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Wealth-Varying Intertemporal Elasticities of Substitution: Evidence from Panel and Aggregate Data

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Abstract

This paper constructs and estimates a model in which the intertemporal elasticity of substitution of consumption expenditure (IES) rises with the level of wealth. The purpose of this paper is to measure the effect that systematic variation in the IES of poor and rich consumers has on the intertemporal elasticity of substitution of aggregate consumption expenditure. The principal innovation embodied in our method of measuring the differences in the IES of poor and rich consumers is that our model for estimation allows us to directly measure the impact that these differences at the individual level have on the allocation of aggregate consumption expenditure over time. We estimate and test the specification of our model in Indian panel data on the consumption of individual households. We re-estimate the model using aggregate times series data from India and the U.S. and check for aggregation bias by comparing parameter estimates in individual level and aggregate level data and to make an international comparison of preferences. We find economically significant differences in the behavior of poor and rich households within the Indian panel data set as well as in the aggregate data from the United States and India.

I. Introduction

This paper constructs and estimates a model of preferences in which the intertemporal elasticity of substitution of consumption expenditure (IES) can vary systematically between rich and poor consumers. Intuitively, there are at least two reasons why the intertemporal elasticity of substitution might be smaller for the poor than it is for the rich. First, if there are positive subsistence consumption requirements, then poor consumers have a smaller portion of their budget left over after satisfying subsistence requirements to save or consume at their discretion. Second, the consumption of necessary goods (such as food) may be less substitutable across time is than the consumption of luxury goods. Since the poor spend a higher fraction of their total expenditure on necessary goods than do the rich, their IES of total consumption expenditure may be smaller than is true for the rich. Thus the intertemporal elasticity of substitution may rise with the level of wealth.[1]

It is clear that systematic differences in the attitudes of the rich and the poor toward the allocation of their consumption expenditure over time can lead to significant effects of changes in the level and distribution of wealth on the behavior of aggregate consumption. The purpose of this paper is to measure the quantitative importance of the effects that differences in the intertemporal elasticity of substitution at

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[1]This intuition is based entirely on our own introspection. Unless preferences are restricted, there are no theoretical reasons that the intratemporal expenditure elasticity of demand for a good and the intertemporal substitutability of expenditure on that good should be related.

the individual level have on the behavior of aggregate consumption.[2]

The model that we use to formalize the intuition given above and to measure the differences in the intertemporal elasticity of substitution (IES) of consumption expenditure for poor and rich consumers is based on the extended addilog utility function. The IES varies with wealth when consumers have this utility function because this utility function is not homothetic. The principal advantage of the extended addilog preference model as a model for measuring the effects of systematic differences in the IES of poor and rich consumers on the behavior of aggregate consumption is that this utility function aggregates in a useful way. Despite the fact that the extended addilog utility function is not quasi-homothetic, we can demonstrate that these preferences aggregate in the following sense: in an economy composed of individuals with identical extended addilog preferences who trade in complete markets, equilibrium aggregate consumption can be modeled as if it were chosen by a representative consumer with extended addilog preferences. Furthermore, certain preference parameters of the representative consumer that are critical for the measurement of the aggregate IES are common to the representative and the individual consumers.[3]

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[2]Blundell, Browning, and Meghir (1990) estimate a different model in which the IES rises with level of wealth. Their main focus is on aggregation bias and they do not discuss the aggregate effects of systematic differences in the consumption growth of the rich and the poor.

[3]Because the extended addilog utility function is not quasi-homothetic, the distribution of individual wealth enters into the parameters of the representative consumer's utility function. The important aspect of the aggregation result here is that the distribution of individual wealth does not affect the parameters which are critical in determining the aggregate IES.

This aggregation result is useful in two respects. First, with this aggregation result, we can use both individual and aggregate level data to estimate the key parameters of the model. Specifically, in this paper we use both individual level panel data and aggregate time series data to obtain sharper estimates of the parameters of interest. Second, with this aggregation result, it is very easy to take individual or aggregate level data and draw implications about the effects that changes in the distribution of individual consumption or changes in the level of aggregate consumption will have on the behavior of aggregate consumption from the micro foundations of the model.

There are two principal drawbacks of the extended addilog utility function as a model for measuring the IES at the individual and the aggregate level. The first drawback of the model is that the extended addilog utility function has only a few parameters for use in explaining consumption behavior. As a result, the measurements derived from the model are subject to specification error. The second drawback of the model is that the aggregation result depends upon an assumption of complete markets. Thus, the aggregate measurements that we derive from the model are subject to aggregation bias.

The principal problem with specification error in this model is that the model makes strong assumptions of additive separability of preferences among goods both intratemporally and intertemporally. To test our model for specification error due to this assumption of separability, we estimate the curvature parameters of the model with panel data from Indian households, using both intertemporal and intratemporal information. Then we form a specification test *a la* Hausman (1978) for our model by comparing estimates of the same parameters from intertemporal and intratemporal regressions.

We test our model for aggregation bias by re-estimating the model in aggregate data. We do this using Ogaki and Park's (1990) cointegration approach in aggregate U.S. and Indian time series data. The purpose is to assess the quantitative importance of the aggregation bias by comparing estimates from individual level panel data and those from aggregate data. We undertake two sets of tests of the model based on these various estimation results. First we compare the estimates obtained in the Indian panel data and the Indian aggregate data to asses the quantitative importance of aggregation bias on our measurements. Second, we compare the estimates obtained in Indian aggregate data with those obtained in U.S. aggregate data to assess the extent to which these preference parameters appear to be consistent across a wide range of consumption expenditure levels.

There are at least two alternative explanations of the differences in consumption growth between poor and rich consumers that are not subsumed in the model that we present. One possible explanation of the differences in consumption growth between poor and rich consumers is based on the assumption that it is the rate of time preference (RTP) rather than the intertemporal elasticity of substitution of consumption expenditure varies with wealth.[4] One can distinguish the wealth-varying IES model and this wealth-varying RTP model in data that contains periods in which rates of

[4]There has been a great deal of theoretical work in this vein (See, e.g., Uzawa (1968), Epstein (1983)). A recent paper by Lawrance (1991) tests this hypothesis in data from the Panel Study of Income Dynamics. Her test maintains an assumption that the intertemporal elasticity of substitution is constant across consumers.

return are high and aggregate consumption is growing and periods in which rates of return are low and aggregate consumption is shrinking in the following manner. The rate of time preference model would imply that if rich consumers were more patient than poor consumers, then rich consumers consumption would always grow faster than poor consumers consumption. The opposite, of course, would be true if rich consumers were less patient than poor consumers. The wealth varying IES model, on the other hand, implies that, if poor consumers have a lower IES than do rich consumers, then consumption growth rates for the poor are simply less responsive to changes in rates of return than is true for rich consumers. That is, consumption should grow more slowly for poor consumers than it does for rich consumers when rates of return are high and aggregate consumption is growing and should shrink more slowly for poor consumers than it does for rich consumers when rates of return are low and aggregate consumption is shrinking.[5] We compare these hypotheses in the Indian panel data.

A second possible explanation of the differences in consumption growth is based on the assumption of incomplete markets such as borrowing constraints.[6] We will argue that the bias of our measurement coming from this source does not seem to be serious for our purpose.

The remainder of this paper is organized as follows. In Section II, we

[5]In our model, the intertemporal elasticity of substitution of consumption expenditure is the reciprocal of the coefficient of relative risk aversion. Thus poor households are more risk averse than rich households. Thus one interpretation of more stable consumption growth of poor households implied by our model is that poor households bear less risk than rich households, who are less risk averse, in equilibrium.

[6]Hayashi (1985), Zeldes (1990), Morduch (1990), Deaton (1991), Runkle (1991) are examples of recent work in the area.

present our model. In Section III we present the aggregation result for the extended addilog utility function that makes it so useful. In Section IV, we show evidence that consumption growth is systematically different for rich and the poor households in the ICRISAT panel data and demonstrate that these differences in consumption growth are better explained with a model based on differences in the intertemporal elasticity of substitution among consumers rather than on differences in the rate of time preference among consumers. In section V, we use the ICRISAT panel data set to estimate the curvature parameters of the extended addilog preference model using both the intratemporal and intertemporal first order conditions of the model. We test for the misspecification that a single set of parameters is not adequate to summarize both sets of first order conditions. In Section VI, we use the cointegration approach to estimate the same curvature parameters in aggregate consumption data for India and U.S. In Section VII, we present some measurements of the range of individuals intertemporal elasticities of substitution of consumption expenditure in our Indian panel data and some measurements of the differences in the intertemporal elasticity of substitution at the aggregate level implied by the data for India and the United States. Section VIII contains our concluding remarks. Appendix A discusses the data, Appendix B presents an approximation result.

II. The Allocation of Consumption Expenditure Over Time

In this section, we present the model of consumers intertemporal allocation of consumption expenditure that we use for estimation and measurement. In particular, we discuss the different implications for consumption growth that can be derived from the model when the rate of time preference is assumed to vary with the level of consumption expenditure (and thus with wealth) and when the intertemporal elasticity of substitution is assumed to vary with the level of consumption expenditure (and thus with wealth).

A. The Model

Consider an economy with H households, each of which consume n goods in each of T time periods. Let the consumer h, $h=1,\dots,H$, have time and state separable utility with an intratemporal utility function $u(C^h(t))$, where $C(t)=(C_1(t),\dots,C_n(t))'$ and let β denote the consumer's discount factor. Let a vector $s(t)$, $s(t) = 1,2,\dots,S$, denote the state of the world in each period and the vector $e(t)=[s(0),s(1),\dots,s(t)]$ be the history of the economy. The consumer h maximizes

$$U^h = \sum_{t=0}^{T} \sum_{e(t)} \beta^t Prob(e(t)\,|\,e(0))u(C^h(t,e(t)) \tag{1}$$

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

$$\sum_{t=0}^{T} \sum_{e(t)} (\prod_{\tau=0}^{t} R(\tau-1,e(\tau-1),e(\tau))^{-1} P(t,e(t))' C^h(t,e(t)) \le W^h(0), \tag{2}$$

where $W^h(0)$ is the consumer h's initial wealth and T can be a finite number as in the life-cycle model or infinity as in the dynasty model. Here $P(t,e(t))=(P_1(t,e(t)),\dots,P_2(t,e(t)))'$ is the intratemporal prices with good one as the numeraire in each period and state $(P_1(t)\equiv 1)$ and $R(t-1,e(t-1),e(t))$ is the (gross) asset return of the state contingent security for the event $e(t)$ in terms of the first good in the event $e(t-1)$ at period $t-1$. We will often suppress e(t)to simplify the notation below.

