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This article develops a method for testing the risk-sharing hypothesis (RSH) against various versions of the permanent income hypothesis (PIH) while allowing for heterogeneity in risk preferences across households. Using 1-year and longer differences in household total nondurable consumption data from Indian villages, we find evidence that favors the RSH over the PIH at the village level.

KEY WORDS: Consumption smoothing; Generalized method of moments; Panel data; Permanent income hypothesis; Risk-sharing hypothesis.

1. INTRODUCTION

Many studies, whether based on household-level or on aggregate-level data, have documented that consumption is less volatile than income both over time and across households. This is known as consumption smoothing; that is, consumption tends to smooth out fluctuations in income. Among the different theories proposed to explain this important phenomenon are the prominent life cycle-permanent income hypothesis (PIH) and the relatively new risk-sharing hypothesis (RSH). The smoothing of consumption studied by the PIH is sometimes called intertemporal smoothing, because it focuses mainly on consumption variation over time. The form of smoothing scrutinized by RSH, on the other hand, can be appropriately called cross-sectional smoothing, because it emphasizes the effect of resource pooling by agents through formal and informal mechanisms in coping with risk. A natural question then is which of these two smoothing mechanisms is the dominant one behind the volatility pattern that we see in the data. Many authors have investigated each of these two hypotheses separately, but so far only a few have considered both of them altogether. Two prominent exceptions are Attanasio and Davis (1996) (A&D henceforth) and Hayashi, Altonji, and Kotlikoff (1996) (HAK henceforth). We join their efforts and develop a method to test RSH against PIH in panel data. To give each mechanism an equal opportunity, we also test the null hypothesis of the Friedman PIH as formulated by Hall and Mishkin (1982) against RSH. Conceptually, if RSH is true, then the implied consumption Euler equations will be identical to the PIH Euler equations for a household. Therefore, we cannot test the null hypothesis of RSH against PIH Euler equations. However, we can test it against some other implications of PIH (e.g., whether or not idiosyncratic income shocks affect consumption).

The two articles cited earlier report evidence against the RSH. The test of HAK (1996) relies on the crucial assumption that the rational expectations forecast error of individual consumption is the sum of a macro component and an idiosyncratic component, and the idiosyncratic component is not correlated with the variables in the current information set across the households. We refer to this assumption as the error decomposition assumption. This assumption has been used in many articles, sometimes implicitly, to overcome the Chamberlain (1984) critique, and it allows HAK to turn the time series orthogonality conditions implied by the Hall (1978) martingale version of PIH into cross-sectional moment conditions. The latter is then used by HAK as the alternative hypothesis in testing. The RSH null hypothesis is rejected under this alternative using the Panel Studies of Income Dynamics (PSID) data. A&D (1996), using a test similar to that of Cochrane (1991) on synthetic panel data drawn from the Current Population Survey (CPS) and Consumer Expenditure Survey (CEX), found that over a longer time horizon, cohort average consumption growth is positively correlated with the growth of the cohort average relative wage rates. This is consistent with PIH, but not with RSH.

Our method differs from theirs in two ways. First, in light of the evidence reported by Ogaki and Zhang (2001) (O&Z henceforth), it is important to allow for decreasing relative risk aversion (DRRA) in testing RSH. Thus the natural starting point of our method is the framework used by O&Z. Second, we use various versions of the PIH as the alternative hypothesis in our tests. The Friedman (1957) version of the PIH is that consumption is a linear function of permanent income with an additional error term (temporary consumption). Such a consumption function can be derived from an optimization problem that assumes a quadratic utility function and equal rates of interest and time preference (see, e.g., Hall and Mishkin 1982; Deaton 1992, p. 82). We refer to this version of the PIH as the Hall and Mishkin formulation in this article, although the two assumptions that these authors used are not necessary for the derivation of the Friedman PIH. On the other hand, as was shown by Hall (1978), an intertemporal choice model featuring a household that borrows and lends at equal rates without any constraint can also imply consumption smoothing over time. The smoothing holds as long as the precautionary saving is not large enough to reverse intertemporal smoothing of consumption. We consider this form of the PIH modified with DRRA preferences and the Hall and Mishkin formulation in our tests.

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However, our test does resemble those of HAK (1996) and A&D (1996) in the sense that we also pay particular attention to testing with longer-differenced data. This is because, as those authors pointed out, tests based on 1- or 2-year changes in consumption and wage rates can be misleading: short-term changes in wage rates might be dominated by transitory or anticipated changes, or by measurement errors.

Using the household total nondurable consumption and income data from the International Crops Research Institute in Semi-Arid Tropics (ICRISAT) dataset, we find no evidence against RSH at the village level when allowing DRRA and find evidence against PIH. In other words, the smoothing of consumption can be explained by RSH combined with the hypothesis that households exhibit DRRA. At least three other articles also studied the consumption smoothing in ICRISAT villages using both RSH and PIH. Lim (1993) uncovered that the amount of consumption smoothing found in the data is larger than that implied by PIH. Ligon (1998) examined the fit of PIH and RSH against a private information model and found that the latter seems to garner the most support from the data. Lim and Townsend (1998) analyzed the household asset data and found little support for RSH and PIH. However, these authors did not consider the effect of DRRA on consumption growth in their complete markets model.

The ICRISAT dataset is suitable for our purpose for two reasons. First, although it is a short panel dataset, it is nevertheless long enough to enable us to distinguish RSH from PIH. A&D (1996, p. 1252) reported finding sizable and statistically significant departures from the RSH for differing intervals of 3 or more years, but not for differing intervals of 1 year and 2 years. HAK (1996) also found evidence against RSH in one of the PSID panels that spans 1985–1987. Therefore, if PIH is really behind consumption smoothing, then 3 or 4 years of data may be sufficient to expose it. Because the ICRISAT total consumption data we use cover 6 years, it should be long enough for sorting out the two competing theories on consumption smoothing.

Second, the use of food data in testing the RSH requires the assumption that food and nonfood consumption are separable in the utility function. But Atkeson and Ogaki (1996) rejected the separability assumption at the 8% significance level in the ICRISAT data. In addition, Attanasio and Weber (1995) argued that food consumption cannot replace total nondurable consumption in studying intertemporal consumption behavior. The ICRISAT dataset contains both food and nonfood consumption data, so it is more desirable than datasets in which only food data are available, such as PSID.

The article is organized as follows. In Section 2 we review the theory of consumption risk sharing and its testable implications under hyperbolic absolute risk aversion (HARA) preferences. In Section 3 we explain our tests for the null hypothesis of RSH against different versions of PIH and those for the null hypothesis of PIH with quadratic preferences against RSH in short panel data. We present the empirical results in Section 4 and conclude the article in Section 5. We describe the ICRISAT data used in our empirical analysis in the Appendix.

