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Decreasing Relative Risk Aversion and Tests of Risk Sharing

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DECREASING RELATIVE RISK AVERSION AND TESTS OF RISK SHARING

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1. INTRODUCTION

THE RELATIVE RISK AVERSION (RRA) coefficient of a household whose consumption is close to the subsistence level may be very high. For example, if consumption is exactly at the subsistence level, the household may not be willing to bear any risk, and both the absolute and relative risk aversion coefficients could be infinite. If this is the case, then the RRA coefficient must be a decreasing function of wealth for poor households. Therefore we should allow the possibility of decreasing RRA (DRRA) in testing the full risk sharing hypothesis.[2] However, existing tests in the empirical literature are derived using preferences that exhibit either increasing or constant RRA even when they are applied to data containing low-income households (see, e.g., Altug and Miller (1990), Deaton (1990), Morduch (1990), Cochrane (1991), Mace (1991), Townsend (1994), Hayashi, Altonji, and Kotlikoff (1996), and Sawada (1996)). We therefore use a Hyperbolic Absolute Risk Aversion (HARA) utility, which implies increasing, constant, and decreasing RRA as special cases, to test full risk-sharing hypothesis.

Using the International Food Policy Research Institute (IFPRI) data for Pakistani households and the International Crops Research Institute of the Semi-Arid Tropics (ICRISAT) data for Indian households, we find evidence in support of the DRRA hypothesis, along with evidence favoring the full risk-sharing hypothesis at the village level, and evidence against the hypothesis at the inter-village level. When RRA is restricted to be constant, we replicate the previous results in the literature: reject the full risk-sharing hypothesis at both levels. Our tests, however, reject this restriction and favor DRRA in almost all cases. These results suggest that it is important to allow for DRRA in testing the full risk-sharing hypothesis when data containing low-income households are investigated.

A unique feature of the IFPRI data set is that food 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 and their measurement errors. The ICRISAT data set has been used extensively in development economics. Townsend (1994) uses it to test the

[1] We are thankful for helpful comments from four anonymous referees and seminar participants at University of Illinois at Urbana-Champaign, Johns Hopkins, University of Kentucky, Ohio State, the Fourth Midwest Macroeconomics Conference at Notre Dame, and the Tow Conference on Macroeconomics at University of Iowa. Qiang Zhang acknowledges the financial support from a Charles Dice Fellowship and a SGAES research grant from Ohio State. This paper is based on a part of Zhang's dissertation, but we each made equal contribution.

[2] This intuition is based on our introspection. Arrow (1965), based on the boundedness of the utility function, argues that RRA should increase with wealth over the whole domain of utility function, though some fluctuations in RRA are likely.

full risk sharing hypothesis.[3] It is encouraging that our test results are qualitatively similar for these two data sets.

Existing evidence is consistent with the idea that the RRA is decreasing for low-income households. For example, Guiso, Jappelli, and Terlizzese (1996) find that the share of risky assets in a household portfolio is positively correlated with income and wealth level in data for Italian households. Kessler and Wolff (1991) find that the share of wealth invested in risky assets is, in general, increasing with wealth for French and American households. Rosenzweig and Binswanger (1993) report that Indian farmers are in general risk-averse, but the wealthier they are, the less their investment portfolios are affected by increasing weather risk.[4]

We present our model and derive its testable implication in Section 2. In Section 3 we explain our tests for full risk sharing. The empirical results are in Section 4. Section 5 concludes the paper. The IFPRI data are described in the Appendix.

2. THE MODEL

Consider an economy with H households participating in a risk-sharing pool.[5] We assume that preferences of each household are additively separable between consumption and leisure and are described by a discounted sum of expected felicities over time. We also assume that households have the identical time discount factor and probability beliefs on the states of the world. The Pareto-efficient consumption sharing rule is given by a programming problem (similar to that of Townsend (1994, pp. 556–557)) of maximizing the weighted sum of discounted household expected utilities, subject to the resource constraint that the sum of household consumption is less than or equal to the aggregate consumption available to the pool in each period and state history. We denote the aggregate consumption by $C_a(t)$.[6]

How RRA varies with wealth has important implications on consumption under full risk sharing. Given the welfare weights, the consumption sharing rule depends only on $C_a(t)$, according to the well-known Mutuality Principle. Let $C_h^*(C_a(t))$ be the sharing rule for household h. Because of the time separability assumption for preferences, Wilson's (1968) Theorem 5 holds in each period in the programming problem above. Hence dC_h^*/dC_a is inversely proportional to household h's absolute risk aversion coefficient. Therefore, we obtain the following proposition.

