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Secular Trends and Cyclical Behavior of Income Distribution in the United States: 1944–1965

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THE OBJECTIVE of this study is to estimate secular and cyclical change in the size distribution of personal income in the United States and investigate its causes. These changes are interpreted first in the aggregate, measuring secular trends and cyclical behavior, and then in the component groups that constitute fundamentally different types of income units as distinguished by sex and age. Both the aggregate and disaggregated approaches are pursued here utilizing income data for persons 14 years and over, derived from postwar Current Population Surveys.

The following section discusses the income data and the techniques used for their analysis. In section II annual estimates of income inequality are interpreted in the context of an aggregate model that incorporates the short-run factors that displace the distribution of income from the long-run equilibrium trend. Results of time series analysis for the United States and the Netherlands are contrasted. Section III deals with component groups in the population that compete in largely different markets for their current income. Analysis of variance is employed to evaluate how the role of these groupings has changed in accounting for over-all income inequality.

I. Income Data and Measures of Income Inequality

To compare income distributions of two or more periods, the income data must be conceptually alike, and be representative of comparable groups in the population. *Current Population Surveys* (CPS) of the U. S. Bureau of Census provide, since 1944, a good foundation for studying year-to-year changes in the distribution of personal income in the United States. While the surveys periodically have adopted new features into their design and have grown in size, such changes do not seriously impair the over-all comparability of the series of annual income estimates.[1] Sampling variability, however, becomes an unavoidable problem when this data source is used for estimating the distribution of income among relatively small groups in the population, such as that among certain sex and age cohorts. From these surveys the percentage of families, unrelated individuals, and all persons with income are estimated and distributed according to as many as seventeen money income classes. Only the percentage distributions (to one decimal place) are published and used in this study. To test some of the hypotheses set forth in this study, data are needed measuring earnings from labor participation, but the current population surveys are consistently tabulated only by total money income, exclusive of capital gains, and this concept of income is adopted in later empirical analyses.

Estimating the mean and median incomes and calculating measures of income inequality from these data requires some assumption as to the distribution of income units within income classes. Rather than follow the usual convention of taking the midpoint as the average income level in each closed income interval, it is assumed that the geometric mean of the interval is the average income level in accord with the approximately log-normal distribution of the income size variate.[2] The

[1] A review of the evolution and design of the *Current Population Survey* is provided in the U. S. Bureau of Census, *The Current Population Survey—A Report on Methodology,* Technical Paper No. 7, Washington, D. C., 1963; or briefly surveyed in T. Paul Schultz, *The Distribution of Personal Income,* Joint Economic Committee, Congress of the United States, December 1964.

[2] The size distribution of the logarithms of personal income conforms approximately to a normal frequency distribution, neglecting nonpositive incomes or conveniently adding them to the first positive income interval. Since Gibrat observed this pattern many studies of income distribution data have confirmed its prevalence. Using the chi-square test of goodness of fit, it was found that the fit between the frequencies of income units in each income interval implied by the estimated log-normal function and those observed in each income interval was distinctly better if the geometric rather than the arithmetic midpoints of the income intervals were assumed to be the mean incomes. The difference between the geometric and arithmetic midpoints is greatest, of course, for the first income interval, in which the geometric mean may be unrealistically low. It is also interesting to note, that when Sheppards correction is applied to estimation of the moments of the log-normal distribution from grouped income data the goodness of fit, according to the chi-square test, deteriorates. For discussion of Sheppards correction and goodness of fit tests see M. G. Kendall and A. Stuart, *Advanced Theory of Statistics,* London, 1963 and 1961, Vol. I, p. 46, and Vol. II, pp. 419+.

average income level in the open-ended interval is estimated by assuming that the frequencies of income units in the last two intervals conform to a Pareto distribution unless the frequencies in the last interval are equal to or greater than those in the next to the last interval. Where the Pareto estimate is inappropriate by these standards, fixed average income levels are attributed in accordance with the estimates used by H. P. Miller in his compendium of these income data.[3]

To analyze and to order different sets of income size data that are distributed in a skewed but regular way, it is convenient to define a summary measure to represent the degree of skewness, inequality, or concentration. The task of devising an appropriate measure of income inequality is largely arbitrary, for no established analytical framework defines or implies the relevant concept of inequality. Two measures of income inequality, the Gini concentration coefficient and the variance of the natural logarithms of income, are employed here and are hereafter designated as the concentration and variance of income. The concentration coefficient is calculated by applying absolute weights to the differences in income between all pairs of observations standardized over the mean, whereas the variance of the logarithms of income is calculated by applying relative weights to deviations in the income from the geometric mean or estimated median.[4] Both measures of income inequality tend to move together, and if the size distribution of income is in fact log-normal, then one measure is a unique function of the other. Where the distribution of income by size is not log-normal no one-to-one correspondence exists between income inequality and a size distribution; or, in other words, two different size distributions could evidence the same degree of income inequality. Inequality within income size classes, which is neglected in these measures of inequality estimated from grouped data, may have declined somewhat over time as the number of income classes increased from thirteen to seventeen and the frequency distribution of income units across classes became more

[3] H. P. Miller, *Trends in the Income of Families and Persons in the United States, 1947–1960,* Technical Paper No. 8, U. S. Bureau of the Census, Washington, D. C., 1963, pp. 24–25. \$44,000 was used for the income interval over \$25,000; \$24,000 for the interval over \$15,000, and \$20,000 was used for the interval over \$10,000.

[4] For a discussion of both measures of inequality refer to Kendall and Stuart, Vol. I, p. 47; and J. Aitchison and J. A. C. Browth, *The Lognormal Distribution,* Cambridge, England, 1957, Chapter II and Table A-1.

uniform.[5] Since the concentration coefficient is less sensitive than the variance of the logarithms of income to these changes in data grouping, the concentration coefficient is more appropriate for year-to-year analysis of time series.

II. Aggregate Change in the Distribution of Income

TIME SERIES AND SECULAR CHANGE

Classic studies by Kuznets show that in the more developed countries the size distribution of income among persons and among families has become less unequal during the twentieth century.[6] The composition of family units has changed markedly in the postwar United States, complicating the task of analyzing change in distribution of earnings, consumption, and welfare among families. For this reason, the focus here is on the distribution of income among persons with income, 14 years of age or older. As a working hypothesis, it is assumed initially that the universe of individuals with income in our sample is not changing. Thus change in the measure of income inequality among these persons does not reflect a change in the composition of income recipients but an actual change in the distribution of income among persons. The observed change in personal income inequality is interpreted as a secular trend in the functional equilibrium distribution of income as perhaps modified by cyclical factors.

Estimates of the concentration and variance of income among families and unrelated individuals, and all persons for the postwar years are presented in Tables 1 and 2, derived according to the procedure outlined in the preceding section. Only in one instance is there a statistically significant tendency for these estimated measures of income inequality to change linearly in time from 1945 to 1965; and in that case, the concentration of income among persons appears to have increased at

[5] This problem of estimating income inequality with income classes is discussed in greater detail in T. Paul Schultz, "The Distribution of Income: Case Study the Netherlands," unpublished Ph.D. dissertation, Massachusetts Institute of Technology, 1965, Chapter V.

[6] "Economic Growth and Income Inequality," *American Economic Review,* Vol. 45 (March 1955); and "Quantitative Aspects of the Economic Growth of Nations: Part VII, Distribution of Income by Size," *Economic Development and Cultural Change,* Vol. 11, No. 2, Part 2 (January 1963).

