Rochester Center for
Economic Research

The Rate of Time Preference, the Intertemporal Elasticity of Substitution, and the Level of Wealth

Ogaki, Masao and Andrew Atkeson

Working Paper No. 363
October 1993

University of
Rochester

The Rate of Time Preference, the Intertemporal Elasticity of Substitution, and the Level of Wealth

Masao Ogaki and Andrew Atkeson

Rochester Center for Economic Research
Working Paper No. 363

The Rate of Time Preference, the Intertemporal Elasticity of Substitution, and the Level of Wealth

Masao Ogaki
University of Rochester

Andrew Atkeson
University of Chicago

First Draft: May 1993
This Revision: July 1993

Some of the results presented were included in earlier versions of Atkeson and Ogaki (1993a). We thank seminar participants at many institutions for comments at various stages of this work and R. Anton Braun for his comments that motivated us to write this paper. We gratefully acknowledge financial support by National Science Foundation grant no. SES-9213930.

The Rate of Time Preference, the Intertemporal Elasticity of Substitution, and the Level of Wealth

Abstract

The rate of time preference (RTP) and the intertemporal elasticity of substitution (IES) are two important factors for intertemporal consumption decisions. Empirical and policy implications for economic development, growth, and income distributions of the models in which the RTP and/or the IES change systematically between rich and poor households are different from those of standard models in which these are constant. In this paper, we estimate a model in which both RTP and IES are allowed to be different between rich and poor households with Indian household level panel data. Our empirical results are consistent with the view that the RTP is constant across poor and rich households, but the IES is larger for the rich than it is for the poor.

Masao Ogaki
Department of Economics
University of Rochester
Harkness 231
Rochester, NY 14627-0156

Andrew Atkeson
Department of Economics
University of Chicago
1126 E. 59th St.
Chicago, IL 60637

I. Introduction

The rate of time preference (RTP) and the intertemporal elasticity of substitution (IES) are two important factors for intertemporal consumption decisions. In the theoretical literature, many authors havfe studied models in which which the RTP changes with the level of wealth (see, e.g., Uzawa (1968) and Epstein (1983)). We call these wealh-varying RTP models. There is little empirical work in which a wealth-varying RTP model is estimated with an exception of Lawrance (1991), who estimated a wealth-varying RTP model with the Panel Study of Income and Dynamics (PSID) data set. In Lawrance's model, the IES is assumed to be constant as in the standard macroeconomic models with isoelastic utility functions. The models in which the IES varies with the level of wealth are called wealth-varying IES models. For example, the quadratic utility function implies that the IES falls as a household becomes richer. Atkeson and Ogaki (1993a) argue that there are intuitive reasons to believe that the IES should rise rather than fall as a household becomes richer[1] and provide some empirical evidence. In their empirical work, they assume that the RTP is constant.

In this paper, we estimate a model that allows both the RTP and IES to change systematically between rich and poor households with Indian panel data. Our empirical results are consistent with the view that the RTP is constant across poor and rich households, but the IES is larger for the rich than it is for the poor.

Empirical and policy implications for economic development, growth, and

[1]For example, 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.

income distributions of the models in which the RTP and/or the IES change systematically between rich and poor households are different from those of standard models in which these are constant. See, for example, Chatterjee (1991), Easterly (1991), Rebelo (1992), and Ogaki (1992a) for discussions of the implications of wealth-varying IES.

The remainder of this paper is organized as follows. In Section II, we present our model. In Section III, we discuss the econometric method used in this paper. In Section IV, we present our empirical results. Section V concludes.

II. The Model

In this section, we present the model of consumers' intertemporal allocation of consumption expenditure that we use for estimation. In particular, we discuss the different implications of wealth-varying RTP and wealth-varying IES models for consumption growth.

