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ABSTRACT

In an attempt to obtain a measure of rural-farm poverty conceptually more meaningful than those currently available, a measure based on "full income" (defined as "the purchasing power available for consumption in a normal year while keeping wealth intact") was proposed. The main task was estimation of the full income, rather than the money income, of rural-farm families in the United States in 1969 and the size distribution of that full income. Estimates were based on data on human and nonhuman wealth and market rates of return. The inequality of the distribution of full income was estimated by means of data on the distribution of human and non-human wealth. A full income poverty threshold was then applied to the constructed size distribution of full income, and full income poverty was measured as the percentage of rural-farm families and unrelated individuals which fell below this threshold. Results produced an estimation that 5 to 14 percent of the U.S. rural-farm population was poor in terms of full income in 1969. By comparison, the corresponding figure as published in the U.S. Census of Population (using annual money income) was 19.9 percent. (Author/JC)

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A FULL INCOME APPROACH TO THE MEASUREMENT OF RURAL POVERTY

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ABSTRACT

This research is an attempt to obtain a measure of rural-farm poverty conceptually more meaningful than those currently available. The measure proposed is based on "full income," defined as the purchasing power available for consumption in a normal year, while keeping wealth intact. The main task is the estimation of full income of rural-farm families in the United States in 1969, and the size distribution of full income. The estimates are based on data on human and non-human wealth and market rates of return. The inequality of the distribution of full income is estimated by means of data on the distribution of human and non-human wealth. Then a full-income poverty threshold is applied to the constructed size distribution of full income. Full-income poverty is measured as the percentage of rural-farm families and unrelated individuals which falls below this threshold. The result is an estimate that 5 to 14 percent of the U.S. rural-farm population is poor in terms of full income in 1969. The corresponding figure as published in the U.S. <u>Census of Population</u>, using annual money income, is 19.9 percent.

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A FULL INCOME APPROACH TO THE MEASUREMENT OF RURAL POVERTY

I. Introduction

While almost everyone agrees that there are poor people in the United States, there is little agreement on exactly how the poor are to be distinguished from the non-poor. Consequently there is little agreement on the size of the population properly referred to as being in poverty. This state of affairs is masked to some extent by the existence of officially-sanctioned poverty statistics calculated by estimating the number of people having incomes below a designated poverty line. However, theoretical foundations for arriving at a "correct" poverty threshold are practically nonexistent.[1]

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[1]It is not even clear whether income is the correct variable, or the only variable, that the threshold should be measured in terms of. And given an income criterion it is not clear whether income level is sufficient. For example, a family of four with an income of $2075 in Mississippi in 1949 might not perceive itself, or be perceived by others, to be as poor as the same family with the same real income in 1969, because 83 percent of the state's farm population would be poorer than this family in 1949, but only 36 percent would be poorer in 1969. Trying to take "relative status" into account leads to insoluble difficulties -- insoluble in the sense that no objective measurement of the trend of poverty is possible. As Schultz (1968, p. 66) emphasizes, the positive analysis of poverty, not only policy proposals, rests on preferences of both the poor and of the non-poor who decide what counts as poverty.

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This paper will not fill the theoretical gap; instead it follows the customary procedure of making a bow to the difficulties of the subject before rushing off to the numbers. But while the paper does not develop a theory of the poverty line, it does attempt to provide a more meaningful income scale along which to rank individuals than that provided in the official statistics on poverty.[2] The income concept used in generating these official statistics is annual money income. The alternative income concept proposed is "full income" as defined below.

II. U.S. Census Data on Rural Poverty

Published reports of the 1970 Census of Population included for the first time estimates of the "low-income population." The population includes all family members and unrelated individuals whose incomes in 1969, as reported in the census, fell below the "poverty threshold" established by the Social Security Administration. This threshold varies with size of family, sex of head, age of head (over or under 65 for two-person families and unrelated individuals), and farm or nonfarm residence. The threshold also varies over time with the consumer price index used to adjust for cost-of-living changes. Thus the poverty threshold for a rural-farm family of four increased from $2527 in 1959 to $3191 in 1969.[3] The income concept used to measure poverty status is money income from all sources in the year preceding the census.

Table 1 shows the estimated percentage of rural-farm families and unrelated individuals below the poverty threshold by states for 1969, 1959 and 1949. The data show substantial reductions in the incidence of poverty among rural-farm people. The percentage of rural-farm

[2] By "official statistics" are meant those published by U.S. government agencies.

[3] "Rural-farm" throughout this paper refers to the definition of the Bureau of the Census: a household living in a rural area on a place of 10 or more acres with $50 or more sales of agricultural products or on a place of less than 10 acres with $250 or more sales in the preceding year.

Table 1. Fraction of rural-farm families and unrelated individuals below "real" SSA poverty threshold, states of U.S.[a]

State	1969	1959	1949
Maine	.219	.350	.511
New Hampshire	.124	.237	.454
Vermont	.203	.362	.554
Massachusetts	.108	.194	.353
Connecticut	.099	.198	.323
New York	.146	.280	.423
New Jersey	.125	.255	.379
Pennsylvania	.154	.277	.425
Ohio	.134	.283	.381
Indiana	.119	.263	.353
Illinois	.131	.294	.347
Michigan	.130	.257	.396
Wisconsin	.145	.320	.470
Minnesota	.197	.402	.448
Iowa	.131	.353	.349
Missouri	.196	.416	.589
North Dakota	.182	.358	.407
South Dakota	.221	.430	.415
Nebraska	.180	.343	.398
Kansas	.150	.289	.425
Delaware	.141	.258	.508
Maryland	.160	.301	.513
Virginia	.258	.468	.631
West Virginia	.258	.455	.588
North Carolina	.281	.544	.677
South Carolina	.302	.592	.757
Georgia	.239	.534	.758
Florida	.196	.378	.649
Kentucky	.271	.488	.671
Tennessee	.288	.531	.707
Alabama	.275	.574	.781
Mississippi	.370	.661	.838
Arkansas	.273	.555	.751
Louisiana	.323	.556	.723
Oklahoma	.186	.343	.555
Texas	.203	.395	.534
Montana	.169	.258	.378
Idaho	.147	.236	.354
Wyoming	.143	.248	.380
Colorado	.194	.259	.407
New Mexico	.267	.356	.515
Utah	.165	.218	.368
Nevada	.158	.244	.457
Arizona	.288	.480	.660
Washington	.141	.203	.340
Oregon	.136	.211	.339
California	.140	.247	.441
United States total	.20	.39	.55