TABLE 1

INCOME INEQUALITY AMONG FAMILIES AND UNRELATED INDIVIDUALS IN THE UNITED STATES: 1944–65

	Families		Families and Unrelated Individuals	
Year	Variance of the Logarithms of Income	Concentration of Incomes	Variance of the Logarithms of Income	Concentration of Incomes
1944	.5598	.4102	.6582	.4521
1945	.4519	.3773	.5302	.4169
1946 [a]	–	–	–	–
1947	.4774	.3827	.5642	.4245
1948	.4551	.3773	.5249	.4134
1949	.4652	.3852	.5445	.4247
1950	.4639	.3831	.5445	.4242
1951	.4733	.3681	.5636	.4099
1952	.5586	.3726	.7155	.4186
1953	.4588	.3648	.5611	.4126
1954	.5272	.3803	.5759	.4244
1955	.5793	.3752	.6190	.4510
1956	.4697	.3635	.5871	.4100
1957	.4597	.3588	.5224	.4002
1958	.4372	.3598	.5063	.4019
1959	.4552	.3646	.5392	.4102
1960	.4781	.3719	.5642	.4148
1961	.4938	.3805	.5906	.4265
1962	.4302	.3642	.5170	.4102
1963	.4264	.3651	.5202	.4146
1964	.4120	.3607	.5130	.4141
1965	.4316	.3658	.5156	.4136

SOURCE: *Current Population Survey, Consumer Income,* Series P-60.
[a] Not available: No *Current Population Survey* income data for 1946.

TABLE 2

ESTIMATED AVERAGE INCOME AND INCOME INEQUALITY FOR PERSONS 14 YEARS OF AGE AND OVER WITH INCOME, BY SEX: 1944–65

	Both Sexes			Males			Females		
Year	Average Income in Current Dollars [a]	Concentration [b]	Variance of Logs [c]	Average Income in Current Dollars [a]	Concentration [b]	Variance of Logs [c]	Average Income in Current Dollars [a]	Concentration [b]	Variance of Logs [c]
1944	1,819	.505	.939	2,330	.450	.795	1,069	.525	1.029
1945	1,746	.487	.909	2,111	.441	.761	984	.496	.778
1946 [b]	—	—	—	—	—	—	—	—	—
1947	2,161	.483	.866	2,610	.433	.725	1,188	.517	.936
1948	2,212	.474	.770	2,731	.424	.672	1,203	.511	.890
1949	2,142	.489	.822	2,623	.439	.711	1,169	.522	.767
1950	2,348	.498	.851	2,939	.441	.720	1,220	.535	.862
1951	2,577	.480	.864	3,263	.418	.766	1,335	.521	.792

1952	2,671	.470	.834	3,367	.405	.713	1,410	.511	.666
1953	2,776	.484	.831	3,533	.422	.750	1,516	.528	1.029
1954	2,796	.489	.837	3,545	.429	.711	1,508	.519	.760
1955	2,952	.509	1.102	3,774	.436	.844	1,540	.542	1.030
1956	3,076	.506	.942	3,988	.429	.702	1,553	.532	.727
1957	3,144	.506	.926	4,065	.432	.726	1,636	.534	.774
1958	3,149	.503	.759	4,175	.433	.650	1,664	.543	.790
1959	3,443	.515	.936	4,451	.435	.662	1,752	.546	.949
1960	3,494	.517	.869	4,625	.444	.688	1,774	.531	.702
1961	3,667	.527	.891	4,892	.456	.725	1,880	.551	.950
1962	3,722	.514	.808	4,922	.441	.627	1,946	.541	.915
1963	3,823	.516	.797	5,144	.443	.640	1,971	.536	.745
1964	3,999	.511	.756	5,341	.442	.603	2,124	.537	.846
1965	4,181	.512	.772	5,566	.447	.641	2,332	.535	.872

SOURCE: *Current Population Survey, Consumer Income*, Series P-60.

[a] These estimates of average income are particularly sensitive to our procedure for estimating the mean income for the various income classes.

[b] The Gini concentration coefficient of income.

[c] The variance of the natural logarithms of income.

an annual rate of slightly more than one-third of 1 per cent.[7] This apparent contradiction between Kuznets' hypothesis of secular equalization of personal income and the recent U. S. experience deserves investigation to determine whether the original hypothesis requires modification.

One way to reconcile these results is to assume that the degree of income inequality, measured in these terms, reaches a limit below which it does not tend to fall as development progresses. Comparable annual data on the personal distribution of income for developed countries are not readily available to firmly test this hypothesis, but much fragmentary data suggest income inequality has continued to decline in the postwar period in Western Europe, and in some instances to lower levels than observed in the United States.[8] For example, in the Netherlands the concentration of income among the top 85 per cent of the income units declined between 1946 and 1959 at an annual average rate of 1.7 per cent, from .477 to .378, and in the same period the variance declined from .747 to .410.[9]

Another hypothesis to account for the difference in the behavior of income inequality over time and among countries would postulate that cyclical factors associated with the level of demand and the rate of growth of real income contribute to different short-run behavior in the distribution of income. General patterns of postwar growth are consistent with this hypothesis. In the United States the average civilian unemployment rate was 4.5 per cent in the twenty-one years for which income data are available, averaging 3.5 per cent in the first nine years and 5.4 per cent in the last twelve years. By comparison, the unemployment rate in the Netherlands averaged 2.0 per cent for the years when income data are available and does not evidence a systematic change over the period. Rates of growth in real GNP for this period are also

[7] To estimate the association between time and income inequality least squares estimates were computed for the data using a linear time trend. These aggregate results are similar to those derived by Lee Soltow, "The Share of Lower Income Groups in Income," *Review of Economics and Statistics*, Vol. 47, No. 4 (November 1965), p. 429, Table 1.

[8] See Kuznets, "Quantitative Aspects . . ." United Nations, Economic Commission for Europe, *Economic Survey of Europe in 1956*, Geneva, 1957, Chapter IX; and Lee Soltow, *Toward Income Equality in Norway*, Madison, Wisc., 1965.

[9] Estimates of income inequality for the Netherlands and their regression analysis are presented in Schultz, Chapter VI.

markedly higher for the Netherlands than for the United States, particularly in the later years. To examine systematically the evidence for this hypothesis, a conceptual model is developed in the next part of this paper that implies aggregate empirical relationships between change in income concentration, a secular trend, and cyclical variables.

A CONCEPTUAL MODEL OF SECULAR AND CYCLICAL CHANGE

The secular trend toward income equalization is a product of two developments that reflect the increasing scarcity of labor relative to physical capital in the economy as a whole and the redistribution, to some extent, of factor ownership. The secular growth of labor's share, which tends to be more equally distributed among persons than that of profits or property, has undoubtedly worked to reduce over-all inequality of personal income. Redistribution of factor-resource ownership has also contributed to the secular equalization of income, though the institutional origin of these changes is uncertain.[10] Because these simultaneous shifts and redistributions of factor demand and supply cannot now be disentangled or explained on an aggregate level, the resultant secular trend in income inequality is here presumed to be linear in time, and determinants of short-run deviations from this equilibrium trend are sought in economic variables that measure aggregate disequilibrium.