Let $E^h(t)=P(t)'C^h(t)$ be the total consumption expenditure in terms of good one and $v(E^h(t),P(t))$ be the intraperiod indirect utility function that maximizes u(C) subject to the intraperiod budget constraint

$$P'C = E^h. \tag{3}$$

Then the household h maximizes

$$\sum_{t=0}^{T} \sum_{e(t)} \beta^t Prob(e(t)|e(0)) v(E^h(t),\ P(t)) \tag{4}$$

subject to a life-time budget constraint

$$\sum_{t=0}^{T} \sum_{e(t)} \left(\prod_{\tau=0}^{t} R(\tau-1)\right)^{-1} E^h(t)) \le W^h(0), \tag{5}$$

where $W^h(0)$ is the consumer h's initial wealth.

The consumer has the familiar first order condition governing the intertemporal allocation of his consumption expenditure

$$\beta \frac{v_E(P(t+1), E^h(t+1))}{v_E(P(t), E^h(t))} = R(t)^{*-1} \tag{6},$$

where $v_E = \partial v/\partial E^h$ and $R^*(t,e(t),e(t+1)) = R(t,e(t),e(t+1)) Prob(e(t+1)|e(t))$. Approximating v_E as a power function (see Appendix B for a lemma that shows this approximates v_E for any indirect utility function) and ignoring changes in intratemporal prices $P(t)$, we get the result that the growth of total consumption expenditure $(\hat{E}(t) = log(E(t+1)) - log(E(t)))$ approximately satisfies

$$\hat{E}^h(t) \cong \sigma^h(t)\{r(t) - \delta^h\} \tag{7}$$

where $\delta^h = \ln(1/\beta^h)$ is the rate of time preference, $r(t) = \ln(R^*(t))$, and $\sigma^h(t) = -v_E/(v_{EE}E^h)$ is the intertemporal elasticity of substitution. From (7), $\sigma^h(t) \cong \partial\hat{E}^h(t)/\partial r(t)$. If there is no uncertainty, r(t) is the real interest rate.

The distinct implications for consumption growth of models in which the rate of time preference varies systematically with wealth and models in which the intertemporal elasticity of substitution varies systematically with wealth can be seen in equation (7). If δ^h falls systematically as wealth rises, then the consumption growth of the poor is always lower than the consumption growth of the rich. On the other hand, if σ^h rises systematically with wealth, then the consumption growth of the rich will be more volatile than the consumption growth of the poor as r varies around δ.

In our estimation and measurement we will use the extended addilog utility function as the direct utility function:

$$u(C^h) = \sum_{i=1}^{n} \frac{\theta_i}{1-\alpha_i} \left[(C_i^h - \gamma_i)^{(1-\alpha_i)} - 1 \right] \tag{8}$$

where $\alpha_i > 0$ and $\theta_i > 0$ for i=1,...,n. We will refer to the parameters γ_i as subsistence parameters and the parameters α_i as curvature parameters.[7] This utility function contains as special cases two utility functions commonly used in demand studies. If $\alpha_i = 1$ for i=1,..,n, then this utility function yields the linear expenditure system in that the intratemporal demand functions for consumption of each good in excess of subsistence consumption $(\tilde{C}_i = C_i - \gamma_i)$ are linear in expenditure in excess of subsistence expenditure $(\tilde{E} = E - \sum_{i=1}^{n} P_i \gamma_i)$. More generally, if $\alpha_1 = \alpha_2 = \ldots = \alpha_n$, then preferences are

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[7]Several authors have considered the effects of subsistence levels on intertemporal consumption behavior. See Gersovitz (1983), for example of work with subsistence levels in the area of saving and development. Rebelo (1991) studies the effects of subsistence levels in an endogenous growth model (See also Ogaki (1991a)). Christiano (1989) uses a model with a time varying subsistence level to model the behavior of the Japanese saving rate in the post war years. Chatterjee (1991) studies effects of subsistence levels on income distribution as an economy develops. Habit formation models include time varying subsistence levels. See Constantinides (1990), Ferson and Constantinides (1991), Heaton (1991), and Cooley and Ogaki (1991) for examples of empirical work with models of habit formation in aggregate data.

quasi-homothetic. If $\gamma_i = 0$ for $i=1,\ldots,n$, then this utility function is Houthakker's (1960) addilog utility function.[8]

In our empirical work, we will focus on special cases where $n=1$ and $n=2$. We will explain how the IES varies with the level of wealth when preferences are represented by the extended addilog utility function for these cases.

B. Single Good Model

Consider the case where $n=1$. Then the IES is

$$\sigma^h(t) = -u'(C(t))/[C(t)u''(C(t))]. \tag{9}$$

The IES is constant when the intratemporal utility function is isoelastic but depends on wealth in general because the IES may vary with the level of consumption and the level of consumption varies with wealth. In this case the extended addilog utility function reduces to the quasi homothetic Geary-Stone utility function that Rebelo (1991) employs

$$u(C(t)) = \frac{1}{1-\alpha}((C(t)-\gamma)^{(1-\alpha)}-1), \tag{10}$$

where $\alpha>0$ and γ is a subsistence level parameter. Then

$$\sigma^h = \frac{1}{\alpha}\left(1 - \frac{\gamma}{C}\right). \tag{11}$$

If $\gamma>0$, then the IES of the poor is smaller than that of the rich. For a poor household, C is close to γ and σ is close to zero. For a rich household, γ/C is close to zero and σ is close to $1/\alpha$. Thus the

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[8]This paper focuses on the intertemporal implications of the extended addilog utility function. See Atkeson and Ogaki (1991) and Ogaki (1990a) for intratemporal implications of the utility function.

intertemporal elasticity of substitution rises with the level of wealth.[9]

The one good model imposes the condition that total consumption expenditure is quasi-homothetic upon the data. Since we also wish to model the potential impact of changes in the distribution of wealth on aggregate consumption behavior, we also consider the following two good version of the model which allows total consumption expenditure to fail to be quasi-homothetic.

C. Two-Good Model

In much of our empirical work, we will estimate the extended addilog utility function with $n=2$ where C_1 is food consumption, C_2 is nonfood consumption. In order to derive the expression for the IES, note that

$$v_E(E,P) \equiv u_1(C(E,P)), \tag{12}$$

where u_1 is the marginal utility for good one because good one is taken as a numeraire. It can be shown that the extended addilog utility function implies

$$\mu = \{\frac{\alpha_1}{\alpha_2} + \omega_1^h(t)(1-\frac{\alpha_1}{\alpha_2})\}^{-1} \omega_1^h(t) \frac{E^h(t)}{C_1^h(t)}. \tag{13}$$

where $\omega_i^h(t) = P_i(t)\tilde{C}^h(t)/\tilde{E}(t)$ is the budget share of good i after removing subsistence levels, $\tilde{C}_i^h(t)=C_i^h(t)-\gamma_i$ is the consumption over the subsistence level for good i, and $\tilde{E}^h(t)=E^h(t)-\gamma_1-P_2(t)\gamma_2$ is total consumption

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[9]It should be noted that there is no theoretical reason to exclude the case where $\gamma<0$, though if $\gamma<0$ then γ is not interpreted as the subsistence level. If $\gamma<0$, then the consumption growth of the poor will be more volatile that that of the rich.

expenditure in excess of subsistence expenditure. Thus the intertemporal elasticity of substitution for the consumer h with the extended addilog utility function is

$$\sigma^h(t) = (\omega_1^h(t)\frac{1}{\alpha_1} + \omega_2^h(t)\ \frac{1}{\alpha_2})\ (1 - \frac{\gamma_1+P_2(t)\gamma_2}{E^h(t)}). \tag{14}$$

Thus, in the case of addilog preferences with two goods, the IES will vary with the level of expenditure for two reasons. First, if subsistence expenditure $\gamma_1+P_2(t)\gamma_2$ is non-zero, then the second term in (14) will either rise from zero to one as expenditure goes from subsistence expenditure to infinity (when subsistence expenditure is positive) or it will fall from infinity to one as expenditure rises from zero to infinity (when subsistence expenditure is negative). Second, if $\alpha_1>\alpha_2$, then ω_2 rises from zero to infinity as $\tilde{E}$ rises from zero to infinity, changing the first term in (14) from $1/\alpha_1$ to $1/\alpha_2$ as expenditure rises. On the basis of our estimation work in Ogaki and Atkeson (1991), we will assume that subsistence expenditure is positive in this paper, so that the IES is assumed to rise with the level of wealth.

III. Aggregation of Preferences

The extended addilog utility function given above has the convenient property that the aggregate consumption data from a competitive equilibrium in a model with multiple consumers with addilog preferences behaves as if it were chosen by a single representative consumer with addilog preferences with the same parameters α_i and γ_i that the individual agents have.

The aggregation result may be stated as follows. Consider an economy with H consumers numbered $h=1,\ldots,H$. Assume that all of the consumers have identical time additively separable preferences with intratemporal utility function over food consumption and nonfood consumption given by the extended addilog utility function (8).

