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Distribution Effects and the Aggregate Consumption Function

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This paper investigates whether and how the distribution of income affects the fraction of disposable income which is consumed. It is shown that a slight generalization of the permanent income or "life-cycle" model of consumption makes each individual's lifetime marginal propensity to consume a fraction of his lifetime disposable resources unless two taste parameters are equal. In considering what this implies for the aggregate consumption function, the tenuous connection between theoretical constructs and observed facts is stressed. Previous empirical work on the subject is criticized for failing to define properly either "income distribution" or "consumption," and a new test, based on the theory, is outlined. Because of data limitations, a number of compromises with this ideal test must be made, and several alternative models are estimated. On the whole, they suggest that equalizing the distribution of income would either leave aggregate consumption unchanged or diminish it slightly.

I. Introduction

Does the manner in which a given amount of income is distributed affect the fraction of it that is consumed? In the early post-Keynesian days it was commonly assumed, presumably on the basis of Keynes's own intuition, that it does—in particular that equalization of the income distribution would increase consumption. With the publication of Kuznets's (1942) and Goldsmith's (1955) data, and the ascendancy of the Friedman (1957) and Modigliani and Brumberg (1954) models of consumer

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behavior, this view fell into disrepute in academic circles. It was supplanted by the view that marginal, and perhaps even average, propensities to consume are constant over the income distribution. Of course, this "modern" view does not accord very well with intuition. It does, after all, seem "obvious" to most people, especially those not schooled in macroeconomics, that the rich save proportionately more than the poor, even at the margin.

While what is "obvious" is not always true, I was somewhat shocked to discover that the notion that aggregate consumption is independent of the income distribution has never been subjected to a direct empirical test. That is, the hypothesis that the *size distribution* of income does not affect consumption has never been treated as a special case of a more general class of consumption function and tested by standard statistical techniques. I propose to do so in this paper.

Let me begin with a confession. At the outset of this research I hoped to establish:

> *PROPOSITION A:* The marginal propensity to consume of an individual falls as his disposable income rises.

And therefore:

> *PROPOSITION B:* Out of any given total disposable income, a larger share is spent on consumption when income is more equally distributed.

As it turns out, while both theory and empirical evidence lend at least some support to A, they do not support B. In fact, as I explain shortly, Proposition B does not follow from Proposition A. What does follow is the similar-sounding proposition:

> *PROPOSITION C:* If income is taken from one individual and given to another individual who is identical in all relevant respects save that his income is higher, then total consumption will decline.

The next section develops the theoretical equipment necessary to investigate the effects of redistribution on aggregate consumption. I derive a plausible condition on individual utility functions which is sufficient to guarantee Proposition A, and then establish (as should be obvious) that A implies C. Section III explains why I view previous tests of the effect of income inequality on aggregate consumption as inconclusive, and derives a model for testing Proposition B in the context of the permanent income theory. Section IV explains some compromises that had to be made because of weaknesses in the data, and presents the empirical results I have obtained with a compromise model. These results suggest that a rise in income inequality, disposable income held constant, would either have no effect on consumption or would actually *increase* it. That is to say, while Propositions A and C may be true, the obverse of B is given at least mild support by postwar American data. Section V offers a variety of possible explanations for this result, and the last section is a brief summary.

II. The Implications of Pure Theory

A. Optimal Life-Cycle Consumption

By now the derivation of the aggregate consumption function from a Fisherian model of intertemporal utility maximization, as pioneered by Modigliani and Brumberg (1954) and Friedman (1957), has achieved widespread acceptance. In this theory, the consumer chooses the time path for consumption, $c(t)$, which maximizes lifetime utility, subject to the constraint that the present discounted value of all consumption, plus the present discounted value of the bequest (if any), is equal to lifetime disposable resources, W. That is,

