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# DECREASING RELATIVE RISK AVERSION AND TESTS OF RISK SHARING[1]

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**Abstract**

It has often been argued that that the relative risk aversion (RRA) coefficient should decrease as a household becomes wealthier. However, existing tests of full risk sharing hypothesis in the empirical literature are derived using preferences that exhibit either *increasing* or constant RRA (CRRA). In this paper, we model decreasing RRA using Stone-Geary utility, which is the CRRA utility augmented by a "subsistence" parameter. Decreasing RRA implies that consumption of wealthy households will fluctuate more than that of poor ones under full risk sharing in an economy with aggregate risk, because rich households are more willing to bear risk. Using IFPRI and ICRISAT data, we find evidence in support of the hypothesis at the village level, and evidence against it at the inter-village level. When restricting the "subsistence" parameter to be zero, we replicate the previous results in the literature: reject it at both levels. Our tests, however, reject this restriction and favor the decreasing RRA in almost all cases. These results suggest that it is important to allow for decreasing RRA in testing full risk sharing hypothesis when data containing low-income households are investigated.

**Key Words**: Risk Sharing, Consumption Smoothing, Decreasing Relative Risk Aversion, Generalized Method of Moments

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# DECREASING RELATIVE RISK AVERSION AND TESTS OF RISK SHARING

## 1. INTRODUCTION

IN THE EMPIRICAL LITERATURE of risk sharing, parameterized forms of utility functions have been used. When functional forms are misspecified, tests can yield misleading results. Even though one cannot entertain a general form of the utility function because it is always possible to find a utility function that is consistent with any observation, it is important to use a flexible enough functional form to allow for important likely features of preferences over risky consumption.

How the relative risk aversion (RRA) coefficient varies with wealth has important implications on dynamic consumption decisions. It has often been argued that it is very likely that the RRA coefficient decreases as a household becomes richer (see, e.g., Mas-Colell, Whinston, and Green (1995), pp. 192-193, for a discussion). Indeed, Rosenzweig and Binswanger (1993) find that farmers are in general risk averse, but the wealthier they are, the less their investment portfolios are affected by increasing weather risk. This is consistent with, and could be caused by, decreasing RRA. However, isoelastic, exponential, and quadratic utility functions are typically used in the empirical literature of risk sharing. The isoelastic utility functions imply constant RRA, and exponential and quadratic utility functions imply that the RRA coefficient *increases* with the level of wealth. If the RRA coefficient decreases with the level of wealth, then wealthier households are willing to bear more risk. As a result, consumption of richer households could fluctuate more than that of poorer households even under complete risk sharing when aggregate shocks are present. Such an allocation of risk may result from systematically different investment strategies between rich and poor households: the poor may choose to hold lower risk and lower return assets than the rich. Again, Rosenzweig and Binswanger (1993) find empirical evidence for such a systematic difference in portfolio strategies, using the household level panel data from India collected by the International Crops Research Institute in the Semi-Arid Tropics (ICRISAT). A simple way to model decreasing RRA is to introduce a (what we call for convenience) "subsistence parameter" into the isoelastic CRRA utility. If this parameter is positive, then RRA coefficient will be very high when a household is consuming near the value of this parameter.[2] Previously, Atkeson and Ogaki (1996) and Ogaki and Atkeson (1997) have found empirical evidence in support of positive subsistence parameter, also using the ICRISAT data.[3] But tests of full risk sharing have generally ignored this parameter and its effect on the RRA coefficient, even though they have been applied to datasets containing low-income households (see, e.g., Altug and Miller (1990), Deaton (1990), Morduch (1990), Cochrane (1991), Mace (1991), Udry (1994), and Hayashi, Altonji, and

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[2] Zimmerman and Carter (1996) simulates an economic model in which poor consumers decide to hold less riskier assets than rich ones because of a positivie subsistence parameter.

[3] This paper is related to these two papers by Akeson and Ogaki in the sense that we are using almost the same complete markets model. However, the theme of this paper is risk aversion and risk sharing. Three major differences in this paper are: (i). A variable addition test, arguably more powerful than Hansen's ***J*** test against incomplete risk sharing alternatives, is developed, as discussed in Section 3 below; (ii). We investigate the effect of restricting the subsistence parameter to be zero on the testing of full risk sharing hypothesis and are able to replicate the results in the current literature; (iii). A Pakistani dataset with much more villages, as well as the Indian data, are used in the empirical work, and very robust results are obtained.

Kotlikoff (1996)). An exception is Townsend (1994), who reports that his results remain unchanged when some values of the subsistence level are tried for the ICRISAT data. This method, however, ignores the effect of estimating unknown subsistence level on the test statistics for full risk sharing. Townsend finds substantial comovements of consumption within a village, and much less comovements across villages. His formal statistics, however, reject the null hypothesis of full risk sharing within villages as well as across villages. Sawada (1996) applies tests that are similar to Townsend's to the panel data set of rural Pakistani households collected by the Food Security Management Project of the International Food Policy Research Institute (IFPRI), and finds mixed results for the hypothesis of full risk sharing within villages.

In this paper, we first present a model with decreasing RRA. Then we test the hypothesis of full risk sharing, taking into account the effect of estimating unknown subsistence parameter on the test statistics. Subsistence levels have been estimated from intratemporal first order conditions of the linear expenditure system in the literature of consumer demand (see, e.g., Stone (1954), Pollak and Wales (1969), Parks (1969), and Deaton (1974)). For the purpose of testing full risk sharing, however, it is desirable to use intertemporal conditions, which are directly related to risk sharing.

We apply our tests to the IFPRI and ICRISAT panel data of households. These data sets are of interest because they contain data of low income households, for whom decreasing RRA may well be the case. A unique feature of the IFPRI data set is that consumption and income data were collected separately from different members of each household. This feature is attractive for our purpose, because our tests, like most of the existing tests of full risk sharing, require the assumption that the measurement error in consumption is not correlated with income variables. The ICRISAT data set has been used extensively in development economics in general, and the risk sharing literature in particular. The reason to include it here is that we think it is important to compare the results from different datasets. After all, a good model should be able to "survive" as many datasets as possible in the empirical tests. In addition, it should also be very interesting to compare our results based on this dataset with those reported in the literature where the implication of decreasing RRA has been ignored.

Incomplete risk sharing can be caused by moral hazard and adverse selection problems, which are in turn caused by private information. Without private information, various arrangements can be used by the members of the community to share risk even when financial markets are not well developed. Private information is more likely to be present in a large community, or in an economy with more complicated production technologies than in a small community with simple production technologies. Hence full risk sharing can be a good approximation of the consumption growth pattern for a village economy in low income countries. On the other hand, full risk sharing across villages which are farther apart is not likely to be a good approximation. This consideration provides us with a measure of the power of our tests: a powerful test for the null hypothesis of full risk sharing should be able to reject full risk sharing *across* villages.