Consider an economy with H households, each of which consumes a good 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 $u(C^h(t))$. 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

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

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

$$\sum_{t=0}^{T} \sum_{e(t)} \prod_{\tau=0}^{t} R(\tau-1,e(\tau-1),e(\tau))^{-1} C^h(t,e(t)) \leq W^h(0), \tag{2}$$

where $W^h(0)$ is the consumer h's initial wealth and T can be either a finite number as in the life-cycle model or infinity as in the dynasty model. Here $R(t\text{-}1,e(t\text{-}1),e(t))$ is the (gross) asset return of the state contingent security for the event $e(t)$ in terms of the good in the event $e(t\text{-}1)$ at period $t\text{-}1$. We will often suppress e(t) to simplify the notation below.

In (1), β^h is the consumer h's discount factor. We assume that β^h can be different across consumers, but is constant over time for each consumer, following Lawrance (1991). This constant discount factor assumption greatly simplifies the empirical work. One interpretation of this assumption is that the discount factor actually changes as a consumer becomes wealthier as in Uzawa's (1968) model, but is roughly constant for a short sample period (six years) of our data. This interpretation seems valid because the range of consumption across households is generally much larger than the range of consumption fluctuations for each household in our data set.

The growth of consumption $(\hat{C}(t)=log(C(t+1))\text{-}log(C(t)))$ approximately satisfies

$$\hat{C}^h(t) \cong \sigma^h(t)\{r(t)\text{-}\delta^h\} \tag{3}$$

(see Atkeson and Ogaki (1993b)) where $\delta^h=\ln(1/\beta^h)$ is the RTP, $r(t)=ln(R^*(t))$, $R^*(t)=R(t)Prob(e(t+1)/e(t))$, and $\sigma^h(t)=\text{-}u'/(u''C^h(t))$ is the IES. From (3), $\sigma^h(t) \cong \partial\hat{C}^h(t)/\partial r(t)$. If there is no uncertainty, $r(t)$ is the real interest rate.

The distinct implications for consumption growth of models in which the RTP varies systematically with wealth and models in which the IES varies systematically with wealth can be seen in equation (3). If δ^h falls systematically as wealth rises as Lawrance's (1991) estimates suggest, then the consumption growth of the poor is always lower than the consumption growth of the rich. As long as σ is constant, there will be no systematic difference in the volatility of consumption growth between the rich and the poor. On the other hand, if σ^h rises systematically with wealth, then the consumption growth of the rich will be higher than that of the poor in the period in which $r(t)>\delta$ and the consumption of the rich will shrink faster than that of the poor in the in the period in which $r(t)<\delta$. Hence 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, we will use the quasi homothetic Geary-Stone utility function:

$$u(C^h) = \frac{1}{1-\alpha}\,[(C^h-\gamma)^{(1-\alpha)} - 1] \tag{4}$$

where $\alpha>0$. We will refer to the parameters γ as the subsistence parameter and the parameter α as the curvature parameter. Then the IES is

$$\sigma^h = \frac{1}{\alpha}\,(1 - \frac{\gamma}{C}). \tag{5}$$

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 intertemporal elasticity of substitution rises with the level of wealth. On the other hand, the IES falls with the level of wealth if $\gamma<0$ as the quadratic utility function.[2]

III. Econometric Method

The intertemporal first order condition is

[2] It should be noted that there is no theoretical reason to exclude the case where $\gamma<0$, although 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 than that of the rich.

$$\left(\frac{C^h(t,e(t))-\gamma}{C^h(t+1,e(t+1))-\gamma}\right)^{-\alpha} = \beta^h R^*(t,e(t),e(t+1)). \tag{6}$$

We assume that consumption $C^h(t)$ is measured with error in the following form:

$$C_m^h(t)-\gamma = (C^h(t)-\gamma)\varepsilon^h(t), \tag{7}$$

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

$$ln(\beta^h) = \beta_0 + \beta_1 y_c^h + \varepsilon_a^h, \tag{8}$$

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 $\varepsilon^h(t)$. Then from (6)-(8), we get

$$ln(C_m^h(t+1)-\gamma)-ln(C_m^h(t)-\gamma)-(\phi(t)+\beta_y y_c^h) = v^h(t), \tag{10}$$

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

$$v^h(t)=ln(\varepsilon^h(t+1))-ln(\varepsilon^h(t))+(1/\alpha)\varepsilon_a^h. \tag{14}$$

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 the 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 a (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

$$E_H[\xi(p^0)] = \operatorname*{plim}_{N\to\infty} (1/N)\sum_{h=1}^{N} [\xi^h(p^0)] = 0, \tag{11}$$

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).[3] 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

[3]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 (see Ogaki (1993b)) 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 (1993a, Section 4.3) provides a more detailed explanation as to how the data for villages are pooled.