$$\int_0^T c(t)e^{-rt} + K_T e^{-rT} = W, \tag{1}$$

where t is age, r is the rate of interest, T is the length of life, K_T is the bequest, and W is defined as the sum of the inheritance plus the present discounted value of earned income.[1] In order to get the typical "life-cycle" or "permanent income" result that consumption at each instant is proportional to W, the lifetime utility functional must be of the form

$$U = \int_0^T \frac{c(t)^{1-\delta}}{1-\delta} e^{-\rho t}\, dt + \frac{bK_T^{1-\beta}}{1-\beta}: \quad \delta, \quad \beta > 0; \quad b \geq 0, \tag{2}$$

with the further stipulation that $\delta = \beta$.[2] Thus (2) represents a minor generalization of the Modigliani-Brumberg-Friedman (henceforth MBF) model—a generalization with important consequences for the question of distribution effects.

The maximization of (2) subject to (1) is a well-known problem, which was first studied by Strotz (1955–56). A heuristic method of solution is

[1] Equation (1) is the budget constraint only under the assumption of a perfect capital market. This same assumption allows me to collapse all gifts to heirs into a single number, K_T. That is, an *inter vivos* gift of G given at age t is *equivalent*—for both donor and recipient—to a gift of $Ge^{r(T-t)}$ at death.

[2] This is implied by some results of Yaari (1964) and is developed in some detail in Blinder (1974, Chap. 2).

given in the Appendix. For present purposes, it suffices to note that the optimal plan is given implicitly by the following equations:

$$c(t) = c_0 e^{gt}, \qquad \text{where } g \equiv (r - \rho)/\delta; \tag{3}$$

$$c_0 = \phi(r, \rho, \delta, T)(W - K_T e^{-rT}); \tag{4}$$

$$K_T = (be^{rT})^{1/\beta} c_0^{\delta/\beta}, \tag{5}$$

where $\phi(\cdot)$ is a known function specified in the Appendix.

The strict MBF model holds that consumption at each instant is proportional to W, with the constant of proportionality dependent on age (t), the length of life (T), the rate of interest (r), and tastes. This result follows from (3)–(5) under two sets of circumstances. The first is $b = 0$. This is the strict life-cycle model of Modigliani-Brumberg-Ando (MBA). If there is no utility from bequests, then the optimal K_T will be zero for every person, as is clear from (5). By (3) and (4), then, c_0 (and hence all $c[t]$) will be proportional to W. Specifically, $c(t) = \phi e^{gt} W$. The second condition is $\delta = \beta$. This is a modification mentioned by Modigliani and Ando (1960). In this case K_T is proportional to c_0 by (5), so that (3)–(5) can be solved to yield

$$c(t) = \left(\frac{\phi e^{gt}}{1 + \phi b^{1/\delta} e^{rT(1-\delta)/\delta}}\right) \cdot W.$$

It is important to note that these are the *only* two cases which give rise to strict proportionality, that is, a constant lifetime marginal propensity to consume (MPC). Extending the life-cycle model to allow for the bequest motive destroys the proportionality property unless $\delta = \beta$. The lifetime MPC in the more general case can be found by implicit differentiation in (3)–(5). The answer turns out to be

$$\frac{\partial c^*}{\partial W} = \left(1 + \frac{\delta}{\beta} b^{1/\beta} e^{rT(1-\beta)/\beta} \phi^{\delta/\beta} c^{*(\delta/\beta)-1}\right)^{-1},$$

where

$$c^* \equiv \int_0^T c(t) e^{-rt}\, dt$$

is lifetime consumption. In words, the lifetime MPC is smaller than unity as long as $b > 0$, and is decreasing with W if $\delta > \beta$ or increasing with W if $\beta > \delta$.

What do these conditions mean? If β, the elasticity of the marginal utility of bequests, exceeds δ, the elasticity of the marginal utility of consumption, then consumption is the luxury good, that is, has a wealth elasticity greater than unity. Conversely, if $\delta > \beta$, then bequests are the luxury good. It seems plausible, to me at least, that bequests should be the luxury good;[3] but, in the absence of the requisite empirical evidence, each reader is free to make his own judgment. My only purpose here is to establish that it is possible, within the basic MBF model, to have an MPC which either rises or falls with income.