The rest of the paper is organized as follows. In Section 2, we present our model and discuss implications of a subsistence parameter on the RRA coefficient and consumption growth. We describe the data in Section 3. In Section 4, we explain our tests for full risk sharing which take into account the effect of estimating subsistence parameter. We present empirical results in Section 5. Section 6 concludes the paper.

## 2. THE MODEL

### 2.1 *An Arrow-Debreu Economy*

Consider an economy with *H* households. Let household *h, h=1,...,H,* have time and state separable utility with an intratemporal utility function $u\left(C_h(t,e(t))\right)$, where $C_h(t,e(t))$ is per capita consumption. In this paper household size is measured by the number of male-adult equivalents. So the per *capita* here, and henceforth, is actually per *male-adult equivalent*. The advantage of doing so is that all demographic variations, both across households and over the time, will have been taken care of. Let a vector *s(t), s(t)=1,2,...,S,* denote the state of the world in each period. The history of the economy can then be denoted by the vector *e(t)=[s(0),s(1),...,s(t)]*. Let $\beta$ denote the common discount factor in the economy. Then household *h*'s problem is to maximize

(1) $$U_h = \sum_{t=0}^{T_h} \sum_{e(t)} \beta^t \Pr(e(t) \mid e(0))\, u\left(C_h(t,e(t))\right),$$

where $\Pr(e(t)|e(0))$ denotes the conditional probability of $e(t)$ given $e(0)$, subject to a life-time budget constraint

(2) $$\sum_{t=0}^{T_h} \sum_{e(t))} \left( \prod_{\kappa=0}^{t} R\left(\kappa-1,e(\kappa-1),e(\kappa)\right) \right)^{-1} C_h(t,e(t)) \leq W(0),$$

where $T_h$ is the life-time of the household, and can either be finite as in the life cycle model or infinite as in Barro (1974); *W(0)* is household *h*'s per male-adult equivalent initial wealth; $R(\kappa-1,e(\kappa-1),e(\kappa))$ is the gross asset return of the state contingent security for the event $e(\kappa)$ in terms of the good in the event $e(\kappa-1)$ at period $\kappa-1$.[4] ($e(\kappa)$ is suppressed below for the sake of simplicity.) Now assume that

(3) $$u(C_h(t)) = \frac{(C_h(t)-\gamma)^{1-\alpha}-1}{1-\alpha},$$

where γ is the preference parameter that governs whether the RRA coefficient increases or decreases with the level of wealth. The RRA coefficient, $\theta_h$, of a household *h* implied by (3) is

(4) $$q_h = a\left(\frac{C_h}{C_h - g}\right)$$.If γ is zero, then the RRA coefficient is $a$, since (3) reduces to the CRRA case. If γ is positive, then the RRA coefficient is large for a poor household whose consumption is close to γ. But as the household becomes richer, the RRA coefficient falls and approaches $a$. If γ is negative, then the RRA coefficient rises with the level of wealth. For positive γ, one interpretation of this preference parameter is that it is subsistence consumption. Therefore we will call γ subsistence consumption in this paper, though other interpretations are also possible. The intertemporal first order condition for the optimization problem of household *h* includes

(5) $$\frac{C_h(t+1)-\gamma}{C_h(t)-\gamma} = \phi(t+1),$$

for any state of the world, where $\phi(t+1)$ = [ β *R(t,e(t),e(t+1))Pr(e(t+1)|e(t))*]$^{1/\alpha}$.

Equation (5) holds for each *h*. Since $\phi(t+1)$ is independent of *h*, $C_h - g$, the consumption in excess of subsistence level, should grow at the same rate for all

[4] *R(-1, e(-1),e(0))* is assumed to be 1.

households in any state of the world. This is because idiosyncratic risk is insured away through the complete asset markets in the model.

2.2 *Consumption Growth of the Rich and the Poor* The existence of wealth-varying relative risk aversion coefficient implies consumption growth differs systematically between the rich and the poor in our model. Intuitively, households with higher RRA coefficients will be more willing to bear risk and will experience more volatile consumption growth than those with lower RRA coefficients. One way to see implications on consumption growth is to examine the exact solution for consumption growth. Let $\overline{C_h} = C_h - \gamma$. Let $\hat{x}$ be the growth rate of any variable $x$, and $\mathbf{ln}\, x$ be the natural log of $x$. Then

(6) $$\hat{C}_h(t) = \mathbf{ln}[\overline{C_h}(t) \exp(\hat{\overline{C}}(t)) + \gamma] - \mathbf{ln}(\overline{C_h}(t) + \gamma),$$

where $\hat{\overline{C}}(t)$ is the common growth rate of $\overline{C_h}(t)$ at $t$. (6) implies

$$sign(\hat{C}_h(t)) = sign(\hat{\overline{C}}(t)).$$

Differentiating the right hand side of (6) with respect to $\overline{C_h}(t)$ yields

(7) $$\frac{\partial \hat{C}_h(t)}{\partial \overline{C_h}(t)} = \frac{\gamma(\exp(\hat{\overline{C}}(t)) - 1)}{(\overline{C_h}(t) \exp(\hat{\overline{C}}(t)) + \gamma)(\overline{C_h}(t) + \gamma)}.$$

Assuming that $\gamma$ is positive, (7) implies

$$sign\left(\frac{\partial \hat{C}_h(t)}{\partial \overline{C_h}(t)}\right) = sign(\exp(\hat{\overline{C}}(t)) - 1)$$

(8) $$= sign\ (\hat{\bar{C}}(t)).$$

Hence the rich households' consumption grows at faster rates in a state in which aggregate consumption grows, declines at faster rates in the states in which aggregate consumption declines.

Note that $\hat{\bar{C}}(t)$ is in fact $\phi(t)-1$. Therefore, a positive $\hat{\bar{C}}(t)$ implies

$$\beta R(t-1,e(t-1),e(t))\,\Pr(e(t)|e(t-1)) > 1,$$

i.e.

$$\ln\ R(t-1,e(t-1),e(t))\,\Pr(e(t)|e(t-1)) > -\ln\beta$$

The right-hand side of the inequality is just the rate of time preference, $\delta$, while the left-hand side is the real rate of return of the state-contingent security for state $e(t)$, $r(t,e(t-1),e(t))$. With this said, it is easy to see that (8) can be furthered to

(8') $$sign\ \left(\frac{\partial \hat{C}_h(t)}{\partial \overline{C_h}(t)}\right) = sign\ (\phi(t)-1)$$

$$= sign\ (r(t,e(t),e(t+1))-\delta).$$

The second equality of (8') says that in states in which the real return of assets is higher than the rate of time preference, the consumption growth rates of the rich households are higher than those of the poor households; but in states in which the real rate of return falls below the time preference rate, the consumption of rich households declines faster than poor households. Since the rate of time preference is a constant, the implication of the model is very clear: accompanying the fluctuations in real returns of assets around the time preference rate, the fluctuations in consumption of the rich households are larger than those of the poor households. Therefore the rich households bear more risk than the poor households in the equilibrium.