Table 1 (continued)

[a]The 1969 figures are taken directly from the 1970 Census of Population, State Reports, Table 58 (U.S. Bureau of the Census, 1972). The 1959 figures apply the revised SSA poverty thresholds for 1959 (U.S. Bureau of the Census, 1969b, p. 5) to the census size distributions of family income as given for each family size (U.S. Bureau of the Census, 1961, Table 141). The 1949 published data do not contain separate income distributions for different family sizes. In order to get an estimate of poverty for 1949 comparable to the 1959 data, an average rural-farm poverty threshold for 1949 was estimated by calculating the average number of persons per household for each state. The poverty threshold appropriate to this average family size was then taken from a set of 1949 poverty thresholds constructed by deflating the 1959 SSA poverty threshold by the 1959/1949 change in the consumer price index. This procedure yields an overall weighted average poverty line for each state that was applied to the size distribution of income for all families and unrelated individuals in the state (U.S. Bureau of the Census, 1952, Table 32).

households below the poverty line in 1949 in Mississippi was about 84 percent, a figure one might think more typical of an underdeveloped country than of the United States. By 1969 this figure had been reduced to 37 percent. In three other southern states (Alabama, Georgia, and Louisiana) the reduction in rural-farm poverty amounted to more than half the rural-farm population.

In the non-southern states, though not so many people moved over the poverty threshold between 1949 and 1969, the reductions in poverty are dramatic. In every state but South Dakota the incidence of poverty was at least halved. In 1949, no non-Southern state had less than 30 percent of its rural-farm households below the poverty line. In 1969, none have more.

III. The Full Income Approach

Full income is defined in this paper as the normal returns to owned resources. It is equal by definition to the purchasing power available for consumption and savings (including unrealized capital gains) in a normal year. Of course, a normal year may never occur. Current money income as measured by the Bureau of the Census will typically not be equal to full income. This is true not only because full income excludes transitory fluctuations, but also because some returns to owned resources are not current money returns. Examples are services provided by owned housing, home-grown food, other do-it-yourself services, and unrealized appreciation of property values. All of these items make full income greater than money income, and all are probably quantitatively more important for farm than nonfarm people. Moreover, they are probably quantitatively greater for poor (as measured in terms of money income) than for non-poor farm people.

Full income is related to, but distinct from, the income concept used in Carlin and Reinsel (1973). Following Weisbrod and Hansen (1968), they add to current money income an annuity which consumes estimated net worth in equal installments over the expected lifetime of the wife of the family. This concept of "well-being" has the defect of excluding non-marketed returns to labor and including transitory income. The

most interesting feature of the Weisbrod-Hansen approach is the annualizing of net worth. This idea provides a systematic way to bring wealth into income accounting, but can be misleading in that it increases measured well-being of older people simply because they have fewer years to live. The full income concept used in the present paper does not annualize net wealth, but only counts as income market returns and capital gains accruing in the current period, keeping wealth constant. This approach follows Hicks' (1946, p. 176) "central criterion" that "a person's [weekly] income is what he can consume during the week and still expect to be as well off at the end of the week as he was at the beginning."

There are two main reasons for expecting less income inequality and consequently less poverty when full income rather than money income is used as a standard. The first is that the class of people observed to have the lowest money incomes will tend to include those with the largest negative transitory component of income in the particular year in which the observation is made. For these people, normal income will be substantially higher than observed income (Reid, 1966). Consider, for example, the lowest income class reported in the U.S. Department of Agriculture (1965) Consumer Expenditures Survey of 1961 (Table 11B). The 5.8 percent of north central U.S. rural-farm families in this category had a mean money income of -\$792, yet had expenditures for current consumption of \$3,058, 27 percent <u>higher</u> than the next-highest income class (which had a mean income of \$1,598). Indeed this "poorest" group had consumption expenditures only \$680 below the U.S. rural-farm population mean. Obviously these families are not so poor as the money income data indicate.

Second, even if there were no transitory fluctuations in income, the use of money income as an indicator can exaggerate the incidence of poverty and inequality. Consider a population in which everyone had equal full income but varying ratios of money to full income. To take the simplest possible case, suppose that everyone has \$5000 full income, but there are two groups within the population, one of which earns its \$5000 totally in the form of money income, while the other gets \$2500 of money income and \$2500 unrealized capital gains and other income in

kind. In such a case, inequality measured in terms of money income would indicate inequality where in fact there is none, and poverty measured by money income would reveal a class of poorer people when in fact there is none.

It is possible, however, to construct cases in which adding non-marketed income increases inequality. This would occur if non-marketed income were an increasing fraction of total income as income increases. This situation may in fact pertain for unrealized capital gains, though it seems unlikely for other forms of home-produced, non-marketed income.[4]

The net effect of a more comprehensive income measure cannot be predicted a priori, though the preceding discussion suggests that inequality and hence poverty will be reduced when measured in terms of full income. To find out whether this is in fact the case, and how much difference it makes, is the aim of the remainder of this paper. There are two steps to the procedure. First, the inequality of the size distribution of full income is estimated. Second, a poverty threshold is applied to this distribution to estimate "full-income poverty."

IV. Estimating the Size Distribution of Full Income

There exist no reliable distributional data that allow piecemeal corrections to be made to census income data by adding in the various non-marketed items left out and by taking out transitory gains and losses. Instead, the procedure followed is to return to the definition of full income as normal returns to owned resources. Data are available pertaining to the distribution of resources owned by rural-farm households. An estimate of the inequality of the distribution of owned resources can be transformed into an estimate of the inequality of full income by means of factor shares. Let the full income of a farm family be the sum of returns to human resources (H), land (L) and capital (K):

(1) $$Y_f = r_H H + r_L L + r_K K,$$

where r_i is the rental return to each resource.

[4]Carlin and Reinsel (1973) find that adding annualized net worth reduces income inequality of farm operators.