Our aggregation result is that, if this economy has a competitive equilibrium, then there exists parameters θ_i^*'s for which the equilibrium prices and the aggregate consumption vector satisfy the equilibrium first order conditions for a single representative consumer who has time separable preferences with an addilog intratemporal utility function given by

$$u(C^*) = \sum_{i=1}^{n} \frac{\theta_i^*}{1-\alpha_i} [(C_i^*-\gamma_i)^{(1-\alpha_i)} - 1] \tag{15}$$

where $C_i^* = \sum C_i^h/H$.

Before we prove this result, it is worthwhile to discuss its implications. This representative consumer has utility with the same parameters α_i and γ_i for i=1,...,n as the individual consumers. Knowledge of these parameters together with aggregate data on food consumption, and total consumption expenditure is sufficient to calculate the intertemporal elasticity of substitution for the representative consumer. The parameters θ_i^*'s will be shown to reflect the distribution of initial wealth in society, and would thus in principle be difficult to measure directly. But, because the individual and the representative consumer share common values of α_i and and γ_i, the impact of these parameters θ_i^*'s on the evolution of aggregate saving behavior over time is completely summarized in the aggregate equilibrium food consumption and total consumption expenditure data.

To prove this aggregation result, begin with the assumption that there exists a competitive equilibrium for the original economy with H consumers. Denote individual consumption in period t by $C_i^h(t)$, i=1,...,n, h=1,2,...,H, and aggregate consumption per capita by $C_i^*(t)$, i=1,...,n. Individual consumption satisfies the following first order conditions in equilibrium. The consumers intertemporal first order conditions are

$$\left(\frac{C_i^h(t,e(t))-\gamma_i}{C_i^h\ (t+1,e(t+1)-\gamma_i}\right)^{-\alpha_i} = \beta R^*(t,e(t),e(t+1))\frac{P_i(t,e(t))}{P_i(t+1,e(t+1))} \tag{16}$$

for i=1,..,n, and all pairs of states of the world $e(t)$, $e(t+1)$. The consumers intratemporal first order conditions are

$$\frac{\theta_i\ \ (C_i^h\ (t,e(t))-\gamma_i)^{-\alpha_i}}{\theta_1\ \ (C_1^h\ (t,e(t+1))-\gamma_1)^{-\alpha_1}} = \frac{P_i(t,e(t))}{P_1(t,e(t+1))} \tag{17}$$

for i=1,2,...,n. To show that the equilibrium prices and the aggregate consumption vector are also a competitive equilibrium in this economy with a representative consumer, we must find values for the parameters θ_i^*'s such that the aggregate consumption per capita vector satisfies the first order conditions (16) and (17) for the representative consumer.

We find the appropriate values of θ_i^*'s as follows. The intertemporal first order conditions (16) for the individual consumers indicate that consumption of each good in excess of subsistence consumption of that good grows at the same rate for all consumers in equilibrium.[10] Note that these intertemporal first order conditions (16) will be satisfied by the representative consumer at the equilibrium prices and the aggregate consumption per capita vector regardless of the values of the parameters θ_i^*'s. These intertemporal first order conditions also imply that each consumer's food consumption in excess of subsistence food consumption is a

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[10] Note that total expenditure growth need not be the same for all consumers since consumption of the two different goods may grow at different rates and consumers can spend different fractions of total expenditure on the two goods.

constant fraction over time of aggregate food consumption per capita in excess of the individual's subsistence consumption level. Thus, we can index the equilibriuminitial wealth of each consumer by the fraction δ_h defined by $\delta_h = (C_1^h(t)-\gamma_1)/(H(C_1^*-\gamma_1))$ where $C_1^* = \sum C_1^h/H$.

The individual consumer's intratemporal first order condition (17) implies that

$$(C_i^h(t)-\gamma_i)^{\alpha_i} = \frac{P_1(t)}{P_i(t)} \frac{\theta_i}{\theta_1} (C_1^h(t)-\gamma_1)^{\alpha_1}$$

for i=1,...,n. Substituting $(C_1^h(t)-\gamma_1) = \delta^h H(C_1^*(t)-\gamma_1)$ into the expression above and summing across consumers we get the condition that in the aggregate

$$H^{\alpha_i}(C_i^*(t)-\gamma_i)^{\alpha_i} = \frac{P_1(t)\,\theta_i}{P_i(t)\,\theta_1}\Big(\sum_h (\delta^h)^{\alpha_1/\alpha_i}\Big)^{\alpha_i} H^{\alpha_1}(C_1^*(t)-\gamma_1)^{\alpha_1}$$

From this expression we can see that the representative consumer's intertemporal and intratemporal first order conditions are satisfied at the equilibrium prices and the equilibrium aggregate consumption vector when $\theta_i^* = \theta_i\big(\sum_h \delta_h^{\alpha_1/\alpha_i}\big)^{\alpha_i} H^{(\alpha_1-\alpha_i)}$. This parameter $\theta_i^* = \theta_i$ if $\alpha_1 = \alpha_i$. When $\alpha_1 \neq \alpha_i$ for some i, then the utility function of the representative consumer depends upon the distribution of initial wealth in society. But, as we mentioned earlier, the impact of this distribution of wealth in society on the evolution in aggregate expenditure growth is completely summarized in the evolution of aggregate consumption data over time.[11]

[11] Our aggregation result is a special case of the aggregation result under complete markets in Ogaki (1990b), who also discusses the relation of our aggregation result with other aggregation results.

IV. Are There Systematic Differences in Consumption Growth of The Rich and the Poor?

In this section, we examine the panel data in India collected by the Institute for Crop Research in the Semi-Arid Tropics (ICRISAT) for evidence of systematic differences in the consumption growth of rich and poor households that would not be explained simply on the basis of systematic differences in income growth across households. In particular, we estimate a model of household consumption growth which encompasses the hypotheses that there is systematic variation in the rate of time preference (RTP) and the intertemporal elasticity of substitution (IES) across households. We use panel data for three villages (Aurepalle, Shirapur, and Kanzara) for the period from the fiscal year 1975-76 to the fiscal year 1984-85. (We denote each fiscal year in India by its first calendar year below). Since construction of food consumption was changed in 1976 and the data for nonfood consumption are missing for most categories after 1982, we set our sample period to be 1976-1981.

These Indian panel data are attractive for several reasons and have been used to study consumption smoothing and risk sharing models by many authors.[12] First, saving behavior of less developed countries are of interest. Second, this is the only panel data set that includes food consumption and nonfood consumption data that covers a period that is longer than two years for same households to the best of our knowledge. Because much of consumption fluctuations within a year are likely to be caused by

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[12] See, e.g., Bhargava and Ravallion (1991), Lim (1990), Morduch (1990, 1991), Ravallion and Chaudhuri (1991), Rosenzweig (1988), Rosenzweig and Binswanger (1990), Rosenzweig and Stark (1989), Rosenzweig and Wolpin (1989), and Townsend (1991).

seasonal shifts that are not of interest for our purpose, it is desirable to study panel data that cover a substantial time period.[13]

This section is organized as follows. In the first subsection, we explain a method of using the single-good version of the wealth-varying IES model, modified to allow for the possibility of a wealth-varying RTP, for investigating the nature of the differences in the consumption growth of the poor and the rich. The second subsection presents our empirical results.

A. Econometric Method

We consider the single good version of the model with wealth-varying RTP. The intertemporal first order condition (16) becomes

$$\left(\frac{C^h\ (t,e(t))-\gamma}{C^h(t+1,e(t+1))-\gamma}\right)^{-\alpha} = \beta^h R^*(t,e(t),e(t+1)) \tag{18}$$

where β^h can depend on the level of wealth. We assume that consumption $C^h(t)$ is measured with error in the following form:

$$C^h_m(t)-\gamma = (C^h(t)-\gamma)\epsilon^h(t), \tag{19}$$

where $C^h_m(t)$ is measured consumption and $\epsilon^h(t)$ is a multiplicative measurement error, which can be serially correlated but is assumed to be independent across households. We assume that $\epsilon^h(t)$ has mean one and is positive. We assume that β^h satisfies

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[13]For example, the Consumer Expenditure Survey in the U.S. includes food and nonfood consumption but keeps track of individual households only up to four consecutive quarters (see, e.g., Mace (1991)). An alternative way to obtain panel data is to construct synthetic panel data from a series of cross-sectional data as in Blundell, Browning, and Meghir (1990) among others. It is, however, better to use real panel data to avoid extra noise that comes from construction of synthetic panel data.

$$\ln(\beta^h) = \beta_0 + \beta_1 y_c^h + \epsilon_a^h, \tag{20}$$

where y_c^h is a proxy of permanent income and ϵ_a^h is also a measurement error that is assumed to be independent across households and independent of $\epsilon^h(t)$. Then from (18)-(20), we get

$$\ln(C_m^h(t+1)-\gamma)-\ln(C_m^h(t)-\gamma)-(\phi(t)+\beta_y y_c^h) = v^h(t), \tag{21}$$

where $\phi(t)=(1/\alpha)(\ln R^*(t)+\beta_0)$, $\beta_y=\beta_1/\alpha$, and

$$v^h(t)=\ln(\epsilon^h(t+1))-\ln(\epsilon^h(t))+(1/\alpha)\epsilon_a^h. \tag{22}$$

Let y_p^h be another proxy of permanent income of household h, $y^h(t)$ be the current income of household h at date t, and $z^h(t)=(1,\ln(y_p^h),\hat{y}^h(t))'$ be a vector of instrumental variables. We assume that $v^h(t)$ is uncorrelated with $z^h(t)$ across households. This choice of instrumental variables is determined by the purpose of the present paper. The growth rate of current income of each household, $\hat{y}^h(t)$, is included as an instrument because we seek to find systematic differences of consumption growth that are not simply explained by differences in household income growth. We need to include a measure of permanent income to make sure that the estimated utility function is consistent with consumption growth of both poor and rich households. In our empirical work, we use the average real income over the last three years in the data that are not included in the sample as y_p^h and the average real food consumption over the last three years in the data as y_c^h.