2. THE MODEL

In this section we derive the dynamic implication of DRRA on consumption growth to make the discussion more relevant for the study of longer changes in consumption. Consider an economy with potentially $H$ households that will ever exist. These households participate in a risk-sharing pool, in which a Pareto-efficient sharing rule is carried out by a social planner. The sharing rule is established at period $t = 0$ and lasts until $T$. Let $s(t) \in \{1, 2, \ldots, S\}$ denote the state of the world in period $t$. The history of the economy can then be denoted by the vector $\mathbf{e}(t) = \{s(0), s(1), \ldots, s(t)\}$. Let household $h$, $h = 1, 2, \ldots, H$, have time and state separable utility $u(C_h(t, \mathbf{e}(t)))$, where $C_h(t, \mathbf{e}(t))$ is the per-person–adult equivalent consumption in the history $\mathbf{e}(t)$. Let $\beta$ denote the common discount factor in the economy, and let $\Pr(e(t)|e(0))$ denote the common conditional probability of $\mathbf{e}(t)$ given $e(0)$. The objective of the risk-sharing pool is to maximize the weighted lifetime discounted utility of the participating members. Therefore, the optimal allocation of consumption among the members can be characterized by the following social planner problem:

$$
\max_{\mathbf{C}_e} \sum_{h=1}^{H} \lambda_h \left[ \sum_{t=0}^{T} \beta^t \Pr(e(t)|e(0)) u(C_h(t, \mathbf{e}(t))) \right],
$$

subject to the resource constraint, for every $t$ and every $\mathbf{e}(t)$,

$$
\sum_{h} C_h(t, \mathbf{e}(t)) \leq C_a(t, \mathbf{e}(t)),
$$

where $\lambda_h$ denotes the welfare weight assigned to household $h$ and is positive. The expression inside the square brackets in (1) is the lifetime utility function of household $h$, and $C_a(t, \mathbf{e}(t))$ is the aggregate consumption good available to the pool at $t$ in the history $\mathbf{e}(t)$. Here we assume that the utility function is common across all households. Given the welfare weights and the probability distribution of possible histories, the solution to this problem (i.e., the consumption sharing rule) depends only on $C_a(t)$. This solution is known as the mutuality principle (see Gollier 2001, chap. 21). For ease of exposition, $\mathbf{e}(t)$ is henceforth suppressed wherever possible. In addition, it can be shown that household $h$’s consumption depends on $t$ and $\mathbf{e}(t)$ only through $C_a$. Therefore, we let $C^*_h(C_a(t))$ denote the sharing rule.

Built on theorem 5 of Wilson (1968), the proposition of O&Z (2001) says that the static consumption growth of rich households fluctuates more than that of poor households, if RRA decreases with wealth. In other words, due to DRRA, the value of $(dC^*_h/dC_a)/C^*_h$ for a rich household is higher (lower) than that of a poor household, when aggregate consumption $C_a$ increases (decreases) in the comparative static sense. This static implication of DRRA can be extended as follows to dynamic settings where $C_a(t)$ varies with time. Suppose that $C_a(t)$ is higher than $C_a(t)$. Under DRRA, $(dC^*_h/dC_a)/C^*_h$ is higher for rich households for any value of $C_a$ between $C_a(t)$ and $C_a(t + 1)$. Because the consumption growth rate is a monotone function of $\ln C^*_h(t + 1) - \ln C^*_h(t)$ and

$$
\ln C^*_h(t + 1) - \ln C^*_h(t) = \int_{C_a(t)}^{C_a(t+1)} \frac{dC^*_h}{dC_a} \frac{1}{C^*_h} dC_a,
$$

we have

$$
\frac{dC^*_h}{dC_a} = \frac{\ln C^*_h(t + 1) - \ln C^*_h(t)}{\int_{C_a(t)}^{C_a(t+1)} \frac{1}{C^*_h} dC_a}.
$$


it will be higher for a rich household if DRRA holds. Therefore, the consumption growth rate for a rich household is higher. But if \( C_a(t + 1) \) is lower than \( C_a(t) \), because under DRRA \( (dC_h/dC_a)/C_h^w \) is lower for a richer household for any value of \( C_h \) between \( C_a(t + 1) \) and \( C_a(t) \), the consumption growth rate decreases with wealth. If instead IRRA holds, then we just need to reverse the foregoing conclusions. Under CRRA, as is well known in the literature, the consumption growth rate is identical for all households in the pool. Therefore, the testable implication of the full RSH clearly depends on what is assumed about RRA. To the extent that our theoretical discussion so far is rather general—in the sense that it does not depend on the functional form of the expected utility—current tests of the complete RSH seem to share a common potential problem: By ignoring the possibility of DRRA, their test results may not be valid if CRRA or IRRA does not correctly characterize households’ risk preferences.

We assume that the common utility function across households is

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

which implies that

\[\text{RRA} = \frac{\alpha}{1 - \gamma/C_h}\]  

where \( \gamma \) is the constant preference parameter that governs whether the RRA coefficient increases or decreases with the level of consumption. A positive (negative) \( \gamma \) implies that the RRA coefficient decreases (increases) with consumption. A positive \( \gamma \) is usually called subsistence consumption in the literature (see Chatterjee and Ravikumar 1999 and references therein for implications of a positive subsistence parameter on growth and distribution). Sethi and Presman (1997) studied the behavior of RRA implied by HARA utility in consumption/investment problems that allow for the possibility of bankruptcy. Zimmerman and Carter (1997) simulated an economic model in which poor consumers decide to hold less-risky assets than rich ones because of a positive subsistence constraint.

With this utility function, the solution to the model is

\[C_h(t) = \frac{\gamma^{1/\alpha} (C_a(t) - H\gamma) + \gamma}{\sum_k \lambda_k^{1/\alpha}}\]  

Hence consumption growth for household \( h \) over a \( j \)-year period is

\[
\frac{\Delta C_h(t + j)}{C_h(t)} = \frac{\Delta C_a(t + j)}{(C_a(t) - H\gamma) + \gamma \sum_k \lambda_k^{1/\alpha}/\lambda_h^{1/\alpha}}.
\]  

where \( \Delta C_h(t + j) = C_h(t + j) - C_h(t) \) and likewise for \( \Delta C_a(t + j) \). If \( \gamma \) is a positive constant so that DRRA holds, then when \( C_a \) increases (decreases), the consumption growth rate of a poor household will be lower (higher) than that of a wealthy one because of the lower welfare weight assigned to the former. Here a poor household is one with a low welfare weight and therefore a low consumption level, because the welfare weight depends positively on a household’s wealth level. Obviously, these results are consistent with our aforementioned theory based on a general utility function.