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 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.

IV. Empirical Results

In this section, we report the results from the panel data in India collected by the Institute for Crop Research in the Semi-Arid Tropics (ICRISAT). We explain the data in the Appendix. 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.[4] First, the saving behaviors of less developed countries are of

---

[4]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), and Townsend (1991), and

interest. The dependence of the IES on the level of wealth is more likely to be important for very poor households whose consumption level is near the subsistence level. For this reason, it is desirable to examine intertemporal consumption behaviors of households less developed countries for the purpose of studying wealth-varying IES. Second, to the best of our knowledge, 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. Because much of consumption fluctuations within a year are likely to be caused by seasonal shifts that are not of interest for our purpose, it is desirable to study panel data that cover a substantial time period.[5]

In Table 1, we report average, minimum, maximum consumption per equivalent adult in terms of 1983 rupees for each of the three villages. These numbers are reported to facilitate the interpretation of the estimates of the subsistence levels that are reported below. From Table 1, we can see that average consumption fluctuates substantially over time in each village and that maximum and minimum consumption levels are substantially different in our data. These features are is desirable for the purpose of our research.

In Table 2, we report results for real total consumption expenditure per equivalent adult. In the first panel, we report estimates of γ and β_y

---

Rosenzweig and Wolpin (1993).

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, which are called likelihood ratio type test statistics.[5] The *C* statistic in 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 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.[6]

---

[6]See, e.g., Ogaki (1993a) for an explanation of the likelihood ratio type 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.

[7]This result is consistent with the hypothesis that there is full risk sharing among the households of each village in the data set. An arguably more powerful test for this hypothesis is to include income growth as an explanatory variable for consumption growth as in Ogaki and Atkeson (1992). When they include income grwoth, the coefficients for income were statistically insignificant and estimates for subsistence levels changed little. See Altug and Miller (1990), Cochrane (1991), Deaton (1991), Hayashi, Altonji, and Kotlikoff (1991), Morduch (1991), Ravallion and

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.[8]

We report estimates of $\phi(t)$'s for Aurepalle, Shirapur, and Kanzara in the second, third, and fourth panels of Table 2 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 3 and results when C(t) is taken as nonfood in Table 4. 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.

IV. Conclusions

In this paper, we estimated a model in which the RTP and the IES can

---

Chaudhuri (1991), and Townsend (1991) for other tests for complete markets.

[8]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.

rise or fall as a household become richer. Our empirical results are consistent with the view that the IES rises with the level of wealth, while the RTP is constant.

Appendix A

We explain the data in this appendix.

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, and alcoholic beverages; clothing, sewing of cloth, other tailoring expenses, thread, needles, chap pals and other footwear etc; 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, 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, food, and nonfood are the consumer price index, the price index for food, and the price index for nonfood, respectively. These real variables are valued at 1983 prices.

There are about forty households for each year in each of the three villages in the data. Some households drop out of the sample and others are added to the sample over years in the ICRISAT data. We exclude these households from our sample. 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.

REFERENCES

Altug, Sumru and Miller, Robert A. "Household Choices in Equilibrium" *Econometrica* 58 (May 1990): 543-70.

Atkeson, Andrew and Ogaki, Masao. "Wealth-Varying Intertemporal Elasticities of Substitution: Evidence from Panel and Aggregate Data," Rochester Center for Economic Research Working Paper No. 303R. 1993a.

Atkeson, Andrew and Ogaki, Masao. "A Note on a Definition of the Intertemporal Elasticity of Substitution," manuscript. 1993b.

