (13) with $\Delta \bar{e}_t$ replaced by village-time dummies, to assess the robustness of our results to the specific choice of common shock measures. Detailed results are given in Kurosaki (1999), which demonstrates that the significance level is enhanced but qualitative results remain the same.

Second, equation (14) is re-estimated for a shorter period of 1976-81, for which data are the most reliable. Detailed results are given in Kurosaki (1999), which shows that the overall pattern is similar but that the relationship between land related variables and parameters a_i and γ_i is more significant, reinforcing the previous results.

4 Summary and Discussion

In this paper, a model of full-information intra-village risk sharing is extended to the case where participating households have different risk and time preferences. The resulting rule of consumption allocation is characterized through the decomposition of individual consumption into fixed and variable parts. The degree of bearing common risk should decrease with households' risk aversion relative to other households in the village economy, a result analogous to what the sharecropping literature predicts. Those households with stronger preferences for immediate consumption should be allocated a higher consumption in earlier periods as a fixed part. An empirical implication of the allocation rule is that, even when first difference variables are used in testing the full insurance hypothesis, household specific effects remain as intercept dummies under the assumption of heterogeneous time preferences and as household slope dummies under the assumption of heterogeneous risk preferences.

As an illustrative application, the empirical models proposed in this paper are applied to the ICRISAT household panel data from rural India. Since the empirical model of this paper generalizes Townsend's (1994) framework, it is not surprising that our major findings regarding the extent of households' vulnerability to idiosyncratic shock and who are more vulnerable are similar to his ¯ndings. This paper further shows that allowing heterogeneity improves the explanatory power of the model in a statistically signi¯cant way. Regression results with structural shifters show that land or caste characteristics are signi¯cantly related with consumption smoothing parameters. Especially, estimation results strongly support that risk preferences are heterogeneous and their distribution depends on households' social positions in the village.

In contrast, little evidence is found in favor of the hypothesis that consumption is reallocated among households intertemporally according to di®erences in time preferences. This ¯nding seems to contradict experimental results from Pender (1996) with signi¯cant heterogeneity in discount rates among households from a semi-arid Indian village. Since what he elicited is discount rates, which are a®ected both by pure time preference and by the curvature of utility functions, heterogeneity in discount rates and homogeneity in pure time preference could co-exist in theory. Furthermore, Pender's (1996) experiment was implemented in 1989, implying a time lag from the period covered in this paper, which could be a reason for the di®erence.

However, as Pender (1996) vividly demonstrates, credit markets in the study area are highly incomplete. Given such an environment, our results might suggest that the existing risk sharing mechanisms are not e®ective in smoothing consumption intertemporally over the long run, even they are able to achieve some inter-state consumption smoothing over the short run, re°ected in the lower value of the excess sensitivity parameter of individual consumption to idiosyncratic income shocks. Since the intertemporal resource allocation according to di®erences in time preferences is very long-run by nature, its enforcement might face more di±culty. It is possible that, because of this shortcoming of risk sharing mechanisms, heterogeneity in time preferences are not re°ected in the observed allocation of consumption. If this is the case, more myopic households with an investment opportunity that yields higher return after a long gestation period such as education [Jacoby and Skou¯as (1997)] cannot utilize the opportunity because they want to consume more today.

Strictly speaking, the parameters estimated in the empirical part of this paper can be regarded as known functions of preference parameters only under the maintained hypothesis of full insurance, which is not supported by the ICRISAT data. This paper does not o®er theoretical explanations regarding the sensitivity to idiosyncratic shocks, either. Therefore, our major task in the future is to construct a theoretical model that is consistent with the regression results in this paper and allows identi¯cation of preference parameters under alternative hypotheses with incomplete insurance. Regarding discount rates, Pender (1996) ¯nds that they are not only diverse but also extremely low compared with developed country data. This issue could not be addressed by the present paper as the average discount factor is not identi¯ed. An alternative model would be necessary to identify it. These theoretical extensions are left for further research.

Among theoretical extensions of Townsend's (1994) full insurance model, recent literature of risk sharing with limited information [Ligon (1998)] and limited commitment [Ligon et al. (1999)] seems to be promising. Both of these studies apply their theoretical models to the ICRISAT data to show that their models explain the data better than Townsend's (1994) model. The limited commitment theory by Ligon et al. (1999), however, could predict that consumption response to idiosyncratic component is higher for large farmers than small farmers,[9] which is the opposite of what is found in this paper. To theoretically justify the ¯ndings of this paper that consumption of households with larger assets responds less to idiosyncratic shocks and responds more to aggregate shock, liquidity constraints should be explicitly incorporated, such as the one modeled by Deaton (1991).

It is left for future research also to investigate the robustness of our empirical results and to relate them to detailed, actual functioning of rural credit and insurance institutions. This paper shows that, by allowing heterogeneous consumption smoothing parameters and by combining estimation results with information on household characteristics, rich insight can be obtained. It is worthwhile to apply the extended model in this paper to recent panel data sets from developing countries, some of which are with longer time horizon and well controlled quality [Udry (1997); Grosh and Glewwe (1998)].