We fix the state of the world and treat $\phi(t)$ as a parameter to be estimated. Let $p = (p_1,\ldots,p_{T+2})$ be (T+2)-dimensional vector of unknown parameters. The true value of p is $p^0 = (\phi(1),\ldots,\phi(T),\gamma,\beta_y)'$. We define a

3-dimensional vector $\xi_t^h(p)$, so that $\xi_t^h(p^0)=z_h(t)v^h(t)\exp(-\gamma/A)$, where A is a constant. Here we normalize the disturbance by $\exp(-\gamma/A)$ to avoid a trivial solution $\phi(t)=0$ for $t=1,\ldots,T$, $\gamma=-\infty$, $\beta_y=0$. Let $\xi^h(p)=(\xi_1^h(p),\ldots,\xi_T^h(p))'$. Then we have 3T orthogonality conditions

$$E_H[\xi(p^0)] = \operatorname*{plim}_{N\to\infty} (1/N)\sum_{h=1}^{N} [\xi^h(p^0)] = 0 \tag{23}$$

where E_H is the expectation operator over households. A subscript H is attached to emphasize that the expectation is taken over households. We have these 3T orthogonality conditions for each village. We pool these orthogonality conditions for the three villages and estimate p for each village with the generalized method of moments (GMM).[14] In pooling the data for the three villages, we allow incomplete markets in the form of different asset returns in different villages. Thus we allow $\phi(t)$ to be different in different villages but restrict preference parameters γ and β_y to be identical across the villages.

Our specification allows consumption growth to depend on the level of wealth in a variety of ways. Consider the case where $\beta_y=0$. If $\gamma>0$ then the consumption of the rich grows faster than that of the poor when aggregate consumption grows and consumption of the rich shrinks faster when aggregate consumption shrinks. If $\gamma<0$, then the reverse is true. On the other hand, when $\gamma=0$, if $\beta_y<0$, then the consumption of the rich always grows faster than

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[14]See, e.g., Hansen (1982) and Gallant and White (1988). We assume that the regularity conditions of Gallant and White are satisfied. Hansen/Heaton/Ogaki's GAUSS GMM package is used for the GMM and the minimum distance estimations in the present paper. In pooling the data for three villages, we allow $\xi(p^0)$ to have different covariance matrices in different villages. Ogaki (1991b) provides a more detailed explanation as to how the data for villages are pooled.

that of the poor and, if $\beta_y>0$, then the reverse it true. In the case where both γ and β_y are nonzero, these effects are combined. If $\beta_y=\gamma=0$, then there is no systematic difference in the consumption growth of the poor and rich.

B. Empirical Results

In Table 1, we report results for real total consumption expenditure per equivalent adult. In the first panel, we report estimates of γ and β_y and test statistics. The first row reports results when no restriction is imposed; the second row, when one restriction $\beta_y=0$ is imposed; the third row, when two restrictions $\beta_y=\gamma=0$ are imposed. The *J* statistic reported in each row is Hansen's (1982) χ^2 test for the overidentifying restrictions. The *C* statistics reported in the second and third rows are the difference between the *J* of each row and the *J* of the first row.[15] The *C* statistic in the the third row tests the restrictions $\beta_y=\gamma=0$ which corresponds with the hypothesis that there is no systematic difference in the consumption growth of the rich and the poor. The *C* test provides strong evidence against this hypothesis. The *C* statistic in the second row tests the restriction $\beta_y=0$. There is little evidence against this hypothesis. The *J* statistics in the

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[15]See, e.g., Eichenbaum, Hansen, and Singleton (1988) for an explanation of the *C* test in the GMM procedure. In order to compare *J* statistics with the *C* test, the same distance matrix needs to be used for unrestricted and restricted estimations. The distance matrix used is based on the estimation with the restriction $\beta_y=0$. The initial distance matrix is an identity and the GMM estimation is iterated three times. The constant *A* for normalization was set to 200 for total consumption expenditure and food in Tables 1 and 2 and to 50 for nonfood consumption in Table 3. The final results were virtually the same when *A* was increased to 300 for total consumption and food and to 100 for nonfood but convergence for the initial distance matrix needed more iterations.

second row tests the hypothesis that there exists no systematic component in consumption growth that can be explained by the income variables in the instruments once the subsistence level γ and aggregate consumption in each village summarized by $\phi(t)$'s are taken into account. We do not reject this hypothesis.[16] Consistent with the *C* test results, γ is estimated to be statistically significantly positive, but β_y are not (statistically) significantly different from zero. Thus our results are more in favor of the wealth-varying IES model than for the wealth-varying RTP model.[17]

We report estimates of $\phi(t)$'s for Aurepalle, Shirapur, and Kanzara in the second, third, and fourth panel of Table 1 when β_y is restricted to be zero. In this case, $\phi(t)$ is the growth rate of $C(t)-\gamma$, which is common to all households. We have both significantly positive values of $\phi(t)$ and significantly negative values of $\phi(t)$. This is important because the wealth-varying IES and the wealth-varying RTP models can be discriminated sharply only when the data contain both periods in which aggregate consumption grows and those in which it shrinks as discussed above.

We report results when C(t) is taken as food in Table 2 and results

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[16] This result is consistent with the hypothesis that there is full risk sharing among the households of each village in the data set. This result does not, in isolation, constitute a full test of that hypothesis (see Ogaki and Atkeson (1991) for more detailed discussions). See Altug and Miller (1990), Cochrane (1991), Deaton (1991), Hayashi, Altonji, and Kotlikoff (1991), Morduch (1991), Ravallion and Chaudhuri (1991), and Townsend (1991) for other tests for complete markets.

[17] Our result is consistent with Lawrance's (1991) result that consumption growth is higher for the rich than the poor in the PSID over 1974-1982 even though Lawrance uses the wealth-varying RTP model. Since positive γ implies more volatile consumption growth for the rich than that for the poor, our result is also consistent with Mankiw and Zeldes's (1991) finding that consumption growth is more volatile for stockholders than nonstockholders in the PSID.

when C(t) is taken as nonfood in Table 3. These provide estimates of γ_1 and γ_2 for the two good model under the separability assumption. The results for food and nonfood are qualitatively similar to those for total consumption. For each of these categories of consumption, consumption of the rich tends to grow faster than that of the poor when aggregate consumption grows while consumption of the rich tends to shrink faster than that of the rich when aggregate consumption shrinks.

We use estimates of the subsistence parameters for food and non-food conumption obtained in this manner in our estimation of the curvature parameters of the two good model. One danger in using these estimates is that our estimates of subsistence levels may be severely biased by market imperfections such as borrowing constraints. We argue that our estimates are not likely to be severely biased for three reasons. First, there is little systematic fluctuation in household consumption growth left that can be explained by income variables after subsistence levels and aggregate consumption are taken into account. This result is particularly useful in justifying the application of our aggregation result. Since the aggregation result depends on the assumption that consumption of each good in excess of subsistence consumption of that good grows at the same rate for all consumers, a finding that there was systematic variation in the rate of time preference or the rate of growth of consumption in excess of subsistence consumption would invalidate that result. Second, our estimates of the subsistence consumption levels are restricted to be the same over time and across households in each village and across different villages. We find little evidence against this restriction across villages.

In the next three sections of the paper, we will estimate the two good model maintaining the assumption that consumers all share the same rate of time preference. We have chosen to make this assumption so as to allow ourselves to make use of the aggregation result of section III.

V. Estimation of Curvature Parameters with Panel Data

This section derives an analytical solution for the growth rate of real consumption expenditure in the two good addilog preference model. This solution is used to develop a method to estimate parameters $1/\alpha_2$ and α_1/α_2 from panel data using both intratemporal and intertemporal restrictions of the model. We then present our empirical results.

A. A Solution for Expenditure Growth

We employ the normalization $P_1(t)\equiv\theta_1\equiv1$. Define the ratio of budget shares by $\omega^h(t)=P_2^h(t)(C_2^h(t)-\gamma_2)/(C_1^h(t)-\gamma_1)$. Then, from (17)

$$\omega^h = \theta_2^{1/\alpha_2} P_2^{h\,(1-1/\alpha_2)} (C_1^h-\gamma_1)^{(\alpha_1/\alpha_2-1)} . \qquad (24)$$

This implies

$$\hat{\omega}^h(t) = (1-1/\alpha_2)\hat{P}_2^h(t) + (\alpha_1/\alpha_2-1)\hat{\tilde{C}}_1^h(t), \qquad (25)$$

where a ^ over a variable indicates the first difference of the logarithm of that variable. Note that (25) implies that ω^h grows the same rate for all households who face the same relative prices for food and non-food consumption and the same real asset returns because $\tilde{C}_1^h$ grows at the same rate for such households. Once $\hat{\omega}^h$ and $\hat{\tilde{C}}_1^h$ are determined, $\hat{\tilde{E}}^h$ for the household with $\omega^h(t)$ is determined by the identity $\hat{\tilde{E}}^h = \hat{\tilde{C}}_1^h + \ln(1+\omega^h(t)\exp(\hat{\omega}^h))-\ln(1+\omega^h(t))$. We will assume that all households face the same relative prices and same asset returns in each village in the ICRISAT data, but we will allow these prices and asset returns to vary across villages. Under this assumption, the growth rate of food consumption in excess of subsistence consumption should be the same across all consumers in each village. Thus we can use equation (25) to test the intertemporal implications of the model by estimating this equation after substituting the growth of aggregate food consumption in excess of subsistence expenditure in each village into equation (25) for the growth of individual food consumption in excess of subsistence expenditure.

B. Econometric Method

We use the intraperiod FOC (24) and Relation (25) to estimate curvature parameters. These relations should be exact when all the variables are measured exactly. We assume that the log of the relative price and $\ln(C_i^h(t)-\gamma_i)$ are measured with errors as in (17). Measurement errors are assumed to be independent across households.