Equations (5) and (6) both involve the unknown welfare weights in ways that are not suitable for empirical analysis. To estimate \( \gamma \) and test the model of full risk-sharing in this article, we use another implication of the model—the intertemporal first-order condition for the social planner problem in our empirical analysis, which imply that for any state of the world,

\[
\frac{C_h(t + j)}{C_h(t)} - \gamma = \phi(t, t + j),
\]  

\[j = 1, 2, \ldots, T - 1, t = 1, 2, \ldots, T - j, \]  

where \( \phi(t, t + j) = [\beta/\Pr(t + 1)\mu(t)/\mu(t + 1)]^{1/\alpha} \) and \( \mu(t) \) is the Lagrange multiplier associated with the resource constraint (2). Equation (7) should hold for each household in the risk-sharing pool. It says that in any state of the world, \( C_h - \gamma \) should grow at the same rate for all households regardless of the time horizon.

To summarize, the testable implication of the full RSH usually depends on what is assumed about households’ RRA. In the HARA framework that we adopt, the CRRA assumption implies that consumption growth should be equalized across households with complete markets. In contrast, the DRRA (IRRA) assumption implies that consumption growth will be positively (negatively) correlated with the wealth level even there is full risk-sharing. The DRRA assumption also implies that with complete risk-sharing, the growth rate of consumption net of subsistence level will be equalized across households. These implications are still valid if we generalize our model to include nontraded goods such as leisure, provided that they are additively separable in the utility function. However, such separability is not necessary for (7) to hold if these goods, like consumption, are freely transferable by the social planner across households, as Cochrane (1991, p. 965) pointed out.

In the next section we consider tests of RSH against various versions of PIH.

3. THE TESTS

In this section we develop a method to test the implications of RSH against various versions of PIH. It extends the method of O&Z (2001) for testing RSH against the Keynesian consumption function with HARA preferences. We first describe how to test the RSH over 2-year and longer time periods. Then we show how the PIH with quadratic or HARA preferences can be nested in the same consumption equation for RSH. We not only test the RSH null against the PIH alternative hypothesis, but also test a version of the PIH null against the RSH alternative, to give each hypothesis an equal opportunity.

3.1 Tests of the Risk-Sharing Hypothesis

Consumption is assumed to be measured with additive error with mean 0 across households,

\[C_h^{\text{emp}}(t) = C_h(t) + \xi_h(t),\]  

where \( C_h^{\text{emp}}(t) \) is measured consumption in per-male–adult equivalent terms and \( \xi_h(t) \) is the measurement error.Combining (7) and (8), we obtain

\[C_h^{\text{emp}}(t + j) = \gamma[1 - \phi(t, t + j)] + \phi(t, t + j)C_h(t) + \nu_h(t, t + j),\]  

where \( v_h(t, t+j) = \xi_h(t+j) - \phi(t, t+j)\xi_h(t) \).

Because \( C^m_h(t) \) is correlated with \( v_h(t, t+j) \), an instrument variable method must be used to estimate the model. In addition, there is a nonlinear restriction between the parameters in the model, so we use Hansen’s (1982) generalized method of moments (GMM) to estimate the model. Under the null hypothesis of full risk-sharing, one good instrument for consumption level is a household’s wealth status, which is directly linked to the household’s welfare weight in (5). The time average of household income in per-adult-equivalent terms can be used as a proxy for wealth status. Let \( \bar{y}_h(t, i, t+k) \) be the average income between years \( t+i \) and \( t+k \) for household \( h \). Now assume that \( \xi_h(t) \) is uncorrelated with household \( h \)’s average income and its measurement error. Let \( Z_h(t) = (1, \bar{y}_h(t, i, t+k)) \) be the vector of instrumental variables. The first-stage regression results based on these two instruments, reported in Table 2 (see Sec. 4), indicate that these two instruments are reasonably relevant for \( C^m_h(t) \).

Let \( \psi_j = (\phi(1, j+1), \phi(2, j+2), \ldots, \phi(T-j, T), \gamma) \) be the \((T-j+1)\)-dimensional vector of unknown parameters, and let \( \psi_{j,0} \) be the vector for the true values of \( \psi_j \). In addition, let \( f_{h,t,t+j} \) be the two-dimensional vector

\[
\text{f}_{h,t,t+j} = Z_h \left[ C^m_h(t+j) - \phi(t+j)C^m_h(t) - \gamma + \gamma \phi(t,j+j) \right]
\]

and

\[
f_{h,j}(\psi_j) = (f_{h,1,1+j}, \ldots, f_{h,T-j,T})', \quad j = 1, \ldots, T-1.
\]

Under the null hypothesis of full risk-sharing and the assumption that consumption measurement errors are not correlated with income and its measurement errors, we have the following 2 · (T - j) orthogonality conditions for \( j = 1, 2, \ldots, T-1 \):

\[
E_H(f_{h,j}(\psi_{j,0})) = \lim_{N \to \infty} \frac{N}{N} \sum_{h=1}^{N} f_{h,j}(\psi_{j,0}) = 0, \quad (11)
\]

where \( E_H \) is the expectation operator over the population of households. Hansen’s test of overidentifying restrictions (the J test) can be used to test (11). Our GMM procedure will no longer be valid if the assumption on consumption measurement errors does not hold. But the violation of such an assumption should also lead to false rejection of the RSH even if it does hold.

3.2 Tests of the Permanent Income Hypothesis

The other type of test is directed at the PIH. HAK (1996) were particularly concerned about the power of risk-sharing tests against the martingale version of PIH. They found overwhelming evidence against full risk-sharing under such an alternative hypothesis. A&D (1996) also found that tests based on longer differences in log consumption (i.e., from the third difference on) seem to favor Friedman’s (1957) version of PIH. We therefore add the income difference term to (9) to obtain

\[
C^m_h(t+j) = \gamma[1 - \phi(t+j)] + \phi(t+j)C^m_h(t) + \eta \Delta y_h(t+j) + v_h(t, t+j), \quad (12)
\]

where \( \Delta y_h(t+j) = y_h(t+j) - y_h(t) \). Under the null hypothesis of full risk-sharing, \( \eta = 0 \) should hold. This is called the variable addition test by O&Z (2001), where \( j = 1 \) is maintained. We now explain how this equation can be used to test RSH against various versions of PIH for \( j = 1 \) or otherwise.