B. The Effect of Redistribution on Aggregate Consumption

I shall now prove that, if the MPC declines with W, an increase in income inequality must reduce consumption. Conversely, if the MPC actually rises with W, a rise in inequality will increase consumption.

Consider a population of individuals identical in every respect save permanent income. Let y denote permanent income, defined as the flow equivalent of the stock of lifetime resources, so that y is proportional to W. Then the model of consumption behavior just developed implies $c = c(y)$, $1 > c'(y) > 0$, and $c''(y) \gtreqless 0$, according as $\delta \gtreqless \beta$. Let the distribution of permanent income be given by a density function $f(y, d)$, where d is a very general indicator of inequality to be explained shortly; and let $F(y, d)$ be the corresponding cumulative distribution function. Finally, let μ, a, and b denote, respectively, the average, lowest, and highest permanent income in the population.

It is easily established that[4]

$$\mu = b - \int_a^b F(y, d)\, dy. \tag{6}$$

The parameter d represents what Rothschild and Stiglitz (1970) have termed a "mean preserving spread." That is, a rise in d signifies a sequence of transfers from poorer persons to richer ones (called "regressive transfers")

[3] Becker's recent analysis of intergenerational transfers (1974) provides some theoretical support for the notion that $\delta > \beta$. Consider, for simplicity, a two-generation family. Letting c_i denote lifetime consumption, W_i denote lifetime wealth, and m_i denote the part of W_i not inherited, the budget constraints for generations $i = 1, 2$ would be $c_1 + k = W_1$ and $c_2 = W_2 \equiv m_2 + k$, where k is the bequest from generation 1 to generation 2, and all quantities are discounted to a common date ($c_2 = W_2$, since generation 2 leaves no bequest). Becker observes that the two budget constraints can be collapsed to $c_1 + c_2 = W_1 + m_2 \equiv S$, where S is what he calls "social income." He then shows that if the elasticity of c_1 (and therefore also of c_2) with respect to S is approximately unity, the elasticity of k with respect to W_1 (which is my δ/β) must exceed unity. The proof is almost immediate. Let η_{xy} denote the elasticity of x with respect to y. The budget constraint for generation 1 implies that a weighted average of $\eta_{c_1W_1}$ and η_{kW_1} equals unity. But $\eta_{c_1S} = 1$ implies that $\eta_{c_1W_1} = (W_1/S)\eta_{c_1S} < 1$. It is thus clear that $\eta_{kW_1} > 1$. QED.

[4] Proof: By definition

$$\mu = \int_a^b yf(y, d)\, dy.$$

Integrating this by parts yields

$$\mu = yF(y, d)\Big|_a^b - \int_a^b F(y, d)\, dy.$$

Equation (6) follows by noting that $F(b, d) = 1$, $F(a, d) = 0$.

which leave the mean unchanged. I add the further stipulation (solely for convenience) that the maximum and minimum incomes are also unaffected, so that a change in d must satisfy the following:

$$\frac{\partial a}{\partial d} = \frac{\partial b}{\partial d} = 0; \tag{7a}$$

$$F_d(y, d) \text{ is continuous on the interval } a \leq y \leq b; \tag{7b}$$

there is some y^* in the interval (a, b) such that

$$\begin{aligned} F_d(y, d) &\geq 0 \quad \text{for} \quad a \leq y \leq y^*, \quad \text{and} \\ F_d(y, d) &\leq 0 \quad \text{for} \quad y^* \leq y \leq b; \end{aligned} \tag{7c}$$

$$\frac{\partial \mu}{\partial d} = -\int_a^b F_d(y, d)\, dy = 0. \tag{7d}$$

The last requirement, that shifts in d leave the mean unchanged, follows from (6).