The variance of Y_f is:

$$(2) \qquad Var\ (Y_f) = r_H^2\ Var\ (H) + r_L^2\ Var\ (L) + r_K^2\ Var\ (K)$$
$$+ 2r_H r_L\ Cov\ (H,L) + 2r_H r_K\ Cov\ (H,K) + 2r_L r_K\ Cov\ (L,K).$$

To put the measure of inequality in relative terms, convert (2) to an expression in terms of the squared coefficient of variation of full income, $C^2(Y_f)$, dividing through by mean income, $\overline{Y}_f^2$, and simplifying. For example, the first r.h.s. term is treated as follows:

$$\frac{r_H^2\ Var\ (H)}{\overline{Y}_f^2} = \frac{r_H^2 \cdot \overline{H}^2}{\overline{Y}_f^2} \cdot \frac{Var\ (H)}{\overline{H}^2}$$
$$= S_H^2\ C^2\ (H)$$

where S_H is the relative share of returns to human resources in full income.

The covariance terms are treated analogously:[5]

$$\frac{2r_H r_L\ Cov\ (H,L)}{\overline{Y}_f^2} = \frac{2r_H r_L\ \rho_{HL}\ \sqrt{Var(H)}\ \sqrt{Var(L)}}{\overline{Y}_f^2}$$
$$= \frac{2r_H\overline{H}\ r_L\overline{L}}{\overline{Y}_f\ \overline{Y}_f} \cdot \frac{\rho_{HL}\ \sqrt{Var(H)}\ \sqrt{Var(L)}}{\overline{H}\ \overline{L}}$$
$$= 2S_H S_L\ \rho_{HL}\ C(H)\ C(L).$$

Doing similar manipulations of the other right-hand side terms of equation (2) yields an equation which expresses the inequality of full income in terms of the inequality of factor ownership:

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[5]The derivation uses the definition of the correlation coefficient between H and L:

$$\rho_{HL} = \frac{Cov\ (H,L)}{\sqrt{Var(H)}\ \sqrt{Var(L)}} .$$

$$C^2(Y_f) = S_H^2\, C^2(H) + S_L^2\, C^2(L) + S_K^2\, C^2(K) + 2S_HS_L\, \rho_{HL}\, C(H)\, C(L) + 2S_HS_K\, \rho_{HK}\, C(H)\, C(K) + 2S_LS_K\, \rho_{LK}\, C(L)\, C(K). \tag{3}$$

Despite its cumbersome appearance, equation (3) has a straightforward interpretation. The first three terms on the right-hand side sum the measured inequality of H, L, and K, respectively, each being weighted by the square of its relative share in full income. The last three terms are weighted relative covariances of resource ownership for each pair of resources.[6] Equation (3) is useful because while there are no data that allow direct estimation of the inequality of full income, there are data that may be used to estimate most of the right-hand elements. Equation (3) is not a behavioral or explanatory function; it is an accounting identity. It is not intended to explain full income inequality, but only to facilitate its measurement.

The reason it can be claimed that equation (3) provides a measure of full rather than money income inequality lies in the use of market rates of return rather than market returns for the r's in equation (1). This point is conveniently illustrated with reference to the $r_L L$ term of equation (1). L is measured as the values of real estate owned by the household, while r_L is the market rate of return as measured by the interest rate on farm real estate debt. This measure of r_L is substantially greater than the rate of money returns to farm real estate. Estimates of money returns to land in the 1960's have been in the neighborhood of 3 to 4 percent of land value (Kost, 1968), while the mortgage interest rate for this period (1969) is more like 6 to 7 percent. Therefore, $r_L L$ as measured in equation (1) is almost twice as large as money returns to land. The difference is an estimate of anticipaced or normal capital gains to real estate. "Anticipated" means

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[6]Note that the sum of the share weights is unity:

$$S_H^2 + S_L^2 + S_K^2 + 2S_HS_L + 2S_HS_K + 2S_LS_K = 1.$$

expected by the market in the sense that capital gains of this magnitude must be expected; otherwise land would not sell at such a high price that cash returns only yield 3 percent. In short, the rates of return and factor shares used in equations (1) to (3) yield estimates of full returns to owned resources, whether the product of the resources is actually marketed or not.

The choice of the coefficient of variation (squared) as a measure of inequality is based on computational considerations. There is no good economic reason to choose it over the Gini coefficient, the standard deviation of the log of income, or other measures which have been proposed. On the other hand, there is no economic reason not to use the coefficient of variation. In practice the correlation coefficient between other measures of inequality and the coefficient of variation is quite high -- an income distribution which is unequal by an alternative criterion is very likely also to have a high coefficient of variation.[7]

Human Resources. The first and most important element of equation (3) for which data can be generated is $C^2(H)$, the inequality of human resource ownership. However, the value of human resources owned is not directly observable, so that $C^2(H)$ cannot be estimated directly. Instead, the approach followed is to estimate the distribution of income from human resources. Under the assumption of the same rate of return for all comparable human resources,[8] the coefficient of variation of human resource ownership and of income from human resources will be equal.[9]

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[7] For a comparison of the coefficient of variation with five other indexes of inequality, see Champernowne (1974). More detailed comparisons in the context of a log-normal distribution are available in Aitchison and Brown (1957).

[8] Not that all workers get the same wage rates, but that all workers with the same human capital get the same wage rates.

[9] Letting Y_H be income from H and w the equilibrium return per unit H,

$$Y_H = wH$$

$$\text{Var } Y_H = w^2 \text{ Var } H$$

$$\frac{\text{Var } Y_H}{w^2 H^2} = \frac{\text{Var } H}{H^2}, \quad \text{i.e., } C^2(Y_H) = C^2(H).$$

This assumed equality of returns implies that equilibrium is attained in labor markets. This assumption fits in well with the idea of estimating the distribution of full income in a normal year as opposed to current money income. Full income is essentially an equilibrium concept.

The variance of full income from human resources is estimated by weighting the coefficients of an estimated income generating function by the distribution of characteristics of the rural-farm population in each state in 1969. The income generating function is:

$$(4) \qquad Y_i = \beta_0 + \beta_1 R + \beta_2 E + \beta_{jk} D_{jk} + e_i.$$

Y_i is family income, the observations being the 1151 rural-farm families in which a male head was present in the 1970 Census of Population User Tapes from the 15 percent sample of the population. All the independent variables are dummies. R is zero for whites, 1 for nonwhites. E is zero where only the male head worked, 1 where other family members worked. The D_{jk} represent 30 age-education cells of the male head, a cross-classification of 6 age classes (18-24, 25-34, 35-44, 45-54, 55-64, 65+) and 5 education classes (less than 5, 5-8, 9-11, 12, 13+ years of schooling). Table 2 shows the resulting age-education coefficients.