PROPOSITION: *$(dC_h^*/dC_a)/C_h^* = \psi(1/\theta_h)$, where θ_h is the RRA coefficient and ψ only depends on $C_a(t)$.*

[3] Townsend (1994) experimented with subsistence consumption and found it to be insignificant. It should be noted that the exponential utility function he uses implies increasing RRA even with a positive subsistence parameter.

[4] Using survey data on a representative U.S. age group (51–61), Barsky et al. (1997) find that RRA tends to increase slightly from the first to the third wealth quintiles and decrease slightly from the third to the fifth quintiles. Barsky et al. are interested in relatively low values of RRA relevant for asset pricing; hence their survey questions are not designed to detect possible very high values of RRA for poor respondents.

[5] Based on a model built upon *individual* utility functions, Zhang (1998) reports test results that are quantitatively similar to ours.

[6] The aggregate consumption depends on the history of states, but we suppress it throughout this paper to ease the exposition.

That is, consumption growth is inversely related to RRA. Suppose that RRA decreases with wealth. Then the consumption growth rates of the rich households are higher than those of poor ones when $C_a(t)$ increases, and are lower than those of the poor ones when $C_a(t)$ decreases. Hence the growth rate of consumption of a rich household will fluctuate more than that of a poor one. This is in contrast with the result under constant RRA that the consumption growth rate is identical for all households in the pool.

To parameterize flexible (i.e. increasing, decreasing, or constant) RRA in a parsimonious way, we now assume that the utility function for household h takes the common form

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

where γ is the preference parameter that governs whether the RRA coefficient increases or decreases with the level of wealth. To take into account the demographic effects in simple form, $C_h(t)$ is defined as total household consumption divided by the adult equivalent household size. The RRA coefficient, θ_h, of household h implied by (1) is

$$(2) \qquad \theta_h = \alpha\left(\frac{C_h}{C_h - \gamma}\right).$$

If γ is zero, (2) reduces to the constant RRA case. If γ is positive, RRA decreases with consumption and approaches α asymptotically. A negative γ, on the other hand, implies that the RRA coefficient rises with consumption. A usual interpretation for positive γ is subsistence consumption. Hence we will call it the subsistence parameter.

The intertemporal first order conditions for the programming problem imply that, for any state of the world,

$$(3) \qquad \frac{C_h(t+1) - \gamma}{C_h(t) - \gamma} = \phi(t+1),$$

where $\phi(t+1) = [\beta \operatorname{prob}(t+1|t)\mu(t)/\mu(t+1)]^{1/\alpha}$, β is the common time discount factor, and $\mu(t)$ is the Lagrange multiplier associated with the resource constraint at period t.

Our econometric method will be based on (3), which holds for each household h. Since $\phi(t+1)$ is independent of h, $C_h - \gamma$, the consumption in excess of the subsistence level, should grow at the same rate for all households in any state of the world. This is because idiosyncratic risk is insured away through the optimal risk sharing arrangements.

3. ECONOMETRIC METHOD

We assume that consumption is measured with error.[7]

$$(4) \qquad C_h^m(t) = C_h(t) + \xi_h(t),$$

where $C_h^m(t)$ is measured consumption in per adult-equivalent terms, and $\xi_h(t)$ is measurement error. Then combining (3) and (4), we obtain

$$(5) \qquad C_h^m(t+1) - \phi(t+1)C_h^m(t) - \gamma + \gamma\phi(t+1) = \nu_h(t+1),$$

where $v_h(t+1) = \xi_h(t+1) - \phi(t+1)\xi_h(t)$.

[7] Ogaki and Atkeson (1997) discuss the choice between additive and multiplicative measurement errors. They suggest that an additive specification would be better for the purpose of testing risk sharing.