From (25), we have the interperiod regression

$$\hat{\omega}^h(t) = b_0 + b_1\hat{P}_2(t) + b_2\phi(t) + \epsilon_1^h(t) \tag{26}$$

for h=1,...,N and t=1,...,T, where $b_1=1-1/\alpha_2$ and $b_2=\alpha_1/\alpha_2$. Here $\phi(t) = \hat{\tilde{C}}_1^*(t)$ is the common growth rate of $C_1-\gamma_1$. We estimate $\phi(t)$ and γ_1 jointly with the curvature parameters by applying the method used in Section IV to food consumption with β_y restricted to be zero. The disturbance term $\epsilon_1^h(t)$ consists of measurement errors of the relative price and consumption of the two goods. To construct $\hat{\omega}^h(t)$, we also need an estimate of γ_2. We use the estimate of γ_2 obtained by applying the method in Section IV.[18] The

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[18] We also tried a joint estimation of γ_2 with the other parameters. The covariance matrix of the disturbance was close to singular and we were not

impact of this first step estimation on our inferences is taken into account by the two-step GMM procedure described in Ogaki (1991b).

We also estimate the curvature parameters of our model using the information in its intratemporal restrictions. From (17), we obtain the intraperiod regression

$$\ln(C_2^h(t)-\gamma_2) = b_3 + b_4\ln(C_1^h(t)-\gamma_1) + b_5\ln(P_2(t)) + \epsilon_2^h(t) \tag{27}$$

for h=1,...,N and t=1,...,T_1, where $b_4 = \alpha_1/\alpha_2$ and $b_5 = -1/\alpha_2$. It should be noted that the intratemporal FOC (17) is valid even if the intertemporal restriction that the growth of food consumption in excess of subsistence food consumption is the same for all individuals is not valid. The disturbance term $\epsilon_2^h(t)$ also consists of measurement errors.

Let $y^h(t)$ be real income of household h at t and define

$$Y_h = [\hat{\omega}^h(1),\ldots,\hat{\omega}^h(T_1-1),\ln(C_1^h(1)-\gamma_1),\ldots,\ln(C_1^h(T_1)-\gamma_1)]',$$

$$X_h' = \begin{bmatrix} \hat{P}_2(1) & \hat{\tilde{C}}_1^*(1) & 0 & 0 & 0 \\ \cdot & \cdot & \cdot & \cdot & \cdot \\ \hat{P}_2(T_1-1) & \hat{\tilde{C}}_1^*(T_1-1) & 0 & 0 & 0 \\ 0 & 0 & 1 & \ln(C_1^h(1)-\gamma_1) & \ln(P_2(1)) \\ \cdot & \cdot & \cdot & \cdot & \cdot \\ 0 & 0 & 1 & \ln(C_1^h(T_1)-\gamma_1) & \ln(P_2(T_1)) \end{bmatrix},$$

$\epsilon_h = [\epsilon_1^h(1),\ldots,\epsilon_1^h(T_1-1),\epsilon_2^h(1),\ldots,\epsilon_2^h(T_1)]'$, and

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able to numerically invert the matrix.

$$Z_h' = \begin{bmatrix} 1 & \hat{P}_2(3) & \hat{y}^h(1) & 0 & 0 & 0 \\ . & . & . & . & . & . \\ 1 & \hat{P}_2(T_1+1) & \hat{y}^h(T_1-1) & 0 & 0 & 0 \\ 0 & 0 & & 1 & \ln(y^h(1)) & \ln(P_2(2)) \\ . & . & . & . & . & . \\ 0 & 0 & & 1 & \ln(y_1^h(T_1)) & \ln(P_2(T_1+1)) \end{bmatrix}.$$

Here Z_h is a matrix of instrumental variables that are assumed to be uncorrelated with ϵ_h across households for. Then we obtain six orthogonality conditions

$$E_H[Z_h\epsilon_h] = \underset{N\to\infty}{\text{plim}}\ (1/N)\sum_{h=1}^{N}[z_h\epsilon_h] = 0 \tag{28}$$

for each village where E_H is the expectation operator over households. Thus we have $z_h y_h = z_h x_h' b + z_h \epsilon_h$, where $b=[b_1,\ldots,b_5]'$ with two linear restrictions

$$b_1\text{-}1 = b_5 \text{ and } b_2+1 = b_4 \tag{29}$$

We estimate b, γ_1, and $\phi(t)$ from the orthogonality conditions (28) and the orthogonality conditions (23) for food with the GMM as in the last section, imposing the linear restrictions (29). As in the Section IV, the data for three villages are pooled. We restrict the preference parameters to be the same across the three villages but allow the other parameters to be different in different villages.

D. Empirical Results

Table 4 reports our results. The *J* statistic tests the overidentifying restrictions when the restrictions (29) are imposed. The *C* statistic is the difference of the *J* statistics for the unrestricted and restricted cases, which tests the restriction (29). This is a specification test for our model. If preferences are not separable for food and nonfood, then the intraperiod regression (27) still consistently estimates $1/\alpha_2$ and α_1/α_2. This is because the intraperiod relation (17) holds for any monotonic transformation of the intraperiod utility function. On the other hand, the interperiod regression (26) does not estimate $1/\alpha_2$ and α_1/α_2 consistently if either preferences are not separable or the assumption of market completeness is seriously in error. Thus our *C* test is a specification test a la Hausman (1978). The *K* statistic is the Wald test for the restriction $\alpha_1=\alpha_2=1$ implied by the linear expenditure system.

The curvature parameters are estimated with the theoretical correct positive sign. We estimate α_1/α_2 to be significantly larger than one at the five percent level (in the sense of one-sided test for $\alpha_1/\alpha_2>1$). The *J* test indicates that the instruments are not correlated with the disturbance. The *C* test does not reject the separability assumption. There is overwhelming evidence against the linear expenditure system in terms of the *K* test.

VI. Estimation with The Cointegration Approach

In this section we estimate the curvature parameters of the model with aggregate Indian and U.S. time series data using Ogaki and Park's (1990) cointegration approach. The reader is referred to their paper for a detailed explanation of the technique and terminology. The data are explained in the Appendix.

A. The Cointegration Restriction

Aggregate consumption satisfies condition (27) for the fictitious representative consumer with θ_2 replaced by θ_2^*. We assume that $\ln(C_i^*(t)-\gamma_i)$ is difference stationary with a positive drift and a positive long-run variance for i=1,2 and measurement errors for $\ln(C_i^*(t)-\gamma_i)$ and $\ln(P_2(t))$ are stationary, so that $\epsilon_3(t)$ is stationary. Then $\ln(C_2^*(t)-\gamma_2)$, $\ln(C_1^*(t)-\gamma_1)$, and $\ln(P_2(t))$ are cointegrated with the deterministic cointegration restriction (defined in Ogaki and Park (1990, section 2)). Parameters $\phi_2=\alpha_1/\alpha_2$ and $\phi_3=-1/\alpha_2$ are estimated consistently by a cointegrating regression.[19]

B. Econometric Method

We use Park's (1990) Canonical Cointegrating Regression (CCR) to estimate (27), imposing the deterministic cointegration restriction. Some correlation parameters including the long-run covariance matrix of the system are estimated by Park and Ogaki's (1991) prewhitening method. The CCR estimator is asymptotically efficient. There are other asymptotically efficient methods to estimate cointegrating vectors, but there are several reasons to use the CCR. First, Monte Carlo studies by Park and Ogaki (1991) show that the CCR estimator with prewhitening performs better than Johansen's (1989) ML and the CCR estimator without prewhitening when the sample size is small. Second, the CCR does not need strong distributional assumptions such as the Gaussian VAR assumption. Third, the CCR can be used to test the null hypothesis of the deterministic cointegration restriction

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[19]Even if one of the terms $\ln(C_i^*(t)-\gamma_i)$ is trend stationary, it is still possible to estimate curvature parameters by cointegrating regressions (see Ogaki and Park (1990)). If both of $\ln(C_i^*(t)-\gamma_i)$'s are trend stationary, it is no longer possible to identify curvature parameters only from trends. Deterministic trends, however, still contain useful information about these parameters that can be exploited (see Ogaki (1988, 1989)).

and that of stochastic cointegration.

As a specification test, we test the deterministic cointegration restriction by testing if the coefficient of a spurious linear trend term is zero. Park's (1990) variable addition tests based on spurious deterministic trends are used to test the null of stochastic cointegration.

We use estimates of γ_1 and γ_2 from the ICRISAT data in these cointegrating regressions. Because we assume $\ln(C_i^*(t)-\gamma_i)$ is difference stationary with a positive drift and because γ_i is a fixed number, $\ln(C_i^*(t)-\gamma_i)$ is asymptotically equal to $\ln(C_i^*(t))$ for i=1,2. Thus estimation of γ_i's does not affect asymptotic distribution. The distribution is affected in small samples, however. Therefore, we report sensitivity analysis with respect to the choice of γ_i's.

C. Empirical Results

1. Indian Time Series

For the Indian aggregate time series data, we use two measures of nonfood consumption: NFC1 and NFC2. The differences between these two series is explained in Appendix A. The aggregate series NFC1 corresponds more closely with the measure of nonfood consumption data we use in the ICRISAT, but NFC2 contains broader categories of consumption and is closer to the aggregate data we use for the United States.

For subsistence levels, we use two sets of estimates. One set of estimates is from Table 2 and Table 3 in the present paper. In the present paper, we restrict our estimates of γ to be below the minimum consumption level observed in the data by assuming a positive multiplicative error as in (17). This assumption is necessary for the joint estimation with the curvature parameters in Section V and for the calculation of the IES for all households for all periods in the ICRISAT data. Since the purpose for using NFC1 is to compare results in the ICRISAT data with those in the aggregate data, we use our estimates in Tables 2 and 3 (γ_1=101.5, γ_2=26.8) for results for NFC1. If the assumption if positive additive measurement error is misspecified, however, it can bias our estimates of γ downward. Ogaki and Atkeson (1991) estimate subsistence levels with an additive measurement error model to allow estimates of γ to be greater than the minimum consumption level observed. Their estimate of γ_1 is 64.2 with standard error of 9.1 and that of γ_2 is 206.5 with standard error of 37.2. These estimates are much larger than our estimates of the present paper, which suggests that our estimates may be biased downward.[20] Therefore we use Ogaki and Atkeson's (1991) estimates for the results reported in the tables for NFC2. These estimates of γ_1 and γ_2 are converted to appropriate units as described in Appendix A.