With the preliminaries thus established, the proof is quite direct.[5] Aggregate consumption is defined as

$$C = \int_a^b c(y) f(y, d)\, dy$$

so that the effect of an increase in inequality on aggregate consumption is

$$\frac{\partial C}{\partial d} = \int_a^b c(y) f_d(y, d)\, dy.$$

Integrating this by parts yields[6]

$$\frac{\partial C}{\partial d} = -\int_a^b c'(y) F_d(y, d)\, dy. \tag{8}$$

First consider (8) in what I take to be the more plausible of the two cases—the case where $c'(y)$ is falling. By (7), F_d is a continuous function which is positive when y is "low," is negative when y is "high," and integrates to zero over its entire range. The integral in (8) attaches higher weights to the (positive) values of F_d which occur when y is low than it does to the (negative) values of F_d which occur when y is high. Thus the integral must be positive, and $\partial C/\partial d$ must be negative. Conversely, if $c'(y)$ were a rising function of y, $\partial C/\partial d$ would be positive. Of course, if $c'(y)$ is constant, (7d) immediately implies that $\partial C/\partial d = 0$. In words, I have established:

> *PROPOSITION D:* A mean-preserving spread in the income distribution will decrease, leave unchanged, or increase aggre-

[5] The proof follows a suggestion by Rothschild and Stiglitz (1970, p. 237 n.).

[6] Again I use the facts that $F(b, d) = 1$, $F(a, d) = 0$.

gate consumption according as δ is greater than, equal to, or less than β.

Of course, Propositions A and C follow from D only with the added assumption that $\delta > \beta$.

It is worth pausing at this juncture to consider what has *not* been proven. Proposition D refers only to transfers within a group which is identical in all relevant respects save income. It is not applicable to transfers from one socioeconomic group to another (e.g., whites to blacks; men to women) if there is any reason to believe that tastes may differ in the two groups. Nor is it applicable to transfers where the age distribution of the donors differs from the age distribution of the recipients. Finally, the fact that the population consists of many age cohorts poses still another problem. Suppose $\delta > \beta$, and there is a decline in inequality in the older (donor) cohorts. As just noted, this would lead to greater average consumption within these cohorts. However, it would lead to lower average bequests, and hence to reduced consumption within the younger (recipient) cohorts. The macro consumption function will reflect both these changes.[7]

The practical implication of all this, of course, is that Proposition D gives no basis for predicting the effect on aggregate consumption of most real-world redistributions.[8] That is, it certainly does *not* establish Proposition B.

III. Testing for Distribution Effects

A. The Definition of Income Distribution

In the typical test for distribution effects in the aggregate consumption function, the income variable in the model is disaggregated into two or more components, and the hypothesis that the two (or more) regression coefficients are equal is tested. Suppose, for example, that the maintained hypothesis is represented by the estimating equation

$$C_t = aY_t + bC_{t-1} + u_t, \tag{9}$$

where Y_t is current disposable income (real or nominal; total or per capita) and C_t is consumer expenditures (defined symmetrically). A typical test is to divide Y_t into labor income (L_t) and property income (P_t), reformulate the model as

$$C_t = a_1L_t + a_2P_t + bC_{t-1} + u_t,$$

[7] I owe this last point to Robert Barro.

[8] The theory does, however, have testable implications that are not given in Proposition D. For example, *ceteris paribus*, a transfer from the young to the old will increase consumption (assuming g is positive).