[9] This is because, in their framework, if a household is hit by an extremely positive income shock, it should be provided a reasonably large consumption in that time to avoid reneging the contract.

References

[1] Binswanger, H. P. (1981), \Attitudes toward Risk: Theoretical Implications of an Experiment in Rural India," Economic Journal 91, pp.867-890.

[2] Cochrane, J. H. (1991), \A Simple Test of Consumption Insurance," Journal of Political Economy 99, pp.957-976.

[3] Crucini, M. J. (1999), \On International and National Dimensions of Risk Sharing," Review of Economics and Statistics 81, pp.73-84.

[4] Deaton, A. (1991), \Saving and Liquidity Constraints," Econometrica 59, pp.1221-1248.

[5] Grosh, M. and P. Glewwe (1998), \The World Bank's Living Standards Measurement Study Household Surveys," Journal of Economic Perspectives 12, pp.187-196.

[6] Hayami, Y. and K. Otsuka (1993), The Economics of Contract Choice: An Agrarian Perspective, Oxford, Oxford Univ. Press.

[7] Hayashi, F. (1997), Understanding Saving, Cambridge, MIT Press.

[8] Jacoby, H. and E. Skou¯as (1997), \Risk, Financial Markets, and Human Capital in a Developing Country," Review of Economic Studies 64, pp.311-335.

[9] Kochar, A. (1999), \Smoothing Consumption by Smoothing Income: Hours-of-Work Responses to Idiosyncratic Agricultural Shocks in Rural India," Review of Economics and Statistics 81, pp.50-61.

[10] Kurosaki, T. (1998), Risk and Household Behavior in Pakistan's Agriculture, Tokyo: Institute of Developing Economies.

[11] Kurosaki, T. (1999), \Risk Sharing Arrangements and the Structure of Risk and Time Preferences: Theory and Evidence from Village India," IER Discussion Paper Series A No.383, November 1999, Hitotsubashi University.

[12] Kurosaki, T. and M. Fafchamps (forthcoming), \Insurance Market E±ciency and Crop Choices in Pakistan," Journal of Development Economics.

[13] Lense, S. H. (2000), \Using Consumption and Asset Return Data to Estimate Farmers' Time Preferences and Risk Attitudes," American Journal of Agricultural Economics 82, pp.934-947.

[14] Ligon, E. (1998), \Risk Sharing and Information in Village Economies," Review of Economic Studies 65, pp.847-864.

[15] Ligon, E., J. P. Thomas, and T. Worrall (1999), \Mutual Insurance with Limited Commitment: Theory and Evidence from Village Economies," Mimeo, U. C. Berkeley, March 1999.

[16] Mace, B. J. (1991), \Full Insurance in the Presence of Aggregate Uncertainty," Journal of Political Economy 99, pp.928-956.

[17] Pender, J. L. (1996), \Discount Rates and Credit Markets: Theory and Evidence from Rural India," Journal of Development Economics 50, pp.257-296.

[18] Ravallion, M. and S. Chaudhuri (1997), \Risk and Insurance in Village India: Comment," Econometrica 65, pp.171-184.

[19] Stiglitz, J. (1974), \Incentives and Risk Sharing in Sharecropping," Review of Economic Studies 41, pp.219-255.

[20] Stiglitz, J. (1988), \Economic Organization, Information and Development," in H. Chenery and T. N. Srinivasan (ed.), Handbook of Development Economics, Vol. I, Amsterdam: Elsevier Science, pp.93-160.

[21] Townsend, R. M. (1994), \Risk and Insurance in Village India," Econometrica 62, pp.539-591.

[22] Townsend, R. M. (1995), \Financial Systems in Northern Thai Villages," Quarterly Journal of Economics 110, pp.1011-1046.

[23] Udry, C. (1994), \Risk and Insurance in a Rural Credit Market: An Empirical Investigation in Northern Nigeria," Review of Economic Studies 61, pp.495-526.

[24] Udry, C. (1997), \Recent Advances in Empirical Microeconomic Research in Poor Countries: An Annotated Bibliography," Journal of Economic Education, Winter 1997, pp.58-75.

[25] Walker, T. S. and J. G. Ryan, J. G. (1990), Village and Household Economies in India's Semi-arid Tropics, Baltimore, Johns Hopkins Univ. Press.

Table 1: De¯nition and Summary Statistics of Empirical Variables

		mean	std.dev.	minimum	maximum
c_{it}	Real household consumption per capita = household consumption expenditure in real Indian Rupees (1983 Rs.) divided by the total adult equivalent units of household members.	1109.1	563.7	112.4	5384.2
y_{it}	Real household income per capita = total household income (a sum of crop income, labor income, and pro¯ts from other self-employed activities) in real Rs. divided by the total adult equivalent units of household members.	1498.6	1267.4	2.4	10098.8
Z_i	A vector of variables that approximate households' social positions, including:				
LANDD	A dummy variable for ownership of agricultural land in 1975.	0.788	0.410	0	1
LANDPC	The value of owned agricultural land in 1975 per capita (10,000 Rs.).	0.339	0.422	0	1.857
CHILDR	The share of children in household size using adult equivalents in 1975.	0.289	0.147	0	0.622
SCHOOL	Education status of the household head in years of complete education.	2.519	3.312	0	12
JGRRANK	Caste rank index compiled by J. G. Ryan with 1 for the socially highest castes and 4 for the lowest ones.	2.356	1.173	1	4

Note: The number of observations (NOB) is 1,040 (=104 households x 10 years) for c_{it} and y_{it}, and 104 for Z_i.