We first test the null of a unit root for $\ln(C_i^*(t)-\gamma_i)$. We use Park and Choi's (1988) J(1,5) test to perform this test. This test has an advantage that no correction for the long-run variance is necessary. This is attractive because results for the Augmented Dickey-Fuller (ADF) tests (Dickey and Fuller (1979)) are often very sensitive to the order of AR used and because Phillips and Perron's (1988) nonparametric correction seem to lead to serious size distortion in some cases. Park and Choi's Monte Carlo studies show that J(1,5) test performs well compared with these tests in

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[20]Rosenzweig and Wolpin (1991) estimate γ in a single consumption good model without using consumption data by analyzing investments in bullocks in the ICRISAT data. The sum of Ogaki and Atkeson's (1991) estimates of γ_1 and γ_2 multiplied by the average family size of 6 is closer to Rosenzweig and Wolpin's estimate of γ than that of our estimates in Tables 2 and 3.

terms of size distortion and size adjusted power. The values of J(1,5) are 1.17 for ln(food-206.5) and 1.19 for ln(food-101.5), 0.92 for NFC1, and 6.21 for NFC2. We do not reject the null of a unit root at the ten percent level for any of these series.

The first panel of Table 5 reports results for NFC1 and the second panel, those for NFC2. The results for NFC1 are similar to those for NFC2. There is little evidence against the model in terms of the H(0,1) test that tests the null of the deterministic cointegration restriction and the H(1,2), H(1,3), and H(1,4) tests that test the null of stochastic cointegration. All estimates of curvature parameters $1/\alpha_2$ and α_1/α_2 have the theoretically correct positive sign. The K statistic tests the restriction $\alpha_1=\alpha_2=1$ implied by the LES, and finds overwhelming evidence against this restriction. The estimates for α_1/α_2 are significantly greater than one.[21]

In Table 6, we compare estimates of the curvature parameters from Indian time series for NFC1 reported in Table 5 with those from the ICRISAT data. We apply the method of Minimum Distance estimation (see, e.g., Chamberlain (1984)) to estimates the curvature parameters from these two sets of estimators. We assume that measurement errors in these two data sets are independent and thus the two sets of estimators are uncorrelated. We impose the restrictions implied by our aggregation result that the curvature and subsistence parameters of the model should be the same in

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[21]As a sensitivity analysis, we also used estimates of γ_1 and γ_2 reported in Tables 2 and 3 for NFC2. The results were little affected by this choice of subsistence levels. For example, $1/\alpha_2$ and α_1/α_2 were estimated to be 0.443 and 2.197, respectively.

individual and aggregate level data. We minimize a quadratic form for the difference between unrestricted estimates and restricted estimates with the weighting matrix equal to the inverse of the covariance matrix for the unrestricted estimates. We test the restriction by the minimized value by J, which has asymptotic chi-square distribution with two degrees of freedom under the null that preferences are identical. For estimates from the Indian time series, we use results for NFC1 because NFC1 corresponds to our nonfood consumption data in the ICRISAT data.

The first panel of Table 6 reports results with the restriction that the value of α_1/α_2 in the time series data is identical with that in the panel data. We do not find much evidence against this restriction. The second panel of Table 6 reports results when two restrictions that $1/\alpha_2$ and α_1/α_2 in the time series data are identical with those in the panel data. We do find evidence against these two restrictions. Thus, our estimates of the level of the IES are not consistent across these two data sets, but, as we shall discuss in the next section, we can compare the ratio of the IES between two agents with confidence that the results of such a comparison are invariant to the data set used to estimate the parameters of the model.

2. U.S.Time Series

We first test the null of a unit root for $\ln(C_i^*-\gamma_i)$ with the J(1,5) test as before. The values of the J(1,5) are 1.20 for food, 1.05 for nonfood consumption. We do not reject the null of a unit root at the ten percent level for any of these series. Since we compare results from the U.S. time series with those for NFC2 in the Indian time series, we convert the estimates of subsistence levels used for the regression with NFC2 into appropriate units as described in Appendix A.

Table 7 reports results. There is evidence against the model in terms of the H(0,1) test that tests the null of the deterministic cointegration restriction. This restriction is rejected at the one percent level, but the evidence against this restriction is not overwhelming. There is little evidence against the model in terms of the H(1,2), H(1,3), and H(1,4) tests. All estimates of curvature parameters $1/\alpha_2$ and α_1/α_2 have the theoretically correct positive sign. The K statistic tests the restriction $\alpha_1=\alpha_2=1$ implied by the LES, which finds overwhelming evidence against this restriction. The estimates for α_1/α_2 are significantly greater than one, implying that preferences are not quasi-homothetic.

3. An International Comparison of Preferences

In this section, we compare our estimates of curvature parameters for Indian households and our estimates of those for U.S. households. We apply the method of Minimum Distance estimation to estimates from the Indian time series data and estimates from the U.S. time series data as we did to estimates from the Indian time series and Indian panel data. For estimates from the Indian time series, we use results for NFC2 because NFC2 corresponds with our nonfood consumption data in the U.S. time series data. Subsistence levels used are estimates in Ogaki and Atkeson (1991). Table 8 reports results. There is little evidence against the hypothesis that preferences are identical for Indian and U.S. households.[22]

VII. Measuring Intertemporal Elasticities of Substitution

In this section, we use our parameter estimates to compare the measured

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[22]This result is consistent with that of Rhee and Rhee (1990) who do not find cultural effects on saving.

intertemporal elasticity of substitution of total consumption expenditure across households and countries of different wealth levels. We use Relation (14) for this purpose. While precise measurement of any individual's intertemporal elasticity of substitution of consumption expenditure requires precise estimates of all four parameters $\alpha_1, \alpha_2, \gamma_1$, and γ_2, the ratio of the intertemporal elasticity of substitution for any two individuals can be accurately measured with knowledge only of α_1/α_2 and the levels of subsistence expenditure γ_1 and γ_2. In order to see this, note that (14) implies

$$\sigma^h(t) = (1/\alpha_1)\left[\frac{\alpha_1}{\alpha_2} + \tilde{\omega}_1^h(t)\left(1-\frac{\alpha_1}{\alpha_2}\right)\right] \left(1 - \frac{\gamma_1 + P_2(t)\gamma_2}{E^h(t)}\right) \qquad (14').$$

Thus the ratio of the IES for any two consumers depends only on our estimates of α_1/α_2, γ_1, and γ_2.

Table 9 presents our results for the Indian panel data. In the first panel, we report the range of intertemporal elasticities of substitution for individual households in the Indian Panel Data implied by estimates of $1/\alpha_2$=0.498, α_1/α_2=1.252 reported in Table 6 and γ_1=205, and γ_2=67 reported in Tables 2 and 3. For each year, we calculate the IES for all households in the sample used in our econometric work and report the mean of the IES of the six households with the smallest IES and that of the six households with greatest IES. We also report the ratio of the greatest and the smallest IES. The differences in the IES of the rich and the poor in the Indian data are substantial. On average, the IES of the richest six households is about 1.6 times that of the poorest six households.

Using the aggregation result presented in Section III, we can also measure the IES that is relevant for aggregate consumption. The second panel of Table 9 reports the IES of the fictitious representative consumer for each of the three villages in the Indian panel data. We calculate the IES for each village, using all households in our sample. In the last column, we report the average IES over the period 1976-1981 for each village. The IES of Shirapur and that of Kanzara are very similar while that of Aurepalle is smaller than the IES of the other two villages. The IES fluctuates over time but the IES of Aurepalle is always smaller that of the other two villages. For comparison, with the same parameter values and the time series data with NFC1, the average IES of the aggregate India is 0.57 over the period 1960-1987.

Table 10 reports measurements of the IES obtained from the aggregate data for the U.S. and India. The estimates reported in Table 8 and the estimates of Ogaki and Atkeson (1991) are used for preference parameters. The IES is estimated to be stable over the period 1929-1988 in the U.S: it rises from 0.38 in 1929 to 0.41 in 1988. For India, we use the series NFC2. The estimated IES is also stable, rising gradually from 0.27 in 1960 to 0.29 in 1987.

Since the IES in the U.S. implied our estimates has been stable in the long-run while the IES among individuals in the Indian panel data are very different, our model is consistent with the empirical puzzle that the time series for the aggregate saving rate is relatively stable in the long-run in the U.S. while the saving rate rises in cross-sectional data. It is well known that the life-cycle hypothesis and the permanent income hypothesis were motivated in part by this empirical puzzle. In our model, the fictitious representative consumer for the U.S. economy has been rich enough that changes in the IES implied by growth has been negligible. On the other hand, we should be able to observe that the saving rate rises as permanent income rises among poor households for whom subsistence levels are not negligible. Though it is not easy to discriminate between low saving rates induced by a temporary drop in income and low saving rates induced by low level of wealth (or permanent income), there exists some evidence from micro data that the saving rate rises as the permanent income increases (see, e.g., Bhalla (1980) and Paxson (1992)).