TABLE 1

U.S. INCOME DISTRIBUTION IN 1962 BY COMPONENTS

DECILE GROUP	SHARE (%) IN: Wages and Salaries	Business and Property Income	Pensions and Annuities	Other Income
Lowest	0	0	11	4
Second	1	2	19	16
Third	3	4	16	24
Fourth	5	4	17	13
Fifth	8	6	4	9
Sixth	10	6	9	4
Seventh	13	6	5	7
Eighth	15	9	6	9
Ninth	18	13	4	9
Highest	27	51	8	5

SOURCE.—Projector, Weiss, and Thoresen (1969, table 4, p. 128).

and test $a_1 = a_2$ against the alternative $a_1 > a_2$. Generally the null hypothesis of no distribution effects cannot be rejected.[9]

What is the rationale for this test? The theory outlined in Section II suggests that MPCs might differ by income bracket, not by source of income. Presumably there is no reason for an individual to spend a different fraction of the marginal dollar, depending on whether it accrues in the form of wages or dividends. Suits (1963), in a review of the early literature on this question, suggested one possible justification: "... since functional shares vary by income bracket, taking account of functional distribution makes some allowance for the curvature in the consumption function...." That is, distributive shares might be a proxy for the distribution of income by size. This assumes, for example, that an increase in labor's share is reliably associated with an equalization in the size distribution.

How accurate is this assumption? The reader familiar with American income distribution statistics since World War II will be immediately suspicious, since labor's share has steadily increased while most conventional measures of inequality in the size distribution have either been constant or exhibited some upward drift. In point of fact, the division of national income between labor and capital has only a tenuous relation to the size distribution. Table 1 presents the distribution of four components of total income in the United States in 1962. Except in the highest decile the distributions of wages and salaries and of business and property income are not radically different in the sense that knowing whether a given dollar went to "labor" or to "capital" conveys relatively little information

[9] The most recent example of this is Taylor (1971), who cannot reject $a_1 = a_2$ but does find significantly different MPCs out of transfers and other types of income.

about where that dollar landed in the size distribution. In fact, it is not unambiguously clear that "nonlabor income" is distributed more unequally than "labor income," for the Lorenz curves cross. If pensions and annuities are grouped with business and property income, the resemblance is even stronger. Thus, testing whether aggregate consumption is sensitive to the *factor share* distribution is not a fair test of whether aggregate consumption is sensitive to the *size* distribution. This is the first error I set out to correct.

B. *The Definition of Consumption*

The second error is the utilization of a theory of consumption to explain the behavior of *consumer expenditures*.[10] It is quite conceivable that the marginal propensity to *consume* could be falling with rising income while the marginal propensity to *spend on consumer goods and services* could be constant, or conversely.

To distinguish the various concepts of consumption, I introduce the following notation: C = consumption; CE = consumer expenditures, as defined in the national income accounts; CD = expenditures on consumer durables; and UD = use value of the stock of consumer durables, defined as the sum of depreciation plus imputed income. Then the following relation holds: $CE = C + CD - UD$, from which it follows that

$$\frac{\partial^2 CE}{\partial Y^2} = \frac{\partial^2 C}{\partial Y^2} + \frac{\partial^2 CD}{\partial Y^2} - \frac{\partial^2 UD}{\partial Y^2}. \tag{10}$$

It is clear from this equation that a theoretical model which implies $\partial^2 C/\partial Y^2 < 0$ carries no obvious prediction about the sign of $\partial^2 CE/\partial Y^2$; in particular, it is possible for $\partial^2 CE/\partial Y^2$ to be zero or even positive. One objective of the present research was thus to test for distribution effects using aggregate *consumption*, rather than *consumer expenditures*. Fortunately, such a time series (complete with a consistent definition of disposable income)[11] has been constructed by the builders of the MIT–Penn–SSRC (MPS) econometric model.[12]

C. *A Statistical Model Deduced from the Theory*

While the theoretical bases of Friedman's permanent income model and Modigliani-Brumberg's life-cycle model are identical, the empirical

[10] This difficulty is noted by Mayer (1972, pp. 12–16), who nonetheless uses consumer expenditures in his tests.

[11] As Mayer (1972, p. 15) notes, this is important. The few previous studies of the consumption function which used C, rather than CE, as the dependent variable failed to use a consistent concept of disposable income.

