#526-78

INSTITUTE FOR RESEARCH ON POVERTY DISCUSSION PAPERS

INCOME STABILITY IN THE UNITED STATES

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December 1978

This research was supported in part by funds granted to the Institute for Research on Poverty at the University of Wisconsin-Madison by the Department of Health, Education, and Welfare pursuant to the Economic Opportunity Act of 1964.

ABSTRACT

This study applies a new index of income mobility which exploits the fact that changes in relative income positions over time are reflected in the degree to which incomes are equalized as the accounting period is extended. This enables the characteristic features of income dynamics to be portrayed in the form of "rigidity curves" or "stability profiles." There will be one such curve corresponding to each measure of income inequality.

Using the Michigan panel data, stability profiles are constructed for different age, sex, occupation, and education groups, and the extent to which the level and type of income mobility varies across these groups is discussed. The paper also compares the stability of earnings with that of total family income, and examines the degree to which the index values are sensitive to the choice of the measure of inequality.

Income Stability in the United States

1. INTRODUCTION

The recent availability of large-scale comprehensive panel data in the United States has opened up new vistas for research into the determinants of income and income dynamics. This paper is concerned with two particular issues: first, the degree to which measured income inequality is affected by the choice of accounting period; and second, the problem of characterizing income movements over time in a way that highlights the major influences. In the approach we shall follow, these two aspects are intimately related.

The potential importance of the income accounting period in any analysis of individual or family welfare is well documented in the literature. The basic argument is that a comparison of, say, annual incomes will exaggerate the true degree of inequality, since some people will have incomes that are temporarily higher or lower than usual and these short-run fluctuations tend to increase income dispersion. Taken to its extreme, this line of reasoning leads to the suggestion that the appropriate income concept in welfare comparisons is that of lifetime income, in which case we would also need to compensate for income differences arising from differences in age and variations in the shapes of age-income profiles. A suggestion along these lines was made by Paglin (1975). The argument against the use of the lifetime income concept points out that it may be little consolation for someone currently poor to know that they can expect to be better off in the future (if they have no way of borrowing against this expectation) or to reflect on their spendthrift behavior in past years. This paper takes no particular position in this debate. But clearly the importance attached to the issue depends to a large extent on the comparative magnitudes of short-run and long-run income inequality. If there were little difference between, say, weekly and lifetime income inequality, then the question of which accounting period is appropriate would not be of such great concern.

A number of earlier studies have dealt with the empirically observed relationship between annual inequality and inequality over several years. However, these studies have not made the appropriate comparison between short-run and long-run inequality and, for this reason, can produce counter-intuitive results. This point is pursued further in Section 2.

The general description and analysis of income dynamics are still in their infancy. In Britain the usual approach has been to assume some simple stochastic process representation of income movements and to estimate the relevant parameters (see Hart, 1976a, 1976b; Shorrocks, 1976). Recent research in the United States has generated the equations for estimation from the conventional earnings equations, and added complex error structures designed to capture various aspects of income changes (see Lillard and Willis, 1978). The problems faced by all this work are quite severe, since income changes occur for a multitude of reasons and the specification errors resulting from the imposition of any particular structure can cause serious distortions in the parameter estimates. This may lead to an inappropriate interpretation of the results.

Shorrocks (1978b) describes a method for summarizing some of the main features of income dynamics without imposing any theoretical structure on the process underlying income mobility. Essentially the approach views income movements in terms of their impact on the inequality value as the accounting period is extended. The degree of income mobility is associated with the extent to which equalization occurs as incomes are cumulated over longer intervals, and the major features of relative income movements can be visually represented by "rigidity curves" or "stability profiles." Section 2 describes these ideas in more detail.

In Sections 3-8 the approach is applied to the Michigan panel data on income dynamics. Certain pronounced and consistent overall patterns emerge, and these provide a useful guide to the process governing income changes. Particular emphasis is placed on the quantitative importance of short-run income fluctuations, which can be loosely equated with variations in "transitory income," relative to longer-term income mobility, which can be regarded as changes in "permanent incomes."

The empirical sections begin with an analysis of earnings and family income data disaggregated by age. I then examine the results for earnings in various occupations and schooling categories.

2. INCOME MOBILITY AND THE INCOME ACCOUNTING PERIOD

The effect of extending the accounting period on measured income inequality has been examined in several studies: for example, Benus and Morgan (1975), Hanna (1948), Kohen et al (1975), Kravis (1962, chap. 8), and Vandome (1958). The comparison has almost always been one between annual inequality and inequality over two or more years, and the broad conclusion is that a longer accounting period reduces inequality very little. In fact some of the empirical results seem to suggest that inequality can increase with the longer measurement interval. Kravis (1962, p.272), for example, reports evidence for a sample of Delaware taxpayers over the period 1925-36. The Gini coefficient G is given for the individual years, and also for incomes cumulated over the first two years, the first three years and so on. The results show that the value of G increased monotonically with the length of the accounting period. However, the trend is reversed if incomes are cumulated backward from the last year, instead of forward from the first.

This apparent paradox was explained correctly by Kravis. Most income data, including the Delaware sample, show a tendency for relative inequality to increase as cohorts of individuals get older. This increasing dispersion within age groups can offset the degree to which cumulating incomes over time reduces inequality values. When the data is cumulated backward from the last year, both factors operate in the same direction.

Recent studies have compared the longer-run inequality value with the means of the annual values. This clearly overcomes the problem of obtaining contradictory results depending on whether incomes are cumulated forward or backward, but it is not an entirely satisfactory solution. With the Gini coefficient, for example, inequality of 2-year incomes can exceed the mean of the annual Gini values. The appropriate comparison is between long-run inequality and a weighted average of the annual inequality values, with the weights proportional to the mean income received in each of the years.

The justification for this last remark rests on a more detailed analysis of the way in which an extension of the accounting period tends to reduce measured inequality. Consider a population of n individuals with incomes recorded in m consecutive time periods. Let $\underset{\sim}{Y}\,(t_{k-1}, t_k)$ be the vector of incomes in the time interval $[t_{k-1}, t_k)$ and $\mu\ (t_{k-1}, t_k)$ be the corresponding mean income. Now suppose we restrict ourselves to measures of relative inequality (i.e., those that are "mean independent," or homogeneous of degree zero in incomes) that are convex functions of relative incomes. Then

$$I\ [\,Y(t_o,t_m)\,] = g\,(\Sigma_k\ \underset{\sim}{Y}\ (t_{k-1},t_k)/\mu(t_o,t_m)) \tag{1}$$

$$\leq \Sigma_k \frac{\mu(t_{k-1},t_k)}{\mu(t_0,t_m)}\ g(\underset{\sim}{Y}(t_{k-1},t_k)/\mu(t_{k-1},t_k))$$

So $$I[Y(t_0,\ t_m)] \leq \Sigma_k w_k\ I[Y(t_{k-1},\ t_k)] \tag{2}$$

where

$$w_k = \frac{\mu(t_{k-1},\ t_k)}{\mu(t_0,\ t_m)}\ . \tag{3}$$

Equation (2) indicates that m-period inequality can never exceed a weighted sum of the single period inequality values, where the weights w_k represent the proportion of the aggregate m-period income received in period k (and consequently the w_k sum to unity). It is therefore the formal statement corresponding to the notion that cumulating incomes over time tends to reduce inequality. The assumption that $g(\cdot)$ is a convex function of relative incomes covers most of the common inequality measures, including the Gini coefficient, the coefficient of variation, the Theil coefficient, and the Atkinson family of indices. The notable exception is the variance of the logarithm of income, which is widely used despite the fact that it violates the Dalton-Pigou principle of transfers and for this reason should not be called an index of inequality (although it can be regarded as an index of dispersion).

The relationship given by (2) holds with equality if the relative incomes of all individuals remain constant through time. In other words ,

$$Y(t_{k-1}, t_k)/\mu(t_{k-1}, t_k) \text{ is independent of } k \tag{4}$$

implies

$$I[Y(t_0, t_m)] = \Sigma_k w_k I[Y(t_{k-1}, t_k)]. \tag{5}$$

The converse of this statement is also true when $g(\cdot)$ is a strictly convex function of relative incomes. (This excludes the Gini coefficient from the above list.) Thus, for this subclass of indices, m-period inequality will always be **less** than the weighted annual average unless relative incomes remain fixed.

The above discussion suggests that the impact of the accounting period on inequality is best described by computing the ratio

$$R = \frac{I[Y(t_0, t_m)]}{\Sigma_k w_k I[Y(t_{k-1}, t_k)]}. \tag{6}$$

This ratio has a maximum value of 1 for indices of the form given in equation (1), and deviations of R from unity indicate the extent of equalization that occurs as the accounting period is lengthened.[1] Clearly the values of R have intrinsic interest in themselves for those concerned with the welfare implications of selecting different accounting intervals, and the later sections present the results of computations made for various population subgroups.