The IES in India is estimated to be lower than that in the U.S.[23] In order to compare our estimate of the ratio of the IES between the U.S. and India with a measure of the average difference in consumption growth between low and high income countries, we examine the Summers and Heston (1991) data set. We regress the average annual growth of real consumption per adult equivalent on a constant and $1/y_p$, where y_p is a measure of permanent income constructed as an average of real GDP per equivalent adult over the sample period.[24] We exclude the countries with negative average consumption growth in the sample; the four countries with historically planned economies, China, Hungary, Poland and Yugoslavia; and the four major oil producing countries with high y_p, Bahrain, Kuwait, Saudi Arabia,and United Arab Emirates. We use $1/y_p$ rather than y_p because our model implies that the IES

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[23]This result is consistent with that of Giovannini (1985), who estimates the IES in low income countries in aggregate data to be lower than estimates of the IES in U.S. aggregate data (see, e.g., Hansen and Singleton (1982, 1983). Hall (1988), however, estimates the IES to be low in aggregate U.S. data.

[24]Real consumption per equivalent adult, RCPEA, is calculated as RGDPEA (Real GDP per equivalent adult) $\times$ (Real consumption as a percentage of RGDPCH). The average annual growth rate of RCPEA from period 1 to period T is calculated as (ln(RCPEA(T))-ln(RCPEA(1)))/(T-1).

increases and then stabilizes as the level of wealth increases. Table 11 reports our results. Statistically significantly negative coefficients on $1/y_p$ imply that the consumption growth of low income countries tends to be slower than that of high income countries. Using the fitted curve for the 1960-85 sample period and y_p for the U.S. and India, the fitted consumption growth for the U.S. is about 2.6 percent and that for India is about 1.3 percent. Thus the ratio of consumption growth of a high income country as the U.S. and that a low income country as India is about 2.[25]

The IES in India is estimated to be about 0.27 and that in the U.S., about 0.40 over the period 1960-87. Thus the ratio of the IES in the U.S. and that in India is estimated to be about 1.5. Thus the difference in the IES between India and the United States appears to be economically significant, but it does not seem to be large enough to explain the entire portion of the systematic differences in consumption growth between low income and high income countries.

VIII. Conclusion

In this paper, we have developed and estimated a model of preferences which allows for systematic variation in the intertemporal elasticity of substitution of consumption expenditure for poor and rich consumers. This model formalizes the intuition that poor consumers have a lower intertemporal elasticity of substitution than do rich consumers because poor consumers spend a higher fraction of their total expenditure on subsistence

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[25] The average real total consumption expenditure growth (deflated by the implicit deflator for total consumption expenditure) over the period 1960-85 in national accouts explained in Appendix A for the U.S. is about 2.3 percent and that for India is about 1.2 percent.

expenditure and necessary goods than do rich consumers and these expenditure inelastic goods are intrinsically less substitutable over time than are expenditure elastic goods.

We estimated our model in panel data on consumption of Indian households using both intratemporal and intertemporal restrictions of the model. We checked the specification of the model by comparing the parameter estimates obtained with these two different restrictions of the model. We re-estimated the parameters of the model in aggregate time series data from India and compared these parameter estimates in aggregate data with the estimates obtained in the household level panel data. We found that the key parameters for comparing the ratio of poor and rich consumer's (or country's) IES were consistent across the two data sets. We also estimated the parameters of the model in U.S. time series data and found that the estimates in aggregate data from India and the U.S. were consistent. We conclude from these results that specification error and aggregation bias do not invalidate the use of this model for measuring the differences in the intertemporal elasticity of substitution across poor and rich consumers at either the individual or the aggregate level.

We find economically significant differences in the intertemporal elasticity of substitution of poor and rich consumers in the Indian data. We also find significant differences in the intertemporal elasticity of substitution measured in aggregate data for the U.S. and India.

Appendix A

We explain the data in this appendix.

A. The ICRISAT Panel Data

We use food including milk, sweets and spices as the measure of food consumption. For nonfood consumption, we subtracted food and ceremonial expenses from total consumption expenditure. Ceremonial expenses are removed because they often jump from zero to large amounts. Nonfood consumption consists of narcotics, tea, coffee, tobacco, pan, alcoholic beverages; clothing, sewing of cloth, other tailoring expenses, thread, needles, chappals and other footwear etc; travel and entertainment; and 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 ghee); and others, including complete meals in hotels, school and educational materials, stamps, stationery, grinding and milling charges, etc. Unfortunately, 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 per male adult equivalent, nominal consumption at t is divided by the family size measure constructed by Townsend (1991) and the corresponding price index at t for each village. The price index for total consumption expenditure for the single good model is the consumer price index. The price index for food is used for food and the price index for nonfood is used for nonfood. These real variables are valued at 1983 prices. Nominal income variables are deflated by the consumer price index for the single good model and by the food price index for the two good model.

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. There is one household in the village of Aurepalle with zero income in 1980. Because we take the log of income, this household is excluded. The number of households in our sample for the village of Aurepalle is 35; that for Shirapur, 33; and that for Kanzara, 36.

B. Indian Time Series Data

Annual data in the National Accounts (India (1989, 1990)) that cover the fiscal years from 1960-61 to 1987-88 are used for the Indian time series data. We use food as the first good. Two measures of real nonfood consumption are constructed in the following way. The first measure is called NFC1 and consists of beverages and tobacco; clothing and footwear; fuel and power; services (in category number 4 in the National Account), recreation, education, and cultural services; and medical care and health services. The second measure is called NFC2 and defined as private consumption expenditure minus food. Price index series for each consumption series are constructed by dividing the nominal consumption series by the real consumption series (valued at 1980 prices). Relative price data are obtained by dividing the food price index by the nonfood price index of NFC1 or NFC2. Real consumption per equivalent adult series are obtained by dividing real consumption series by the equivalent-adult population for India in Summers and Heston (1991).

The difference between NFC2 and NFC1 include housing and transportation. For NFC1, the estimate for the subsistence level for nonfood consumption from the ICRISAT data valued at 1983 prices is converted into the level valued at 1980 prices with the average of the nonfood price indexes for the three villages in the ICRISAT data. For NFC2, the subsistence level for NFC1 is multiplied 1.7, which is the average ratio of real NFC2 and real NFC1.

C. U.S. Time Series Data

Annual data in the National Income and Product Accounts (NIPA) that cover 1929-1988 are used for the U.S. time series data. The proximate source is the 1989 NIPA diskette that extends U.S. (1986). We use food excluding alcoholic beverages (food for short) as the first good. We use nonfood consumption, defined as personal consumption expenditure minus food as the second good. Price index series for each consumption series are constructed by dividing the nominal consumption series by the real consumption series valued at 1982 prices. Relative price data are obtained by dividing the food price index by the nonfood price index. We divide real consumption by the equivalent adult population for the U.S. in Summers and Heston's (1991) data to obtain real consumption per equivalent adult.

We need to convert subsistence level estimates in terms of 1983 rupees into those in terms of U.S. dollar. For this purpose, we use the purchasing power parity (PPP) data in Kravis, Heston, and Summers (1982) for 1975 rupees and 1975 U.S. dollars. First, we convert estimates of subsistence levels from the ICRISAT data valued at 1983 prices into the level valued at 1975 prices with the average price indexes for the three villages in the ICRISAT data. The estimate for nonfood is multiplied by 1.7 as we did for NFC2 in the Indian data. Second, we convert 1975 rupees into 1975 dollars by dividing subsistence levels in 1975 rupees by the PPP's reported in Table 6.3 of Kravis, Heston, and Summers. We construct the PPP for nonfood by total consumption expenditure (SNA concept) minus food in terms of rupees in Table 6.1 of Kravis, Heston, and Summers bu that in terms of 1975 U.S. dollar that is obtained by dividing total consumption and food by the corresponding PPP. We then convert 1975 dollars into 1982 dollars using the price index of each category.

Appendix B

In this Appendix, we show that (7) approximately holds. This result immediately follows from the following lemma that a strictly increasing or decreasing function mapping the positive real line to the positive real line can be approximated by a power function which takes on the original function's value and slope at a point.

Lemma 1: Let $y=f(x)$ be a differentiable function mapping positive x into positive y. Assume that $f'(x)\neq 0$. Then for any $\hat{x}>0$, we can find parameters $A(\hat{x})$ and $\mu(\hat{x})$ such that $A\hat{x}^{\mu} = f(\hat{x})$ and $A\mu\hat{x}^{(\mu-1)} = f'(\hat{x})$.

Proof: Case 1. Assume $f'(\hat{x})>0$. Then we must have $\mu>0$. We can solve for A and μ from equations

$$log(A) + \mu\ log(\hat{x}) = log(f(\hat{x}))$$

$$log(a) + log(\mu) + (\mu-1)\ log(\hat{x}) = log(f'(\hat{x}))$$

The first equation gives

$$\mu = \frac{log(f(\hat{x})) - log(A)}{log(\hat{x})}$$

Since $\mu>0$, we can use this expression and the second equation to derive a single equation in μ

$$log(\mu)+log(f(\hat{x}))-log(\hat{x})-log(f'(\hat{x}))=0$$

Thus this equation has a unique solution in A and μ. Case 2, when $f'(\hat{x})<0$ and $\mu<0$, can be handled in the same fashion by solving for $-\mu$ using $-f'(\hat{x})$.

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Table 1

REAL TOTAL CONSUMPTION EXPENDITURE GROWTH IN THE ICRISAT PANEL DATA

γ	s.e.	β_y	s.e.	J[a]	d.f.	p-value[b]	C[c]	d.f.	p-value[b]
177.6	6.70	-0.023	0.053	34.46	28	18.6	. . .		. . .
177.2	7.45	0	. . .	34.70	29	21.5	0.237	1	62.6
0	. . .	0	. . .	98.89	30	0.0	64.428	2	0.0

$\phi^a(1)$	s.e.	$\phi^a(2)$	s.e.	$\phi^a(3)$	s.e.	$\phi^a(4)$	s.e.	$\phi^a(5)$	s.e.
0.017	0.198	0.163	0.041	0.475	0.052	-0.020	0.053	-0.150	0.068

$\phi^s(1)$	s.e.	$\phi^s(2)$	s.e.	$\phi^s(3)$	s.e.	$\phi^s(4)$	s.e.	$\phi^s(5)$	s.e.
-0.124	0.034	0.050	0.049	-0.129	0.062	0.147	0.057	-0.095	0.062

$\phi^k(1)$	s.e.	$\phi^k(2)$	s.e.	$\phi^k(3)$	s.e.	$\phi^k(4)$	s.e.	$\phi^k(5)$	s.e.
-0.008	0.036	-0.087	0.052	0.358	0.045	-0.139	0.034	-0.143	0.036

[a]Chi-square test statistics for the overidentifying restrictions.
[b]In percentage.
[c]Chi-square test statistics for the restrictions imposed.