Table 2: Summary of Estimation Results of Time Series Estimation for Each Household

Parameter:	Distribution of coe±cient estimates mean	 std.dev.	 minimum	 maximum	Rejection ratio for $H_0 : b_i = 0$ @10%	 @5%
CARA model						
b_i	-7.4	87.9	-396.3	301.3	2.9 %	1.9 %
a_i	0.735	1.238	-2.544	4.173	23.1 %	16.3 %
³$_i$	0.203	0.448	-1.050	1.344	24.0 %	18.3 %
CRRA model						
b_i	-0.013	0.078	-0.325	0.196	1.9 %	1.9 %
a_i	0.558	0.802	-1.258	2.609	18.3 %	10.6 %
³$_i$	0.233	0.439	-0.906	1.607	28.8 %	18.3 %

Notes: 1) The estimated equation is (13).
2) NOB is 104.
3) Raw results that generate this table are given in Kurosaki (1999, Appendix Table 1).

Table 3: Bivariate Correlation Coefficients between Parameter Estimates and Household Characteristics

	LANDD	LANDPC	CHILDR	JGRRANK	SCHOOL
CARA model					
$\hat{\beta}_i$	-0.018	-0.058	0.153	0.065	-0.010
$\hat{\alpha}_i$	0.118	0.182 *	-0.072	-0.128	0.093
$\hat{\gamma}_i$	0.021	-0.049	0.032	0.086	-0.067
CRRA model					
$\hat{\beta}_i$	0.009	-0.057	0.181 *	0.065	0.034
$\hat{\alpha}_i$	0.099	0.216 **	-0.037	-0.106	0.079
$\hat{\gamma}_i$	-0.056	-0.104	0.078	0.126	-0.048

Note: Significant at 1% = ***, 5% = **, and 10% = *.

Table 4: Estimation Results with Household Structural Shifters

CARA model	LANDD		LANDPC		CHILDR		SCHOOL		JGRRANK	
b_0	-11.16		-5.92		-27.21		-6.05		-18.32	
	(0.30)		(0.27)		(0.72)		(0.28)		(0.47)	
b_1	6.48		-2.99		73.42		0.17		5.16	
	(0.15)		(0.07)		(0.63)		(0.03)		(0.35)	
a_0	0.498	**	0.589	***	0.928	***	0.692	**	1.098	***
	(0.88)		(4.61)		(4.10)		(5.49)		(4.72)	
a_1	0.269		0.552	**	-0.562		0.030		-0.137	
	(1.10)		(2.25)		(0.80)		(0.96)		(1.60)	
γ_0	0.236	**	0.151	***	0.079	*	0.119	*	0.091	**
	(2.09)		(4.20)		(1.90)		(3.58)		(2.08)	
γ_1	-0.136		-0.081	*	0.105		-0.003		0.008	
	(1.18)		(1.72)		(0.73)		(0.53)		(0.38)	
R^2	0.087		0.091		0.086		0.085		0.087	
$\bar{R}^2$	0.082		0.087		0.081		0.081		0.082	
Homogeneity test	0.795		2.423	*	0.473		0.373		0.886	

CRRA model	LANDD		LANDPC		CHILDR		SCHOOL		JGRRANK	
b_0	-0.021		-0.011		-0.041		-0.012		-0.020	
	(0.76)		(0.66)		(1.46)		(0.77)		(0.72)	
b_1	0.011		-0.007		0.103		0.000		0.004	
	(0.36)		(0.23)		(1.20)		(0.08)		(0.33)	
a_0	0.417	***	0.392	***	0.578	***	0.483	***	0.672	***
	(2.73)		(4.34)		(3.60)		(5.48)		(4.30)	
a_1	0.170		0.534	***	-0.055		0.032		-0.052	
	(0.99)		(3.02)		(0.11)		(1.39)		(0.88)	
γ_0	0.308	***	0.188	***	0.115	***	0.178	***	-0.009	
	(4.46)		(6.46)		(2.84)		(6.27)		(0.25)	
γ_1	-0.229	***	-0.217	***	-0.079		-0.026	***	0.056	***
	(3.20)		(4.14)		(0.59)		(3.79)		(3.45)	
R^2	0.110		0.121		0.101		0.114		0.111	
$\bar{R}^2$	0.105		0.116		0.097		0.109		0.107	
Homogeneity test	3.499	**	7.473	***	0.604		4.965	***	4.054	***

Notes: 1) NOB = 936.
2) "Homogeneity test" gives $F(3, 930)$ statistics for testing the joint hypothesis that $b_1 = a_1 = \gamma_1 = 0$.
3) Significant at 1% = ***, 5% = **, and 10% = * (2-sided test for t statistics whose absolute value is shown in parenthesis).
4) The estimated equation is (14).