A society in which relative incomes are constant through time is appropriately called completely (income) immobile. Thus, for inequality measures that are strictly convex functions of relative incomes, R = 1 if and only if the society is completely immobile. In fact there are reasonable grounds for taking R to be a measure of income rigidity (and M = 1 - R to be a measure of mobility). For instance, consider the m-period income distribution to be fixed. In a more mobile society, income changes will be more frequent and/or more pronounced. These greater fluctuations will tend to inflate the inequality values in the subperiods, and the value of R will fall as a consequence. From this viewpoint, income mobility is seen as the degree to which incomes are equalized as the accounting period is extended.[2] It has the particular attraction of integrating the static aspects of income inequality, as represented by the measure of inequality, and the dynamic aspects, as reflected in the measure of mobility. In later sections I shall comment on the amount and type of income mobility suggested by the data.

Before any computations can be made, two important dimensions of choice need to be decided. The first concerns the selection of an inequality measure I of the type given in (1). As might be expected, the exact extent of the tendency toward equalization depends on the particular index chosen, since inequality measures vary in their sensitivity to different portions of the distribution. By examining values of R for alternative indices, it is possible to get some idea of the degree to which equalization is occurring at different places in the distribution. For example, the particular characteristics of the Gini coefficient accounts for the relatively small trend towards equality found in the earlier studies that used this index.

Four different indices are used in this paper: G, the Gini coefficient; C_2, the square of the coefficient of variation; C_1, the Theil coefficient; and C_0, another "entropy" index proposed by Theil. These indices are given by

$$C_2 = \frac{1}{n} \Sigma_i [(\frac{y_i}{\mu})^2 - 1]$$

$$C_1 = \frac{1}{n} \Sigma_i \frac{y_i}{\mu} \log \frac{y_i}{\mu}$$

$$C_0 = \frac{1}{n} \Sigma_i \log \frac{\mu}{y_i} . \tag{7}$$

C_0, C_1 and C_2 are members of a single-parameter family C_β, whose members are all mean independent, additively decomposably, and strictly convex functions of relative incomes (see Shorrocks, 1978c). Moreover, it can be shown that the members of the family become more sensitive to the bottom of the distribution as β decreases. Thus the indices used range from C_2, which is extremely sensitive to the distribution of top incomes, to C_0, which is particularly sensitive to lower incomes.

The second major dimension of choice in the computation of R is the number of individual subperiods over which incomes are cumulated-- the parameter m in the earlier notation. The ability to choose different overall time horizons is particularly useful in the discussion of mobility, since the amount of mobility observed depends on both the mobility inherent in the income structure and the length of time during which potential income movements can be translated into actual changes. Even in an extremely mobile society, little income movement can be expected in a very short period of time.[3]

One method of highlighting the impact of time on observed mobility is to compute the values of R for m = 2, 3, . . . up to the maximum that the data allow. The degree of income rigidity or income stability can then be represented diagramatically by a graph of R against m, to obtain <u>rigidity curves</u> or <u>stability profiles</u>. The reference curve is the horizontal line R = 1, representing a completely immobile structure. The rigidity curves for other structures will lie below this horizontal and are normally expected to decline with the value of m, although they can slope upwards (and do so in the results below, for reasons that will be explained).

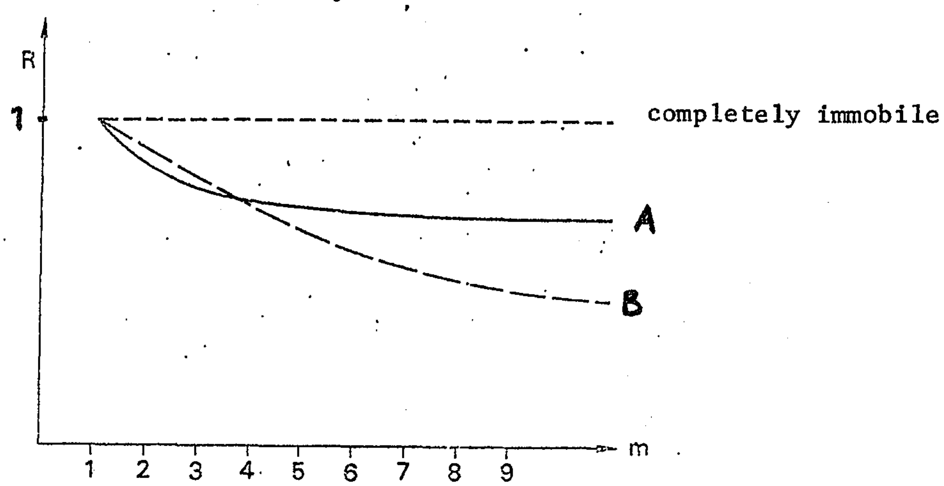


Figure 1. Rigidity Curves.

Two different types of rigidity profiles are shown in Figure 1. Curve B shows a continuous egalitarian trend as the time horizon is stretched. This is the kind of curve we might expect if, for example, the underlying mobility was such that lifetime incomes were indentical for all individuals: $I(Y(t_0, t_m))$ would tend to zero as t_m approaches the end of the lifetime, and R therefore would eventually become zero. Curve A shows a substantial initial fall, but then becomes more or less horizontal. This is the type of curve we should expect if there is a good deal of "transitory" fluctuation in income, but little or no variation in "permanent" income.

To illustrate this point, consider the extreme case in which permanent incomes y^p_{it} are fixed through time and an additive random transitory component ε_{it} is the only source of income movements. Thus

$$y^p_{it} = y^p_i$$

$$y^a_{it} = y^p_i + \varepsilon_{it}. \tag{8}$$

Assume that ε_{it} is independently distributed across individuals and through time, with mean 0 and variance $\sigma^2(\varepsilon)$. Then it (roughly) follows that

$$\bar{y}^p_t = \bar{y}^a_t = \bar{y}$$

$$\sigma^2(y^a_t) = \sigma^2(y^a) = \sigma^2(y^p) + \sigma^2(\varepsilon). \tag{9}$$

Furthermore, if

$$\begin{aligned} Y^a_{im} &= \sum_{t=1}^{m} y^a_{it} \\ &= \sum_{t=1}^{m} [y^p_i + \varepsilon_{it}] \\ &= my^p_i + \sum_{t=1}^{m}\varepsilon_{it} , \end{aligned} \tag{10}$$

then

$$\bar{Y}^a_m = m\bar{y}^p$$

and $$\sigma^2(Y^a_m) = m^2\sigma^2(y^p) + m\sigma^2(\varepsilon). \tag{11}$$

For the square of the coefficient of variation, C_2, we have

$$C_2(y^a_t) = \frac{\sigma^2(y^a_t)}{\bar{y}^2} = C_2(y^p) + \frac{\sigma^2(\varepsilon)}{\bar{y}^2}$$

and $$C_2(Y^a_m) = C_2(y^p) + \frac{\sigma^2(\varepsilon)}{m\bar{y}^2}. \tag{12}$$

Thus the rigidity value R_m corresponding to m periods will be

$$R_m = \frac{C_2(Y^a_m)}{C_2(y^a)} = 1 - \frac{(1 - \frac{1}{m})\sigma^2(\varepsilon)}{C_2(y^p)\bar{y}^{2} + \sigma^2(\varepsilon)} . \tag{13}$$

Taking the limit as $m \to \infty$ gives

$$R_\infty = 1 - \frac{\sigma^2(\varepsilon)}{C_2(y^p)y^{-2} + \sigma^2(\varepsilon)} \tag{14}$$

and R_m may be rewritten in the convenient form

$$R_m = 1 - (1 - \frac{1}{m})(1 - R_\infty)$$

or
$$1 - R_m = (\frac{m-1}{m})(1 - R_\infty). \tag{15}$$

The value of R_m therefore converges to its limiting value fairly rapidly: one-half the distance when two periods are aggregated, two-thirds the distance when m = 3, and so on. For example, if R_∞ were .88 this income structure would generate the sequence $\{R_m\}$ = 1.00, .94, .92, .91, .90, .90, .90, . . . ;

A detailed examination of the shape of the rigidity curves may consequently help to distinguish between income changes that are relatively short lived and those that continue to exert their influence.[4]

3. LABOR INCOME OF MALES

Data for the labor incomes of male head of households are summarized in Table 1,[5] using the Gini coefficient as the index of inequality. For each of the ten-year age groups there is a somewhat erratic trend toward increasing inequality with age. The stability index R is computed from equation (6) with I = G and t_k equal to the beginning of the year 1967 + k. R was calculated for nominal incomes, real incomes and real incomes discounted by 2%. It soon became clear that these three income concepts

Table 1

Real Labor Incomes, Males, by Age Group: Gini Coefficient

	Age Group in 1967					
	< 20	20-29	30-39	40-49	50-59	60 +
Mean Income ($)						
1967	2582	5904	7906	8284	7329	6510
1968	3357	6574	8496	8628	7415	6017
1969	3882	7045	8941	8754	7476	6210
1970	4715	7268	8709	8796	7800	5718
1971	5291	7576	8868	8947	7473	4995
1972	6478	8237	9480	9237	7644	4297
1973	6699	8558	9861	9414	7367	4350
1974	6454	8424	9207	9250	6760	3693
1975	6295	8314	8946	8683	5976	3557
Gini Coefficient (individual years)						
1967	.29	.29	.30	.31	.38	.47
1968	.26	.25	.31	.31	.39	.52
1969	.26	.25	.32	.31	.38	.49
1970	.32	.26	.30	.32	.42	.56
1971	.22	.25	.30	.32	.42	.57
1972	.20	.26	.31	.32	.43	.61
1973	.22	.27	.32	.32	.42	.59
1974	.23	.29	.33	.34	.45	.55
1975	.23	.30	.35	.34	.47	.56
Stability Index R						
1967	1.00	1.00	1.00	1.00	1.00	1.00
1967-68	.91	.95	.98	.98	.98	.97
1967-69	.82	.92	.97	.97	.97	.97
1967-70	.75	.90	.96	.97	.97	.95
1967-71	.71	.88	.95	.96	.96	.94
1967-72	.68	.86	.95	.95	.96	.94
1967-73	.69	.85	.94	.95	.94	.92
1967-74	.70	.85	.94	.94	.93	.91
1967-75	.71	.84	.93	.94	.92	.90
Sample Size						
	33	432	526	443	226	33

made virtually no difference to the values of R, and so throughout this paper, to avoid duplication, R is reported only for real incomes.