Some relevant characteristics of rural-farm families are omitted from the income generating function, notably the age, schooling, and race of the wife.

The dependent variable ideally should be full labor income. Both earnings reported in the census and family income have deficiencies as a measure of it. Earnings exclude self-employed labor returns on farms as well as in other home production. Family income, on the other hand, includes labor earnings on farms but also includes some returns to non-human resources. In choosing the latter for the estimation of equation (4), it is implicitly assumed that returns to non-human wealth do not vary systematically with age and schooling (or that, if they do, these variations in income are properly counted as attributable to variations in human resources).

Table 2. Partial effects of male head's age and education on rural-farm family income[a]

	Years of schooling				
Age	0-4	5-8	9-11	12	13+
18-24	-4320 (0)[b]	-4320 (4)	-4320 (1)	-1312 (8)	-6520 (2)
25-34	-4320 (0)	-3340 (17)	-3850 (24)	-2143 (82)	-2027 (29)
35-44	-3682 (5)	-1823 (45)	- 902 (40)	-1706 (118)	6965 (26)
45-54	-3836 (8)	-2050 (106)	0 (58)	- 62 (85)	1560 (30)
55-64	-4980 (15)	-3680 (123)	-2960 (54)	-3166 (64)	-1145 (23)
65 +	-6950 (19)	- 550 (104)	-5747 (37)	-5662 (15)	-1390 (9)

[a]Relative to those with 9-11 years of schooling aged 45-54. Estimated β_{jk} from equation (4).

[b]Number of observations in each class in parentheses. Classes with zero or one observation were combined with neighboring class in estimation.

To obtain an estimate of the variance of income accounted for by the characteristics included in the income generating function, the coefficients of equation (4) are applied to the distribution of the income-generating characteristics among the rural-farm population of each state. For example, the computation of the variance of full income attributable to age and schooling is:

$$\text{Var}\ (Y_H) = \sum_{j=1}^{6} \sum_{k=1}^{5} f_{jk}\ (\hat{\beta}_{jk} - \bar{Y})^2$$

where the f_{jk} are the fraction of a state's rural-farm males in the j^{th} age and k^{th} schooling class and the $\hat{\beta}_{jk}$ are the corresponding coefficients from the income generating function. Data on the joint distribution of age and schooling, number of earners, and race are available in U.S. Bureau of the Census (1972, Tables 148 and 200).

The sum of the estimated variances of income attributable to race, earners per family, and age-schooling jointly for each state is used to estimate $C^2(H)$ for equation (3).

Because some important aspects of human resources are left out, the estimates may overstate or understate $C^2(H)$. Whether the estimate is understated or overstated depends on the covariance between left out and included items of income. The main left-out item is the wife's full earnings.

If the wife's contribution relative to the husband's is the same at each level of husband's earnings, then family earnings are understated by a constant percentage. This leaves $C^2(H)$ unchanged, so that the omission of the wife's earnings has no effect on estimated inequality. However, if the wife's earnings increase (decrease) as a percentage of family earnings as the husband's earnings increase, then omitting the wife's earnings understates (overstates) $C^2(H)$ for the household. Since full earnings (including productive time at home) rather than labor market earnings is wanted for $C^2(H)$, the ratio of wife's to husband's market earnings is less relevant than the ratio of wife's to husband's productive characteristics. Thus, better evidence that $C^2(H)$ is understated (overstated) would be data showing that the

ratio of wife's to husband's schooling increases (decreases) as husband's schooling increases. In fact, in the census user tape sample (as well as in a sample of 1100 North Carolina families), the ratio of wife's to husband's schooling decreases as the husband's schooling increases, although not dramatically.[10] This crude evidence suggests that $C^2(H)$ is overstated to some extent.

Real Estate. The other distributional item in equation (3) for which detailed state data are available is $C^2(L)$, the relative variance of real estate ownership. U.S. Bureau of the Census (1969a, Parts 1-50, State Table 9) gives data on farms classified by value of land and buildings per farm.[11] Land and buildings includes the residence so that when wealth data are converted to income flows, the implicit rental value of owned housing is included in the full income measure. Although there are no state data on what fraction of this real estate is owned rather than rented by the farm operator, the census does provide separate data for full owners, part owners, and tenants. This information can be used to infer a size distribution of ownership as illustrated in the following paragraph for U.S. part-owner farms in 1964.

First the cumulative value distribution of part-owned farms is plotted on lognormal graph paper. Then using the datum, from U.S. Bureau of the Census (1964, Vol. II, p. 754), that the value of real estate owned by part owners is 55 percent of the value of farms operated by that tenure group, each value on the distribution of farms operated is multiplied by .55 (which on log scale yields a constant horizontal shift). This yields an inferred size distribution of real estate owned by part owners. For example, although 24 percent of part owners operate farms worth over $100,000, only 9 percent are estimated to own that value of real estate.

This procedure makes two important assumptions. The first is that the 55 percent applied to farmers in all value classes. The second is

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[10] In the North Carolina sample, as the husband's schooling increases from 8-11 to 12+, the wife's mean schooling increases from 10.1 to 13.5; the ratio of wife's to husband's schooling decreases from about 1.1 to 0.9.

[11] Ten value classes are given: less than $10,000, $10-20,000, $20-40,000, $40-70,000, $70-100,000, $100-150,000, $150-200,000, $200-500,000, and $500,000 or more.

that land ownership is distributed lognormally. The first assumption is required for the shift by a (logarithmic) constant to be legitimate; the second is required to switch without distortion from a log-of-income to a percentage-of-farmers scale. In fact, the approximation to log-normality, as judged by visual inspection, appears quite good except at the tails of the size distribution. The errors at the tails may not be too serious because there are so few observations at extreme levels. For example, the 1964 U.S. data give .006 of U.S. part owners possessing more than $500,000 of real estate. Even a rather large error in the estimated logarithmic shift would change this estimate by only a few tenths of 1 percent.