Cyclical change in the distribution of personal income is more tractable to economic analysis than secular change, for short-run change can be attributed to the level of aggregate demand, where the supply and distribution of factors are constant. Aggregate disequilibrium propagates its effects on the distribution of income through a variety of mechanisms having different time lags and dimensions. From the theoretical and empirical literature on the behavior of income shares and cyclical change in wage and unemployment rates, a number of relationships are plausible, linking the personal distribution of income to

[10] Evidence of resource redistribution is numerous but regrettably fragmentary. Inequality in the distribution of personal physical wealth has decreased in both the United States and the Netherlands, and much evidence suggests that educational and economic opportunities are more equally distributed in mobile industrial societies today than they were in the past. For example, see Robert J. Lampman, *The Share of Top Wealth-Holders in National Wealth, 1922–1956,* Princeton for NBER, 1963.

fluctuations in excess demand. Emphasis is placed here on the role of the rate of change in prices and real output and that of the rate of unemployment.

The model has one unusual feature: it deals explicitly with demand for labor as a nonhomogeneous factor of production. The level of demand determines the level of output and the optimum long-run level and composition of employment, subject to the constraint that in the short run, at the given structure of real wages, capital will not be used unless it earns a nonnegative profit.[11] The level and structure of wages are largely functions of the long-run trend in labor productivity and the relative scarcity of various skills at (equilibrium) full employment levels. The costs to employers of hiring and training labor for specific firm functions is assumed to be positively related to skill and experience levels.[12] Involuntary unemployment occurs whenever demand is inadequate at existing wage levels and turnover costs to warrant "full" employment.

The total supply of labor measured in man-hours is relatively elastic in the short run, but because of cyclical changes in the composition of labor demand, the composition of unemployment and the structure of wage rates among skill and experience levels fluctuates. Labor turnover costs induce employers to concentrate cyclical hiring and firing among

[11] This discussion follows closely the integration of neoclassical production theory and a macroeconomic theory of unemployment advanced by Edwin Kuh in two recent contributions. His concentration on postwar developments in the United States may explain his neglect of the distributional effects of price changes. He also makes no attempt to deal with labor as a nonhomogeneous factor of production since he is not interested in how the labor share of income is distributed among persons. To complement the higher quality of labor available in a period of deficient demand, the firm first withdraws from production the least skill intensive processes (capital). Skills embodied in a firm's labor force are thus regarded as a fixed cost just as are capital costs, giving the firm the incentive to utilize these resources to the greatest possible extent in periods of low demand and add to them reluctantly in periods of high demand. The changing composition of the employed labor force with respect to capital stock offers an explanation for the cyclical pattern of labor productivity. See "Unemployment Production Functions and Effective Demand," *Journal of Political Economy,* Vol. 74 (June 1966), pp. 238–249; and "Income Distribution and Employment over the Business Cycle" in J. S. Duesenberry, *et al.* (eds.), *The Brookings Quarterly Economic Model of The United States,* Chicago, 1965, pp. 227–280.

[12] This argument is developed in different contexts both by Becker and Oi. See Gary Becker, *Human Capital,* New York, NBER, 1964, p. 23+; and Walter Oi, "Labor as a Quasi-fixed Factor," *Journal of Political Economy,* Vol. 70, No. 6 (December 1962), pp. 538–555.

the least skilled and least experienced workers. This hypothesized relationship between the business cycle and the composition of labor demand provides an explanation for the observed behavior of labor productivity, profit shares, and the structure of wage and unemployment rates. Labor productivity rises most rapidly in the early phases of an expansion when the quality of the mix of skills and capital stock retained by the employer provides him with the potential to increase output substantially without incurring large variable costs. Consequently, the rate of change in real output is a good predictor of the profit share in national income.[13] As a rule, wage differentials associated with skill and experience levels narrow in periods of excess demand and widen in periods of deficient demand. Thus, the least skilled and experienced would appear to be subject to the widest cyclical fluctuations in wage rates and also suffer the most pronounced changes in unemployment rates.[14]

Price movements can also affect the distribution of income between income shares (wages and profits) and between particular groups debtors and creditors). If prices of output vary in response to factor price and wage movements, changes in price level may have the opposite effect on the profit share in the short run. Where fluctuations in demand for final output generate the changes in price level, and wages are assumed to be money determined in the short run, the profit share will move in the same direction as price changes.

[13] Charles L. Schultze, "Short-Run Movement in Economic Shares," in *The Behavior of Income Shares: Selected Theoretical and Empirical Issues,* Studies in Income and Wealth, Vol. 27, Princeton for NBER, 1964.

[14] In the United States, Daniel Creamer has shown that the relative increase in labor's share of national income in recessions is not widely spread among members of the labor force. He finds in the manufacturing sector of the U. S. economy executive and professional salary payments fluctuate least during the various phases of the business cycle, general salary compensations fluctuate somewhat more, and wage payments experience the widest movement (pp. 111–113). *Personal Income During Business Cycles,* Princeton for NBER, 1956. The concentration of unemployment among unskilled groups in the labor force as unemployment increases has been observed in many studies of the U. S. labor market: Edward D. Kalachek, "The Determinants of Higher Unemployment Rates, 1958–1960," unpublished Ph.D. dissertation, Massachusetts Institute of Technology, 1963, pp. 166–167; W. G. Bowen and T. A. Finegan, "Labor Force Participation and Unemployment," in A. M. Ross (ed.), *Employment Policy and the Labor Market,* Berkeley, Calif., 1965; and R. M. Solow, *The Nature and Sources of Unemployment in the United States,* Wicksell Lectures, April 14–16, Stockholm, Sweden, 1964.

In summary, short-run fluctuations in excess demand exert a variety of influences on the distribution of income. First, the behavior of income shares appears to be a function of the rate of change in real output and labor productivity, and it may also be altered by unanticipated price changes, depending upon the origin of the change in price level. Second, cyclical change in the composition of labor demand leads to a redistribution of labor's share of income. A tightening of the labor market contributes to a narrowing of wage and unemployment differentials associated with skill or experience levels, which reduces income inequality within the active labor force.

AN EMPIRICAL MODEL

The secular component of the model is assumed to take a linear form in time, representing the concomitant effect of secular changes in factor scarcity and distribution among the population, an effect that determines a long-run functional distribution of personal income. Deviations from the secular trend are hypothesized to be a linear function of the rates of change in prices, real output, the unemployment rate, and the effects of a random error term.

$$C_t = B_1 + B_2\dot{P}_t + B_3\dot{Y}_t + B_4U_t + B_5T_t + e_t \tag{1}$$

where C_t is the concentration of income at time t

$\dot{P}_t$ is the rate of change in wholesale prices at time t

$\dot{Y}_t$ is the rate of change in real output at time t

U_t is the unemployment rate in the civilian labor force at time t

T_t is a linear time trend equal to the number of years elapsed from 1943.

The B's represent a constant term and parameters, and e_t is an error term with constant variance distributed independently over time. According to the earlier discussion, B_3 and B_4 are thought to be positive, and B_5 to be negative if income concentration is secularly declining. Where price changes are due to initial changes in factor prices and wages, B_2 is likely to be negative, or where demand induced changes in output prices lead the price level adjustment, positive.