The first point of interest is the fairly pronounced age effect on R. As might be expected, income instability is greatest for the youngest group, followed by those aged 20-29. The pattern then stabilizes for the next three age groups and shows a slight tendency to increase again for those over 60. (Of course only a small proportion of those over 60 have positive earnings throughout the 9-year period (approximately 15%) and those that do are likely to move to new, lower paying jobs; hence the decline in mean income recorded for this group). For those above age 30 it seems that the Gini drops by about 5% when a 5-year accounting period is used, and by about 8% over 9 years. These are fairly modest amounts, indicating a reduction in the Gini of around .02 to .04. This is generally smaller than the positive age trend in inequality over the 9-year period. Since the two effects work in opposite directions, for any cohort of individuals the Gini coefficient for earnings in the current year should be a reasonable guide to the inequality in their aggregate earnings over the next 5 to 10 years. This is not the case for the younger age groups, where the 5-year Gini is 16% below the average annual value for those aged 20-29 in 1967 and 29% below for those under 20. This last figure suggests that the annual Gini for the youngest age group exaggerates longer-run inequality by approximately .07.

The shapes of the stability profiles are also of considerable interest. As anticipated, R generally falls monotonically as the overall time horizon is extended. In addition, there is a tendency for the initial decline to be greater than that recorded subsequently. This leveling off is particularly pronounced in the two youngest age groups, as can be seen from Figure 2. Since the curves for the age groups 30-39, 40-49, and 50-59 are very similar, only one "typical" curve has been drawn.[6]

The stability profile of the youngest age group is of special interest since the value of R shows a small increase at the end of the 9-year period. This means that the income movements in 1974 and 1975 have partially counteracted the earlier income changes and tended to reestablish the distribution of relative incomes over, say, the first 5 years. The likely explanation for this phenomenon is the timing of promotions (interpreted in a broad sense to cover any significant income change that tends to persist). For if some members of the age cohort are promoted one year they may have to wait some time before they are considered for subsequent promotion. In the meantime others in the cohort may be promoted and the distribution of relative incomes may therefore move back toward that observed earlier.

We now turn to the impact of the choice of inequality measure. In Table 2 rigidity values are given for the three alternative indices C_o, C_1, C_2. For each of these indices the general pattern is similar to that found using the Gini coefficient : a marked age effect, and a tendency for the profiles to level off--particularly at younger ages. There is some variation across the indices in the precise value of R, but the figures are broadly comparable (except, perhaps, for the C_o values above age 60). However, they are substantially lower than the Gini figures

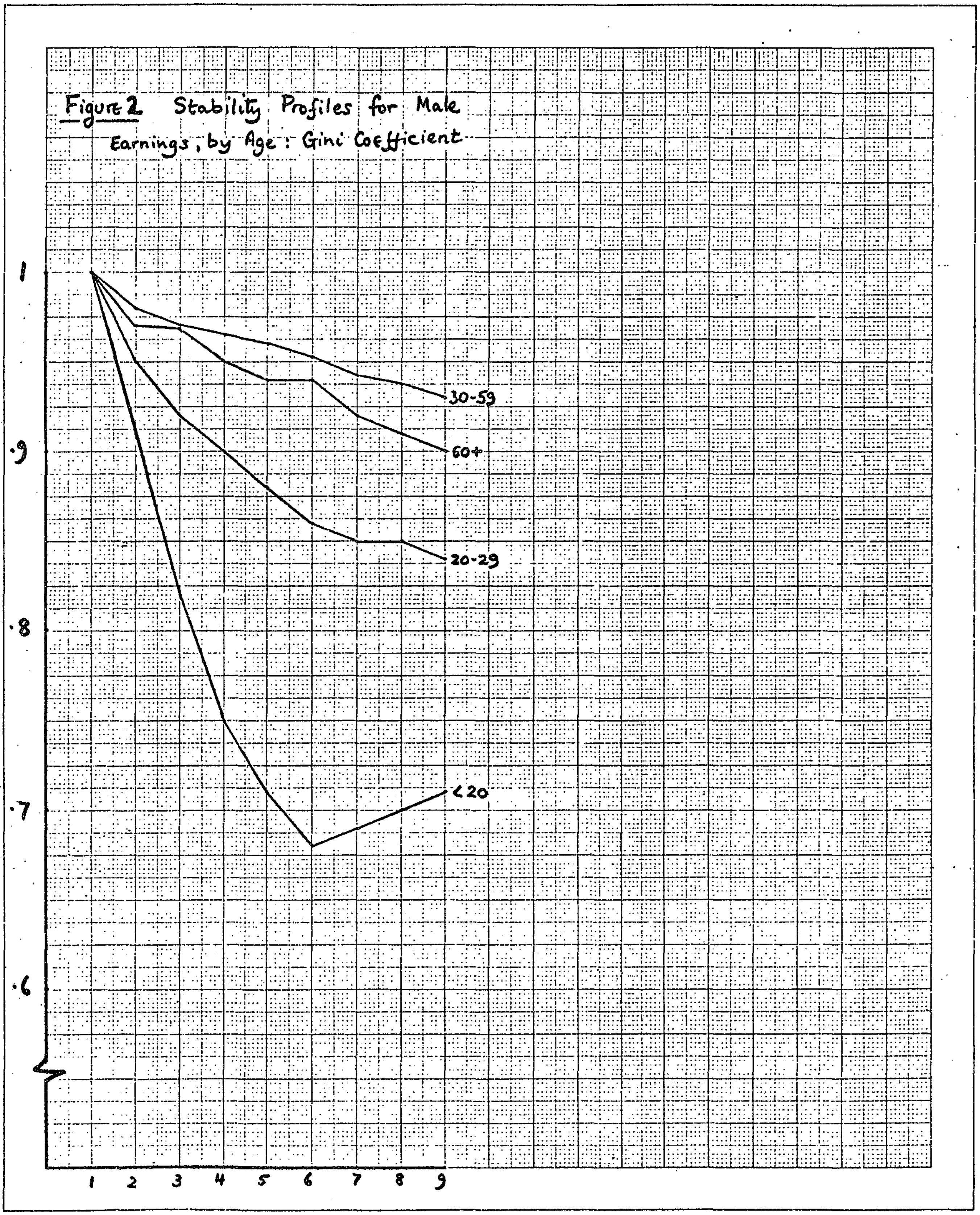


Figure 2 Stability Profiles for Male Earnings, by Age: Gini Coefficient
1
·9
·8
·7
·6
30-59
60+
20-29
<20
1 2 3 4 5 6 7 8 9

Table 2

Stability of Real Labor Incomes, Males, by Age Group: Indices C_0, C_1, C_2

		Age Group in 1967					
Index	Period	< 20	20-29	30-39	40-49	50-59	60 +
C_0	1967	1.00	1.00	1.00	1.00	1.00	1.00
	1967-68	.78	.86	.93	.93	.93	.85
	1967-69	.66	.78	.90	.91	.92	.81
	1967-70	.44	.73	.88	.89	.90	.78
	1967-71	.40	.69	.85	.86	.88	.75
	1967-72	.38	.65	.84	.85	.85	.74
	1967-73	.39	.65	.83	.83	.82	.71
	1967-74	.39	.63	.82	.82	.80	.70
	1967-75	.40	.62	.80	.81	.77	.69
C_1	1967	1.00	1.00	1.00	1.00	1.00	1.00
	1967-68	.82	.90	.95	.96	.95	.93
	1967-69	.67	.84	.93	.95	.94	.92
	1967-70	.54	.80	.90	.93	.92	.90
	1967-71	.49	.77	.88	.91	.90	.87
	1967-72	.45	.73	.87	.90	.89	.85
	1967-73	.46	.72	.86	.89	.87	.82
	1967-74	.46	.71	.85	.88	.85	.81
	1967-75	.47	.70	.84	.87	.84	.79
C_2	1967	1.00	1.00	1.00	1.00	1.00	1.00
	1967-68	.82	.89	.95	.96	.95	.96
	1967-69	.67	.85	.92	.95	.93	.96
	1967-70	.57	.81	.88	.94	.90	.93
	1967-71	.52	.79	.86	.92	.88	.88
	1967-72	.48	.76	.84	.91	.87	.85
	1967-73	.48	.71	.82	.90	.85	.81
	1967-74	.49	.70	.81	.88	.83	.80
	1967-75	.49	.69	.80	.87	.82	.79
Sample Size		33	432	526	443	226	33

given in Table 1. With few exceptions the difference between R and one is at least double the corresponding difference for the Gini coefficient. At ages 30-59, 9-year inequality is 13-23% below the average annual value compared with 6-8% for G. For the age group 20-29 the range is 30-38% compared with 16%; and for the youngest group, 9-year inequality is a massive 51-60% less than the average annual value, compared with a 29% fall registered for the Gini. Furthermore, more or less all this fall is recorded in just the first 5 years.