A similar adjustment was made for tenant and full-owner farms. In the case of full-owner farms, adjustment was necessary because on the average they own real estate worth 16 percent more than the value of real estate operated.

Non-Commercial Farms.[12] These present special problems because the census did not publish size distributions of value of farm operated for each tenure class, rather there is only one size distribution for all non-commercial farms. However, it is possible to make some inferences about the size distribution of ownership by using the tenure classification of non-commercial farms given in U.S. Bureau of the Census (1964, Ch. 8). Assuming that the distribution of farms by tenure class as one moves from low to high value-of-farm classes changes in the same way for commercial and non-commercial classes, a tenure classification can be estimated for non-commercial farms, as shown for three illustrative classes in Table 3.

Using the preceding approach, and assuming that the ratio of owned to operated land is the same for non-commercial as for commercial farms

[12] Non-commercial farms as defined in the 1964 Census of Agriculture are those having sales of less than $2,500, and whose operators either worked off the farm 100 days or more or were over 65 years of age. In 1964 approximately 1.0 million of the 3.6 million farms in the U.S., or 32.7 percent, were non-commercial farms (U.S. Bureau of the Census, 1964, Vol. II, p. 749 and 792). The 1969 census does not use the term "non-commercial," but has 1.0 million, or 37 percent, of all farms having sales less than $2,500 (U.S. Bureau of the Census, 1969a, Vol. II, Ch. 3, p. 20).

Table 3. Non-commercial farms by tenure and value of land and buildings, U.S., 1964[a]

Value classes	Commerical			Non-commercial		
	Percent full owner	Percent part owner	Percent tenant	Percent full owner	Percent part owner	Percent tenant
All	.474*	.314*	.212*	.807*	.107*	.096*
< $10,000	.56*	.15*	.29*	.84	.02	.03
70-99,000	.30*	.45*	.25*	.51	.29	.20
200-499,000	.21*	.55*	.24*	.36	.43	.21

[a]The starred numbers are taken from the census; the others are estimated from them.

of a given tenure class, the fraction of value of operated real estate owned for each value class can be calculated. This figure is used to shift the size distribution of value for non-commercial farms as was done above for commercial part-owner farms. The only difference is that in this case there is not a constant shift. The shift is greater for higher-value farms because fewer of these are operated by full owners. The relevant figures are shown in lines 7 and 8 of Table 4. This table shows illustrative calculations of the size distribution of farm real estate ownership for the U.S. (bottom line). A table of this kind was constructed for all the states of the U.S. from 1969 census data. The coefficient of variation of the value of real estate ownership is computed from the resulting size distribution of ownership. The estimated values of $C^2(L)$, to be used in equation (3), are shown in Table 5.

Capital. The final relative variance component of equation (3) is that for capital, $C^2(K)$. Included in $C^2(K)$ should be not only owned non-land farm resources (which account for about 30 percent of the value of all farm non-human resources), but also nonfarm capital owned by farmers. This category includes savings accounts, financial assets and nonfarm business property. It does not include farmland owned but not operated, which is already included in $C^2(L)$ as estimated for Table 5.

Unfortunately, there are no state data on the size distribution of ownership of nonfarm wealth of farm operators. Two alternative calculations were made in estimating equation (3). The first alternative was simply to ignore $C^2(K)$ and set the relative share of K equal to zero in all states, i.e., pretend that all income is returns to human capital and land. The second alternative is to let $C^2(K)$ be the geometric mean of $C^2(H)$ and $C^2(L)$. This procedure is as arbitrary as the first alternative, but it does allow the relative share, S_K, to play a rôle.

Covariances. Other important pieces of data for the estimation of equation (3) are the correlation coefficients between income from owned land, capital, and human resources. Unfortunately, there are no state data for any of these. There are only two bits of information on this subject known to me. In one, my own survey data of Sampson County, North Carolina, I found a correlation coefficient between returns to human

Table 4. Estimated size distribution of farm real estate ownership, 1964

Tenure	Value class								
	Less than $10,000	$10,000-20,000	$20,000-40,000	$40,000-70,000	$70,000-100,000	$100,000-150,000	$150,000-200,000	$200,000-500,000	$500,000 +
	(thousands)								
Full owners:									
Operated	187	218	294	178	60	37	15	19	4.6
Owned	148	201	298	199	72	46	19	27	7
Part owners:									
Operated	49	75	144	156	91	76	36	48	12
Owned	156	133	167	118	49	31	11	11	1
Tenants and Managers:									
Operated	99	50	83	94	52	40	18	21	7
Owned	458	3	0	0	0	0	0	0	0
Non-Commercial:									
Operated	413	313	192	53	13	7	3	3	1
Owned	447	273	197	57	12	4	1	1	0
Total U.S.:									
Operated	747	656	713	480	215	160	71	91	24
Owned	1218	609	663	373	133	82	32	38	9

Table 5. Estimated squared coefficient of variation of real estate ownership, by states, 1969

Northeast:	
Maine	2.361
New Hampshire	2.122
Vermont	1.736
Massachusetts	2.637
Connecticut	2.149
New York	2.634
New Jersey	2.393
Pennsylvania	2.400
Midwest:	
Ohio	2.011
Indiana	1.842
Illinois	2.277
Michigan	1.908
Wisconsin	1.865
Minnesota	1.659
Iowa	1.599
Missouri	2.405
North Dakota	1.298
South Dakota	1.763
Nebraska	1.975
Kansas	2.057
South:	
Delaware	3.462
Maryland	2.652
Virginia	4.220
West Virginia	3.291
North Carolina	3.627
South Carolina	4.358
Georgia	3.433
Florida	4.012
Kentucky	3.934
Tennessee	3.261
Alabama	4.504
Mississippi	5.048
Arkansas	4.053
Louisiana	4.642
Oklahoma	2.543
Texas	3.407
West:	
Montana	1.765
Idaho	2.279
Wyoming	2.406
Colorado	2.390
New Mexico	2.934
Arizona	2.586
Utah	2.436
Nevada	2.237

Table 5 (continued)

Washington	2.022
Oregon	2.401
California	2.322
United States	3.2

and non-human capital of .22. A second source of such data, more useful for present purposes, is the User Tape sample of 1151 farmers from the 1969 Census of Population. Arbitrarily allocating 50 percent of farm and business income to human capital, the correlation coefficient between income from human and non-human resources for males is .26. The former estimate is conceptually closer to what is needed, but the latter is from a more representative data set. At least, it is encouraging that both estimates agree that the correlation coefficient is positive -- that farmers with high labor income also tend to have high capital income -- and indeed the estimates are quite close. A value of .25 is used for all states for ρ_{HL} and ρ_{HK} in equation (3).