It is plausible that disequilibrium in the personal distribution of income due to cyclical factors is not eliminated entirely in each year. If the effect of cyclical variables persists for more than a year or adjustment of disequilibrium is only partially accomplished each year, the model is specified somewhat differently. Inclusion of past values of the cyclical variables or the introduction of the lagged dependent variable provides this measure of greater generality to the model. But since this refinement did not receive support from the empirical analysis of postwar data in either the United States or the Netherlands, the more elaborate dynamic model is omitted here.

AGGREGATE EMPIRICAL RESULTS

Because cyclical fluctuations in economic activity are thought to exert systematic effects on the aggregate distribution of income, the secular trend in income concentration can be properly estimated only after time series have been adjusted for these cyclical factors. Estimating by least squares the parameters in equation (1) for U. S. and Dutch data the regression coefficients are of reasonable magnitude, but the sign on the secular trend remains positive for the U. S. and negative for the Netherlands. In the U. S. case only the regression coefficient on the secular trend is statistically significant at standard levels, which may be due partly to sampling variability in the underlying data and the shortcomings of the procedure used to estimate income concentration from the grouped data. For the United States the regression results were:

$$\begin{array}{llllll} C_t = .475 - & .0003\dot{P}_t + & .0006\dot{Y}_t + & .0015U_t + & .0014T_t & R^2 = .543 \\ & (.005) & (.0008) & (.0025) & (.0006) & (1944\text{–}65)\ n = 21 \end{array}$$

where the standard errors are shown below each regression coefficient, and n denotes the number of observations available for the regression. For the Netherlands fewer observations were available, but they are derived from sources in which sampling variability is much less and mean incomes are given.

$$\begin{array}{llllll} C_t = .420 - & .0010\dot{P}_t + & .0012\dot{Y}_t + & .0090U_t - & .0041T_t & R^2 = .993 \\ & (.0000) & (.0001) & (.0017) & (.0004) & (1946\text{–}59)\ n = 9 \end{array}$$

These regression results are consistent with the hypothesis that unemployment or deficient demand for labor adds to income concentration, though the relationship emerges more strongly from analysis of Dutch than U. S. aggregate income data. As hypothesized, the rate of increase in real output, as a proxy for the proportion of profit in national income, is directly associated with income concentration. The current rate of inflation is associated with a decrease in income concentration, as would be expected, if the impetus to price adjustment arose from the side of wage and factor prices rather than final goods prices.

In summary, variation in time series on income concentration is approximated by a sum of a linear trend in time and proxies for excess demand or aggregate disequilibrium. Though the regression coefficients estimated for cyclical variables are much less statistically significant and of smaller magnitude for U. S. than Dutch income data, they agree in sign. The secular trend in the two sets of income data, though reduced in magnitude by the inclusion of cyclical variables, still differ in sign. Aggregate income concentration appears to be increasing in the United States by three-tenths of 1 per cent per year, while in the Netherlands it is decreasing by 1 per cent per year. The paradox persists, and is only partially resolved in the following section by disaggregated analysis of of income inequality.

III. Disaggregate Change in the Distribution of Income

NONCOMPETING GROUPS

The distribution of income can be traced to the distribution of factor ownership among persons and factor scarcity in the market economy, as modified by transfers. Earnings from participation are the major portion of personal income, and individuals for a variety of reasons do not compete in identical markets for current earnings. An individual's talents, acquired training and skills, and experience all differentiate the services he has to offer the market, though imperfections in factor markets and discriminatory barriers may further influence his ability to realize his potential earnings. By this line of reasoning, individuals distinguished by sex and age are, at least to some extent, noncompeting groups in the labor market. It is assumed here that these differences in income and earning levels between groups are largely a reflection of differences in their marginal productivity.[15] Educational attainment might also be

[15] As noted at the end of the paper, the assumption that the variance of incomes represents the variance of individuals' marginal productivities presumes that the distribution of part-time and part-year participation in the aggregate and component groups is not changing. Table A-2 at the end of this paper suggests that this assumption may not be valid, but these data are not sufficient to form any conclusions.

justifiably considered in this context, but breakdowns of the *Current Population Survey* income data by this personal characteristic were not begun until 1958, and the format of these data was later changed.[16] An individual's color is also evaluated as a characteristic associated with differences in income, not because of any established difference in talent between white and nonwhite persons, but because this attribute reflects the effect of present and past imperfection and discrimination in the United States labor market.

Two approaches to the analysis of income inequality within and between these component groups in the population are undertaken here. First, analysis of variance is employed to estimate the proportion of over-all income inequality associated with various groupings of the population and to determine how changes in the relative size of these groups has altered over-all income inequality in the postwar period. Second, regression analysis is used to distinguish secular trends and adjust for cyclical behavior of income distribution within these relatively homogeneous groups. Tables 3, 4, 5, and 6 report the income and income inequality among various sex, age, and color groupings for all years in which income data are available.[17]

ANALYSIS OF VARIANCE

Analysis of variance assumes the normality of populations, and therefore its application to the size distribution of personal income is more justified when the logarithmic transformation of income is used as the variate. The results of the analysis of variance are summarized for a single factor partition in terms of the correlation ratio and the F ratio statistic. The correlation ratio (squared) is equivalent to the proportion

[16] Income size distributions are published divided by sex and education groups for persons 14 years or older with income in 1958, 1961, 1963, and persons 25 years or older with income divided by a smaller number of income classes for 1963, 1964, and 1965. See *Current Population Reports, Consumer Income,* Series P-60, U. S. Bureau of Census, Washington, D. C., various issues.

[17] The relative weights for the sex distribution of the income recipient population is not given in 1944 or 1945 by *Current Population Survey.* Using data on participation rates for men and women and populations 14 years of age or over, it was estimated that there were 29,488,000 women income recipients in 1944 and 23,268,000 in 1945, and 44,654,000 men income recipients in 1944 and 45,347,000 in 1945. For the same reason estimates had to be made as to the proportion of white and nonwhite income recipients in the male and female totals for the years 1949 through 1952. It was assumed that the proportions of nonwhites in the male and female totals remained constant at the 1953 levels of 9.43 per cent and 13.30 per cent respectively.

TABLE 3

VARIANCE OF THE LOGARITHMS OF INCOME FOR MALES 14 YEARS AND OVER BY AGE GROUPS: 1947–65

Year	14–19	20–24	24–34	35–44	45–54	55–64	65+
1947	.867	.337	.420	.602	.692	.690	1.674
1948	1.024	.296	.345	.556	.684	.713	1.405
1949	1.140	.341	.318	.544	.722	.764	1.280
1950	1.229	.351	.297	.564	.687	.814	1.347
1951	1.148	.304	.346	.576	.666	.751	1.243
1952	1.203	.308	.244	.409	.935	1.417	1.411
1953	1.093	.488	.322	.396	.789	.870	1.948
1954	1.077	.713	.345	.482	.655	.795	1.413
1955	1.206	.353	.306	.450	.732	1.100	1.475
1956	1.240	.346	.291	.555	.701	.915	1.434
1957	1.264	.366	.294	.546	.694	.944	1.099
1958	1.438	.397	.305	.391	.680	.609	.849
1959	1.272	.395	.325	.418	.629	.892	1.031
1960	1.284	.403	.361	.514	.604	.742	1.242
1961	1.359	.438	.400	.446	.614	.683	1.739
1962	1.424	.429	.308	.650	.486	.643	1.100
1963	1.337	.450	.294	.368	.531	.784	1.162
1964	1.341	.421	.303	.364	.446	.590	1.477
1965	1.223	.403	.270	.423	.542	.616	.982

SOURCE: *Current Population Survey, Consumer Income,* Series P-60.

of the total population variance accounted for by a particular grouping of the population.[18] In all cases reported the F ratio statistic permits us to reject the null hypothesis that the grouping of the population was random with respect to the income variate. Table 7 summarizes the results of the analysis of variance in terms of the correlation ratio.