Under the assumption of quadratic preferences, the first-order condition for the PIH implies

\[
E_h C_h(t+j) = [1 - \Psi(t+1, t+j)]C^* + \Psi(t+1, t+j) C_h(t),
\]

where \( \Psi(t+1, t+j) = (1 + \delta)^j / \prod_{i=1}^{j} (1 + r(j + i)) \), \( \delta \) is the common time preference rate, \( r(j + i) \) is the interest rate between years \( t+i-1 \) and \( t+i \), and \( C^* \) is the bliss point in the quadratic utility function. (The bliss point \( C^* \) is assumed to be a constant for simplicity. We experimented with a time-varying bliss point in our empirical analysis and found that the assumption of constant bliss point cannot be rejected.) After introducing measurement errors as in (8) and rearranging, (13) becomes

\[
C^m_h(t+j) = [1 - \Psi(t+1, t+j)]C^* + \Psi(t+1, t+j) C^m_h(t) - v_h(t, t+j).
\]

The disturbance term consists of both rational expectations forecast errors to consumption level \( \epsilon_h(t+1), \ldots, \epsilon_h(t+j) \) and the measurement errors in consumption,

\[
v_{h,t,t+j} = \frac{1 + \delta}{1 + r(t+j)} \epsilon_h(t+j) - \frac{(1 + \delta)^2}{1 + r(t+j)} \epsilon_h(t+j-1) + \cdots + \Psi(t+1, t+j) \epsilon_h(t+1) - \epsilon_h(t+j) + \Psi(t+1, t+j) \epsilon_h(t).
\]

The difference between (14) and (9) is in \( \Psi(t+1, t+j) \) and the disturbance term. The expectations error decomposition assumption used by HAK (1996) can be invoked for \( v'_{h,j} \) in (14) to set up the cross-sectional orthogonality conditions for the PIH. When the average income \( \bar{y}_h \) is calculated using incomes between years \( t+1 \) and \( t+j \), our test for the null of RSH has power against PIH. This is because under PIH, the rational expectations forecast errors in \( v'_{h,j} \), that \( \epsilon_h \)'s, contains unanticipated changes in income between years \( t+1 \) and \( t+j \), and such shocks cause \( \gamma y_h(t+1, t+j) \) and \( v'_{h,j} \) to be cross-sectionally correlated. So if PIH is true, then we should expect the orthogonality condition (11) to fail. But if RSH is true, then this condition should hold.

We consider three versions of PIH associated with (14). The first of these is the Hall and Mishkin (1982) formulation under the assumption of equal, constant interest rate and time preference rate. The attraction of this formulation is that it delivers a most plausible version of the Friedman (1957) consumption function based on the notion of permanent income. Under this assumption, one testable implication is that \( \Psi(t+1, t+j) = 1 \) for every \( t \) and \( j \). The second version of PIH associated with (14) features constant interest rate and time preference rate, but with \( r \neq \delta \). This is interesting because the assumption of \( r = \delta \) is too stringent and is unlikely to be true in every period in the real world. Under this version of the PIH, \( \Psi(t+1, t+j) \) should be a constant different from 1 over time but changes
with \( j \) by the constant proportion \((1 + \delta)/(1 + r)\). The third version of the PIH associated with (13) allows the interest rate to be time-varying, which is more realistic. In this case \( \Psi(t + 1, t + j) \) should change with \( t \) and \( j \). For all of these three cases, the estimate for the bliss point \( C^* \) should be higher than the highest level of consumption in the data. Although measurement errors and sampling errors in estimates might complicate this comparison, it is reasonable to assume that the estimate for the bliss point \( C^* \) should be higher than the average level of consumption. In addition, the income change term can also be added to (14) to pick up the effect of income shocks in the expectations errors in \( v_h(t, t + j) \). This means that both RSH and these three versions of PIH are nested in (12).

We now consider the PIH with DRRA preferences. This is important because the difference in preferences used above in testing between RSH and PIH may unduly affect the test result. We start from the general PIH Euler equation with a stochastic interest rate for household \( h \),

\[
 u'(C_h(t)) = \frac{1}{1+\delta} E_t \left[ (1+r(t+1))u'(C_h(t+1)) \right],
\]

(15)

where \( u'(C_h(t)) \) is the marginal utility of consumption at time \( t \). Under the assumption of DRRA preferences, when the time horizon is extended to \( t+j \), (15) implies

\[
 C_h(t+j) - \gamma = \Psi_h(t+1, t+j),
\]

(16)

where

\[
 \Psi_h(t+1, t+j) = \left[ \Psi(t+1, t+j) \prod_{i=1}^{j} (1 + \xi_h(t+i)) \right]^{-1/\alpha}
\]

and \( \xi_h(t+i) \) is household \( h \)'s rational expectations forecast error to the marginal utility growth adjusted by \((1+r(t+i))/(1+\delta)\) for period \( t+i \). After we introduce the consumption measurement errors into (16), it can be rewritten as

\[
 C_h^m(t+j) = \gamma [1 - \Psi_h(t+1, t+j)] + \Psi_h(t+1, t+j) C_h^m(t) + v_h(t, t+j),
\]

(17)

where \( v_h(t, t+j) = \xi_h(t+j) - \Psi_h(t+1, t+j) \xi_h(t) \) is the disturbance term. Comparing (9) for RSH with (17) for PIH, we find that the only differences are between \( \phi(t, t+j) \) and \( \Psi_h(t+1, t+j) \) and between \( v_h(t, t+j) \) and \( v_h^m(t, t+j) \). If this version of PIH is true, then (11) should fail. This is because we force \( \Psi_h(t+1, t+j) \) to be the constant (across households) \( \phi(t, t+j) \) in testing the null hypothesis of RSH. As a result, the disturbance term \( v_h^m(t, t+j) \) will include both \( C_h^m(t) \) and \( \Psi_h(t+1, t+j) \) if (17) holds. So the instrument \( v_h, \) the time average of income between year \( t+1 \) and \( t+j \), will be correlated with \( C_h^m(t) \) or the income shocks in \( \Psi_h(t+1, t+j) \) across households. Therefore, our test has power against RSH when this version of PIH holds. Finally, when imposing RSH in (17), the unanticipated income change of a household, being correlated with the expectations errors in \( \Psi_h(t+1, t+j) \) in the disturbance term, will be significant in explaining \( C_h^m(t+j) \) under this version of PIH. So we can add the income change term to (17) to try to capture these changes. It is clear that (12) also nests this equation. Therefore, if PIH with DRRA preferences is true (and RSH is not true), then the GMM test of overidentifying restrictions should reject (11), and the \( \eta \) estimate should be significant if there are unanticipated income changes.

The income change in the short run may be largely anticipated changes or mainly measurement errors. A permanent income consumer will not respond to such changes in consumption decisions. However, the income change in the longer time period may contain more surprises. Such unexpected changes will induce the permanent income consumer to change his level of consumption. Therefore, we should expect the \( \eta \) estimate to be significant for larger \( j \) if PIH holds. A third possibility is that \( \Delta y_h(t+j) \) in (12) may consist of both unanticipated and anticipated changes, regardless of the time horizon \( j \). To detect the alternative PIH in such a context, we should ideally replace \( \Delta y_h(t+j) \) with the unanticipated changes in income. But due to the somewhat short time dimension of the data used in our empirical analysis, it is impossible to decompose the income change into these two components. Thus we must stick to \( \Delta y_h(t+j) \) in estimating (12) and can obtain only a downward-biased estimate of \( \eta \) (relative to its true value \( \alpha \)) in such a case. This bias arises because, from the standpoint of estimation, the anticipated income change can virtually be viewed as a measurement error in the unanticipated income changes. This bias notwithstanding, the probability limit of the \( \eta \) estimate would still be positive.