The rigidity profiles for alternative indices can be compared in Figure 3. The quantitative difference between the Gini and the other indices may come as a surprise, but there is a simple explanation. The Gini coefficient gives little weight to income transfers in either tail of the distribution. In this respect it is more or less unique among the conventional inequality measures. Since the main effect of cumulating incomes is to average out incomes that are temporarily high or low, the strongest egalitarian trend will be found in the tails. The distribution of relative incomes in the middle range is not substantially changed by cumulating incomes over time. Hence the Gini coefficient is not as sensitive to the accounting period as other measures. In some sense this is a justification for using the Gini. For if one felt that a time horizon longer than one year were more appropriate, the annual Gini value would be less biased than other annual inequality values. If one still prefers to use an index other than G, the results in Table 2 suggest that the choice of accounting period may significantly affect the index value.

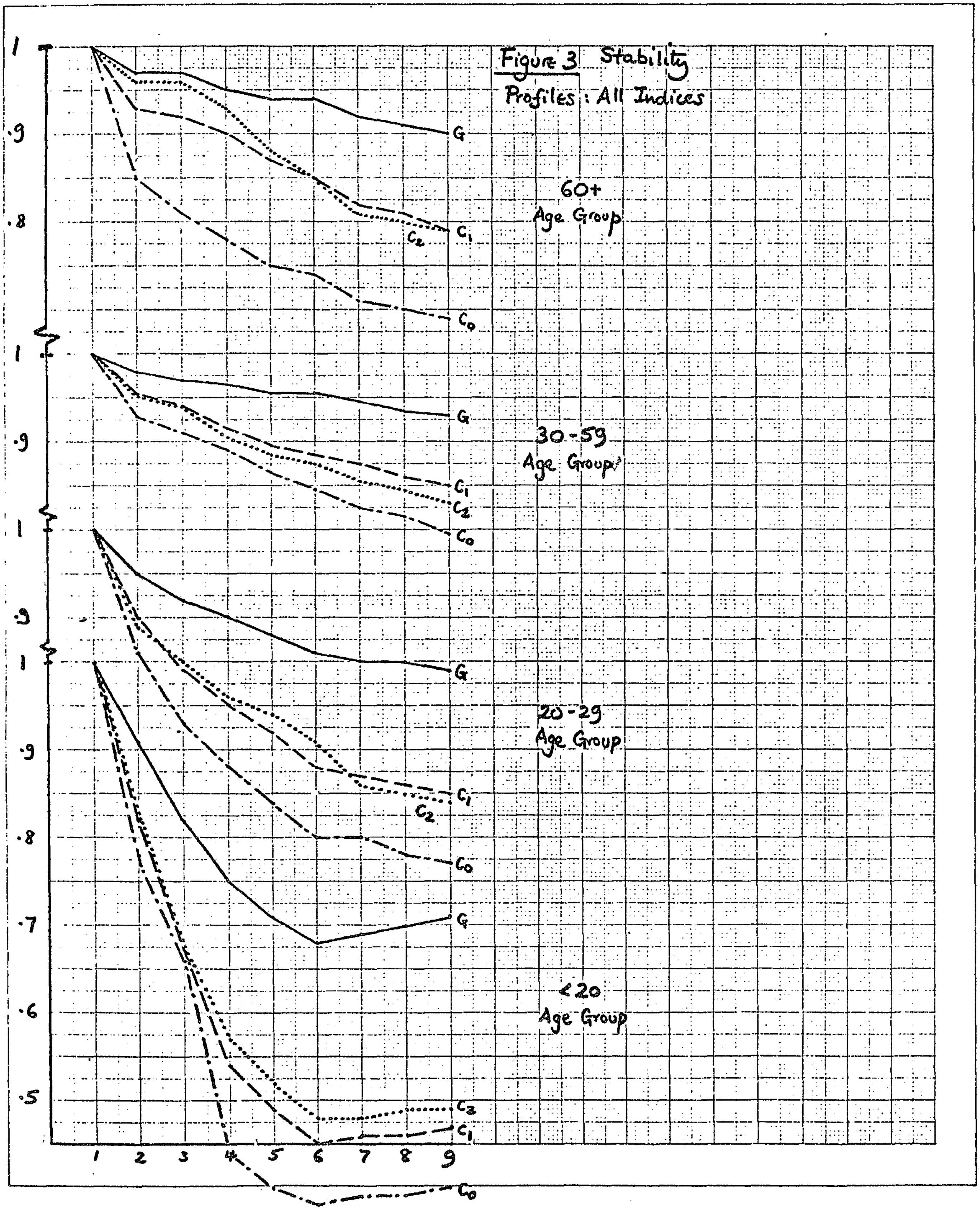


Figure 3 Stability
Profiles : All Indices
60+
Age Group
30-59
Age Group
20-29
Age Group
<20
Age Group
G
C1
C2
C0
1
.9
.8
.7
.6
.5
1 2 3 4 5 6 7 8 9

One final point concerns the type of income mobility observed. In Section 2 the rigidity curve was obtained for the coefficient of variation C_2 when the only source of income fluctuations is an additive transitory component. The curve tended to become horizontal as m increased, and this pattern is observed in Table 2 for at least the youngest age groups. How well, then, does the "hypothetical" rigidity curve given in equation (15) compare with the actual profile for C_2? The answer, it appears, is fairly well for the first 3 or 4 years. In Figure 4 the "hypothetical" curves are compared with the actual curves for the two youngest age groups, taking R_∞ to be .52 for those below age 20 and .76 for those aged 20-29. The conclusion one might draw is that short-run "transitory" changes dominate income movements over a few years (at least at younger ages) and income changes that tend to persist become apparent only when data is available for periods of perhaps 5 or more years. Furthermore, the results suggest that transitory fluctuations are considerably more important at younger ages.[7]

4. LABOR INCOME OF FEMALES

The number of females in the Michigan survey who are heads of households throughout the 9-year period is 736, and of these only 287 record positive earnings in each of the years. The sample sizes in each of the age groups are consequently much smaller than those for males, and the results are more prone to sampling variations. It was felt that the results for samples of fewer than 10 were too unreliable to be reported.

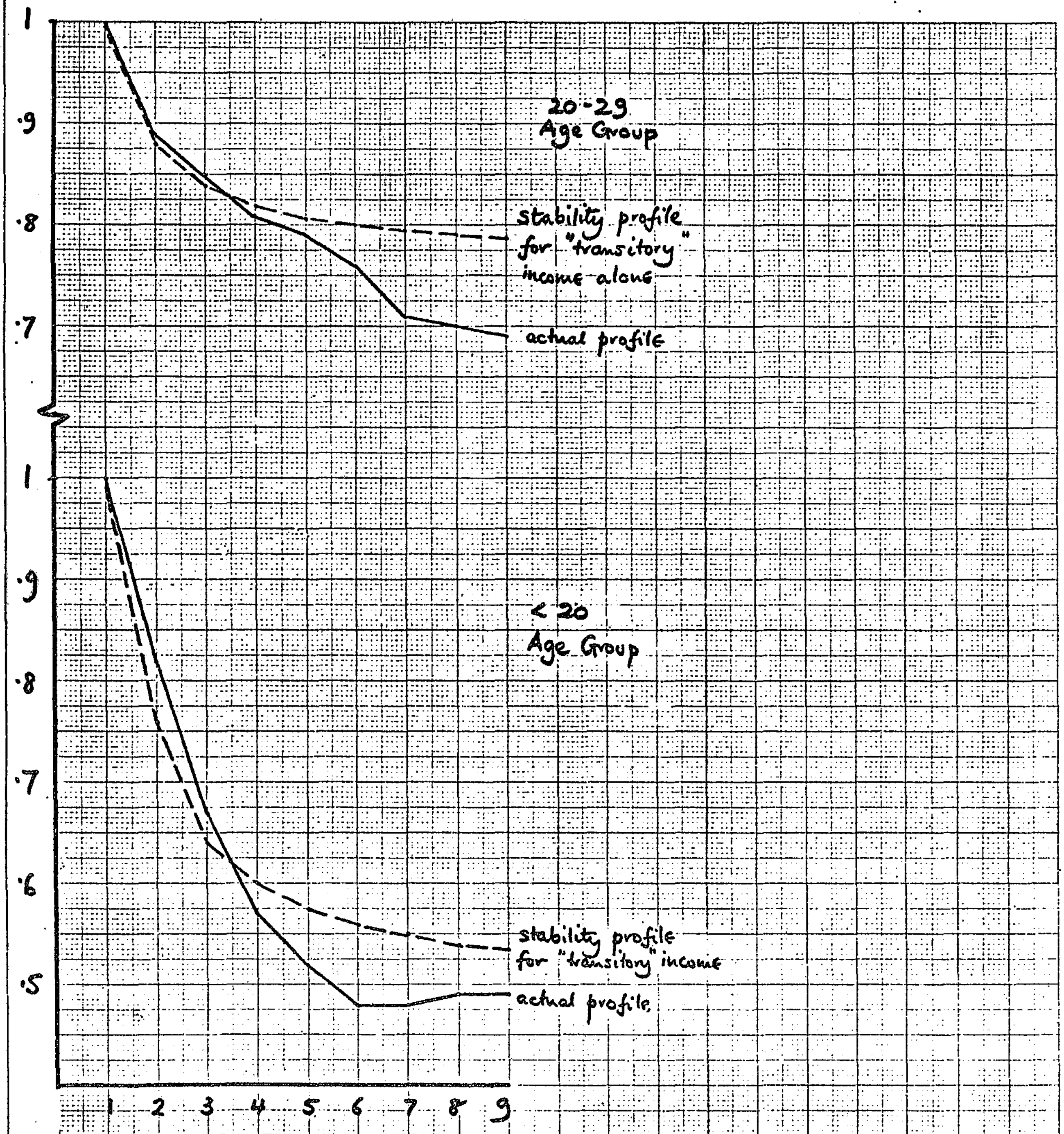


Figure 4 A Comparison of Actual Profiles and Predicted "Transitory" Income Profiles : Index C_2