[18] The analysis of variance based on relative distributions with 1 in 1,000 accuracy (published income distributions have only .1 per cent accuracy) do not yield the normal property of having the total population variance exactly exhausted by the weighted variance between and within component groups. Consequently the correlation ratio is interpreted here as the ratio of the weighted between group means variance to the sum of the weighted between group means and within group variances. Further research is needed to determine if the discrepancy between the components of total variance and total variance itself is due to only the limited accuracy of the data sources or also reflects some methodological error.

The distinction between the sexes accounted for 10 per cent of the variance of incomes in 1945, and rose to about 20 per cent in 1962–65. Fourteen sex and age groups accounted for about 10 per cent of the variance in 1947 and also show a tendency to rise to around 15 per cent by the mid 1960's. The four groups distinguished by sex and color, on the other hand, account for some 10 per cent of the variance of incomes in the first years of the series, 1949–52, but this proportion gradually falls to a low of somewhat less than 6 per cent in 1965. How should these estimates be interpreted?

There are three possible sources for change in the correlation ratio associated with a particular grouping of the population:

(1) The relative size of the groups may change, though variance and

TABLE 4

VARIANCE OF THE LOGARITHMS OF INCOME FOR FEMALES 14 YEARS AND OVER BY AGE GROUPS: 1947–65

Year	14–19	20–24	25–34	35–44	45–54	55–64	65+
1947	.868	.447	.587	.533	.818	.984	2.033
1948	.876	.530	.567	.720	.699	1.003	1.339
1949	1.012	.419	.489	.628	.656	1.133	1.542
1950	1.158	.420	.591	.698	.754	.971	1.772
1951	1.085	.417	.511	.551	.579	1.021	1.815
1952	1.056	.436	.624	.861	.631	1.087	1.623
1953	1.096	.404	.805	.606	1.005	.791	2.151
1954	1.135	.434	.494	.684	.618	1.193	1.396
1955	1.191	.473	.516	.702	.939	1.508	1.673
1956	1.172	.592	.558	.808	.658	.807	1.524
1957	1.267	.439	.596	.949	.591	.726	1.581
1958	1.386	.563	.575	.560	.596	.985	1.094
1959	1.379	.498	.777	.660	.589	.879	1.863
1960	1.306	.760	.591	.534	.575	.703	1.333
1961	1.408	.589	.760	.841	.904	.917	1.302
1962	1.555	.606	.644	.577	.578	1.113	1.103
1963	1.405	.500	.660	.569	.690	.795	1.231
1964	1.375	.491	.552	.677	.546	.762	1.555
1965	1.707	.624	.553	.532	.617	.688	1.189

SOURCE: *Current Population Survey, Consumer Income*, Series P-60.

TABLE 5

ESTIMATED AVERAGE INCOME AND INCOME INEQUALITY FOR MALES BY COLOR: 1949–65

	White			Nonwhite		
Year	Average Income in Current Dollars [a]	Concentration [b]	Variance of Logs [c]	Average Income in Current Dollars [a]	Concentration [b]	Variance of Logs [c]
1949	2,767	.429	.697	1,370	.466	.701
1950	3,074	.433	.704	1,574	.440	.523
1951	3,380	.401	.679	1,750	.398	.496
1952	3,612	.403	.849	1,911	.407	.716
1953	3,683	.409	.665	1,943	.422	.530
1954	3,707	.414	.647	1,911	.459	.746
1955	3,920	.424	.803	2,036	.440	.510
1956	4,218	.424	.740	2,152	.440	.536
1957	4,224	.418	.655	2,234	.447	.520
1958	4,352	.421	.606	2,310	.465	.712
1959	4,715	.426	.669	2,405	.476	.710
1960	4,861	.438	.698	2,637	.466	.579
1961	5,085	.444	.674	2,711	.466	.684
1962	4,885	.429	.647	2,714	.444	.516
1963	5,340	.431	.602	2,981	.451	.556
1964	5,576	.436	.596	3,294	.488	.676
1965	5,465	.418	.474	3,109	.469	.567

SOURCE: *Current Population Survey, Consumer Income,* Series P-60.

[a] These estimates of average income are particularly sensitive to our procedure for estimating the mean income for the various income classes.

[b] The Gini concentration coefficient of income.

[c] The variance of the natural logarithms of income.

means within the groups do not. If those groups that receive particularly low (or high) incomes increase as a proportion of the total population, this will, other things equal, contribute to increasing the correlation ratio and the over-all variance of incomes. The increasing participation of women in the U. S. labor force, and the recent wave of young entrants into the labor force has been responsible for shifting current year group weights toward the outlying income groups, having the predicted effect on over-all income variance in the United States. This inference can be tested by applying base year group weights to final year relative income distributions. For 1965 the sex distinction accounts for .185 of total variance in incomes with 1965 weights, but only .159 of total variance with 1945 weights. The fourteen age and sex groups account for .153 of total variance of incomes with current year weights, but less than .089 with 1947 weights. The breakdown by color and sex accounted for .055 of total variance in 1965 with current year weights, but somewhat less, .053, with 1949 weights. In this last case the change

TABLE 6

ESTIMATED AVERAGE INCOME AND INCOME INEQUALITY FOR FEMALES BY COLOR: 1949–65

	White			Nonwhite		
Year	Average Income in Current Dollars [a]	Concentration [b]	Variance of Logs [c]	Average Income in Current Dollars [a]	Concentration [b]	Variance of Logs [c]
1949	1,240	.501	.712	609	.631	.954
1950	1,296	.514	.801	608	.644	.959
1951	1,424	.497	.728	649	.626	.931
1952	1,549	.486	.682	746	.644	1.015
1953	1,600	.512	.969	931	.559	.771
1954	1,586	.504	.717	889	.585	.844
1955	1,635	.526	.972	844	.600	.876
1956	1,700	.529	.937	950	.583	.818
1957	1,715	.522	.737	996	.578	.814
1958	1,750	.531	.754	1,016	.593	.855
1959	1,853	.537	.929	1,167	.598	1.278
1960	1,894	.531	.880	1,229	.585	.900
1961	1,955	.543	.916	1,319	.572	.831
1962	2,119	.477	.800	1,319	.561	.782
1963	2,056	.532	.731	1,397	.549	.745
1964	2,203	.532	.853	1,606	.557	1.023
1965	2,397	.531	.853	1,828	.554	.868

SOURCE: *Current Population Survey, Consumer Income*, Series P-60.

[a] These estimates of average income are particularly sensitive to our procedure for estimating the mean income for the various income classes.

[b] The Gini concentration coefficient of income.

[c] The variance of the natural logarithms of income.