Although there are no data for the correlation coefficient between farm real estate and other wealth ownership, ρ_{LK}, there is one statistic that will help in at least keeping the crude estimates made consistent with one another. The Federal Reserve national survey of consumer wealth in 1962 included 86 farm families (Projector, 1964). From the size distribution of total non-human wealth for these families (not published but provided to the author in correspondence), the squared coefficient of variation of total wealth ownership is about 2.4. This figure is somewhat smaller than the U.S. squared coefficient of variation of real estate ownership in 1964 as estimated from the data of Table 4 as 3.2.

Using the same approach as was used in deriving equation (3), total wealth inequality can be estimated as,

$$C^2(W) = S_L^{\ 2}\, C^2(L) + S_K^{\ 2}\, C^2(K) + 2S_L S_K\, \rho_{LK}\, C(L)C(K),$$

where $C^2(W)$ is the 2.4 figure, $C^2(L)$ is 3.2, and S_L and S_K are the shares of real estate and other non-human capital, respectively, in total non-human wealth. There are two unknown quantities in this equation: $C^2(K)$ and ρ_{LK}. $C^2(K)$ was taken above as the geometric mean of $C^2(L)$ and $C^2(H)$. Once this arbitrary step is taken, ρ_{LK} is determined. For the 1964 U.S. data, the implied ρ_{LK} is about .6. This figure is used for all states.

<u>Factor Shares</u>. Factor share estimates in this paper are based on estimated full income, not reported money income, and therefore are

quite different from previously reported factor share estimates in U.S. agriculture. For full labor income an estimate of average earnings of full time (50-52 weeks) earners was obtained from the state data on the earnings of men and women who worked 50-52 weeks off the farm in 1969 (U.S. Bureau of the Census, 1972, Part 2, Table 195).[13] For non real estate farm capital, data on the value of machinery and equipment on farms and the value of livestock and poultry inventories in 1969 (U.S. Bureau of the Census, 1969a, Tables 6 and 7) were converted to flows by means of the rate of interest on non real estate debt.[14] Full income to real estate was calculated by multiplying value owned minus real estate debt by an interest rate of .07.[15] This procedure implicitly counts expected or normal capital gains as well as current receipts as part of land income. Receipts from interest, dividends, rents, and one-half of nonfarm self-employment income are counted as income to other non-human resources (K). Social security and pension income are also included in this category on the grounds that they are a return to past savings. The resulting state estimates of mean full income and the three relative shares are shown in Table 6.

Aggregation to Coefficient of Variation of Full Income. The estimated factor shares and variances and covariances of factor ownership

[13] This procedure assumes that farm work yields the same normal annual returns as does off-farm work by a comparably-skilled individual. Implicitly any labor market disequilibria are placed with the transitory income components that are left out of account.

[14] These returns should be adjusted by subtracting out returns to resources on farms owned by nonfarm residents. This fraction will be lower than the almost two-fifths nonfarm ownership of land because tenants typically own some of the equipment they use. I attribute all returns to equipment and inventories (minus interest on non real estate debt) to the incomes of full and part owners, and one-half of these returns to the incomes of tenants. This is admittedly a very rough procedure. However, an error here will make only a small difference in the estimated shares. For example, in a typical state the share in net farm income of these returns would be around 10 percent, and with 30 percent tenant farms an adjustment from one-half to, say, two-thirds as the correct operator share would change our estimated share of land by approximately .002.

[15] Approximately the average interest rate on new farm mortgage loans in 1969.

Table 6. Estimated full income and factor shares, 1969

State	Full income per farm (Household) ($)	Land's share	Labor's share	Capital share
Maine	13516	.150	.636	.213
New Hampshire	16711	.165	.626	.209
Vermont	15484	.221	.559	.220
Massachusetts	18898	.201	.569	.231
Connecticut	23077	.266	.507	.227
New York	16086	.170	.630	.200
New Jersey	24006	.306	.470	.223
Pennsylvania	16374	.188	.653	.159
Ohio	16559	.205	.638	.157
Indiana	17655	.218	.624	.158
Illinois	18475	.282	.563	.155
Michigan	16948	.168	.670	.162
Wisconsin	14275	.173	.640	.187
Minnesota	13776	.238	.576	.186
Iowa	17035	.268	.563	.169
Missouri	13931	.227	.606	.167
North Dakota	15492	.301	.512	.187
South Dakota	14897	.280	.506	.214
Nebraska	15885	.299	.504	.198
Kansas	16231	.280	.535	.185
Delaware	21507	.222	.452	.326
Maryland	19363	.287	.515	.198
Virginia	13084	.215	.605	.180
West Virginia	11859	.128	.685	.188
North Carolina	11949	.192	.661	.147
South Carolina	12500	.216	.640	.144
Georgia	13899	.253	.598	.149
Florida	20956	.378	.455	.167
Kentucky	12023	.183	.666	.151
Tennessee	11469	.180	.668	.153
Alabama	12370	.179	.668	.153
Mississippi	12233	.244	.595	.161
Arkansas	13676	.274	.567	.159
Louisiana	14046	.266	.568	.166
Oklahoma	14915	.265	.559	.176
Texas	16591	.320	.503	.177
Montana	20453	.374	.414	.212
Idaho	17487	.289	.526	.185
Wyoming	20840	.344	.406	.250
Colorado	18420	.321	.484	.194
New Mexico	20924	.398	.413	.189
Arizona	37355	.607	.245	.148
Utah	17039	.251	.583	.165
Nevada	31340	.422	.333	.246
Washington	21306	.292	.539	.168
Oregon	19470	.250	.557	.194
California	27708	.426	.409	.165
United States total	15000	.22	.62	.16

are plugged into equation (3) to get an estimated squared coefficient of variation of full income for each state. Two estimates were made which handle the missing capital ownership data differently as discussed above. The two methods are close enough that when the results are used to generate the estimated full-income poverty data of the following section, it makes no appreciable difference which is used. The estimate used, shown in Table 7, takes $C^2(X)$ as the geometric mean of $C^2(L)$ and $C^2(H)$.