Thus, in Table 3, stability profiles for the labor income of females are only given for those over 20 years of age. To keep the data presentation within reasonable limits, the rigidity profiles in this and subsequent tables are summarized by the values of R over 3-, 6-, and 9-year periods.

On the whole the results are similar to those for men. Somewhat more stability is indicated for the 20-29 age group; somewhat less stability is shown for those aged between 30 and 49. The net effect is that the increase in income stability with age, fairly pronounced in the case of male earnings, more or less vanishes in the female sample. There is some evidence of greater instability in the oldest age group, but the sample is very small.

The shapes of the profiles also resemble those of males, but the "transitory" pattern found for males aged 20-29 is now apparent in the 30-39 group as well, and even occurs in the 40-49 age group with the index C_o. Since C_o is sensitive to the lower part of the distribution, this suggests that transitory fluctuations are significant for female low earners well into middle age. This may be due to the greater prevalence of part-time and seasonal work among women of all ages.

5. TOTAL FAMILY INCOME

Table 4 presents the stability profiles for the total money income of family units, including unearned income, business profits, earnings of other family members, and transfer payments. To preserve comparability with Tables 1 and 2, these were calculated for those households whose male head

Table 3

Stability of Real Labor Incomes, Females, by Age Group

Index	Period	< 20	20-29	30-39	40-49	50-59	60 +
			Age Group in 1967				
G	1967-69	---	.94	.92	.95	.97	.94
	1967-72	---	.91	.89	.91	.96	.86
	1967-75	---	.91	.88	.90	.93	.87
C_0	1967-69	---	.74	.76	.79	.90	.69
	1967-72	---	.68	.69	.75	.86	.60
	1967-75	---	.67	.65	.70	.78	.59
C_1	1967-69	---	.85	.82	.90	.93	.82
	1967-72	---	.80	.78	.84	.91	.73
	1967-75	---	.79	.75	.81	.86	.74
C_2	1967-69	---	.87	.81	.93	.94	.79
	1967-72	---	.83	.79	.87	.92	.75
	1967-75	---	.82	.78	.84	.85	.78
Sample Size		4	42	75	88	66	12

Table 4

Stability of Total Family Income for Households Headed by Males

		Age Group in 1967					
Index	Period	< 20	20-29	30-39	40-49	50-59	60 +
G	1967-69	.80	.91	.96	.96	.98	.98
	1967-72	.74	.84	.93	.93	.96	.97
	1967-75	.73	.83	.91	.91	.94	.95
C_0	1967-69	.63	.76	.90	.89	.93	.94
	1967-72	.53	.67	.82	.83	.88	.90
	1967-75	.48	.62	.80	.80	.85	.86
C_1	1967-69	.64	.82	.91	.92	.95	.96
	1967-72	.54	.71	.84	.86	.90	.91
	1967-75	.51	.68	.81	.83	.87	.88
C_2	1967-69	.63	.83	.89	.92	.93	.97
	1967-72	.51	.71	.80	.86	.87	.87
	1967-75	.51	.67	.76	.82	.83	.84
Sample Size		33	432	526	443	226	33

did not change in the 9-year period and who had positive earnings in each of the years.

Total family income was expected to show greater stability than individual earnings, due to transfer payments and the pooling of family earnings. It is therefore surprising to discover that family income is not significantly more stable than male earnings.[8] If anything, family income shows more instability than the labor income of the head alone.

An obvious explanation is that the number of family members working may vary substantially over time, and the corresponding variation in total earnings may well dominate the impact of pooling incomes and transfer payments. Computations were consequently made for those families with the same number of members throughout the period. This turned out to reduce the sample size dramatically--the family size remains constant for only 23% of families with male heads aged 20-59. Table 5 gives the rigidity computations for the various constant family size categories. To reduce further the volume of reported data, figures are provided only for the Theil index C_1. Values of R for the indices C_o, C_1 should be roughly of the same order of magnitude; and a good predictor of the Gini R value is half the difference between 1 and the R value for C_1. To facilitate comparison with the total family population, the last three rows reproduce the corresponding figures for the index C_1 from Table 4.

The relatively small sample sizes make it difficult to state any conclusion with confidence; but the overall impression obtained is that controlling for family size does not affect the computed values of the stability index in any systematic way. Comparing the figures in the top

Table 5

Stability of Total Family Income for Households Headed by Males
Family Size Constant 1968-76: Index C_1

Family Size 1968-76	Period	Age Group in 1967 20-29	30-39	40-49	50-59
2	1967-69	---	.89	.87	.95
	1967-72	---	.86	.82	.92
	1967-75	---	.88	.81	.89
		(9)	(13)	(32)	(70)
3	1967-69	.81	.92	---	---
	1967-72	.60	.84	---	---
	1967-75	.61	.84	---	---
		(20)	(14)	(9)	(5)
4	1967-69	.92	.95	.96	---
	1967-72	.70	.92	.78	---
	1967-75	.64	.86	.73	---
		(36)	(36)	(15)	(3)
5	1967-69	.92	.91	.93	---
	1967-72	.79	.87	.92	---
	1967-75	.82	.85	.89	---
		(17)	(32)	(11)	(5)
6	1967-69	---	.83	---	---
	1967-72	---	.80	---	---
	1967-75	---	.74	---	---
		(7)	(20)	(4)	(0)
7	1967-69	---	.85	---	---
	1967-72	---	.81	---	---
	1967-75	---	.79	---	---
		(0)	(12)	(0)	(0)
All Families	1967-69	.81	.91	.92	.95
	1967-72	.71	.84	.86	.90
	1967-75	.68	.81	.83	.87
		(432)	(526)	(443)	(226)

Note: Figures in brackets indicate sample size.

part of the table with their corresponding values in the last three rows, the index is found to increase in 19 cases and decrease in 16 cases. If attention is restricted to significant changes in the index value (taken here to be a difference of .05 or more), then there are 9 examples of a significant increase in stability when family size remains constant and 8 examples of a significant reduction. Only for 5-person families does there appear to be a consistent increase in income stability when family size is held constant; and for 6- and 7-person families the evidence points in the opposite direction.

Of course, considering only those families of constant size still does not rule out the possibility of wives being released to work when the children become of school age, and of children entering the work force, both of which factors will contribute to fluctuations in family income. But this argument can hardly be applied to 2-person families where the evidence is equally inconclusive.

The overall conclusion, therefore, is that family income exhibits no more stability than does male earnings, and there is no obvious explanation of this unexpected result. This is clearly an important topic for further research, since the evidence here seems to question the basic rationale of policies directed at income stabilization at the household level.

6. MALE EARNINGS BY OCCUPATION

We now return to the labor incomes of males and disaggregate the population on the basis of their 1-digit occupation code recorded for the first survey year, 1968. Attention is again confined to the Theil index C_1, and the results summarized in Table 6. The last column gives the figures for the combined age groups 20-59. Here the sample sizes are sufficiently large for comparisons to be made with some degree of confidence.

No one occupation dominates the rest in the sense that earnings are more stable over the 3-, 6-, and 9-year periods. Nor is one occupation dominated by all the others. But the partial ordering given by this dominance relation permits a fairly complete ordering to be made, and this ordering by income stability corresponds almost exactly to the ranking by occupation code numbers. It is represented diagramatically in Figure 5. The managerial and professional groups are found at the top, followed by the nonmanual intermediate occupations and the skilled manual category. These all exhibit more stability than operatives, who in turn have more stable earnings than laborers. Farmers comprise the group exhibiting the most income instability, particularly over the 6- and 9-year periods.