Now assume that $\xi_h(t)$ is uncorrelated with household h's permanent and current income variables and their measurement errors. Let y_h^p be a proxy of its permanent income, and $y_h(t)$ be a measure of its current income. Let $Z_h(t) = (1, y_h^p, \Delta y_h(t))'$ be the vector of instrumental variables. Since there are nonlinear constraints on parameters as shown in (5), Hansen's (1982) Generalized Method of Moments (GMM) is used to estimate the model as explained in Section 4.3 of Ogaki (1993). Compared with Townsend's (1994) method for panel data, our method has the advantage of allowing for a general form of serial correlation.

We consider two types of tests. The first type is the χ^2 test of the over-identifying restrictions (i.e., Hansen's J test). Under the null hypothesis of full risk sharing, the disturbance term in (5) is uncorrelated with the income variables in the instrumental variables set. Therefore, the J test statistic has an asymptotic χ^2 distribution. Under the alternative hypothesis of incomplete risk sharing, the disturbance in (5) 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 (5) to obtain

$$C_h^m(t+1) - \phi(t+1)C_h^m(t) - \gamma + \gamma\phi(t+1) - \eta\,\Delta y_h(t+1) = \nu_h(t+1). \tag{6}$$

Under the null hypothesis of full risk sharing, $\eta = 0$, because the model implies that individual income change should play no role in explaining individual consumption growth when the effects of subsistence consumption and aggregate shocks are controlled by the parameters γ and $\phi(t)$. However, under the alternative hypothesis of incomplete risk sharing, individual income variables will affect individual consumption growth even after controlling for these effects. For example, take the alternative of Keynesian consumption function, $C_h = a + by_h$, where $0 < b < 1$. Under this hypothesis, if current income is measured without errors, GMM estimation for (6) would result in $\phi(t) = 1$ and $\eta = b$.[8] In our empirical work, this test is conducted at two levels. At the village level we test if the η estimate is significantly different from zero for each village. At the pooled district level, we test if the η estimates of all 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 J statistic and the unconstrained J statistic (see, e.g., Ogaki (1993, Section 7)). The variable addition test is more directly related to the alternative hypothesis of incomplete risk sharing than Hansen's J test, and therefore is arguably more powerful. The test is similar to the ones based on regression of consumption change/growth on individual shocks frequently seen in the empirical risk sharing literature.

Another experiment that we will do is to examine what happens if we force $\gamma = 0$ in the estimation and testing. 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 full risk sharing at the village level when we impose this restriction, but can not reject it when allowing γ 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 whether or not this restriction itself is reasonable by constructing another likelihood-ratio type test.

[8] If current income is measured with error, the probability limit of our estimator for η is smaller than b, but it is positive. So the variable addition test still has power against this alternative hypothesis.

4. EMPIRICAL RESULTS

The empirical results reported in Tables I–V are based on food consumption data from the IFPRI data set described in the Appendix and the ICRISAT data set described in Townsend (1994).[9] 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 and their measurement errors. This assumption is likely to be valid for Pakistani data as explained in the Introduction. Second, the age-sex weights used to calculate adult-equivalent household size 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. Third, the notion of subsistence consumption is more likely to be important for food than for nonfood consumption. It should be helpful to note, for the purpose of interpreting the estimates of the subsistence parameter, that the three-year whole-sample average for food consumption is 2,646 Pakistani rupee (in terms of 1986 rupee, when 1 rupee = US$.063).

The test results for different districts in the Pakistani data are presented in Tables I to IV. In each table, the first row reports the baseline results in which γ is restricted to be equal across all villages in the sample, and full risk sharing is assumed within each village by restricting $\phi(t)$ to be equal across the households in the same 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. The point estimate of γ is positive and statistically significant in all districts except for Dir. The point estimate for Dir is negative, but it is not significantly different from zero. Because the standard error for γ is much larger for Dir than 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. The degree of freedom for the C test in each row is the difference between the degree of freedom of each row's J test and that of the first row. In the second row, we impose the restriction that $\phi(t)$ is the same *across* villages, an implication of full risk sharing *across* villages. We find overwhelming evidence against the restriction for each district from the C statistic reported in the second row. The third row reports the variable addition test results. A significant coefficient for income change indicates rejection of full risk-sharing for that village. We do not reject this hypothesis for most villages. The exceptions are Village 1 in Table I and Village 22 in Table III.[10] The C test of this row, on the other hand, tests the joint hypothesis that all the coefficients on the income changes are equal to zero. The C test does not reject it, and hence the null hypothesis of full risk sharing within villages.