To summarize, testing RSH against the PIH is conducted at two levels. At the pooled cross-village level, we use the \( J \) test of Hansen (1982) to test whether \( \phi(t, t+j) \) is constant across households for every \( t \) and \( j \) in every village, which is the major implication of RSH. To test RSH against the Hall and Mishkin (1982) formulation of the Friedman (1957) PIH, we test whether \( \phi(t, t+j) = 1 \) holds jointly for all of the villages by imposing the restriction \( \phi(t, t+j) = 1 \) in estimation for every \( j \). At the village level, we test whether or not the \( \eta \) estimate is significantly different from 0 for each village, whether the bliss point estimate is high enough to be interpreted as the bliss point, and whether the \( \phi(t, t+j) \) estimates are constant over \( t \) or different from 1, using the \( t \)-statistic for each estimate.

We now explain how our method can be used to test the PIH with quadratic preferences as the null hypothesis against the alternative RSH. The instruments we use in our tests, \( Z_h = (1, \bar{y}_h(i, t+k)) \), are also valid instruments under the PIH, provided that \( \bar{y}_h \) is calculated with the lagged incomes of a household (e.g., \( i = 1-t \) and \( k = 0 \) for year \( t \) and \( i = 1-t \) and \( k = 1 \) for year \( t+1 \)). The average income obtained in this manner is not correlated with the expectations error in the disturbance terms of (14) and (17) across households; however, it is correlated with household consumption in the cross-section, because it is correlated with the permanent income. Furthermore, the unknown parameters in (14) and (17) can now be consistently estimated with this choice of instruments, if PIH is true. Nonetheless, these instruments remain valid under the alternative RSH as well. Therefore, in this context we should not use the \( J \) test to discriminate between the two hypotheses. Instead, we should use the estimates of \( \Psi(t+1, t+j), C^*, \) and \( \eta \) in (14) to distinguish one theory from the other. For example, if the PIH with quadratic preferences and equal discount rate and interest rate is true, then we should find that the \( \Psi(t+1, t+j) \) estimates are not significantly different from 1, the \( C^* \) estimate is high enough to be interpreted as the bliss point, and the \( \eta \) estimates are significant in a longer-differencing test.
3.3 Further Discussion

As is well known in the literature, a positive and significant \( \eta \) estimate could be consistent with both PIH and RSH with nonseparability between consumption and leisure in the utility function. So how do we tell which hypothesis the data support when we find such an estimate for \( \eta \)? We just need to look at the \( J \) test result of (11) or (12) to tell RSH from the PIH with DRRA preferences, because if it does not reject them, then the restriction of constant \( \phi(t, t + j) \) across households—

the major implication of RSH—is not rejected. Note that under RSH, \( \phi(t, t + j) \) reflects the shocks to aggregate consumption growth and thus should be time-varying. This feature allows us to discriminate RSH from the first two versions of the PIH with quadratic preferences even if the \( \eta \) estimates are significant. Moreover, the size of the \( \gamma \) estimate relative to the actual maximum or average level of consumption can be used to distinguish RSH from PIH with quadratic preferences and stochastic interest rate, as explained in Section 3.2. On the other hand, Ham and Jacobs (2001) reported that the assumption of separable consumption and leisure in the preferences does not affect the evidence against the full risk-sharing in the PSID data; but these authors did not consider the possibility of DRRA preferences.

The test based on longer difference (i.e., higher \( j \)) should be more powerful against the PIH than that based on 1-year difference, because it provides a more ideal setting for the permanent income model to hold by using the longer changes in income and consumption. A caution is in order here, however: Using longer differences means that we will be using fewer data points to estimate \( \gamma \), which usually causes efficiency loss and thus may render the estimate of \( \gamma \) inaccurate. This in turn makes the weighting matrix used in the iterative GMM procedure for the higher-difference testing less efficient than the 1-year difference case. In principle, this problem can be fixed by using the estimates from 1-year difference testing as initial parameter values in the baseline estimation for the higher-difference testing, and constructing the initial weighting matrix with these estimates. Depending on the size of the efficiency loss, however, the fix may not work. An alternative solution is to restrict \( \gamma \) in the longer difference by taking the \( \gamma \) estimate obtained in the first-difference baseline estimation. Imposing such a restriction usually changes the asymptotic distribution of the estimator for other parameters. But a sequential estimation procedure is available to handle this issue (see, e.g., Ogaki 1993, sec. 4.1). Therefore, we also pursue it when the first solution does not yield satisfactory results.

4. EMPIRICAL RESULTS

Here we present the empirical results based on the household level ICRISAT data. This dataset has been used to study consumption smoothing and risk-sharing models by many authors (see, e.g., Bhargava and Ravallion 1991; Rosenzweig 1988; Rosenzweig and Stark 1989; Rosenzweig and Wolpin 1993; Ravillion and Chauduri 1997; Jacoby and Skoufas 1997). For a description of the data, refer to the Appendix.

Table 1 presents the average real consumption data, including real total consumption, real food consumption, and real nonfood consumption, for three villages—Aurepalle, Shirapur, and Kanzara—for the period 1976–1981. These data are useful for the purpose of interpreting the parameter that corresponds to the subsistence consumption in (9) or (12) for RSH and the bliss point in (14) for PIH.

Table 2 reports the results for the first-stage regression. We regress household per–male–adult equivalent total nondurable consumption on per–male–adult equivalent income by pooling three villages together. The \( R^2 \) of 29.2% supports the relevancy of average income as an instrument. Tables 3 and 4 present test results based on the real total nondurable consumption per male–adult equivalent.

Before moving to Tables 3 and 4, we now briefly explain the test results for over-identifying restrictions based on (11) for \( j = 1, \ldots, 5 \) (not reported). The parameter \( \gamma \) is restricted to be constant across villages, and \( \phi(t, t + j) \) is restricted to be constant across all of the households within the same village in our estimation and testing. (It would be interesting to test whether the baseline result is sensitive to the restriction that \( \gamma \) is constant across wealth classes. However, the small sample size does not accommodate such a splitting of the sample.) The data do not reject these restrictions and the orthogonality between \( \psi_k(t + j) \) and the instruments at conventional significance levels: the \( p \) values for these tests are higher than 40% in all cases. In other words, the null hypothesis of full risk-sharing is not rejected for each \( j \). Note that the test of (11) has power against the alternative hypothesis the PIH with DRRA preferences, according to our discussion in Section 3.2. In addition, about one-half to two-thirds of the estimates for the \( \phi(t, t + j) \)'s are significantly different from 1.