There is one outlier to this neat and intuitively plausible classification scheme--the self-employed. This group has a distinctly different rigidity profile, with the lowest degree of stability of all groups over 3 years and one of the highest stability values over 9 years. The rigidity profile therefore shows considerably more levelling off than the other groups, as the aggregate time period is extended. Although this is not entirely unexpected, it is comforting to discover that the computed values of the index R allow the unique characteristics

Table 6

Stability of Real Labor Incomes, Males, by 1968 Occupation: Index C_1

Occupation	Period	Age Group in 1967 20-29	30-39	40-49	50-59	20-59
Professional, Technical	1967-69	.79	.87	.95	.90	.90
	1967-72	.73	.80	.91	.85	.85
	1967-75	.70	.77	.86	.82	.80
		(63)	(87)	(63)	(27)	(240)
Managers	1967-69	.90	.94	.95	.93	.93
	1967-72	.80	.78	.93	.89	.85
	1967-75	.74	.70	.89	.81	.79
		(30)	(49)	(47)	(18)	(144)
Self Employed	1967-69	.84	.81	.73	.85	.81
	1967-72	.81	.81	.73	.83	.80
	1967-75	.76	.80	.76	.75	.76
		(11)	(21)	(22)	(18)	(72)
Clerical, Sales	1967-69	.86	.82	.96	.92	.91
	1967-72	.69	.74	.89	.86	.81
	1967-75	.59	.70	.88	.71	.73
		(56)	(44)	(42)	(18)	(160)
Craftsmen, Foremen	1967-69	.79	.95	.89	.91	.91
	1967-72	.53	.86	.82	.79	.80
	1967-75	.51	.82	.76	.71	.75
		(79)	(121)	(97)	(47)	(344)
Operatives	1967-69	.80	.88	.93	.89	.88
	1967-72	.68	.81	.81	.80	.77
	1967-75	.65	.74	.77	.69	.72
		(109)	(111)	(84)	(32)	(336)
Laborers	1967-69	.81	.83	.84	.89	.84
	1967-72	.62	.78	.81	.81	.76
	1967-75	.58	.75	.77	.72	.72
		(54)	(62)	(66)	(48)	(230)
Farmers	1967-69	.63	.74	.90	.91	.82
	1967-72	.28	.61	.70	.74	.63
	1967-75	.37	.59	.67	.68	.61
		(11)	(14)	(15)	(18)	(58)

Note: Figures in brackets indicate sample size.

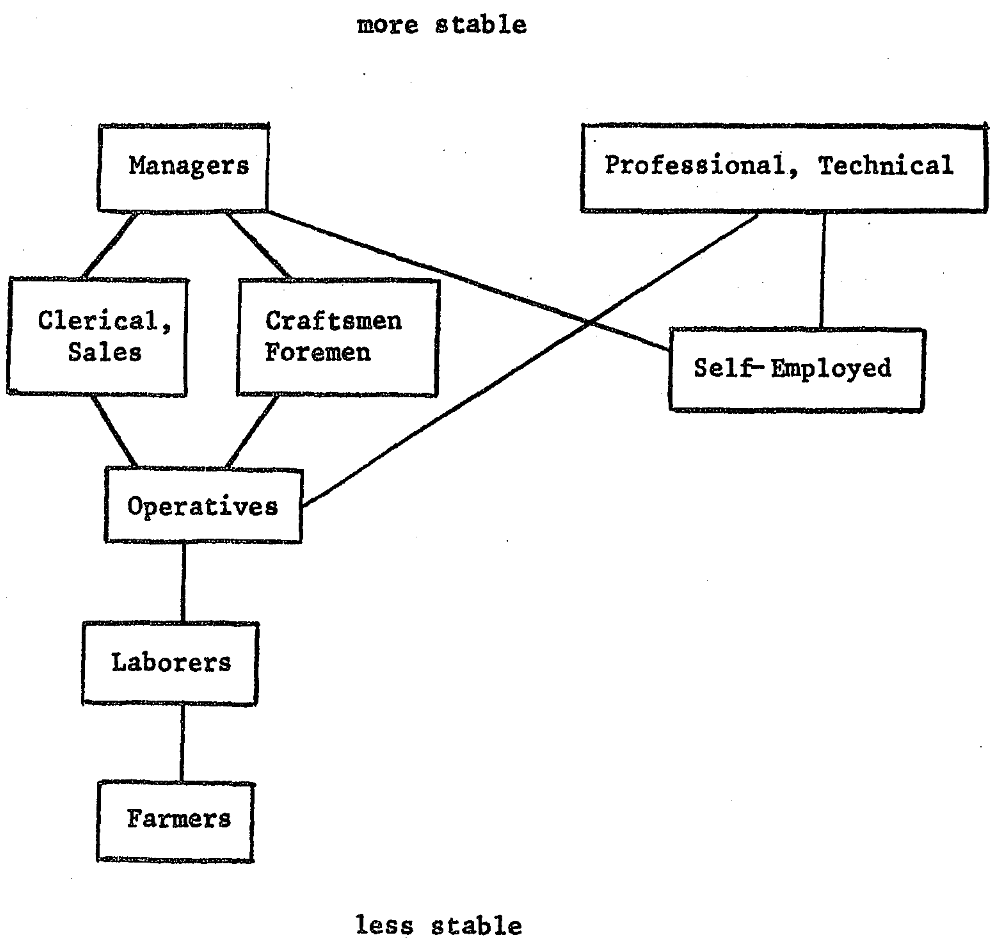


Figure 5. Partial ordering of occupations based on the classification of individuals in 1968.

of this group to be identified.

Within each occupational category the variation in income stability with age generally agrees with that found earlier for all occupations combined--a tendency for stability to initially increase with age and then decline as people approach retirement. The exception to this general rule is again the self-employed, for whom, on the basis of rather small samples, there appears to be little or no systematic age effect.

One of the interesting aspects of income dynamics concerns the extent to which income variations accompany changes in occupations and jobs within an occupation. This is too large a topic to be considered here in any detail, but some indication of the impact of occupational movements on income variability can be obtained by eliminating from the sample those who change occupation. Table 7 presents the results of computations made for those males who classified themselves in the same 1-digit occupation group in each of the 9 survey years. The numbers dropped from the sample are themselves quite revealing. The evidence indicates a very high degree of occupation change, all the more remarkable since changes in occupation code within the 1-digit categories have not been distinguished. Farmers and those in the professional group have the highest probability of remaining in their category for the whole 9 years, but even here the proportion of "stayers" is barely one-half. Some 25-35% of managerial, clerical-sales, craftsmen-foremen, and self-employed persons remain in the same group. The corresponding figure for operatives is a little under 25%, and that for laborers is just below 20%. Thus the ranking by the proportion of stayers in an occupation also corresponds closely to the 1-digit code ranking, with only farmers deviating from this rule. This

Table 7

Stability of Real Labor Incomes, Males, Same Occupation Group 1968-76: Index C_1

		Age Group in 1967				
Occupation	Period	20-29	30-39	40-49	50-59	20-59
Professional, Technical	1967-69	.80	.85	.96	.96	.91
	1967-72	.76	.77	.94	.89	.86
	1967-75	.69	.75	.87	.86	.79
		(34)	(52)	(34)	(14)	(134)
Managers	1967-69	.92	.92	.97	---	.94
	1967-72	.91	.92	.95	---	.93
	1967-75	.90	.81	.94	---	.88
		(11)	(19)	(18)	(3)	(51)
Self Employed	1967-69	---	---	---	---	.73
	1967-72	---	---	---	---	.66
	1967-75	---	---	---	---	.59
		(4)	(5)	(6)	(7)	(22)
Clerical, Sales	1967-69	.91	.95	.96	---	.94
	1967-72	.70	.86	.94	---	.89
	1967-75	.61	.80	.92	---	.85
		(13)	(11)	(17)	(6)	(47)
Craftsmen, Foremen	1967-69	.76	.89	.92	---	.89
	1967-72	.63	.84	.87	---	.83
	1967-75	.63	.78	.83	---	.78
		(20)	(41)	(35)	(9)	(105)
Operatives	1967-69	.79	.92	.91	.93	.90
	1967-72	.77	.86	.70	.91	.80
	1967-75	.79	.81	.66	.88	.77
		(21)	(21)	(26)	(11)	(79)
Laborers	1967-69	---	.89	.82	.89	.88
	1967-72	---	.77	.70	.89	.81
	1967-75	---	.77	.68	.88	.80
		(1)	(13)	(14)	(16)	(44)
Farmers	1967-69	---	.72	---	---	.70
	1967-72	---	.65	---	---	.56
	1967-75	---	.59	---	---	.55
		(5)	(10)	(6)	(8)	(29)