The fourth row reports the results when γ is allowed to be different across villages. The C test does not reject the restriction that γ is the same across villages. The fifth and sixth rows report the results when γ 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

[9] The results based on total consumption are similar; see Zhang and Ogaki (2000).

[10] In the variable addition test, we also used the income net of remittance, which could rise and help smooth consumption in bad times. However, the test results are not qualitatively different. The income in ICRISAT data is already net of remittance, and includes crop profit, labor income, profit from trade and handicrafts, and profit from animal husbandry. See Table A.I. in Townsend (1994).

TABLE I

GMM RESULTS FOR FOOD CONSUMPTION: FAISALABAD (IFPRI-PAKISTAN)

Risk Sharing	γ	Coeff. $\Delta y_1(t+1)^b$	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	C^c
Within	1511	...	...	...	...	...	...	25.7	...
Vil.	(124)[a]							(.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	...[d]	...	...	...	...	...	...	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)

[a] 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-values and degrees of freedom, respectively. Since each village has 6 orthogonality conditions, the total number of orthogonality conditions for the baseline estimation is 36. There are 13 parameters estimated, 12 $\phi(t)$'s and γ. The degree of freedom is therefore 23.

[b] The subscript of income difference term denotes village identification number, e.g. 1 for Vil. 1.

[c] J is a χ^2 statistic, and C is a likelihood-ratio type statistic.

[d] The subsistence estimates for Villages 1 to 6, in the fourth row, are (with standard errors in parenthesis) as follows: 1851 (215), 1467 (240), 1131 (803), 1326 (335), 1738 (206), −512 (2411), respectively.

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 constant RRA. 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 decreasing RRA is ignored.

Table V presents the results for the ICRISTA data. The test results in this table are consistent with those obtained in the IFPRI data set. The first row reports the baseline results, where the J test does not reject the null hypothesis of full risk sharing at any conventional significance level. The point estimate of γ is positive and statistically significant. As in the Pakistani data, the results reported in the last two rows indicate that one can find evidence against full risk sharing when decreasing RRA is ignored.

5. CONCLUSIONS

In this paper, we have tested the full risk sharing hypothesis while taking into account the effect of estimating a parameter that 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. It is not surprising that we have two villages as exceptions, for Townsend (1995) finds that different villages in low-income countries have strikingly different institutional arrangements to cope with risk. We, however, find strong evidence against risk sharing across villages in both data sets. These results confirm the following intuition drawn from the information economics literature. Incomplete risk sharing can be caused by moral hazard

TABLE II

GMM RESULTS FOR FOOD CONSUMPTION: ATTOCK (IFPRI-PAKISTAN)

Risk Sharing	γ	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	1867	...	...	...	...	...	...	...	...	37.8	...
Vil.	(117)									(.187, 31)	
Across	1820	...	...	...	...	...	...	...	...	556.9	519.1
Vil.	(77)									(.000, 45)	(.000, 14)
Within	1749	.0214	.0013	.0785	−.0304	.0221	.0257	−.0618	.0079	28.1	9.7
Vil.	(148)	(.0423)	(.0284)	(.0472)	(.0163)	(.0606)	(.0211)	(.0501)	(.0203)	(.212, 23)	(.290, 8)
Within	...[a]	...	...	...	...	...	...	...	...	24.4	13.4
Vil.										(.441, 24)	(.063, 7)
Within	0	...	...	...	...	...	...	...	...	122.0	82.2
Vil.										(.000, 32)	(.000, 1)
Within	0	.0525	.0437	.0151	−.0573	.0199	.0509	−.0736	.0387	57.8	64.2
Vil.		(.0397)	(.0184)	(.0161)	(.0135)	(.0640)	(.0171)	(.0557)	(.0140)	(.000, 24)	(.000, 8)

[a] The subsistence level estimates for Villages 7 to 14, in the fourth row, 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. See also the notes of Table I for the explanation of the results in this table.