TABLE 7

PROPORTION OF VARIANCE OF THE LOGARITHMS OF INCOME ACCOUNTED FOR BY THE WEIGHTED VARIANCE BETWEEN GROUP MEANS: 1944–65 [a]

Year	Sex (2 Groups)	Age and Sex [b] (14 Groups)	Color and Sex [c] (4 Groups)
1944	.129	n.a.	n.a.
1945	.103	n.a.	n.a.
1946	n.a.	n.a.	n.a.
1947	.137	.098	n.a.
1948	.163	.109	n.a.
1949	.130	.143	.097
1950	.162	.153	.109
1951	.147	.183	.117
1952	.136	.133	.099
1953	.172	.162	.066
1954	.141	.126	.080
1955	.132	.138	.067
1956	.150	.157	.071
1957	.146	.161	.076
1958	.189	.151	.089
1959	.189	.187	.087
1960	.168	.156	.068
1961	.195	.145	.061
1962	.213	.140	.069
1963	.196	.156	.068
1964	.218	.168	.071
1965	.185	.153	.055

SOURCE: *Current Population Survey, Consumer Income,* Series P-60.
n.a. = Data not available by these groupings for these years.
[a] Otherwise known as the correlation ratio squared. Because of rounding of the relative distributions from the CPS to 1 in a 1,000, the variance of the total population is not necessarily exhausted by the variance components between and within the groups. The correlation ratio squared reported here represents the proportion of between group mean variance to the sum of between and within group variances rather than the proportion of variance between group means to total population variance. See R. A. Fisher, *Statistical Methods for Research Workers*, London, 1930, pp. 223–224.
[b] Age classes available were as follows: 14–19, 20–24, 25–34, 35–44, 45–54, 55–64, and those persons over 65 years of age.
[c] Color classes available were white and nonwhite.

in weights has not been responsible for the decline in the color/sex correlation ratio as shown in Table 7.

(2) The second source of change in the correlation ratio can arise if the mean incomes of the groups change. Since the services of different groups are not perfect substitutes in the labor market, the more rapidly growing groups will tend to depress their income status relative to that of other groups when shifts in demand are neutral. If this inference is correct, shifting weights toward low income units will be reinforced by the tendency for these low income groups to experience somewhat slower advances in their income than the population as a whole. Table 8 gives the relative income status of sex and age cohorts in several years. The decline in relative income status is most noticeable for the youngest cohort of both sexes. However, no general deterioration in the income

TABLE 8

INCOME OF SEX/AGE GROUPS RELATIVE TO AVERAGE FOR POPULATION: SELECTED YEARS 1944–65 [a]

Year	14–19	20–24	25–34	35–44	45–54	55–64	65+
			Part A. Male				
1947	29	74	125	154	153	128	85
1950	20	82	130	161	158	132	78
1953	24	82	140	159	158	136	80
1956	18	85	143	170	164	171	77
1959	16	80	144	172	163	149	73
1962	15	79	144	165	173	151	79
1965	18	82	151	182	181	148	75
			Part B. Female				
1947	25	53	58	64	66	55	43
1950	20	55	59	61	62	54	29
1953	20	55	65	63	69	55	34
1956	18	54	58	64	66	58	31
1959	15	51	58	58	67	59	33
1962	16	49	57	64	70	65	34
1965	19	60	62	67	74	66	35

SOURCE: *Current Population Survey, Consumer Income*, Series P-60.

[a] These estimates of average income relatives are particularly sensitive to our procedure for estimating mean income for the various income classes.

status of older cohorts of women is apparent although their proportion in the population receiving income has increased. Theoretically the effects of both (1) and (2) would have operated in the United States to augment the over-all variance of incomes and increase the correlation ratio associated with age and sex groupings of the population.

(3) Finally, the variance of incomes within the groups can change. If these within group variances tend to increase, other things equal, the correlation ratio would decrease and over-all income variance would increase. Investigation of changes in within group variance is postponed to the last part of this section.

In conclusion, there appears to be evidence that the increasing participation of women, as well as a growing rate of entry into the labor force of younger cohorts, added to over-all U. S. income inequality during the postwar period. The number of persons with income was increasing in the United States at more than twice the rate of that recorded in the Netherlands, and the U. S. rate was rising over time.[19] One method of partially correcting for this change in the composition of persons in the population with income is to hold the group weights constant through time, and only allow the relative distributions of income attributed to the groups to vary. Ultimately, however, it may make more sense to view differences in personal income within sex and age cohorts as the fundamental measure of income inequality through time.

SECULAR AND CYCLICAL CHANGE WITHIN SEX/AGE SPECIFIC GROUPS

The secular trends in income variance within sex and age specific groups has differed markedly as seen in Tables 3 and 4. It might be expected that where labor services from a group have been in short supply, income differences would have narrowed according to the conceptual model developed in Section II. Conversely, when the supply of services from a sex/age specific group increased rapidly, income differences might be expected to widen even in the absence of unemployment. This hypothesis reverses the compositional argument from the previous study

[19] In the Netherlands the economically active population grew at about .7 per cent per year in the period analyzed, 1946–59. In the United States the number of persons with income covered by the *Current Population Survey* grew at a 1.8 per cent per year compounded average rate in the period 1945–58, and from 1958–65 at a rate in excess of 2.2 per cent per year.

of cyclical changes in labor demand to investigate the effects of longer-run changes in the composition of labor supply.

Linear secular trends in income variance within the sex/age specific groups are estimated for the period 1947–65 by least squares, as is shown in the first two columns of Table 9 accompanied by the t statistic for each regression coefficient. For the younger cohorts the variance of income has tended to increase, whereas the older cohorts have all experienced decreasing income variance. For women the turning point between increasing and decreasing income variance comes somewhat later than for men, but the similarity in the pattern of coefficients is nevertheless unambiguous.

TABLE 9

ESTIMATED ANNUAL AVERAGE PERCENTAGE CHANGE IN SEX/AGE GROUP INCOME VARIANCE WITH AND WITHOUT ADJUSTMENT FOR DEMAND FACTORS: 1947–65 [a]

Sex/Age	Secular Change Unadjusted	t Statistic [b]	Secular Change Adjusted [a]	t Statistic [b]
Male				
14–19	1.6	5.0	.7	1.5
20–24	1.3	1.4	1.7	.6
25–34	–.6	–1.1	–.2	–.4
35–44	–1.5	–2.2	–1.1	–1.4
45–54	–2.1	–4.1	–2.4	–4.7
55–64	–1.3	–1.3	–2.0	–2.1
65+	–1.4	–1.7	–1.1	–1.2
Female				
14–19	2.8	10.4	2.3	4.5
20–24	2.0	3.0	1.7	1.6
25–34	.9	1.3	1.5	1.6
35–44	–.3	–.4	–.3	–.3
45–54	–1.0	–1.3	–1.0	–1.0
55–64	–1.7	–2.1	–2.1	–2.3
65+	–1.9	–2.8	–1.4	–1.8

[a] Adjusted for current rate of change in prices, real output, and the unemployment rate within the specific sex/age group.

[b] Regression coefficient significant at .01 level if t statistic exceeds 2.9; at .05 level if it exceeds 2.2; at .1 level if it exceeds 1.75.

Because the rate of growth in the twenty-year period studied has declined somewhat over time and the unemployment rate has tended to increase particularly among the youngest and oldest members of the labor force, it was decided to estimate secular trends allowing for the effect of cyclical changes in demand. The same factors are emphasized as in the aggregate model: rate of change in wholesale prices, and real GNP, and the age/sex specific unemployment rate.[20] The annual estimates of the rates of change in income variance are shown in the last two columns of Table 9. Adjusting the secular trend estimates for demand factors reduces the magnitude of the rising secular trend attributed to the younger cohorts and increases the rate of decline of the secular trend for the cohorts between the ages of 55 and 64.