The implications above are obtained with the aid of the assumption that $\gamma$ is positive. It is natural for us to make this assumption, since we interpret it as subsistence consumption. (As a curvature parameter, $\gamma$ can be either positive or negative.) As we shall see in the empirical part below, the estimates of $\gamma$ are indeed all positive, statistically significant, and economically meaningful, except for one district in the Pakistani data.[5]

These results suggest that consumption growth is correlated with the level of wealth and hence current and lagged income across households in our model. The direction of correlation depends on whether or not aggregate consumption grows or declines. Our results also suggest that consumption growth can be correlated with income level and income growth even under complete risk sharing. For example, consider the case where the economy in our model grows over time. In a growing economy, rich households' consumption grows at a faster rate. In general, higher consumption growth of rich households are attained by higher saving rates that result in higher income growth. Thus consumption growth will be positively correlated with income growth across households. It is possible that such nonzero correlation of consumption growth and income variables are misinterpreted as evidence for liquidity constraints or incomplete markets in our model economy if subsistence levels are ignored. This problem may be alleviated by examining correlation of consumption growth and labor income growth rather than total income growth. This, however, is not a final solution if faster rates of human capital

[5] Even for this district (Dir), the estimate of $\gamma$ becomes positive and significant when we pool the data from all the districts in the sample. Please refer to Tables V and VI.

accumulation and resulting higher labor income growth are used to achieve higher consumption growth of rich households. If we make an additional assumption that the real risk free interest rate is constant as in Hall (1978), then the common growth rate of $\overline{C}$ is a martingale difference with possibly a drift (using a log normal assumption or a linear approximation). However, we do not assume that the real risk free interest rate is constant.

**3. ECONOMETRIC METHOD**

As discussed in the previous section, complete risk sharing implies that the growth rate of $C_h - \gamma$ is identical for all households in each state of each period (see Equation (5)). We now assume that consumption is measured with error:[6]

(9) $$C_h^m(t) = C_h(t) + \xi_h(t),$$

where $C_h^m(t)$ is measured consumption in per adult-equivalent terms, and $x_h(t)$ is measurement error. Then combining (5) and (9), we obtain

(10) $$C_h^m(t+1) - \phi(t+1)C_h^m(t) - \gamma + \gamma\phi(t+1) = e_h(t+1),$$

where $e_h(t+1)$ $= \xi_h(t+1) - \phi(t+1)\, \xi_h(t)$. Now assume that $\xi_h(t)$ is uncorrelated with household *h's* permanent and current incomes at time *t*. Let $y_h^p$ be a proxy of its permanent income, and $y_h(t)$ be its current income. Let $Z_h(t) = (1, y_h^p, \Delta y_h(t))'$ be a vector of instrumental variables. Let $\psi = (\phi(2),\ldots,\phi(T),\gamma)$ be the *T*-dimensional vector of unknown parameters and $\psi_0$ be the corresponding vector of their true values. *T* is the number of the time periods of the sample. In addition, let $f(C_h^m(t+1), y)$ be the 3-dimensional vector

[6] Ogaki and Atkeson (1997) discuss the choice between additive and multiplicative measurement errors.

(11) $$f(C_h^m(t+1), y) = Z_h(t+1)[C_h^m(t+1) - f(t+1)C_h^m(t) - g + gf(t+1)]$$

$$\equiv f_{h,t+1}$$

and

$$f_h(\psi) = (f_{h,2}(\psi), f_{h,3}(\psi), \ldots, f_{h,T}(\psi))'.$$

Then we have *3(T-1)* orthogonality conditions

(12) $$E_H\left(f_h(y_0)\right) = p\lim_{N\to\infty} \frac{1}{N}\sum_{h=1}^{N} f_h(y_0) = 0,$$

where $E_H$ is the expectation operator over households. Here *H* is attached to emphasize that the expectation is taken over households. Assume a central limit theorem applies to $f_h(\psi_0)$ so that $(1/N)^{1/2}\sum_{h=1}^{N} f_h(\psi_0)$ converges to a normal random vector with mean zero and covariance

(13) $$\Omega = plim_{N\to\infty} \frac{1}{N}\sum_{h=1}^{N} f_h(y_0) f_h(y_0)'.$$

$\frac{1}{N}\sum_{h=1}^{N} f_h(y_0) f_h(y_0)'$ can be estimated by Hansen (1982)'s generalized method of moments (GMM) using (12) as othorgonality conditions. In the estimation, we first pool the orthogonality conditions for all the households in each village, then stack the matrix of orthogonality conditions for all the villages in the same district. $f_h(\psi_0)$ is allowed to have different covariance matrices for different villages when the data are pooled.[7] We consider

They suggest that an additive specification would be better for the purpose of testing risk shairng.

[7] We assume that the regularity conditions of Gallant and White (1988) are satisfied. Hansen/Heaton/Ogaki's GAUSS GMM package (see Ogaki (1993b)) is used for the GMM estimation in this paper. Ogaki (1993a, Section 4.3) provides a more detailed explanation as to how the data for villages are pooled.

two types of tests. One type of test is the $\chi$-square test of the overidentifying restrictions, which is called Hansen's ***J*** test. Under the null hypothesis of full risk sharing, the disturbance term in (10) is uncorrelated with the income variables in the set of the instrumental variables. Therefore, the ***J*** test statistic has an asymptotic $\chi$-square distribution. Under the alternative hypothesis of incomplete risk sharing, the disturbance in (10) will be correlated with income variables. Hence the ***J*** test statistic will tend to be large. The other type of test is based on variable addition. We add the income difference term to (10) to obtain:(14)

$$C_h^m(t+1) - f(t+1)C_h^m(t) - g + gf(t+1) - h\Delta y_h(t+1) = e_h(t+1)$$

Under the null hypothesis of full risk sharing, $\eta = 0$, because the first order condition of the model indicates that individual income change should play no role in explaining individual consumption when the effect of $\gamma$ is accounted for. However, under the alternative hypothesis of incomplete risk sharing, individual income variables will affect individual consumption growth even after controlling for the effect of $\gamma$ and the effect of the aggregate shock. For example, take the alternative hypothesis of Keynesian consumption function, $C_h(t) = a + b\,y_h(t)$, where $0<b<1$. Under this hypothesis, a GMM estimation for (14) would result in $f(t) = 1$, and $h = b$, and therefore $\eta$ would be positive. Thus we can test the hypothesis of full risk sharing by testing the null of $h = 0$. This additional term could increase the power of the full risk sharing test, since by augmenting the model this way it is easier to pick up the correlation between the consumption and income variables now if it indeed exists.