Bhargava, Alok and Ravallion, Martin. "Does Household Consumption Behave as a Martingale? A Test of the Permanent Income Hypothesis for Rural South India." Manuscript, 1991.

Blundell, Richard, Martin Browning, and Costas Meghir, "A Microeconometric Model of Intertemporal Substitution and Consumer Demand," manuscript, 1990.

Chatterjee, Satayajit. "The Effect of Transitional Dynamics on the Distribution of Wealth in a Neoclassical Capital Accumulation Model." Manuscript, Federal Reserve Bank of Philadelphia, 1991.

Cochrane, John. "A Simple Test of Consumption Insurance." *Journal of Political Economy* 99 (October 1991): 957-976.

Deaton, Angus. "Saving and Liquidity Constraints." *Econometrica* 59 (September 1991): 1221-48.

Easterly, William, "Economic Stagnation, Fixed Factors, and Policy Thresholds." Manuscript, World Bank, 1991.

Epstein, Larry G. "Stationary Cardinal Utility and Optimal Growth under Unceretainty," *Journal of Economic Theory* 31 (1983): 133-152.

Gallant, A. Ronald and White, Halbert. *A Unified Theory of Estimation and Inference for Nonlinear Models*. New York: Basil Blackwell (1988)

Gomme, Paul and Greenwood, Jeremy "On the Cyclical Allocation of Risk." Federal Reserve Bank of Minneapolis Discussion Paper No. 71 (1992).

Hansen, Lars Peter, "Large Sample Properties of Generalized Method of Moments Estimators." *Econometrica*, 50, (1982), 1029-54.

Hayashi, Fumio; Altonji, Joseph; and Kotlikoff, Laurence. "Risk-Sharing, Altruism, and the Factor Structure of Consumption." NBER Working Paper No.3834, 1991.

Lawrance, Emily C. "Poverty and the Rate of Time Preference: Evidence from Panel Data." *Journal of Political Economy* 99 (February 1991): 54-77.

Lim, Youngjae. "Disentangling Permanent Income from Risk Sharing: A General Equilibrium Perspective on the Rural Market in Developing Countries." Ph.D. disseration, University of Chicago, 1992

Mace, Barbara J. "Full Insurance in the Presence of Aggregate Uncertainty." *Journal of Political Economy* 99 (October 1991) 928-956.

Mankiw, N. regory and Stephen P. Zeldes. "The Consumption of Stockholders and Nonstockholders." *Journal of Financial Economics* 29 (March 1991), 97-112.

Mendoza, Enrique G. "Real Business Cycles in a Small Open Economy." *American Economic Review* 81 (June 1991), 797-818.

Morduch, Jonathan. "Risk, Production, and Saving: Theory and Evidence from Indian Households." Manuscript, Harvard University, 1990.

——. "Consumption Smoothing Across Space: Tests for Village-Level Responses to Risk." Manuscript, Harvard University, 1991.

Ogaki, Masao. "Growth in Open Economies: A Comment," *Carnegie-Rochester Conference Series on Public Policy* 36, 1992b.

——. "Generalized Method of Moments: Econometric Applications," forthcoming in *Handbook of Statistics, Vol. 11: Econometrics*, edited by G.S. Maddala, C.R. Rao, and H.D. Vinod. Amsterdam: North-Holland, 1993a.

——. "GMM: A User's Guide." Manuscript, 1993b. Rocehster Center for Economic Research Working Paper No. 348, 1993b.

Ogaki, Masao and Atkeson, Andrew. "Estimating Subsistence Levels with Euler Equations in Indian Panel Data." Manuscript, 1992.

Ravallion, Martin and Chaudhuri, Shubham. "Testing Risk-Sharing in Three Indian Villages." Manuscript, 1991.

Rebelo, Sergio. "Growth in Open Economies," *Carnegie-Rochester Conference Series on Public Policy* 36 (1992).

Rosenzweig, Mark R. "Risk, Implicit Contracts and the Family in Rural Areas of Low-Income Countries." *The Economic Journal* 98 (December 1988):1148-70.