TABLE III

GMM RESULTS FOR FOOD CONSUMPTION: BADIN (IFPRI-PAKISTAN)

Risk Sharing	γ	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	...[a]	...	...	...	...	...	...	...	...	...	...	...	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)

[a] The subsistence level estimates for Villages 21 to 39, in the fourth row, 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. Please also see the notes of Table I for the explanation of the results in this table.

TABLE IV

GMM RESULTS FOR FOOD CONSUMPTION: DIR (IFPRI-PAKISTAN)

Risk Sharing	γ	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	...[a]	...	...	...	...	...	...		
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)

[a] No convergence obtained for the fourth row. Please also see the notes of Table I for the explanation of the results in this table.

TABLE V

GMM RESULTS FOR FOOD CONSUMPTION
(ICRISAT-INDIA)

Risk Sharing	γ_A[a]	γ_S[a]	γ_K[a]	Coeff. $\Delta y_A(t+1)$	Coeff. $\Delta y_S(t+1)$	Coeff. $\Delta y_K(t+1)$	J	C
Within	237.3	237.3	237.3	...	...	...	46.2	...
Vil.	(15.2)						(.507, 47)	
Across	269.7	269.7	269.7	...	...	...	1239.4	1193.2
Vil.	(14.0)						(.000, 63)	(.000, 16)
Within	238.4	238.4	238.4	.024	.011	.005	41.3	4.88
Vil.	(16.1)			(.013)	(.011)	(.016)	(.589, 44)	(.181, 3)
Within	237.9	233.1	240.5	...	...	...	46.1	.03
Vil.	(21.4)	(30.3)	(30.6)				(.425, 45)	(.985, 2)
Within	0	0	0	...	...	...	114.5	68.3
Vil.							(.000, 48)	(.000, 1)
Within								
Vil.	0	0	0	.012	.033	.012	102.3	12.2
				(.011)	(.010)	(.016)	(.000, 45)	(.007, 3)

[a] Subscripts A, S, and K denote Aurepalle, Shirapur, and Kanzara, respectively. Please also see the notes of Table I for the explanation of the results in this table.

or adverse selection, which are in turn caused by private information. It is more likely to be present in a large community, or in an economy with complicated production technologies. Without private information, various arrangements can be used by the community to share risk even if financial markets are not well developed. 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, especially those far away from each other, is less likely to occur because private information may become a problem.

When we restrict the RRA coefficient to be constant, our tests replicate the well-known results of rejecting the full risk sharing hypothesis even within villages in both data sets, except for one district, Dir, in the Pakistani data. However, except for this district, our tests always reject this restriction in both data sets. Thus, this paper shows that misleading results may be obtained when decreasing RRA is ignored in testing the 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 constant RRA or increasing RRA, the test results based on these preference specifications need to be interpreted with caution.

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APPENDIX

The IFPRI data used in this paper cover the period from April, 1986 to September, 1989. Although the original survey covered at least six years starting with 1986, the data set does not contain consumption data for the fourth year. Hence we use the data up to 1989. 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, various revenues, and nonfood expenditures, and so forth. The female questionnaire had around 120 questions, and mainly included demographic and food consumption information. 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 in footnote 12 in Townsend (1994). In each round of the survey, the status of each member was recorded: present, traveling, or moved to a new household. This information is incorporated into the calculation of the household size. 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 IFPRI are used in the empirical analysis. The income measure includes nine subcategories: rental earnings in crops, net crop profits, farm wage income, nonfarm income, net livestock profits, returns to capital, remittances, pension, and *zakat* (private loan association). Since data on total consumption are not readily available, we test risk sharing for food consumption. Assuming that food consumption is separable from other consumption categories, the model in Section 2 applies to food consumption. We calculate a village-specific food price index using the food prices provided by IFPRI. Then we use the index to deflate food consumption and income to obtain real food consumption and real income. 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. The data for Village 52 and Villages 15 to 20 are missing. We exclude from our sample 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.

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