[12] It should be noted that the MPS model does not classify residential structures as a consumer durable. Thus neither consumption nor disposable income includes the imputed yield on owner-occupied houses.

formulations differ. I have followed Friedman's model (which expresses consumption as a function of current income and lagged consumption), rather than MBA's (which expresses consumption in terms of current *labor* income and net worth) solely because of data limitations. While there are several annual time series on the distribution of income, there are no time series on the distribution of wealth, and the distribution of labor income can only be guesstimated from the available data. (Example: How do you decompose income of the self-employed into "labor" and "property" components?)

The model of Section II*A* implies that permanent consumption, c^*, is some (nonlinear and complicated) function of permanent income: $c^* = \psi(y^*)$. To make the problem tractable, I assume that $\psi(\cdot)$ is approximately linear *within* each income class. That is, if i is the index of income class, I assume $c_{it}^* = \gamma_i + k_i y_{it}^*$, where the γ_i's and k_i's may depend on interest rates. Allowing for some transitory consumption which is uncorrelated with permanent income, the expression for measured consumption in the ith income group is

$$c_{it} = \gamma_i + k_i y_{it}^* + u_{it}, \tag{11}$$

where u_{it} is the transitory component. Summing over i to obtain aggregate consumption yields:

$$c_t = \sum_i c_{it} = \gamma + \sum_i k_i y_{it}^* + v_t, \tag{12}$$

where

$$\gamma \equiv \sum_i \gamma_i \qquad \text{and} \qquad v_t \equiv \sum_i u_{it}.$$

Were data on permanent income by income class available, (12) could be estimated directly. However, in the absence of such data, it is necessary to have proxies for the permanent income of each group. Following Friedman's suggestion, I assume

$$y_{it}^* = (1 - \lambda_i)[y_{it} + (1 + m_i)\lambda_i y_{i,t-1} + (1 + m_i)^2 \lambda_i^2 y_{i,t-2} + \dots],$$

where m_i is an extraneously estimated growth rate and $0 \leq \lambda_i \leq 1$. Under this assumption, (11) can be expressed as:

$$c_{it} = \gamma_i[1 - (1 + m_i)\lambda_i] + k_i(1 - \lambda_i) y_{it} + (1 + m_i)\lambda_i c_{i,t-1} + \eta_{it}, \tag{13}$$

where $\eta_{it} = u_{it} - (1 + m_i)\lambda_i u_{i,t-1}$. Lacking data on consumption by income class, I am forced to assume that λ is the same for each income group. While m_i could easily be made different for each income class, the actual differences are so trivial that I ignore them. Summing (13) over i with $\lambda_i = \lambda$ and $m_i = m$ leads to

$$C_t = \gamma[1 - (1 + m)\lambda] + (1 - \lambda) \sum_i k_i y_{it} + (1 + m)\lambda C_{t-1} + \eta_t,$$

where

$$\eta_t = \sum_i \eta_{it}.$$

Note that if the model is well specified, the u_{it} in (11) will not display much serial correlation, but η_t will.[13] Also, it seems likely that the u_{it} (and hence η_t) would be heteroskedastic. If the standard deviation of the u_{it} grows proportionately with aggregate disposable income, a more efficient estimating equation would be

$$\frac{C_t}{Y_t} = \frac{\gamma^*}{Y_t} + (1 - \lambda) \sum_i k_i \left(\frac{y_{it}}{Y_t}\right) + (1 + m)\lambda \frac{C_{t-1}}{Y_t} + \varepsilon_t, \quad (14)$$

where $\gamma^* \equiv \gamma[1 - (1 + m)\lambda]$, $\varepsilon_t \equiv \eta_t/Y_t$, and Y_t is aggregate disposable income. In words, (14) requires regressing the average propensity to consume (APC) on the inverse of disposable income, the shares of each income group,[14] and the lagged APC adjusted for growth, $C_{t-1}/Y_t = (C_{t-1}/Y_{t-1})(Y_{t-1}/Y_t)$. If each k_i also depends on the rate of interest, r_t, then interaction terms between r_t and each income share should also be included. Since with five quintile shares this would involve nearly as many coefficients as observations, the constraint that $k_i = w_i + \xi r_t$ is imposed. That is, w, but not ξ, is permitted to vary across income classes.