In our empirical work, we pool data for villages, and the variable addition testing is conducted at two levels. At the village level, we test whether or not $h$ estimate is significantly different from zero for each village. At the pooled district level, we test whether or not the $h$ estimates of the villages in the same district are jointly significant. This is done by computing the likelihood-ratio-type test statistic, which is the difference between the constrained Hansen's $\boldsymbol{J}$ statistic and the unconstrained $\boldsymbol{J}$ statistic.[8] The variable addition tests are (arguably) more powerful than Hansen's $\boldsymbol{J}$ test against the alternative hypothesis of incomplete risk sharing. Hansen's $\boldsymbol{J}$ test tests against any correlation of the instrumental variables and the disturbance term. The variable addition tests are specifically directed toward the positive correlation between consumption and income which incomplete risk sharing implies. In addition, the variable addition test based on the income coefficient in each village can be used to test against incomplete risk sharing for each village even when data are pooled for many villages. Indeed, our empirical results are consistent with this argument. The results from the variable addition tests and Hansen's $\boldsymbol{J}$ tests are consistent in most cases. The only exceptions are for two Pakistani villages, where the variable addition test based on the income-difference coefficient rejects the null hypothesis of complete risk sharing, but Hansen's $\boldsymbol{J}$ test does not reject it. Another experiment that we will do is to examine what happens if we force $\gamma=0$ in the estimation and testing. As pointed out in the Introduction, the current literature generally ignores the role of $\gamma$. Forcing $\gamma=0$ is equivalent to what other researchers have done in their tests. If we can replicate the result of rejecting the null of

[8]See, e.g., Ogaki (1993) for a description of the likelihood-ratio-type test.

full risk sharing at village level when we impose this restriction, but can not reject it when allowing $\gamma$ to be estimated, then we can be confident that it is the restriction $\gamma=0$ that is driving the rejection of the model. In turn, we can test if this restriction itself is "reasonable" or not. If it is decisively rejected, then we conclude that it should not have been imposed in the estimation and testing, i.e. $\gamma$ should have been allowed to be different from zero. Then, if by taking into account the effect of estimating $\gamma$, indeed we do not reject the null hypothesis, we conclude that the theory presented in Section 2 is not rejected by the data.

**4. DATA** In this section, we describe the two sets of household level panel data we use: the data collected in Pakistan by the IFPRI and the data collected in India by the ICRISAT.

4.1 *The Pakistani Data* The IFPRI data used in this paper cover the period from April 1986 to September 1989.[9] During this period, 12 rounds of interviews were conducted at each sampled household. In each interview, a male questionnaire and a female questionnaire were used separately for collecting different data. The male questionnaire consisted of about 170 questions and was mainly about production, marketing, financing, various revenues, male labor supply and hiring, and nonfood expenditures. The female questionnaire had around 120 questions, and was mainly about demographics, food consumption, health status of children, and female labor supply. Food consumption data included purchases of 37 food items, and consumption from gifts and own production. The original survey started with 974 households at 52 villages in four districts. These four districts were distributed in three provinces in Pakistan: Punjab, Sind, and Northwest Frontier Province. Following Townsend (1994), we use demographic information in each household to calculate male-adult-equivalent household size, according to the equivalence scales provided by Ryan, Bidinger, Pushpamma, and Rao (1985). These scales are weights assigned to each member of a household according to his/her age and sex. Specifically, the scales are: 1.0 for males 19 or older, and .9 for females of the same age group; .94 and .83 for males and females, respectively, aged between 13 and 18; for children of either gender between 7 and 12 years old, .67; for children aged 4-6, .52; for those aged 1-3, .32; and for infants, .05. In each round of the survey, the status of each member was recorded: present, traveling, or moved to a new household. The annual household size used in this paper is the weighted average of male-adult-equivalent household sizes of all rounds in a year.

The annual income and food consumption expenditure data calculated by IFRPI are used in empirical analysis. The income measure includes nine subcategories: rental earnings in crops, net crop profits, farm wage income, non-farm income, net livestock profits, returns to capital, remittances, pension, and zakat. Since data on total consumption are not readily available, we test risk sharing for food consumption.[10] Assuming that food consumption is separable from other consumption categories, the

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[9]Although the survey covered at least six years, the data set does not contain consumption data for the fourth year. Hence we use the data up to 1989.

[10]Although some nonfood consumption data were collected, they were in nominal terms. It is not clear to us how to obtain real nonfood consumption since prices for nonfood items were not recorded.

model in Section II applies to food consumption. For our purpose, using food consumption is attractive for three reasons. First, our tests assume that the measurement error in reported food consumption is uncorrelated with income variables. Because food consumption and income variables are essentially collected from different household members, this assumption is more likely to be valid. Second, subsistence consumption is more likely to be important for food than for nonfood. Third, the aforementioned age-sex weights were obtained from dietary studies, and are more appropriate if used only for food consumption. It is not clear how to obtain appropriate adult-equivalent scales for nonfood consumption. The data for Village 15 to Village 20, and Village 52 are missing. From our sample, we exclude the households with incomplete information on any of the following: the age-sex composition, the food consumption and the income level for each of the three years. Concerned about sample size, we also exclude the villages with less than 11 households. As a result, in our sample, we have 633 households in 31 villages. In upper panel of Table I, we report district and village annual average real per male-adult-equivalent food consumption. These numbers are reported to facilitate the interpretation of the estimates of the subsistence level reported in the ensuing tables.