Table 2

REAL FOOD CONSUMPTION GROWTH IN THE ICRISAT PANEL DATA

γ	s.e.	β_y	s.e.	J[a]	d.f.	p-value[b]	C[c]	d.f.	p-value[b]
101.4	4.30	-0.083	0.360	30.69	30	33.1	. . .	. . .	
101.5	3.70	0	. . .	32.28	29	30.8	1.597	1	20.6
0	. . .	0	. . .	56.93	28	0.0	26.247	2	0.0
$\phi^a(1)$	s.e.	$\phi^a(2)$	s.e.	$\phi^a(3)$	s.e.	$\phi^a(4)$	s.e.	$\phi^a(5)$	s.e.
0.362	0.077	0.057	0.034	0.383	0.050	-0.090	0.050	-0.274	0.049
$\phi^s(1)$	s.e.	$\phi^s(2)$	s.e.	$\phi^s(3)$	s.e.	$\phi^s(4)$	s.e.	$\phi^s(5)$	s.e.
-0.101	0.044	0.146	0.059	-0.193	0.063	0.158	0.058	-0.216	0.075
$\phi^k(1)$	s.e.	$\phi^k(2)$	s.e.	$\phi^k(3)$	s.e.	$\phi^k(4)$	s.e.	$\phi^k(5)$	s.e.
-0.025	0.040	-0.190	0.053	0.375	0.035	-0.051	0.043	-0.152	0.063

[a] Chi-square test statistics for the overidentifying restrictions.
[b] In percentage.
[c] Chi-square test statistics for the restrictions imposed.

Table 3

REAL NONFOOD CONSUMPTION GROWTH IN THE ICRISAT PANEL DATA

γ	s.e.	β_y	s.e.	J^a	d.f.	p-value[b]	C^c	d.f.	p-value[b]
28.8	1.44	-0.014	0.059	28.17	30	45.6	. . .		. . .
26.8	1.44	0	. . .	28.22	29	50.6	0.053	1	81.9
0	. . .	0	. . .	35.69	28	0.0	35.687	2	0.0

$\phi^a(1)$	s.e.	$\phi^a(2)$	s.e.	$\phi^a(3)$	s.e.	$\phi^a(4)$	s.e.	$\phi^a(5)$	s.e.
-0.970	0.124	0.828	0.083	0.294	0.072	0.047	0.065	0.124	0.118

$\phi^s(1)$	s.e.	$\phi^s(2)$	s.e.	$\phi^s(3)$	s.e.	$\phi^s(4)$	s.e.	$\phi^s(5)$	s.e.
-0.141	0.047	0.051	0.071	0.050	0.068	0.021	0.066	0.060	0.079

$\phi^k(1)$	s.e.	$\phi^k(2)$	s.e.	$\phi^k(3)$	s.e.	$\phi^k(4)$	s.e.	$\phi^k(5)$	s.e.
-0.027	0.039	0.043	0.054	0.234	0.056	-0.211	0.025	-0.153	0.039

[a] Chi-square test statistics for the overidentifying restrictions.
[b] In percentage.
[c] Chi-square test statistics for the restrictions imposed.

TABLE 4

ESTIMATION OF CURVATURE PARAMETERS FROM THE ICRISAT PANEL DATA

α_1/α_2[a]	$1/\alpha_2$[a]	γ_1[a]	γ_2[a]	J[b]	C[c]	K[d]
1.151 (0.080)	5.409 (0.795)	86.10 (20.81)	26.8 (1.44)	43.39 (28.9)	1.887 (38.9)	37.48 (0.0)

[a]Standard errors are in parentheses.

[b]Chi-square test statistic for the overidentifying restrictions with 39 degrees of freedom.

[c]Chi-square test statistic with two degrees of freedom for the restriction across intratemporal and intertemporal regressions. P-value in percentage is in the parenthesis.

[d]Wald test statistic with two degrees of freedom for the restriction $\alpha_1=\alpha_2=1$. P-value in percentage is in the parenthesis.

TABLE 5

CANONICAL COINTEGRATING REGRESSION RESULTS FROM INDIAN TIME SERIES

$1/\alpha_2$[a]	α_1/α_2[a]	K[b]	$H(0,1)$[c]	$H(1,2)$[d]	$H(1,3)$[d]	$H(1,4)$[d]
NFC1						
0.582 (0.185)	2.098 (0.336)	18.14 (0.0)	0.403 (52.5)	0.947 (33.0)	1.711 (42.5)	3.238 (35.6)
NFC2						
0.483 (0.129)	2.119 (0.250)	34.17 (0.0)	0.448 (50.3)	0.003 (95.8)	0.396 (82.0)	2.696 (44.1)

[a]Standard errors are in parentheses.

[b]Chi-square test statistic with two degrees of freedom for the restriction $\alpha_1=\alpha_2=1$. P-values in percentage are in parentheses.

[c]Chi-square test statistic with one degree of freedom for the deterministic cointegration restriction. P-values are in parentheses.

[d]Chi-square test statistics for stochastic cointegration. P-values are in parentheses.

TABLE 6

MINIMUM DISTANCE ESTIMATION FOR INDIAN PANEL AND INDIAN TIME SERIES DATA

$1/\alpha_2$	s.e.	α_1/α_2	s.e.	J[a]	p-value
One Restriction on α_1/α_2					
. . .	. . .	1.202	0.078	7.519	2.3
Two Restrictions on $1/\alpha_1$ and α_1/α_2					
0.713	0.176	1.262	0.077	43.27	0.0

[a]Chi-square test statistics with one degree of freedom in the first panel and two degrees of freedom in the second panel.

TABLE 7

CANONICAL COINTEGRATING REGRESSION RESULTS FROM U.S. TIME SERIES

$1/\alpha_2$[a]	α_1/α_2[a]	K[b]	$H(0,1)$[c]	$H(1,2)$[d]	$H(1,3)$[d]	$H(1,4)$[d]
-0.075	2.283	29.13	8.102	0.002	0.336	2.834
(0.501)	(0.250)	(0.0)	(0.4)	(96.1)	(85.6)	(41.8)

[a]Standard errors are in parentheses.
[b]Chi-square test statistic with two degrees of freedom for the restriction $\alpha_1=\alpha_2=1$. P-value is in the parenthesis.
[c]Chi-square test statistics with one degree of freedom for the deterministic cointegration restriction. P-values are in parentheses.
[d]Chi-square test statistics for stochastic cointegration. P-values are in parentheses.

TABLE 8

MINIMUM DISTANCE ESTIMATION FOR INDIAN AND U.S. TIME SERIES DATA

$1/\alpha_2$	s.e.	α_1/α_2	s.e.	J[a]	p-value
0.450	0.125	2.156	0.171	1.199	54.9

[a]Chi-square test statistic with two degree of freedom. This statistic tests the restrictions that curvature parameters are the same for households in two countries.

TABLE 9

INTERTEMPORAL ELASTICITY OF SUBSTITUTION IN THE ICRISAT PANEL DATA

	1976	1977	1978	1979	1980	1981	Mean
Individual Households							
Poorest 6 households	0.33	0.36	0.40	0.42	0.41	0.36	0.38
Richest 6 households	0.60	0.58	0.59	0.60	0.59	0.59	0.59
Ratio of the IES	1.82	1.60	1.47	1.44	1.43	1.65	1.57
Aggregate at the Village Level							
Aurepalle	0.47	0.44	0.47	0.51	0.56	0.50	0.48
Shirapur	0.54	0.54	0.55	0.53	0.56	0.55	0.55
Kanzara	0.54	0.54	0.53	0.55	0.54	0.53	0.27

NOTE: The intertemporal elasticities of substitution for total consumption expenditure are calculated. Subsistence levels used are γ_1=101.5 and γ_2=26.8 reported in Table 2 and Table 3. Estimates of the curvature parameters used are $1/\alpha_2$=0.713 and α_1/α_2=1.262 in the second panel of Table 4.

TABLE 10

INTERTEMPORAL ELASTICITY OF SUBSTITUTION, AGGREGATE DATA

U.S.								
1929	1939	1949	1959	1969	1979	1988	Mean	Mean over 1960-87
0.38	0.38	0.38	0.39	0.40	0.40	0.41	0.39	0.40

India							
1960	1965	1970	1975	1980	1985	1987	Mean
0.27	0.27	0.27	0.27	0.28	0.29	0.29	0.27

NOTE: The intertemporal elasticities of substitution for total consumption expenditure are calculated. Subsistence levels used are based on estimates of Ogaki and Atkeson (1991) as explained in the text. The curvature parameters used are $1/\alpha_2$=0.450 and α_1/α_2=2.156 reported in Table 8.

TABLE 11

REAL TOTAL CONSUMPTION EXPENDITURE GROWTH IN THE SUMMERS-HESTON DATA

Sampler Period	Sample Size	$1/y_p$	t-statistic[a]	R^2
1960-1985	97	-11.11	-4.69	0.74
1960-1972	98	-9.06	-3.32	0.77
1973-1985	83	-2.06	-2.06	0.69

NOTE: Average annual growth of real per equivalent adult consumption in the Summers-Heston (1991) data is regressed on a constant and the reciprocal the permanent income measure ($1/y_p$) explained in the text.

[a]White's (1980) heteroskedasticity-consistent covariance matrix estimator is used to calculate these t-statistics.