Table 3 presents the results based on (12) for 1-, 3-, 4-, and 5-year differences. In this table the estimates of \( \gamma \) are all positive and statistically significant and are below the average total nondurable consumption and average food consumption reported in Table 1 for each village for the sample period. So

| Table 1. Real Consumption Per–Male–Adult Equivalent: ICRISAT Data |
|----------------|----------------|----------------|----------------|----------------|----------------|
|-----------------|-----|-----|-----|-----|-----|-----|----------|
| Average household total consumption |
| Aurepalle        | 502 | 490 | 544 | 750 | 738 | 660 | 614 |
| Shirapur        | 1,063 | 980 | 749 | 869 | 787 | 664 | 852 |
| Kanzara         | 852 | 847 | 758 | 993 | 937 | 815 | 867 |
| Average food consumption |
| Aurepalle        | 313 | 381 | 408 | 538 | 502 | 423 | 408 |
| Shirapur        | 604 | 555 | 644 | 543 | 623 | 521 | 582 |
| Kanzara         | 490 | 489 | 418 | 578 | 571 | 479 | 504 |
| Average nonfood consumption |
| Aurepalle        | 190 | 101 | 156 | 214 | 240 | 236 | 158 |
| Shirapur        | 337 | 313 | 345 | 352 | 329 | 364 | 235 |
| Kanzara         | 369 | 359 | 353 | 426 | 364 | 345 | 267 |

NOTE: The data reported here are in 1983 Indian rupees.

| Table 2. First-Stage Regression Results for Total Consumption |
|----------------|----------------|----------------|
| \( C_{1t} = \beta_1 + \beta_2 y_{2t} + v_{1t} \) |
| \( \beta_1 \) | 566.02 (20.57) |
| \( \beta_2 \) | .20 (.01) |
| \( R^2 \) | 29 |
| \( F(1,513) \) | 211.53 |
| Sample size | 515 |

between 1975–1977 and 1975–1980 for  j, which is imposed, but results do not vary much after inserting a constant or time-varying intercept term into the equation.

It is reasonable to interpret  as subsistence consumption, but not bliss point. Based on our discussion in Section 2, a positive  implies DRRA. For each panel of the table, we find that the Hansen statistic for testing the overidentifying restrictions, reported at the end of the first row under the letter  , does not reject the orthogonality between the instruments and the disturbance term  in (12) and the restriction that  is constant across households at conventional significance lev-

Table 4. Results for Testing the Hall and Mishkin (1982) Formulation of PIH Against RSH

<table>
<thead>
<tr>
<th>C^a</th>
<th>Time effect</th>
<th>Aurepalle</th>
<th>Shirapur</th>
<th>Kanzara</th>
<th>η_A^c</th>
<th>η_S^c</th>
<th>η_K^c</th>
<th>J</th>
</tr>
</thead>
<tbody>
<tr>
<td>Difference over 1-year period (j = 1)</td>
<td>Ψ^1,2</td>
<td>Ψ^3,4</td>
<td>Ψ^4,5</td>
<td>Ψ^5,6</td>
<td>Ψ^1,2</td>
<td>Ψ^3,4</td>
<td>Ψ^4,5</td>
<td>Ψ^5,6</td>
</tr>
<tr>
<td>467.1</td>
<td>51.9</td>
<td>86.3</td>
<td>-44.7</td>
<td>-1.2</td>
<td>1.07</td>
<td>2.51</td>
<td>1.10</td>
<td>.73</td>
</tr>
<tr>
<td>(43.9)</td>
<td>(14.3)</td>
<td>(53.8)</td>
<td>(36.7)</td>
<td>(32.2)</td>
<td>(24)</td>
<td>(.51)</td>
<td>(.11)</td>
<td>(.15)</td>
</tr>
<tr>
<td>469.7</td>
<td>55.9</td>
<td>94.4</td>
<td>-49.5</td>
<td>-2.4</td>
<td>1.26</td>
<td>2.28</td>
<td>1.19</td>
<td>.72</td>
</tr>
<tr>
<td>(46.3)</td>
<td>(24.5)</td>
<td>(55.2)</td>
<td>(112.4)</td>
<td>(37.2)</td>
<td>(1.11)</td>
<td>(1.39)</td>
<td>(75)</td>
<td>(19)</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>Difference over 3-year period (j = 3)</th>
<th>Ψ^1,5</th>
<th>Ψ^2,6</th>
<th>Ψ^3,6</th>
<th>Ψ^4,5</th>
<th>Ψ^5,6</th>
<th>Ψ^1,5</th>
<th>Ψ^2,6</th>
<th>Ψ^3,6</th>
<th>Ψ^4,5</th>
<th>Ψ^5,6</th>
<th>Ψ^1,5</th>
<th>Ψ^2,6</th>
<th>Ψ^3,6</th>
<th>Ψ^4,5</th>
<th>Ψ^5,6</th>
</tr>
</thead>
<tbody>
<tr>
<td>408.2</td>
<td>108.4</td>
<td>-16.7</td>
<td>2.70</td>
<td>2.00</td>
<td>.94</td>
<td>.84</td>
<td>.95</td>
<td>1.17</td>
<td>.89</td>
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</tr>
<tr>
<td>(56.7)</td>
<td>(73.0)</td>
<td>(65.0)</td>
<td>(.33)</td>
<td>(.22)</td>
<td>(.20)</td>
<td>(.17)</td>
<td>(.16)</td>
<td>(.23)</td>
<td>(60.3)</td>
<td></td>
<td></td>
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<td></td>
</tr>
<tr>
<td>407.0</td>
<td>-583.4</td>
<td>290.5</td>
<td>.22</td>
<td>.25</td>
<td>.25</td>
<td>.41</td>
<td>3.57</td>
<td>-.49</td>
<td>4.6</td>
<td>.87</td>
<td>1.73</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(24.4)</td>
<td>(576.4)</td>
<td>(245.9)</td>
<td>(16.3)</td>
<td>(1.43)</td>
<td>(1.31)</td>
<td>(.39)</td>
<td>(2.07)</td>
<td>(1.26)</td>
<td>(3.81)</td>
<td>(.72)</td>
<td>(1.34)</td>
<td></td>
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<table>
<thead>
<tr>
<th>Difference over 4-year period (j = 4)</th>
<th>Ψ^1,6</th>
<th>Ψ^2,6</th>
<th>Ψ^3,6</th>
<th>Ψ^4,5</th>
<th>Ψ^5,6</th>
<th>Ψ^1,6</th>
<th>Ψ^2,6</th>
<th>Ψ^3,6</th>
<th>Ψ^4,5</th>
<th>Ψ^5,6</th>
<th>Ψ^1,6</th>
<th>Ψ^2,6</th>
<th>Ψ^3,6</th>
<th>Ψ^4,5</th>
<th>Ψ^5,6</th>
</tr>
</thead>
<tbody>
<tr>
<td>423.4</td>
<td>79.9</td>
<td>2.38</td>
<td>.80</td>
<td>.71</td>
<td>1.36</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(85.2)</td>
<td>(104.8)</td>
<td>(.44)</td>
<td>(.26)</td>
<td>(.33)</td>
<td>(24.1)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>
els. Therefore, the RSH is overall not rejected for the three villages. However, because the Hall and Mishkin (1982) formulation of PIH implies that \( \phi(t, t + j) = 1 \) for \( j = 1, 2, \ldots, 5 \), we now check whether the estimates for \( \phi(t, t + j) \) are significantly different from 1. From the first row of each panel, we see that for Aurepalle and Kanzara, the estimates for \( \phi(t, t + j) \), \( j = 1, 3, 4 \), are obviously different from 1 for most of the cases. Meanwhile, the \( \eta_A \) and \( \eta_K \) estimates are not significantly different from 1. Therefore, it is clear that the RSH is not rejected against this version of PIH in these two villages. For Shirapur, the \( \phi(t, t + j) \) estimates do not differ significantly from 1, and none of the \( \eta_S \) estimates is significantly different from 0 at conventional significance levels. (But two estimates of \( \phi(t, t + j) \) in the 1-year difference test of (11) are significantly different from 1 at the 5% level.) These are consistent with PIH when income changes were largely transient or anticipated, or they simply reflect measurement errors in income. So there seems to be some evidence for PIH in Shirapur. However, it is important to note that the estimates of \( \gamma \) in all cases are much lower than the average consumption level reported in Table 1. Therefore, it cannot be interpreted as the bliss point parameter \( C^* \) in (14) of Section 3.2. This suggests that the PIH with quadratic preferences is unlikely to be behind the data. Moreover, because the \( \phi(t, t + j) \) estimates vary over time and the interval of differencing, \( j \), and the restriction of \( \phi(t, t + j) \) being constant across households is not rejected, there is no evidence for any version of PIH outlined in Section 3.2, including PIH with DRRA preferences.