V. Estimating Full-Income Poverty

In order to transfer from a measure of inequality of full income, which equation (3) and Table 7 provide, to a measure of full-income poverty, it is assumed that the log of full income is distributed lognormally. Under the lognormal assumption, the proportion of the population below any income level chosen as the poverty line, Y_p, is the cumulative density of the standardized normal distribution up to $(\ln Y_p - \overline{\ln Y}) / \sqrt{Var(\ln Y)}$. The accuracy of the lognormality assumption for this purpose can be tested on the census money income data for which f_p, the proportion below Y_p, can be measured directly. If these percentages can be predicted accurately using the lognormality assumption, then it does not seem unreasonable to apply the same approach to full income poverty.

Accordingly, the cumulative normal procedure was used to predict the values of f_p for 1969 using as Y_p the Social Security Administration's poverty threshold for a farm family of four, $3191. Var (ln Y) was estimated from the published census state size distributions of money income (U.S. Bureau of the Census, 1972, Part 2, Table 57). The standard error of the predicted f_p was .013. The mean error of the predicted f_p was .004, on the average, there were slightly more households below the poverty threshold than the lognormal procedure predicts. One possible reason is that the log variance of the size distribution of income is understated by estimating it from grouped data. Therefore, the estimated log variances were increased by intervals of 1 percent until the mean squared error between actual and predicted f_p was

Table 7. Estimated relative variance (squared coefficient of variation) of full income of rural-farm households, by states, 1969

Northeast:	
Maine	0.285
New Hampshire	0.280
Vermont	0.266
Massachusetts	0.353
Connecticut	0.358
New York	0.292
New Jersey	0.438
Pennsylvania	0.270
Midwest:	
Ohio	0.257
Indiana	0.249
Illinois	0.334
Michigan	0.228
Wisconsin	0.218
Minnesota	0.231
Iowa	0.249
Missouri	0.294
North Dakota	0.226
South Dakota	0.280
Nebraska	0.312
Kansas	0.309
South:	
Delaware	0.504
Maryland	0.417
Virginia	0.437
West Virginia	0.332
North Carolina	0.319
South Carolina	0.401
Georgia	0.390
Florida	0.775
Kentucky	0.328
Tennessee	0.308
Alabama	0.380
Mississippi	0.549
Arkansas	0.498
Louisiana	0.560
Oklahoma	0.361
Texas	0.539
West:	
Montana	0.380
Idaho	0.363
Wyoming	0.497
Colorado	0.407
New Mexico	0.607
Arizona	0.978
Utah	0.350
Nevada	0.591
Washington	0.333
Oregon	0.382
California	0.565
United States total	0.53

minimized. The optimum adjustment turned out to be 1.04 times the directly estimated log variance of income. The actual and predicted values are shown in Table 8. The lognormal hypothesis produces no great errors.

It would seem appropriate to have a full income poverty threshold higher than that for money income for rural families. One reason is that the official rural-farm threshold is set at 85 percent of the nonfarm threshold because of the relatively great importance of non-money income for farms. So at least this 15 percent should be added on when the estimate of full income is used.

A more difficult problem arises from the fact that the poverty threshold is not an objective, observable subsistence income level but is ultimately a matter of values and tastes.[16] If the threshold were formulated in terms of full income, everyone's income would be higher. Therefore, to the extent that poverty is a relative concept (or, in Schultz's (1968) terminology, that the income elasticity of the poverty line is greater than zero), the poverty threshold should be increased when a full income measure is used.

Unfortunately, there is no way to tell how much the threshold should be increased. Two alternatives are used in this paper: Schultz's rough estimate of .55 as the income elasticity of the poverty line, and an elasticity of one (which increases the poverty line by the same percentage as median income). This latter alternative provides a relative measure of poverty in that the poverty line is defined to be at a given percentage of mean income.

The 1970 census estimate of the mean income of rural-farm families and unrelated individuals is $6,253 (U.S. Bureau of the Census, 1973, Part 1, Table 94). My corresponding estimate of mean full income is about $15,000. The U.S. average rural-farm poverty line for 1970 was $2,577 as measured in terms of money income.[17] Adding back the

[16] See Schultz (1968, pp. 65-79), for an interesting discussion of the role of preferences in the determination of the poverty line.

[17] Weighted average of $2,927 for families (75 percent of rural-farm poor households) and $1,527 for unrelated individuals (25 percent of poor households).

Table 8. Fraction of rural-farm households below census income of $3191, and fraction predicted by lognormal hypothesis

State	Actual < $3191	Predicted < $3191
Maine	.280	.294
New Hampshire	.192	.202
Vermont	.228	.251
Massachusetts	.180	.170
Connecticut	.157	.145
New York	.191	.188
New Jersey	.191	.187
Pennsylvania	.201	.201
Ohio	.202	.182
Indiana	.186	.171
Illinois	.189	.186
Michigan	.190	.171
Wisconsin	.185	.186
Minnesota	.240	.253
Iowa	.171	.182
Missouri	.279	.275
North Dakota	.213	.241
South Dakota	.258	.285
Nebraska	.220	.239
Kansas	.223	.235
Delaware	.199	.207
Maryland	.212	.209
Virginia	.330	.330
West Virginia	.333	.318
North Carolina	.343	.332
South Carolina	.334	.324
Georgia	.292	.290
Florida	.275	.264
Kentucky	.343	.333
Tennessee	.360	.350
Alabama	.336	.316
Mississippi	.443	.438
Arkansas	.347	.342
Louisiana	.366	.358
Oklahoma	.272	.271
Texas	.282	.282
Montana	.219	.241
Idaho	.197	.206
Wyoming	.199	.222
Colorado	.248	.256
New Mexico	.293	.298
Utah	.209	.207
Nevada	.227	.237
Arizona	.316	.312
Washington	.191	.179
Oregon	.192	.191
California	.213	.210

rural-urban differential mentioned above increases this figure to \$3,032. Keeping a constant relative poverty line when moving to full income implies a full income poverty line of 3032/6253 x 15000 = \$7,300. Using the .55 "income elasticity" yields 3032 + (15000 - 6253)/6253 x .55 x 3032 = \$5,400.