4.2 *The Indian Data* In this section, we describe the household level panel data collected in India by the ICRISAT.[11] We use panel data for

[11] Following Townsend (1994), we use the consumption data in ICRISAT's summary data. There are two ways of estimating consumption using the ICRISAT data. The ICRISAT's method is to infer it from transactions. The other method is to retrieve consumption by applying flow accounting identities to the production and storage data, which Ravallion and Chaudhuri (1997) propose. Their consumption data are very different from the ICRISAT's consumption data, and the difference is correlated with income. We

three villages (Aurepalle, Shirapur, and Kanzara) from fiscal year 1975-76 to fiscal year 1984-85. (In what follows, we denote each fiscal year by its first calendar year). In this paper, we report the results for food consumption.[12] Because the construction of food consumption was changed in 1976, food consumption data for 1975 are not used. These Indian panel data have been used to study consumption smoothing and risk sharing models by many authors.[13]

We use food including milk, sweets, and spices as the measure of food consumption. To construct real consumption per male adult equivalent, nominal consumption at *t* is divided by the family size measure constructed by Townsend (1994) and the corresponding food price index at *t* for each village. These real variables are valued at 1983 prices. There are about forty households for each year in each of the three villages in the data. Some households drop out of the sample and others are added to the sample over years in the ICRISAT data. We exclude these households from our sample. The number of households in our sample for the village of Aurepalle is 35; that for Shirapur, 33; and that for Kanzara, 36. We report in the lower panel of Table II village level annual average real per male-adult-equivalent food consumption.

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refer the reader to Ravallion and Chaudhuri (1997) and Townsend (pp. 554-555) for discussions of the suitability of these consumption data. It does not seem clear which consumption data set is more reliable.

[12] Nonfood Consumption are missing for most categories after 1982. Results for total consumption are similar to those for food consumption, and can be found in Ogaki, Atekeson, and Zhang (1997).

[13] See, e.g., Bhargava and Ravallion (1993), Jacoby and Skoufias (1993, 1995), Lim (1992), Rosenzweig (1988), Rosenzweig and Stark (1989), and Rosenzweig and Wolpin (1993).

**5. EMPIRICAL RESULTS** In this section, we report our empirical results for the two panel data sets.

5.1 *Results for the Pakistani Data* The test results for different districts in the Pakistani data are presented in Tables II, III, IV, and V. In each table, the first row reports the baseline results in which subsistence consumption (γ) is restricted to be equal across all villages in the sample, and full risk sharing is assumed within each village by restricting $f(t)$ to be equal across the households in each village. In each district, the ***J*** test in the first row does not reject the null hypothesis of full risk sharing at the five percent level. For the baseline case, the point estimate of γ is positive and statistically significant in all districts except for Dir. For Dir, the point estimate is negative, but it is not significantly different from zero. Because the standard error for γ is larger for Dir than is for the other three districts, the data for Dir do not seem to contain much information about γ. In the second, third, fourth, and fifth rows, the likelihood ratio type test statistic, ***C***, is reported, which is the difference between the ***J*** value for each row and that for the first row. In the second row, we impose the restrictions that $f(t)$ is the same *across* villages for each $t$ =1,2. If full risk sharing is achieved *across* villages, then these restrictions must be satisfied. We find overwhelming evidence against these restrictions for each district from the ***C*** statistic reported in the second row. Because full risk sharing is not likely to be achieved across villages given that private information is more likely to be a problem there, this indicates that the ***J*** test's power against incomplete risk sharing is at least not zero. The third row reports the variable addition test results. If the coefficient on the income change is significant for a village, we reject the full risk sharing hypothesis for that village. The ***C*** test tests the joint hypothesis that all the coefficients on the income changes are equal to zero. We do not reject the full risk sharing hypothesis for most villages except for Village 1 in Table II and Village 22 in Table IV. The ***C*** test does not reject the null hypothesis of full risk sharing within villages. The fourth row reports the results when $\gamma$ is allowed to be different across villages. The ***C*** test does not reject the restriction that the subsistence level is the same across villages. The fifth and sixth rows report the results when $\gamma$ is restricted to be zero. This corresponds to Townsend's (1994) model, except that he uses an exponential utility function. The ***C*** test strongly rejects this restriction in the fifth row for all districts except for Dir. The ***C*** test statistic reported in the sixth row is the difference between the ***J*** value in this row and that in the fifth row. The ***J*** test in the fifth row, and the ***C*** test in the sixth row test the null hypothesis of full risk sharing with $\gamma = 0$. These tests reject the null hypothesis of full risk sharing in all districts except for Dir. These results indicate that one can find evidence against full risk sharing when subsistence consumption is ignored. Table V reports the results when the data for all four districts are pooled. The first row reports the baseline results in which subsistence consumption ($\gamma$) is restricted to be equal across all the villages in the whole sample, and full risk sharing is assumed within each village. In the second row, we find overwhelming evidence against the null hypothesis of full risk sharing *across* districts. The ***C*** statistic in the fourth row tests the hypothesis that $\gamma$ is the same across all the households in the whole sample. This hypothesis is rejected at the 5 percent level, though not at the 1 percent level. The *p* values reported here, however, are based on the asymptotic approximation. It is possible that the approximation error is larger when the data for Dir are included, given the large standard error for $\gamma$ found for Dir. Indeed, when the data for Dir are excluded from the sample, we do not reject the hypothesis that $\gamma$ is the same across all the households at the 10 percent level.

5.2 *Results for the Indian Data* Table VII presents the results for the ICRISTA data. The first row reports the baseline results. The ***J*** test in the first row does not reject the null hypothesis of full risk sharing at any conventional significance level. The point estimate of $\gamma$ is positive and statistically significant. In the second row, we impose the restrictions that $f(t)$ is the same across villages for each *t*. As in the Pakistani data, we find overwhelming evidence against these restrictions for each district from the ***C*** statistic reported in the second row. This result is expected because full risk sharing is not likely to be achieved across villages given that they are far apart. The third row reports the variable addition test results. No coefficient on the income change is significant at the 5 percent level, and the ***C*** test does not reject the null hypothesis of full risk sharing within villages. The fourth row reports the results when $\gamma$ is allowed to be different across villages. The ***C*** test does not reject the restriction that the subsistence level is the same across villages.

The fifth and sixth rows report the results when $\gamma$ is restricted to be zero. As before, this corresponds to Townsend's (1994) model. The ***C*** test strongly rejects this restriction in the fifth row. The ***J*** test in the fifth row, and the ***C*** test in the sixth row test the null hypothesis of full risk sharing with $\gamma = 0$. The ***J*** test in the fifth row rejects the null hypothesis of full risk sharing within each village. The ***C*** test in the sixth row does not reject the null hypothesis at the fifteen percent level, but the income coefficient for Shirapur rejects it for the village at the one percent level. As in the Pakistani data, these results indicate that one can find evidence against full risk sharing when subsistence consumption is ignored.