The negative sign of the \( \eta \) estimates for Shirapur and Kanzara could be caused by either measurement errors in income (change) or taste shifters that are swept into the disturbance term \( v_h(t + j) \) but are correlated with income change. In principle, these two problems could also invalidate the inference on the \( \phi(t, t + j) \) estimates. However, to assess how large an impact they create, we can compare the \( \phi(t, t + j) \) estimates reported in Table 3 with those obtained in estimating (9). This is because (9) does not include the term \( \Delta \gamma_t(t + j) \) and thus is not affected by these two problems. To conserve space, the results for (9) are not reported. The \( \phi(t, t + j) \) estimates in Table 4 also can be compared with those in the first row of the first panel of Table 3. We have done such comparisons and found that the two sets of \( \phi(t, t + j) \) estimates are remarkably close for each \( j \). More important, we also found that the statistical significances of these estimates do not change much for each \( j \) whether or not \( \Delta \gamma_t(t + j) \) is included in the equation. Hence we conclude that the inference on \( \phi(t, t + j) \) estimates is valid.

As we explained in Section 3, to further test between RSH and the Hall and Mishkin (1982) formulation of PIH, we can impose the restriction \( \phi(t, t + j) = 1 \) in the estimation and testing (with due attention paid to the \( \eta \) estimates). The test results are presented in the second row of each panel in Table 3. Note that as a result of the restriction \( \phi(t, t + j) = 1 \), the \( J \) statistics become substantially larger than those in the upper rows for \( j = 1, 3, 4 \), and the \( p \) values for these statistics are virtually 0 for these three cases. To test the validity of the restriction \( \phi(t, t + j) = 1 \), we can subtract the \( J \) statistic of the first row of each panel from that of the second row to obtain a chi-squared statistic with degrees of freedom equal to the difference of the degrees of freedom of the two \( J \) statistics involved. This is the so-called “C test” in O&Z (2001). Although we do not report these statistics due to space limitations, it can be easily seen that they strongly reject this implication of PIH for \( j = 1, 3, 4 \). These results reinforce those of the variable addition test. Furthermore, the \( J \) test statistic remains very large after allowing for a time-varying or a constant intercept term in the equation, except for \( j = 4 \), for which the \( J \) statistic drops to such a low level that the rejection of \( \phi(t, t + j) = 1 \) at conventional significance level disappears. This test is necessary because the lack of intercept term in (12) after imposing \( \phi(t, t + j) = 1 \) could potentially explain the rejection (i.e., the high \( J \) statistic) in the second row of each panel from a purely statistical standpoint.

A&D (1996) reported that the empirical evidence favored the PIH when they looked into time horizons longer than 3 years in the synthetic panel data constructed from CPS and CEX. Hak (1996) were also able to reject RSH under the alternative of PIH in one set of their results based on a 1985–1987 panel of PSID data. Our results contrast with theirs in the sense that we do not find much evidence for PIH, and we cannot reject RSH even when we look at a longer time horizon.

We now turn to the last panel of Table 3, which presents the results on 5-year differencing based on both (9) and (12). The reason for presenting both sets will become evident as we move on. First, note that the estimate of \( \gamma \) becomes much smaller than the previous estimates, and it now has such a large standard error that it is insignificantly different from 0. Even when we use the parameter estimates from the 1-year difference baseline estimation as the initial values and construct an initial weighting matrix using these estimates, it is still insignificant. This insignificance may simply reflect the loss of data points because of the use of longer differences. Indeed, consistent with this observation, we find that the standard errors of the \( \gamma \) estimates increase as we move from a 1-year difference to a 5-year difference in estimating (9). Second, we find indecisive evidence regarding our two theories on consumption smoothing. On the one hand, RSH is not rejected, as we see from the \( J \) statistic of the first row of this panel. On the other hand, the restriction of \( \phi(1, 6) = 1 \) is no longer rejected for each village, and at least one of the three \( \eta \) estimates is positive and statistically significant. So one might be led to conclude that PIH seems to be emerging from the data. But it is important to notice that the ambiguity could be due to the insignificant \( \gamma \) estimate, which may in turn affect the estimates of \( \phi(1, 6) \) and the statistical significance of the restriction \( \phi(1, 6) = 1 \). In addition, even if we forget the insignificance for a moment, there are still other interpretations of the same evidence that are consistent with the RSH. For example, the highly significant \( \eta_A \) estimate (for Aurepalle) of .75, obtained when \( \phi(1, 6) = 1 \) is imposed in the estimation, may simply reflect the covariance of income change and preference shifts across households, and hence may not be interpreted as evidence against full risk-sharing, a point made by Cochrane (1991). Another possible reason for the significant \( \eta_A \) estimate is the way in which we handle demographic variables. Attanansio and Browning (1995) have argued that the significance disappears when demographics are included explicitly in the utility function. As for the other two \( \eta \) estimates that are not significantly different from 0, for them to be compatible with the PIH, we need to assume that the income changes over the 5-year horizon were anticipated or transitory,
or that simply they reflected the measurement errors in income. But for such a time horizon, an income change could well reflect changes in permanent income. If this is the case in our data, then the insignificant \( \eta \) estimates are consistent with the RSH.