Rosenzweig, Mark R. and Stark, Oded. "Consumption Smoothing, Migration, and Marriage: Evidence from Rural India." *Journal of Political Economy* 97 (August 1989): 905-26.

Rosenzweig, Mark R. and Kenneth I. Wolpin. "Credit Market Constraints, Consumption Smoothing and the Accumulation of Durable Production Assets in Low-Income Countries." *Journal of Political Economy* 101 (April 1993): 223-244.

Townsend, Robert M., "Risk and Insurance in Village India," manuscript, 1991, University of Chicago.

Uzawa, H. "Time Preference, the Consumption Function, and Optimum Asset Holdings." in *Value, Capital and Growth: Papers in Honour of Sir John Hicks.* Edited by James N. Wolfe. Edinburgh: Edinburgh University Press, 1968.

White, Halbert. "A Heteroskedasticity-Consistent Covariance Matrix Estimator and a Direct Test for Heteroskedasticity." *Econometrica* 48 (May 1980): 817-838.

TABLE 1

CONSUMPTION PER EQUIVALENT ADULT

	1976	1977	1978	1979	1980	1981	1976-81
			Average 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
			Minimum Total Consumption				
Aurepalle	179	308	334	359	249	229	. . .
Shirapur	304	491	485	364	423	242	. . .
Kanzara	393	370	294	460	419	432	. . .
			Maximum Total Consumption				
Aurepalle	1,300	913	984	1,574	1,945	1,510	. . .
Shirapur	1,703	1,723	1,974	1,525	1,596	1,676	. . .
Kanzara	2,694	1,509	1,290	2,189	2,085	1,945	. . .
			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
			Minimum Food Consumption				
Aurepalle	221	178	173	214	288	187	. . .
Shirapur	238	137	274	221	308	102	. . .
Kanzara	221	178	173	214	288	187	. . .

TABLE 1 - *Continued*

	Maximum Food Consumption						
Aurepalle	646	658	766	1,044	1,132	829	. . .
Shirapur	1,133	1,063	1,166	888	1,075	1,088	. . .
Kanzara	1,441	870	704	1,284	1,409	1,081	. . .
	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
	Minimum Nonfood Consumption						
Aurepalle	40	27	83	68	67	32	. . .
Shirapur	65	112	136	129	114	48	. . .
Kanzara	134	151	113	133	131	133	. . .
	Maximum Nonfood Consumption						
Aurepalle	908	377	415	711	831	688	. . .
Shirapur	681	698	894	870	777	1,013	. . .
Kanzara	1,019	1,274	759	836	929	885	. . .

TABLE 2

GMM RESULTS FOR TOTAL CONSUMPTION

γ	s.e.	β_y	s.e.	J^*	d.f.	p-value**	C^{***}	d.f.	p-value**
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

*Chi-square test statistics for the overidentifying restrictions.

**In percentage.

***Likelihood ratio type test statistics for the restrictions imposed.

TABLE 3

GMM RESULTS FOR FOOD CONSUMPTION

γ	s.e.	β_y	s.e.	J^*	d.f.	p-value**	C^{***}	d.f.	p-value**
101.4	4.30	-0.083	0.360	30.69	28	33.1	. . .		. . .
101.5	3.70	0	. . .	32.28	29	30.8	1.597	1	20.6
0	. . .	0	. . .	56.93	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.
0.362	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

*Chi-square test statistics for the overidentifying restrictions.

**In percentage.

***Likelihood ratio type test statistics for the restrictions imposed.

TABLE 4

GMM RESULTS FOR NONFOOD CONSUMPTION

γ	s.e.	β_y	s.e.	J^*	d.f.	p-value**	C^{***}	d.f.	p-value**
28.8	1.44	-0.014	0.059	28.17	28	45.6	. . .		. . .
26.8	1.44	0	. . .	28.22	29	50.6	0.053	1	81.9
0	. . .	0	. . .	35.69	30	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

*Chi-square test statistics for the overidentifying restrictions.

**In percentage.

***Likelihood ratio type test statistics for the restrictions imposed.