**6. CONCLUSIONS** In this paper, we have tested full risk sharing hypothesis while taking into account the effect of estimating a parameter which allows the RRA coefficient to vary with the level of wealth. For 29 out of 31 villages in the Pakistani data and every village in the Indian data, we do not reject the hypothesis of full risk sharing within each village. Townsend (1995) finds that different villages in low-income countries have strikingly different institutional arrangements to cope with risk. Hence, it is not surprising that we find evidence against full risk sharing for two villages while we do not find such evidence for others. We, however, find strong evidence against risk sharing across villages in both data sets. Because it is more difficult to cope with the private information problem across villages, this result is also sensible. When we restrict subsistence consumption parameter to be zero (which implies CRRA), our tests replicate the well-known results of rejecting the full risk sharing hypothesis even within villages in both datasets, except for one district, Dir, in the Pakistani data. However, except for this district, our tests always reject the restriction of $\gamma = 0$ in both datasets. Thus, our test results show that misleading results may be obtained when decreasing RRA is ignored in testing full risk sharing hypothesis. In the empirical risk sharing literature, isoelastic, quadratic and exponential utility functions are often used. Because these utility functions imply either CRRA or increasing RRA, the test results based on these preference specifications need to be interpreted with caution.

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## Table I  Real Food Consumption Per Male-Adult Equivalent

1. IFPRI-Pakistan

*District Average*

| Year | Faisalabad | Attock | Badin | Dir | Whole Sample Average |
|---|---|---|---|---|---|
| 1986-87 | 2849 | 3545 | 2184 | 2841 | 2820 |
| 1987-88 | 2900 | 3108 | 1973 | 3519 | 2822 |
| 1988-89 | 2142 | 2293 | 2017 | 2806 | 2296 |
| 1986-89 | 2630 | 2982 | 2058 | 3055 | 2646 |

*Village Average*

| | *N* | 1986-87 | 1987-88 | 1988-89 | 1986-89 |
|---|---|---|---|---|---|
| *Faisalabad* | | | | | |
| Vil. 1 | 25 | 3572 | 2908 | 2212 | 2897 |
| Vil. 2 | 26 | 2212 | 2113 | 1857 | 2061 |
| Vil. 3 | 25 | 2866 | 3572 | 2155 | 2864 |
| Vil. 4 | 20 | 2779 | 3151 | 2362 | 2764 |
| Vil. 5 | 25 | 2661 | 2908 | 2116 | 2562 |
| Vil. 6 | 26 | 3010 | 2833 | 2200 | 2681 |
| *Attock* | | | | | |
| Vil. 7 | 18 | 3591 | 4113 | 2409 | 3371 |
| Vil. 8 | 21 | 3628 | 3193 | 2210 | 3010 |
| Vil. 9 | 16 | 3990 | 2990 | 2327 | 3102 |
| Vil. 10 | 20 | 3632 | 3476 | 2470 | 3193 |
| Vil. 11 | 23 | 3102 | 3000 | 2157 | 2753 |
| Vil. 12 | 19 | 3615 | 2389 | 2024 | 2676 |
| Vil. 13 | 20 | 3131 | 2824 | 2242 | 2732 |
| Vil. 14 | 16 | 3898 | 2889 | 2596 | 3128 |
| *Badin* | | | | | |
| Vil. 21 | 17 | 2131 | 2390 | 2285 | 2269 |
| Vil. 22 | 19 | 2414 | 2018 | 1950 | 2127 |
| Vil. 23 | 18 | 1793 | 1720 | 1875 | 1796 |
| Vil. 24 | 14 | 1807 | 1743 | 1801 | 1786 |
| Vil. 25 | 15 | 2159 | 2073 | 1914 | 2049 |
| Vil. 26 | 13 | 2755 | 2322 | 2064 | 2380 |
| Vil. 29 | 23 | 2439 | 2070 | 2154 | 2221 |
| Vil. 30 | 21 | 2103 | 1699 | 1966 | 1922 |
| Vil. 34 | 12 | 2359 | 1891 | 2119 | 2123 |
| Vil. 37 | 22 | 1896 | 1829 | 1941 | 1888 |
| Vil. 39 | 12 | 2340 | 2095 | 2155 | 2197 |
| *Dir* | | | | | |
| Vil. 41 | 40 | 2878 | 3447 | 3152 | 3159 |
| Vil. 42 | 12 | 2624 | 3359 | 2807 | 2930 |
| Vil. 45 | 27 | 2706 | 3231 | 2386 | 2774 |
| Vil. 47 | 32 | 2842 | 3804 | 2836 | 3161 |
| Vil. 48 | 13 | 3145 | 3853 | 3043 | 3347 |
| Vil. 51 | 23 | 2872 | 3481 | 2518 | 2957 |

## 2. ICRISAT-India

| | *N* | 1976 | 1977 | 1978 | 1979 | 1980 | 1981 | 1982 | 1983 | 1984 | 1976-84 |
|---|---|---|---|---|---|---|---|---|---|---|---|
| Aurepalle | 35 | 313 | 381 | 408 | 538 | 502 | 423 | 414 | 409 | 526 | 423 |
| Shirapur | 33 | 604 | 555 | 644 | 543 | 623 | 521 | 388 | 351 | 351 | 528 |
| Kanzara | 36 | 490 | 489 | 418 | 578 | 571 | 479 | 418 | 578 | 571 | 363 |

*Notes*: 1. The figures here are in 1986 Pakistani Rupee (when 1 Rupee =US$.063), and 1975 Indian Rupee, respectively.

2. *N* indicates the number of households in the sample in each village.

**Table II GMM Results for Food Consumption- Faisalabad**
**(IFPRI-Pakistan)**

| Risk Sharing | $\gamma$ | coeff. $\Delta y_1(t+1)$ ** | coeff. $\Delta y_2(t+1)$ | coeff. $\Delta y_3(t+1)$ | coeff. $\Delta y_4(t+1)$ | coeff. $\Delta y_5(t+1)$ | coeff. $\Delta y_6(t+1)$ | ***J*** *** | ***C*** *** |
|---|---|---|---|---|---|---|---|---|---|
| ***Within*** | 1511* | … | … | … | … | … | … | 25.7 | … |
| ***Vil.*** | (124) | | | | | | | (.316,23) | |
| ***Across*** | 1447 | … | … | … | … | … | … | 115.2 | 89.55 |
| ***Vil.*** | (80) | | | | | | | (.000,33) | (.000,10) |
| ***Within*** | 1474 | .1659 | .0012 | .0361 | -.0168 | .0013 | -.0046 | 19.7 | 5.95 |
| ***Vil.*** | (146) | (.0797) | (.0190) | (.0316) | (.0504) | (.0153) | (.0108) | (.289,17) | (.429,6) |
| ***Within*** | …*** | … | … | … | … | … | … | 19.7 | 7.73 |
| ***Vil.*** | | | | | | | | (.459,18) | (.172,5) |
| ***Within*** | 0 | … | … | … | … | … | … | 71.3 | 45.6 |
| ***Vil.*** | | | | | | | | (.000,24) | (.000,1) |
| ***Within*** | 0 | .2808 | .0273 | .0134 | -.0626 | -.0053 | .0002 | 49.2 | 22.1 |
| ***Vil.*** | | (.0699) | (.0139) | (.0308) | (.0486) | (.0106) | (.0105) | (.000,18) | (.001,6) |