These ambiguities point to the need for a method that can get around the insignificant \( \gamma \) estimate problem. To this end, we use the sequential estimation procedure alluded to in Section 3.3. Specifically, we restrict \( \gamma \) to be equal to its estimate obtained in the 1-year difference estimation of (9), and adjust the standard errors of the \( \phi(t,t+j) \) estimates accordingly. As a result, the \( \phi(1,6) \) estimate for Aurepalle becomes significantly different from 1 (see the third row of the last panel). Therefore, at least the ambiguity for this village now disappears.

Another way of addressing the aforementioned ambiguities is to test the PIH as the null hypothesis with RSH as the alternative, as mentioned at the end of Section 3.2. Table 4 reports test results along this dimension. The instruments used in these tests are valid under both PIH and RSH. So if PIH with quadratic preferences holds, then we should expect the estimates for \( C^\phi \) to be at least larger than the average consumption level in three villages; otherwise, it cannot be interpreted as the bliss point. But the \( C^\phi \) estimates in Table 4 are not much different from those in Table 3, and are far lower than the average consumption level. The estimates for \( \Psi(t+1,t+j) \) at the 1- and 2-year differencing are time varying and do not support the hypothesis of \( \Psi(t+1,t+j) = 1 \) or being constant, although for 3- and 4-year differencing, the estimates for Shirapur and Kanzara no longer differ significantly from 1. The time effects reported here come from the macro component of the rational expectations forecast errors in (14). Although the time effect for 1977 is significant for \( j = 1 \), the overall pattern is that it becomes insignificant as \( j \) increases. This is at odds with PIH with quadratic preferences but is consistent with RSH, because (12) has no time effect under RSH. In addition, the \( \eta \) estimates are all insignificant at conventional significance levels. Thus we conclude that the null hypothesis of the PIH with quadratic preferences is rejected. The \( J \) tests in this table therefore indicate that the alternative RSH is not rejected, given that the instruments are also valid under RSH.

To summarize, we cannot reject RSH when different forms of PIH are used as the alternative hypotheses. As long as DRRA is taken into account, there is little evidence against RSH within each village.

5. CONCLUSIONS

In this article we have developed an econometric method for testing RSH against various versions of PIH. In formulating RSH, we have allowed the possibility of DRRA. We have conducted our test using first and longer differences of the data to increase its power. The empirical results based on ICRISAT total nondurable consumption data confirm the results in O&Z (2001), which were obtained using food consumption data alone, that these households are characterized by DRRA preferences. This result should be of interest to many economists, because risk preferences play a prominent role in many economic models. Our results shed new light on how preferences should be treated in these models. Our tests support full risk-sharing at the village level for two of the three (and possibly all three) villages. Therefore, it seems that cross-sectional smoothing plays the major role in consumption smoothing, despite the lack of sufficient formal market mechanisms to provide risk-sharing opportunities. This result is most likely due to the nonmarket institutions that can function effectively to fill the gap, because of its comparative advantage in monitoring and enforcement capacity. Besley (1995) provided an excellent survey on this issue.

Because our empirical results are obtained by using a dataset that consists mainly of low-income households, it should be safe to say that DRRA is at least important for understanding the consumption behavior of poor households. Many researchers have rejected complete risk-sharing for other datasets; however, most of these datasets do contain low-income households. Moreover, based on the available evidence obtained from asset data, such as that presented by Guiso, Jappelli, and Terlizzese (1996) and Kessler and Wolf (1991), the possibility that DRRA also holds for higher-income households cannot be ruled out. Therefore, it will be of great interest to see whether allowing for DRRA or other forms of heterogeneity in risk preferences may change their test results. The findings of two recent reports echo this point. Guvenen (2002) reported that the RSH cannot be rejected for nonstockholders in the PSID data but can be rejected for stockholders when these two classes of agents are allowed to have different curvature parameters in their utility functions. Dubois (2000) parameterized household utility functions with observable characteristics to allow for heterogeneous risk aversion, and one set of his tests does not reject the RSH at village level in the Pakistani dataset that O&Z (2001) uses. These findings suggest that future efforts along this dimension should produce more insightful results.

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APPENDIX: DATA

ICRISAT conducted intensive interviews in southern rural India between 1975 and 1984. This article uses data from three villages for which the complete consumption data are available: Aurepalle, Shirapur, and Kanzara. Three measures of consumption—food, nonfood, and total consumption—were constructed. Because the construction of food consumption was changed in 1976 and the data for nonfood consumption are missing for most categories after 1982, 1976–1981 is chosen as the relevant sample period for total consumption. The income data that we use is already net of remittance and includes crop profit, labor income, profit from trade and handicrafts, and profit from animal husbandry.

We use food, including grain, milk, vegetables, meat, sweets, and spices, as the measure of food consumption. To obtain nonfood consumption, we subtract food and ceremonial expenses from total consumption expenditure. Ceremonial expenses are removed because they often jump from 0 to large
amounts. Nonfood consumption consists of narcotics, tea, coffee, tobacco, pan, and alcoholic beverages; clothing, sewing of cloth, other tailoring expenses, thread, needles, chap pāls and other footwear and so on; travel and entertainment; medicines, cosmetics, soap, barber service; electricity, water charges and cooling fuels for household use; labor expenses for domestic work; edible oils and fats (other than gee); and others, including complete meals in hotels, school and educational materials, stamps, stationery, grinding and milling charges, and so on. The ICRISAT consumption data do not include housing and transportation, because the market values of these categories of consumption are hard to measure in these villages. Total consumption expenditure is the sum of food and nonfood consumption.

To construct real consumption and real income per–male–adult equivalent, nominal consumption and income at some year $t$ is divided by the family size measure used by Townsend (1994) and the corresponding price index at $t$ for each village. The price index for total consumption expenditure, food, and nonfood are the consumer price index, the price index for food, and the price index for nonfood. These real variables are valued at 1983 prices.

There are about 40 households for each year in each of the 3 villages in the dataset. Some households drop out of the sample, and others are added to the sample over the years. These households were excluded from the sample. The numbers of households in our sample are 35 for the village of Aurepalle, 33 for Shirapur, and 36 for Kanzara.

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