IV. Empirical Results

A. The Data

The average propensity to consume was obtained by dividing consumption (MPS definition) by disposable income (also MPS definition). Continuous time series (annually from 1947 to 1972) on the shares received

[13] Worse yet, η_t is a first-order moving average rather than an autoregressive process. Strictly speaking this makes standard procedures for dealing with autocorrelation inappropriate. However, Shaman (1969) has shown that the inverse of the covariance matrix of a first-order moving average is best approximated by the inverse covariance matrix of a first-order autoregressive. This provides a pragmatic justification for my use of the Cochrane-Orcutt technique in the subsequent estimation. I am indebted to Michael Hurd for a valuable discussion on this point.

[14] Note that using the Koyck representation of permanent income makes it appropriate to use the shares in *measured*, rather than *permanent*, income for the estimation of distribution effects in (14).

by each quintile of families were obtained from the Bureau of the Census.[15]

The rate of interest posed the most complex measurement problem. What is wanted for a consumption function, presumably, is the *real* opportunity cost of a *consumer*. I first constructed a nominal savings rate, annually from 1949 to 1972, as a weighted average of the rates paid by commercial banks on time deposits, by savings and loan associations, and by mutual savings banks.[16] To obtain a real rate, I then subtracted a proxy for the expected rate of inflation. Following conventional procedures, I assumed an adaptive expectations mechanism:

$$\Pi_t = \theta \frac{\Delta P_t}{P_{t-1}} + (1 - \theta)\Pi_{t-1}, \tag{15}$$

where Π_t is the expected rate of inflation and P_t is the actual price level (the deflator for personal consumption expenditures). To be sure that initial conditions did not influence the series, I started the series by assuming $\Pi_t = \Delta P_t/P_{t-1}$ for $t = 1930$, used the recursion formula (15) to generate Π_t from 1931 forward, and then discarded all the data prior to 1949.[17] The value of θ was chosen from the set [0, 0.1, 0.2, . . . , 1.0] so as to minimize the standard error of each regression. This is approximately equivalent to maximizing the likelihood function over θ, and the optimal value of θ is reported (without a standard error) with each regression.

B. Discussion of a Failure

In brief, the regression is of the form

$$\frac{C_t}{Y_t} = \frac{a_0}{Y_t} + a_4 + a_1F_{1t} + a_2F_{2t} + a_3F_{3t} + a_5F_{5t} + a_6r_t + a_7 \frac{C_{t-1}}{Y_t} + \varepsilon_t, \tag{16}$$

[15] I wish to thank the Bureau of the Census for furnishing me with the data prior to publication. The distributions, of course, are based on the Current Population Survey (CPS) definition of total money income. It should be noted that the shares were tabulated from ungrouped data from 1958 through 1972, but from grouped data from 1947 through 1958. In using these series I averaged the two figures for 1958. See U.S. Bureau of the Census (1973, table 16).

[16] The three constituent rates were obtained from the MPS model data file and are only available from 1949 on. The weights used were 7/16 for time deposits, 6/16 for savings and loan shares, and 3/16 for mutual savings bank shares. I adopted this weighting scheme from Springer (1973).