Now applying the methods used in Table 8 yields estimates of the fraction of rural-farm households below each of these poverty thresholds. Under the lognormal assumption Var (ln Y) is obtained from the $C^2(Y)$ data of Table 7 using the relation [Aitchison and Brown, 1957, p. 154]:

$$\text{Var}\ (\ln Y) = \ln\ (1 + C^2(Y)).$$

The results are presented in Table 9, together with census estimates of rural-farm poverty using money income. Since it is not possible to say what the "correct" full income poverty line is, both the "optimistic" \$5,400 and the "pessimistic" \$7,300 poverty line are presented.

Full income poverty under the \$7,300 threshold would be the same as predicted money income poverty under a \$3,032 threshold if the log variance of money income and full income were the same. This occurs under lognormality because log (15000-7300)= log (6253-3032). Predicted full income poverty is lower than the "official" figures even in this case because the estimated log variance of full income is less than that of money income. This difference in variances is to be expected, as discussed above; nonetheless, it should be recognized that if the estimated log variance of full income is underestimated in Table 7, then so will be estimated full income poverty in Table 9.

In order to get an indication of the sensitivity of the estimates to the procedures used, an alternative measure of full income poverty is presented in column (3) of Table 9. This alternative measure uses the assumption of lognormality in the distribution of income at all stages. If equation (1) is replaced by an income generating function of Cobb-Douglas form, an equation identical to (3) results except that logarithmic variances replace all the squared coefficients of variation. Accordingly, the earnings function (4) is specified in natural logarithms and the log variance instead of the squared coefficient of variation is

Table 9. U.S. census and full income estimates of rural-farm poverty, 1970

State	Census Estimate[a]	Full income estimates[b] (1)	(2)	(3)
Maine	.219	.041	.140	.124
New Hampshire	.124	.015	.068	.051
Vermont	.203	.020	.084	.068
Massachusetts	.108	.016	.064	.050
Connecticut	.099	.007	.032	.030
New York	.146	.023	.090	.064
New Jersey	.125	.010	.037	.039
Pennsylvania	.154	.014	.066	.068
Ohio	.134	.016	.076	.077
Indiana	.119	.011	.056	.061
Illinois	.131	.017	.067	.078
Michigan	.130	.010	.059	.057
Wisconsin	.145	.023	.107	.092
Minnesota	.197	.033	.133	.132
Iowa	.131	.013	.062	.084
Missouri	.196	.044	.148	.133
North Oakota	.182	.019	.089	.106
South Oakota	.221	.032	.116	.118
Nebraska	.180	.031	.108	.122
Kansas	.150	.027	.100	.098
Oelaware	.141	.019	.061	.039
Maryland	.160	.022	.073	.070
Virginia	.258	.084	.206	.173
West Virginia	.258	.079	.218	.207
North Carolina	.281	.092	.247	.235
South Carolina	.302	.098	.239	.215
Georgia	.239	.068	.182	.159
Florida	.196	.046	.107	.081
Kentucky	.271	.087	.234	.209
Tennessee	.288	.089	.244	.230
Alabama	.275	.092	.233	.198
Mississippi	.370	.123	.257	.215
Arkansas	.273	.088	.205	.162
Louisiana	.323	.090	.202	.160
Oklahoma	.186	.046	.139	.125
Texas	.203	.056	.141	.114
Montana	.169	.015	.056	.074
Idaho	.147	.026	.088	.077
Wyoming	.143	.021	.064	.065
Colorado	.194	.027	.087	.086
New Mexico	.267	.035	.092	.085
Utah	.165	.015	.039	.073
Nevada	.158	.025	.090	.036
Arizona	.288	.007	.023	.072
Washington	.141	.009	.042	.039

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Table 9 (continued)

State	Census Estimate[a]	Full income estimates[b]		
		(1)	(2)	(3)
Oregon	.136	.017	.062	.048
California	.140	.011	.036	.049
United States total	.20	.05	.13	.14

[a]The census estimate is from [U.S. Bureau of the Census, 1972, Part 2, Table 58].

[b]Full income estimate (1) uses a full income poverty threshold of $5400 and (2) and (3) use $7300. (1) and (2) use the relative variance of full income given in Table 7, while (3) uses the alternative procedure described in the text.

calculated for land ownership in Table 5. Because the log variance puts more weight on inequality at the lower tail relative to the upper tail, the state-to-state pattern of log variances is somewhat different than for the coefficient of variation. However, columns (2) and (3) do not differ a great deal.

The range of full income estimates yields poverty percentages substantially smaller than the census figures for every state. However, there are marked regional differences. Using the "pessimistic" estimates (2) and (3), full income poverty is less than half of census money income poverty in several northeastern and western states. But in several southern states, there remain many rural-farm households in full income poverty. The regional concentration of rural poverty in the South, which is apparent in the census data, is still greater when the full income measure of poverty is used.

VI. Summary and Conclusion

This research is an attempt to provide a measure of rural-farm poverty conceptually more meaningful than that provided by Census of Population money income data. The measure proposed is based on full income, defined as the purchasing power available for consumption and savings in a normal year. Full income is not measured directly in any national surveys, nor is there comprehensive information on many of the components necessary to construct it. Consequently, the generation of data on the personal distribution of full income required many assumptions and short-cuts. The resulting measure is certain to contain substantial error. Nonetheless, the range of estimates of full income poverty may provide a useful first approximation to the order of magnitude of revision that an improved measurement of income would lead to in statistics on rural poverty. For the U.S. as a whole in 1969, the census annual money income data yield 19.9 percent of rural-farm families and unrelated individuals with income below the poverty level. In contrast, the corresponding full income percentage is estimated to be in the range 5 to 14 percent depending on the assumptions about the full income poverty line and the distribution of full income. The estimated

more equal distribution of full income than annual money income accounts for the reduction from 20 to 13-14 percent, while the higher average level of full-income than money income accounts for the further reduction to the neighborhood of 5 percent. Similar data for individual states are presented in Table 9.

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