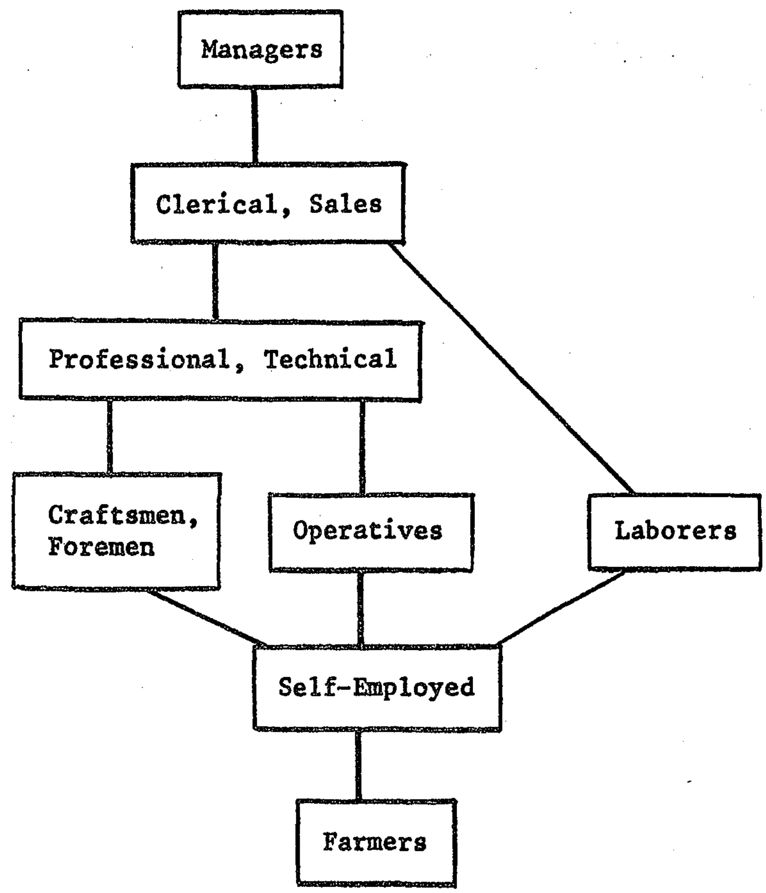
Note: Figures in brackets indicate sample size.

raises the question of whether the ranking of occupations by income stability based on Table 6 is largely accounted for by the fact that those in different occupations in 1968 have different propensities to change occupation. The variation in income stability by occupation previously described might then be more appropriately assigned to the individuals initially in these job categories, rather than to the occupations themselves.

The figures reported in Table 7 are a better reflection of the intrinsic stability associated with the occupational categories. A comparison of the last column with that of Table 6 reveals that the stability index increases over all three periods for 6 of the 8 groups. Exactly the reverse is true of the two remaining categories--the self-employed and farmers--who are now seen to experience significantly less stability than the other groups. Within the set of more income-stable occupations, the increase in stability is not uniform. Managers, clerical-sales, and laborers show somewhat larger improvements, which changes the partial ordering given in Figure 5 to that indicated in Figure 6.[9]

Those who remain in an occupation also tend to have flatter stability profiles than those for all individuals initially recorded in the occupation, although this is not the case for the professional group (for whom the stability values are virtually unchanged between Tables 6 and 7) or the self-employed. In Figure 7, the pairs of profiles are drawn for four of the occupations.

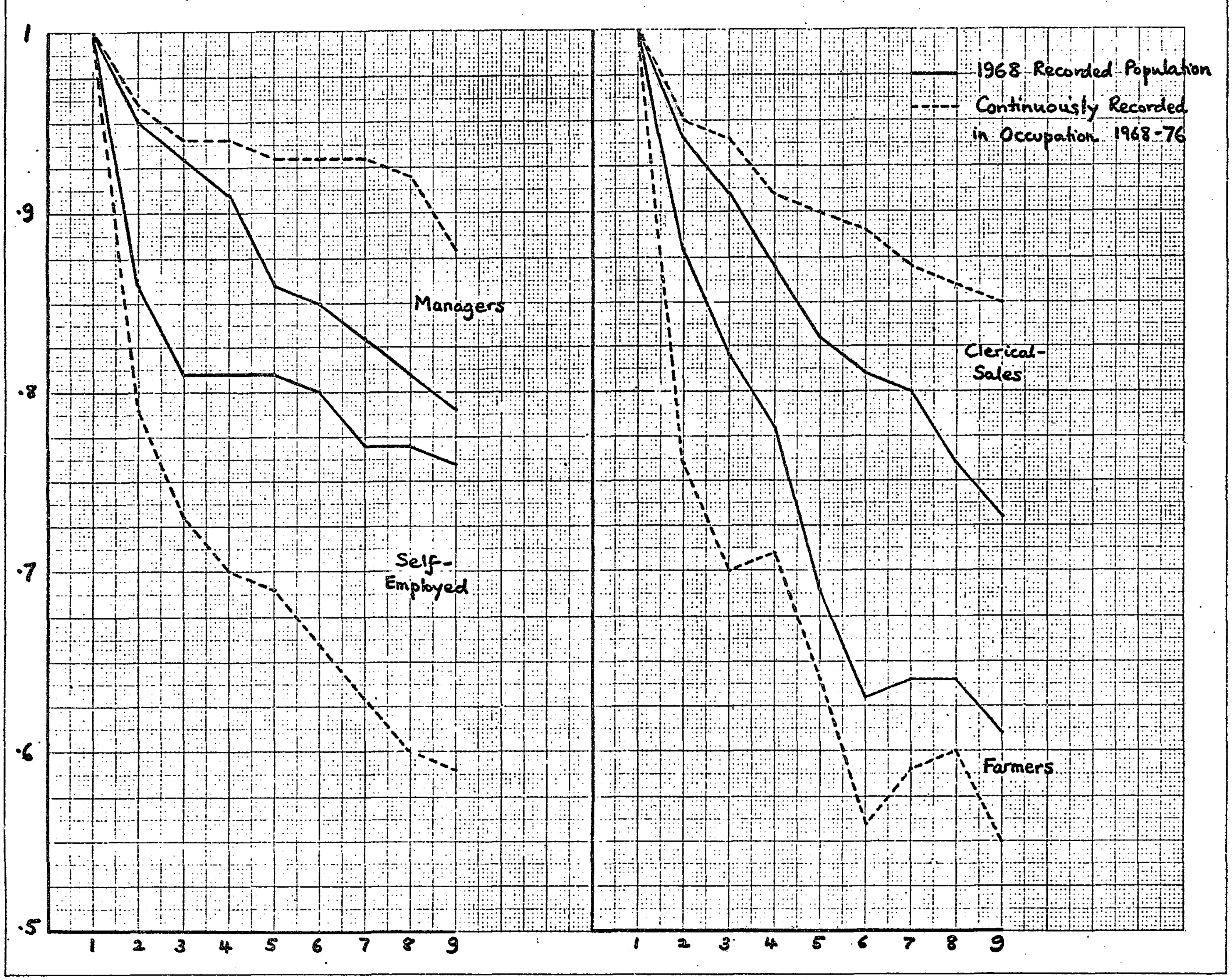
more stable



less stable

Figure 6 Partial ordering of occupations based on those who remain in the same occupation, 1968-76.

Figure 7 Comparison of Stability Profiles for Selected Occupational Groups.



7. MALE EARNINGS BY LEVEL OF EDUCATION

A final set of computations was made for the labor incomes of males disaggregated according to their educational background. The results are summarized in Table 8. The most noticeable feature is the relative similarity of the stability values across education groups. In the last column, for instance, the figures indicate that the degree of equalization occurring as a result of aggregating over 9 years varies between 21% and 29%. This is considerably less than the corresponding range found in the occupational breakdown of Table 6 (20-39%) or Table 7 (12-45%). To the extent that stability does differ across the groups, it generally increases with the level of education, the main exception being the high school + non-academic training category. This group and the group holding advanced degrees experience the most stable incomes, followed by those with bachelor degrees or some college education; those with no more than high school education are in the third and lowest tier. Since the differences are not very pronounced, and may well be attributable to different choices of occupation, it is reasonable to conclude that educational background does not have a major impact on income stability.

8. CONCLUDING REMARKS

One of the main objectives of this paper has been to show that the computation of rigidity values R, and the construction of stability profiles, provide a useful way of summarizing data on income dynamics. The rigidity values have an intrinsic interest for those concerned with the extent to which short-run inequality values exaggerate the degree of inequality found

Table 8

Stability of Real Labor Incomes, Males, by Education Level: Index C_1

		Age Group in 1967				
Education Level	Period	20-29	30-39	40-49	50-59	20-59
Grade 6-8	1967-69	.81	.87	.89	.91	.89
	1967-72	.72	.77	.81	.82	.80
	1967-75	.68	.72	.75	.65	.71
		(34)	(88)	(83)	(59)	(264)
Grade 9-11	1967-69	.80	.86	.90	.91	.87
	1967-72	.66	.75	.82	.73	.76
	1967-75	.61	.71	.79	.61	.71
		(92)	(99)	(84)	(27)	(302)
Completed High School	1967-69	.77	.87	.87	.81	.86
	1967-72	.64	.81	.80	.75	.77
	1967-75	.61	.75	.79	.75	.72
		(107)	(104)	(60)	(34)	(305)
High School + Non-academic training	1967-69	.85	.95	.91	.89	.93
	1967-72	.66	.87	.82	.84	.84
	1967-75	.57	.85	.79	.70	.79
		(37)	(57)	(41)	(15)	(150)
College--No Degree	1967-69	.82	.82	.94	.92	.88
	1967-72	.68	.80	.91	.90	.82
	1967-75	.63	.72	.87	.84	.76
		(90)	(58)	(56)	(19)	(223)
Bachelor Degree	1967-69	.88	.92	.87	.80	.90
	1967-72	.78	.82	.81	.74	.81
	1967-75	.73	.79	.78	.72	.76
		(46)	(51)	(40)	(15)	(152)
Advanced Degree	1967-69	.82	.89	.97	.95	.94
	1967-72	.69	.63	.92	.90	.85
	1967-75	.67	.63	.87	.83	.77
		(19)	(35)	(25)	(14)	(93)

Note: Figures in brackets indicate sample size.

over a longer accounting period. Stability profiles form the basis for comparisons across a variety of population subgroups and permit some of the important features of income dynamics to be captured diagramatically.