*: Standard errors are in parenthesis under the estimates, except for the two columns for the ***J*** and ***C*** statistics, where the numbers in parenthesis are *p*-value and degree of freedom, respectively.
**: The subscript of income difference term denotes village identification number, i.e., 1 for Vil. 1.
***: ***J*** is a $\chi^2$ statistic, and ***C*** is a likelihood ratio type statistic.
****: The subsistence estimates for Villages 1 to 6 are (with standard errors in parenthesis) as follows: 1851 (215), 1467 (240), 1131 (803), 1326 (335), 1738 (206), -512 (2411), respectively.

**Table III GMM Results for Food Consumption- Attock**
**(IFPRI-Pakistan)**

| Risk Sharig | $\gamma$ | coeff. $\Delta y_7(t+1)$** | coeff. $\Delta y_8(t+1)$ | coeff. $\Delta y_9(t+1)$ | coeff. $\Delta y_{10}(t+1)$ | coeff. $\Delta y_{11}(t+1)$ | coeff. $\Delta y_{12}(t+1)$ | coeff. $\Delta y_{13}(t+1)$ | coeff. $\Delta y_{14}(t+1)$ | ***J* ***** | ***C* ***** |
|---|---|---|---|---|---|---|---|---|---|---|---|
| ***Within Vil.*** | 1867*<br>(117) | … | … | … | … | … | … | … | … | 37.8<br>(.187,31) | … |
| ***Across Vil.*** | 1820<br>(77) | … | … | … | … | … | … | … | … | 556.9<br>(.000,45) | 519.1<br>(.000,14) |
| ***Within Vil.*** | 1749<br>(148) | .0214<br>(.0423) | .0013<br>(.0284) | .0785<br>(.0472) | -.0304<br>(.0163) | .0221<br>(.0606) | .0257<br>(.0211) | -.0618<br>(.0501) | .0079<br>(.0203) | 28.1<br>(.212,23) | 9.7<br>(.290,8) |
| ***Within Vil.*** | …*** | … | … | … | … | … | … | … | … | 24.4<br>(.441,24) | 13.4<br>(.063,7) |
| ***Within Vil.*** | 0 | … | … | … | … | … | … | … | … | 122.0<br>(.000,32) | 82.2<br>(.000,1) |
| ***Within Vil.*** | 0 | .0525<br>(.0397) | .0437<br>(.0184) | .0151<br>(.0161) | -.0573<br>(.0135) | .0199<br>(.0640) | .0509<br>(.0171) | -.0736<br>(.0557) | .0387<br>(.0140) | 57.8<br>(.000,24) | 64.2<br>(.000,8) |

*: Standard errors are in parenthesis under the estimates, except for the two columns for the ***J*** and ***C*** statistics, where the numbers in parenthesis are *p*-value and degree of freedom, respectively.
**: The subscript of income difference term denotes village identification number, i.e., 7 for Vil. 7.
***: ***J*** is a $\chi^2$ statistic, and ***C*** is a likelihood ratio type statistic.
****: The subsistence level estimates for Villages 7 to 14 are (with standard errors in parenthesis): 1998 (294), 1978 (172), -1445 (2175), 1936 (347), 1532 (607), 2018 (259), 1.9e+5 (1.9e+7), 2256 (265), respectively.

## Table IV GMM Results for Food Consumption- Badin (IFPRI-Pakistan)

| Risk Sharig | $\gamma$ | coeff. $\Delta y_{21}(t+1)$** | coeff. $\Delta y_{22}(t+1)$ | coeff. $\Delta y_{23}(t+1)$ | coeff. $\Delta y_{24}(t+1)$ | coeff. $\Delta y_{25}(t+1)$ | coeff. $\Delta y_{26}(t+1)$ | coeff. $\Delta y_{29}(t+1)$ | coeff. $\Delta y_{30}(t+1)$ | coeff. $\Delta y_{34}(t+1)$ | coeff. $\Delta y_{37}(t+1)$ | coeff. $\Delta y_{39}(t+1)$ | ***J* ***** | ***C* ***** |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|

| | | | | | | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| ***Within*** | 1441* | … | … | … | … | … | … | … | … | … | … | … | 44.2 | … |
| ***Vil.*** | (84) | | | | | | | | | | | | (.420,43) | |
| ***Across*** | 1701 | … | … | … | … | … | … | … | … | … | … | … | 326.0 | 282.2 |
| ***Vil.*** | (48) | | | | | | | | | | | | (.000,63) | (.000,20) |
| ***Within*** | 1434 | .046 | .110 | -.047 | .002 | .019 | .015 | .005 | .025 | .044 | .046 | .029 | 28.8 | 15.4 |
| ***Vil.*** | (100) | (.038) | (.050) | (.046) | (.067) | (.012) | (.022) | (.013) | (.036) | (.034) | (.036) | (.032) | (.630,32) | (.165,11) |
| ***Within*** | …*** | … | … | … | … | … | … | … | … | … | … | … | 29.4 | 14.8 |
| ***Vil.*** | | | | | | | | | | | | | (.647,33) | (.140,10) |
| ***Within*** | 0 | … | … | … | … | … | … | … | … | … | … | … | 92.9 | 48.7 |
| ***Vil.*** | | | | | | | | | | | | | (.000,44) | (.000,1) |
| ***Within*** | 0 | .040 | .124 | -.079 | .041 | .027 | .051 | -.010 | .047 | .033 | .040 | -.005 | 50.4 | 42.5 |
| ***Vil.*** | | (.038) | (.052) | (.042) | (.066) | (.013) | (.011) | (.012) | (.036) | (.028) | (.035) | (.020) | (.000,33) | (.000,11) |

*: Standard errors are in parenthesis under the estimates, except for the two columns for the ***J*** and ***C*** statistics, where the numbers in parenthesis are *p*-value and degree of freedom, respectively.

**: The subscript of income difference term denotes village identification number, i.e., 21 for Vil. 21, an so on.

***: ***J*** is a $\chi^2$ statistic, and ***C*** is a likelihood ratio type statistic.

****: The subsistence level estimates for Villages 21 to 39 are (with standard errors in parenthesis): 1627 (385), 2018 (74), 1625 (147), 1760 (86), 1850 (18), 1254 (138), 1453 (334), 1405 (174), 2082 (221), 7172 (2e+5), 1768 (355), respectively.