IV. Conclusion

In the aggregate, the relative size distribution of income among persons in the United States has not changed substantially since World War II. In contrast, income inequality in some European countries appears to have diminished. Though undoubtedly this difference between the U. S. and European experiences has been exacerbated by the slow postwar growth of the U. S. economy that permitted the level of unemployment to rise to successively higher plateaus, change in the composition of the economically active population has also played an important role.

The flow of young entrants and married women into the U. S. labor force has increased the proportions of these two low-paid groups among income recipients. Were one to hold constant the weights of age and sex specific groups in the population, aggregate income inequality would decrease over the postwar period in the United States. Considering the income variance within sex and age specific groups over time, all but the younger cohorts exhibit a tendency toward secular equalization of personal incomes.

To end on a discordant note, it should be observed that, for lack of data, one factor has been neglected in this analysis: the extent to which income recipients are full-time, year-round members of the labor force. If the composition of income recipients with respect to annual time spent

[20] Age and sex specific unemployment rates were taken from *Manpower Report of the President,* March 1966, Department of Labor, Washington, D. C., 1966, Table A-12, p. 167.

TABLE A-1

TIME SERIES FOR THE UNITED STATES, 1942–65
(per cent)

Year	Rate of Change in Wholesale Prices [a]	Rate of Change in Cost of Living [a]	Rate of Change in GNP, 1958 Prices [a]	Civilian Labor Force Unemployment Rate
1942	12.97	7.21	12.93	4.71
1943	4.63	6.16	13.20	1.92
1944	.71	1.66	7.18	1.22
1945	1.76	2.28	−1.69	1.93
1946	14.16	8.45	−11.99	3.94
1947	22.84	14.41	−.86	3.56
1948	8.25	7.71	4.45	3.80
1949	−5.01	−.96	.12	5.90
1950	3.95	.96	9.63	5.30
1951	11.41	8.00	7.91	3.30
1952	−2.79	2.21	3.05	3.10
1953	−1.38	.76	4.48	2.90
1954	.22	.43	−1.41	5.60
1955	.32	−.32	7.62	4.40
1956	3.22	1.50	1.85	4.20
1957	2.91	3.48	1.43	4.30
1958	1.41	2.76	1.15	6.80
1959	.20	.79	6.39	5.50
1960	.01	1.58	2.48	5.60
1961	−.40	1.07	1.95	6.70
1962	.30	1.15	6.56	5.60
1963	−.30	1.23	4.00	5.70
1964	.20	1.31	5.26	5.20
1965	1.99	1.67	5.93	4.60

SOURCE: *Economic Report of The President 1966*, Washington, D. C., 1967, Statistical Appendix and earlier issues.

[a] The per cent change for year X is derived by subtracting the value of year X-1 from the value of year X, dividing the difference by the value of year X-1, and then multiplying the result by 100.

TABLE A-2

PER CENT OF CIVILIAN INCOME RECIPIENTS WHO HAD YEAR-ROUND FULL-TIME WORK[a] IN 1955, 1960, AND 1965

Sex/Age	1955	1960	1965
All males	63.1	58.3	59.8
14–19	10.9	7.4	7.2
20–24	47.6	42.1	49.1
25–34	77.0	72.7	77.6
35–44	79.6	77.8	81.9
45–54	77.6	74.4	79.3
55–64	65.3	64.8	68.0
65+	24.5	17.2	15.4
All females	31.1	28.3	29.3
14–19	9.8	8.5	6.8
20–24	37.5	33.8	35.6
25–34	37.2	32.3	36.4
35–44	41.6	40.4	41.4
45–54	41.0	41.6	54.4
55–64	36.6	33.5	37.0
65+	6.2	4.3	4.8

SOURCE: *Current Population Survey, Consumer Income*, Series P-60.

[a] Defined as working 35 hours or more per week for 50 weeks or more during the given year.

in the labor force has changed, then changes in income variance do not necessarily represent changes in the variance of individuals' marginal productivities. For example, if the younger cohorts contain an increasing proportion of students who work only part-time until they finish school, and then rise rapidly to a much higher income level, the observed differences in income within this heterogeneous group do not reflect differences in the productivity of the members of this group, some of whom are investing their time in education while others are realizing the benefits of full-time earnings. To deal with this issue of the duration of participation will require more detailed data than are available from the *Current Population Survey*.[21]

[21] See Table A-2 in the appendix to this chapter for data relating to the proportion of income recipients in age/sex specific groups that were full-time, year-round workers in 1955, 1960, and 1965. Data for this group are not distributed by income size, nor are these proportions given for years before 1955.

COMMENT

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In this paper Dr. Schultz extends his earlier analysis of secular and cyclical change in the size distribution of personal income in the Netherlands to a comparable study of the United States for the period 1944 to 1965. He first presents an aggregate model "that incorporates the short-run factors that displace the distribution of income from the long-run equilibrium trend." He also presents, in Section III, measures of income inequality for the population classified by color and sex, and age and sex. My comments will focus on his analysis of changes in income inequality of the sex-color population groups.

Dr. Schultz has calculated two measures of inequality—the Gini concentration coefficient and the variance of the natural logarithms of income. As he has noted, if the size distribution of income is normal in the logarithms, both measures of inequality move together and one is a function of the other. His data indicate that the income size distributions of some subgroups in the total population in fact are not log-normal.

Tables 5 and 6 of his paper show estimated average incomes and the two sets of coefficients for the seventeen year period 1949–65 for males and females classified by color. The two measures of inequality differ both in level and amplitude of year-by-year variation. The Gini coefficients show smaller change in the short run as well as over the entire time span. In any single year, the level of income inequality, as measured by the Gini coefficients, is lowest for white males, followed sequentially by nonwhite males, white and nonwhite females. (Mean incomes, ranked from high to low, are ordered similarly.) This finding is similar to that of Herman Miller for the period 1947–60.[1] In contrast, the variance of logarithms for nonwhite males tends to be lower (less income inequality) than those for white males.

The two measures of inequality do not display a consistent pattern of either year-by-year change or average change over the entire period. I

[1] Herman P. Miller, *Trends in the Income of Families and Persons in the United States, 1947–1960.* Technical Paper No. 8, U. S. Bureau of the Census, Washington, D. C., 1963, Table 16.

therefore calculated three-year moving averages of the two sets of coefficients and average income and examined the correlations. The data are shown below.