In Section 3 it was shown that the choice of inequality index has a considerable impact on the computed values of R, and that use of the Gini coefficient will tend to indicate much higher stability than alternative indices. A significant age effect was found for male earnings stability which, along with the characteristic shape of the stability profiles, suggests that short-run, transitory income fluctuations dominate income movements at lower ages. Although this may come as no surprise, the actual degree of income instability within the younger age groups is perhaps considerably higher than expected; and the fact that transitory fluctuations have an important age component has obvious implications for the modeling of income dynamics which have not always been given sufficient attention.[10]

The stability of female earnings (for those women who are heads of households) seems to be broadly comparable to that of males, although the variation by age group is less pronounced. Total family income also appears to exhibit roughly the same degree of stability as male earnings, despite the fact that the former includes transfer payments and the incomes of other family members, which one would expect to have a stabilizing influence. The evidence suggests that the explanation of this surprising result is not to be found in variations in the number of persons in the family unit.

A breakdown of male earnings by occupation was considered in Section 6; it was found that income stability varied substantially across occupations. Furthermore, the ranking of occupations is altered when only those individuals who remain in the same occupational category are considered. Within this population of "stayers," Farmers and the self-employed experience significantly more income instability than the other occupations. Finally, disaggregating by the level of educational attainment showed little difference between the categories. This suggests that the level of education is unlikely to be a major determinant of income stability, apart from its indirect effect via the choice of occupation.

APPENDIX

Notes on the Data Used in This Study

The data base for the empirical results is the Michigan Survey of Income Dynamics, and the surveys considered are those for the 9-year period 1968-1976. It has been assumed that reported incomes correspond to the calendar year prior to the survey, and that the age of the individuals in 1967 was one year less than their age given in 1968.

The 9-year tape contains records on 5,862 families, but a large proportion of these are splits from the original panel. I considered only the incomes of those households whose heads did not change during the 9 years, a total of 2,985 families. Some of these (28) were eliminated because the sex of the head in 1976 differed from that given in 1968 (although no change in the head was registered). Another 69 were disregarded because the age of the head in 1976 was inconsistent with their age in 1968, even allowing for a little forgetfulness (families were eliminated if the difference between the age of the head in 1976 and his or her age in 1968 was less than 5 or greater than 11). This left 2,888 families, of which 2,152 had male heads and 736 female heads.

For the computations reported in this study, I also dropped those families whose heads did not have positive earnings in each of the 9 years. This procedure was followed primarily because many inequality indices, including C_0 and C_1, are not defined for zero values. A few tests were made to see whether this was likely to introduce a bias in the results. Table A1 gives the breakdown of the 2,888 families by age and sex of the head, together with the number of those with positive labor income in the first 3, 6 and 9 years. The table shows that about 20% of the male sample is dropped, and the majority of these came from the last two age groups. However, 60% of the females are dropped, so the figures reported for women in Section 4 may be considerably less reliable.

Some indication of the impact of reducing the sample in this way can be gauged from Table A2. Males with positive earnings in the first 3 years 1967-69 total 1,949, more than 90% of the maximum possible sample of 2,152. Stability values for this subsample can be compared with the corresponding values for the smaller sample of those with 6 years of positive earnings, and the even smaller sample of those with 9 years of positive earnings. If this reduction of the sample did introduce a bias in the results, then we might expect the values of R computed from successively smaller samples to vary in some systematic way. Table A2 shows that the corresponding values are virtually identical, so there is no evidence of any significant bias. Only in the oldest age group, where the sample size drops considerably, do the computed values of R differ by more than .01 between the three sets of figures.

Table A1

Analysis of Sample Losses due to the Restriction of Positive Labor Incomes.

		Postive Labor Income in each of the years		
Males	Total Population	1967-69	1967-72	1967-75
< 20	37	35	33	33
20-29	457	446	440	432
30-39	559	546	539	526
40-49	513	488	473	443
50-59	367	325	304	226
60 +	219	109	65	33
All ages	2,152	1,949	1,854	1,693
Females				
< 20	7	6	4	4
20-29	88	50	46	42
30-39	169	97	85	75
40-49	170	113	97	88
50-59	176	108	92	66
60 +	126	39	21	12
All ages	736	413	345	287

Table A2

Analysis of Impact of Sample Losses, Males, Real Labor Income:

Index C_1

Age Group in 1967	Positive Earnings Recorded for	Stability Index R for Period					Sample Size
		1967-8	1967-9	1967-70	1967-71	1967-72	
20-29	1967-69	.90	.84				446
	1967-72	.90	.84	.80	.77	.73	440
	1967-75	.90	.84	.80	.77	.73	432
30-39	1967-69	.94	.92				546
	1967-72	.95	.92	.90	.88	.87	539
	1967-75	.95	.93	.90	.88	.87	526
40-49	1967-69	.96	.94				488
	1967-72	.96	.94	.93	.91	.90	473
	1967-75	.96	.95	.93	.91	.90	443
50-59	1967-69	.95	.93				325
	1967-72	.95	.94	.92	.91	.89	304
	1967-75	.95	.94	.92	.90	.89	226
60 +	1967-69	.95	.92				109
	1967-72	.94	.92	.89	.86	.83	65
	1967-75	.93	.92	.90	.87	.85	33

NOTES

* I am indebted to the Institute for Research on Poverty, University of Wisconsin for financial support and computing assistance which enabled me to carry out this study.

1. Hanna (1948, pp.211-213) suggests that the impact of the accounting period can be assessed by comparing the Lorenz curve for the overall accounting period with the "average" Lorenz curve, which is constructed by summing the lowest income in each subperiod, the two lowest incomes in each subperiod, and so on. This has some similarities with the procedure suggested here. For if the Gini coefficient is taken as the index of inequality, the denominator in (6) is the same as the Gini coefficient for Hanna's "average" Lorenz curve. The value of 1-R computed for the Gini coefficient would therefore represent the area between the "average" Lorenz curve and the Lorenz curve for the overall accounting period, expressed as a proportion of the area between the "average" Lorenz curve and the line of complete equality. As Hanna points out, the maximum reranking of individuals occurs when all incomes are equalized over the longer accounting period. This gives the minimum value of zero for the rigidity index R, and the maximum mobility value $M = 1-R = 1$.

2. Benus and Morgan (1975, pp.216-221) make separate computations for the accounting period variation and income instability. They conclude that "the tendency is for groups with high instability levels to exhibit a greater impact of accounting period variation. The relationship, however, is not strong." This may appear to provide grounds for rejecting the correspondence between mobility and the degree of equalization occurring as the accounting period is lengthened. However, they are not consistent in their use of an inequality index. The impact of the accounting period is defined in terms of the Gini coefficient; income instability is indicated by the average coefficient of variation. Furthermore, their definition of income instability includes any income changes, and this leads to a very restrictive concept of complete income stability (it can only occur when mean income is constant over time).

3. The problem of separating out the impact of the time period between observations is discussed in some detail in Shorrocks (1978a).

4. The term "income mobility" has been used generically in the literature to refer to any kind of changes in income over time. From the evidence presented in the following sections, it appears that a good deal of income mobility is accounted for by short-run fluctuations. In these circumstances it seems more appropriate to use the term "income instability." However, the distinction is not great, and for the most part the two terms are taken to be synonymous.

5. The data base for this study is the 9-year merged tape of the Michigan Survey of Income Dynamics. The same series of surveys has been used by Benus and Morgan (1975) and Lillard and Willis (1978), amongst others. Background information on the precise sample used here is provided in the appendix.

6. Benus and Morgan (1975, p.216) state that "lengthening the accounting period beyond three periods has almost no impact on the distribution of income." This suggests that the profiles become more or less horizontal after 3 years, which conflicts with the evidence in Table 1 and Figure 2. However, they compare the m-year Gini with the Gini for the first year, 1967, rather than the weighted average annual value. For the reasons already discussed, this gives a spuriously low value for the degree of equalization resulting from the longer accounting period. This is recognized on p.215: "...the use of the initial year as a benchmark probably underestimates the impact of lengthening the accounting period."

7. Schiller (1977, pp. 933-34) discusses whether mobility is transitory in nature and comes to the opposite conclusion: "...most of the mobility observed here is... 'permanent'." This conclusion is based on the fact that income changes in the period 1957-62 are only <u>completely</u> reversed in the period 1962-67 for one income group, the bottom 5% of income recipients in 1962. However, this is a very stringent condition to impose, and seems to assume that pure transitory fluctuations would lead to a correlation of -1 between income changes in consecutive periods.

8. It should be emphasised that the discussion here refers to the stability of earnings compared to family income. The change to the latter income concept reduces measured <u>inequality</u> quite substantially.

9. If occupations are partially ordered on the basis of the stability values for the eight periods 1967-68 to 1967-75, as opposed to just the three periods of Tables 6 and 7, there is almost no change in Figures 5 and 6. Figure 5 remains the same, and in Figure 6 the only change is that the self-employed are no longer unambiguously more stable than farmers. This provides some vindication for condensing the reported information into the stability values for just three periods.

10. See, for example, Lillard and Willis (1978). By the standards of the current literature their model is fairly sophisticated. Yet they assume that the pure random component of income has constant variance across all ages.

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