## Table V GMM Results for Food Consumption- Dir (IFPRI-Pakistan)

| Risk Sharing | $\gamma$ | coeff. $\Delta y_{41}(t+1)$** | coeff. $\Delta y_{42}(t+1)$ | coeff. $\Delta y_{45}(t+1)$ | coeff. $\Delta y_{47}(t+1)$ | coeff. $\Delta y_{48}(t+1)$ | coeff. $\Delta y_{51}(t+1)$ | ***J*** *** | ***C*** *** |
|---|---|---|---|---|---|---|---|---|---|
| ***Within*** | -207* | … | … | … | … | … | … | 19.8 | … |
| ***Vil.*** | (905) | | | | | | | (.652,23) | |
| ***Across*** | 976 | … | … | … | … | … | … | 58.3 | 38.48 |
| ***Vil.*** | (387) | | | | | | | (.004,33) | (.000,10) |

| | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| ***Within*** | -317 | .0387 | .0019 | -.0003 | .0024 | -.0055 | .0235 | 15.2 | 4.62 |
| ***Vil.*** | (1555) | (.0340) | (.0324) | (.0110) | (.0064) | (.0202) | (.0261) | (.570,17) | (.593,6) |
| ***Within*** | …*** | … | … | … | … | … | … | | |
| ***Vil.*** | | | | | | | | | |
| ***Within*** | 0 | … | … | … | … | … | … | 19.9 | .06 |
| ***Vil.*** | | | | | | | | (.703,24) | (.806,1) |
| ***Within*** | 0 | .0380 | .0035 | .0008 | .0022 | -.0088 | .0018 | 17.2 | 2.7 |
| ***Vil.*** | | (.0339) | (.0318) | (.0098) | (.0063) | (.0139) | (.0186) | (.511,18) | (.845,6) |

*: Standard errors are in parenthesis under the estimates, except for the two columns for the ***J*** and ***C*** statistics, where the numbers in parenthesis are *p*-value and degree of freedom, respectively.
**: The subscript of income difference term denotes village identification number, i.e., 41 for Vil. 41.
***: ***J*** is a $\chi^2$ statistic, and ***C*** is a likelihood ratio type statistic.
****: No convergence achieved for this case.

## Table VI GMM Results for Food Consumption- The Whole Sample (IFPRI-Pakistan)

| Risk Sharig | $\gamma_F$ * | $\gamma_A$ | $\gamma_B$ | $\gamma_D$ | coeff. $\Delta y_F(t+1)$ * | coeff. $\Delta y_A(t+1)$ | coeff. $\Delta y_B(t+1)$ | coeff. $\Delta y_D(t+1)$ | ***J*** *** | ***C*** *** |
|---|---|---|---|---|---|---|---|---|---|---|
| ***Within*** | 1812 | 1812 | 1812 | 1812 | … | … | … | … | 146.2 | … |
| ***Vil.*** | (27)** | | | | | | | | (.076,123) | |
| ***Across*** | 1767 | 1767 | 1767 | 1767 | … | … | … | … | 1093.2 | 947.1 |
| ***Dist.*** | (29) | | | | | | | | (.000,177) | (.000,54) |
| ***Within*** | 1787 | 1787 | 1787 | 1787 | .0010 | .0007 | .0135 | .0054 | 138.0 | 8.2 |
| ***Vil.*** | (27) | | | | (.0077) | (.0087) | (.0048) | (.0060) | (.112,119) | (.086,4) |
| ***Within*** | 1639 | 1824 | 1851 | 1103 | … | … | … | … | 137.7 | 8.5 |

| | | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| ***Vil.*** | (105) | (119) | (28) | (468) | | | | | (.129,120) | (.037,3) |
| ***Within*** | 0 | 0 | 0 | 0 | … | … | … | … | 406.7 | 260.6 |
| ***Vil.*** | | | | | | | | | (.000,124) | (.000,1) |
| ***Within*** | 0 | 0 | 0 | 0 | .0092 | .0155 | .0160 | .0047 | 388.0 | 18.7 |
| ***Vil.*** | | | | | (.0059) | (.0070) | (.0049) | (.0058) | (.000,120) | (.001,4) |

*: Subscripts *F*, *A*, *B*, and *D* are for Faisalabad, Attock, Badin, and Dir, respectively.

**: Standard errors are in parenthesis under the estimates, except for the two columns for the ***J*** and ***C*** statistics, where the numbers in parenthesis are *p*-value and degree of freedom, respectively.

***: ***J*** is a $\chi^2$ statistic, and ***C*** is a likelihood ratio type statistic.

## Table VII GMM Results for Food Consumption

## (ICRISAT-India)

| Risk Sharing | $g_A{}^*$ | $g_S$ | $g_K$ | coeff. $\Delta y_A(t+1)^*$ | coeff. $\Delta y_S(t+1)$ | coeff. $\Delta y_K(t+1)$ | ***J***** | ***C*** $^{**}$ |
|---|---|---|---|---|---|---|---|---|
| ***Within Vil*.** | 237.3** (15.2) | 237.3 | 237.3 | … | … | … | 46.2 (.507, 47) | … |
| ***Across Vil*.** | 269.7 (14.0) | 269.7 | 269.7 | … | … | … | 1239.4 (.000, 63) | 1193.2 (.000, 16) |
| ***Within Vil.*** | 238.4 (16.1) | 238.4 | 238.4 | .024 (.013) | .011 (.011) | .005 (.016) | 41.3 (.589, 44) | 4.88 (.181, 3) |
| ***Within Vil.*** | 237.9 (21.4) | 233.1 (30.3) | 240.5 (30.6) | … | … | … | 46.1 (.425, 45) | .03 (.985, 2) |
| ***Within Vil*.** | 0 | 0 | 0 | … | … | … | 114.5 (.000, 48) | 68.3 (.000, 1) |
| ***Within Vil*.** | 0 | 0 | 0 | .012 (.011) | .033 (.010) | .012 (.016) | 102.3 (.000, 45) | 12.2 (.007, 3) |

*: Subscripts *A*, *S*, and *K* denote Aurepalle, Shirapur, and Kanzara, respectively.
**: Standard errors are in parenthesis under the estimates, except for the two columns for the ***J*** and ***C*** statistics, where the numbers in parenthesis are *p*-value and degree of freedom, respectively.
***: ***J*** is a $\chi^2$ statistic, and ***C*** is a likelihood ratio type statistic.