THREE-YEAR MOVING AVERAGES, 1949–65

	White Females			Nonwhite Females		
Center Year	Average Income	Gini Coefficient	Variance of Logarithms	Average Income	Gini Coefficient	Variance of Logarithms
1950	$1,320	.504	.747	$ 622	.634	.948
1951	1,423	.499	.737	668	.638	.968
1952	1,524	.498	.793	775	.610	.906
1953	1,578	.501	.789	855	.596	.877
1954	1,607	.514	.886	888	.581	.830
1955	1,640	.520	.875	894	.589	.846
1956	1,683	.526	.882	930	.587	.836
1957	1,722	.527	.809	987	.585	.829
1958	1,773	.530	.807	1,060	.590	.982
1959	1,832	.533	.854	1,237	.592	1.011
1960	1,901	.537	.908	1,238	.585	1.002
1961	1,989	.517	.865	1,289	.573	.838
1962	2,043	.517	.816	1,345	.561	.786
1963	2,126	.514	.796	1,441	.556	.850
1964	2,219	.532	.812	1,610	.553	.879
	White Males			Nonwhite Males		
1950	$3,074	.421	.693	$1,565	.435	.573
1951	3,355	.412	.744	1,745	.415	.578
1952	3,558	.404	.731	1,868	.409	.581
1953	3,667	.409	.720	1,922	.429	.664
1954	3,770	.416	.705	1,963	.440	.595
1955	3,948	.421	.730	2,033	.446	.597
1956	4,121	.422	.733	2,141	.442	.522
1957	4,265	.421	.667	2,232	.451	.589
1958	4,430	.422	.643	2,316	.463	.647
1959	4,643	.428	.658	2,351	.469	.667
1960	4,887	.436	.680	2,584	.469	.658
1961	4,944	.437	.673	2,687	.459	.593
1962	5,103	.435	.641	2,802	.454	.585
1963	5,267	.432	.615	2,996	.461	.583
1964	5,460	.428	.557	3,128	.469	.600

Since the three-year moving averages of income rise consistently over the period under observation, the data can be examined for trend. My reading of the charts yields the following:

CORRELATION/TREND, 1949–65, AVERAGE INCOME AND:

	Gini coefficient	Variance of logarithms
White males	Positive	Negative
Nonwhite males	Positive	None
White females	Positive	None
Nonwhite females	Negative	None

As average income increases, income inequality will decrease if greater gains are made by the lower income groups. If, on the other hand, an increase in the over-all average merely reflects an expansion in over-all income range and relatively small gains achieved at the lower end of the income range, income inequality could increase. Herman Miller's data [2] indicate that relative shares of aggregate income have declined for the lower income quintile ranges, with a corresponding increase at the upper

QUINTILE SHARES OF AGGREGATE INCOME, BY SEX AND COLOR, 1951 AND 1960
(*percentage share in aggregate income*)

	Males		Females	
Quintile Range	White	Nonwhite	White	Nonwhite
1960				
Lowest 2 quintiles	12.6	10.2	9.1	9.5
Middle quintile	17.7	16.8	13.9	12.5
Highest 2 quintiles	70.0	73.0	77.1	77.9
Total	100.3	100.0	100.1	99.9
1951				
Lowest 2 quintiles	15.5	14.7	11.1	12.6
Middle quintile	17.6	19.0	16.3	13.5
Highest 2 quintiles	66.9	66.4	72.7	74.0
Total	100.0	100.1	100.1	100.1

[2] *Ibid.*

quintiles. A comparison of the income share distributions in 1951 and 1960 shows that income shares of the lowest income quintiles of all four sex-color groups have declined.

It is also possible to compare the upper income limits for the first four income quintiles. Of the four sex-color groups, nonwhite females in the lower quintiles have experienced the greatest expansion in income range in 1951 and 1960.

PERCENTAGE INCREASE IN UPPER INCOME LIMIT PER QUINTILE, BY SEX AND COLOR, 1951 AND 1960

Income Quintile	Males		Females	
	White	Nonwhite	White	Nonwhite
Lowest quintile	9.4	1.6	9.8	41.2
Second quintile	30.6	16.9	8.0	51.5
Third quintile	49.8	41.5	18.2	55.7
Fourth quintile	52.6	57.6	37.5	75.3

Clearly, this comparison indicates that nonwhite females have moved more rapidly towards greater income equality than the other population groups during this period, a finding consistent with the trend displayed by the Gini coefficients in Schultz's paper (Table 5).

In calculating the average income data given in Tables 5 and 6, Dr. Schultz, unlike Miller in his volume, did not take midpoints as the average income of each closed income interval. Rather, he "assumed that the geometric mean of the interval is the average income level in accord with the approximately log-normal distribution of the income size variate." By choosing this procedure, he has provided an excellent opportunity to compare average income values for sex-color groups as calculated by the two methods. (Both authors followed the same procedure for estimation of the average value of open-ended intervals.) Over the 1949–60 period a comparison is possible. There are only minor variations in the two sets of estimates for white males, indicating that Schultz's assumption of a log-normal distribution holds reasonably well for this aggregate population group. Variations between the two series increase for nonwhite males, white and nonwhite females, respectively. There is a significant and quite consistent convergence of the two series of income estimates over time for all four population groups, with the greatest relative and absolute convergence occurring among nonwhite females, dropping from a difference of 22 per cent in 1949 to 6 per cent in 1960. It would be of interest to know if the trend towards identity has continued since 1960.

AVERAGE INCOME ESTIMATES: PERCENTAGE DIFFERENCES BETWEEN THOSE OF SCHULTZ AND MILLER,[3] BY COLOR AND SEX, 1949–60
(*Schultz estimates taken as base*)

	Males		Females	
	White	Nonwhite	White	Nonwhite
1949	1.2	4.0	7.8	22.2
1950	1.8	3.1	6.7	21.7
1951	1.4	2.6	5.2	20.0
1952	–	–	–	–
1953	1.0	1.7	4.3	10.0
1954	0.6	4.2	5.3	12.1
1955	0.3	1.6	4.7	12.7
1956	−0.3	2.2	2.9	11.1
1957	1.0	2.3	5.1	9.8
1958	−0.1	0.9	4.2	11.0
1959	−0.2	1.2	3.0	6.8
1960	0.5	1.4	3.1	6.1

[3] *Ibid.*

Among whites, both male and female, the Gini coefficients are positively correlated with level of income. They seem to respond to the *secular* trend in the levels of income rather than year-by-year changes in income. This effect is most pronounced among the males.

Among nonwhite females, the data also appear to reflect the overriding effect of secular trend. In this group, however, a secular rise in average incomes is accompanied by a *decrease* in inequality.

The nonwhite males present a different picture. This population group seems to be more affected by year-by-year changes in the *level,* than by the secular trend of income. (These comments relate to the period commencing in 1951, following the 1948–49 recession and the beginning of the Korean war in 1950.)

The level of inequality of income may be altered by a variety of factors, as Dr. Schultz has stated. Reduction of unemployment, increase in the proportion of full-time workers, etc., will have the effect of raising incomes among the lower income groups. These may account for reduction of income inequality among nonwhite females. And, relatively small increases in income at the lower tail of the distributions and proportionately greater increases at the *upper* income levels for white males and females and nonwhite males could be reflected in a *positive* correlation between mean income and the measure of inequality.

A combination of these two separate effects could result in either an increase or decrease in inequality, depending upon their relative importances. In the case of nonwhite males, both factors seem to be operating—so that a positive correlation between average income and the inequality measure is not as clear cut.

Dr. Schultz has concluded that separation of secular and cyclical effects can best be achieved through standardizing the distributions over time by age and sex and I concur. It probably is also desirable to standardize by place of residence (urban, rural farm and nonfarm), changes in the occupational distribution to take account of differential changes in wage rates, and changes in the proportion of income recipients.

Summary

Both the Gini coefficients and the variance of the natural logarithms of income are influenced by secular and cyclical changes as well as shifts in the relative importance of factors affecting level of income of particular population groups. It would appear necessary to take explicit account of changes in population characteristics (income affecting characteristics) before it is possible to isolate secular changes.