[17] It is hard to defend the notion that expectations are strictly adaptive in historical episodes that include drastic events such as the Great Depression and World War II. However, my estimated θ is always so high that the influence of 1930–47 price behavior on 1949–72 expectations is negligible.

where F_{it} is the share of total income received by the ith quintile (counting from the bottom) of families. Note that since

$$\sum_{i=1}^{5} F_{it} = 1$$

for all t, one quintile has to be omitted in order to avoid exact multi-collinearity. The choice is arbitrary, and I selected the fourth quintile because it had the least variability.[18]

There are three obvious reasons why I was doomed to failure. First, with only 24 annual observations it is asking a lot of the data to estimate eight coefficients (seven parameters in [16] plus the autocorrelation coefficient). Second, as is well known, the income distribution has been relatively stable since World War II; and this means the F's have very modest variances. Finally, even with one share omitted, considerable collinearity remains, since $F_{1t} + F_{2t} + F_{3t} + F_{5t} = 1 - F_{4t}$, for all t, and F_{4t} is nearly constant through time.

For what it is worth, the regression is reported in column 2 of table 2. A constrained regression, which assumes equal MPCs for all income classes, is presented in column 1 for comparison. Even this latter regression allows for distribution effects of sorts. Recall that the MBF model has distribution effects unless $\delta = \beta$, and, when this equality holds, the consumption function should be strictly proportional. Since the constant (i.e., the coefficient of $1/Y$) in column 1 is significant at the 2 percent level (by a two-tail test), we can reject strict proportionality. On the basis of an extraneous estimate of the growth rate of real disposable income (FMP concept) of 3.83 percent per annum, it is possible to identify the underlying parameters. The implied estimate of λ is 0.24, a rather faster speed of adjustment than found by Friedman. Similarly, the long-run MPC (which, in this equation, is smaller than the APC) is 0.81 evaluated at the mean value of r;[19] this is, of course, lower than Friedman's estimate.

Although there is a hint of distribution effects, column 2 shows that equation (16) does not capture them at all well: the standard error of the regression exceeds that for column 1, and the estimated short-run MPCs are very suspect, ranging from 1.06 to 0.07. Clearly there is too much multicollinearity among the F_i's and too little data to get accurate estimates of subtle differences in MPCs.

One obvious approach is to omit one or more of the shares. This already compromises the theoretical model by imposing the constraint that some

[18] Analogous regressions were run using the shares received by unrelated individuals as well. The results were always similar, but slightly inferior to those reported below.

[19] The estimated short- and long-run MPCs are actually quite insensitive to r. The interest elasticity of consumption (evaluated at the means) in col. 1 is $-.0027$ in the short run and $-.0035$ in the long run. Other equations gave similarly trivial elasticities.

TABLE 2

REGRESSIONS WITH QUINTILE SHARES, 1949–72

VARIABLE	COEFFICIENTS (t-RATIOS) (1)	(2)	(3)*	(4)	(5)
$1/Y$	57.51 (2.6)	46.28 (2.1)	42.80 (2.1)	46.36 (2.2)	44.77 (2.3)
Constant	.616 (16.0)	.066 (0.2)	.113 (0.3)	.130 (0.4)	.264 (1.5)
r	−.0019 (−2.2)	−.0022 (−2.1)	−.0021 (−2.5)	−.0018 (−2.4)	−.0022 (−2.9)
C_{-1}/Y	.253 (3.4)	.290 (3.8)	.312 (4.1)	.312 (3.9)	.295 (4.1)
F_1	...	.692 (0.8)	.796 (1.1)	.837 (1.2)	.515 (1.9)
F_2	...	.168 (0.1)	...	...	...
F_3	...	.997 (1.2)	.908 (1.1)	.837 (1.2)	.515 (1.9)
F_5	...	.721 (1.5)	.645 (1.7)	.621 (1.8)	.515 (1.9)
$\hat{\rho}$	.93 (12.1)	.92 (11.5)	.92 (11.4)	.92 (11.6)	.92 (11.5)
$\hat{\theta}$	0.8	0.8	0.9	1.0	0.9
R^2	.944	.953	.952	.952	.952
SE	.00351	.00364	.00356	.00344	.00336

NOTE.—Estimation was by the Cochrane-Orcutt iterative technique; $\hat{\rho}$ is the estimated autocorrelation coefficient; t-ratios are in parentheses; as the standard errors are only valid under the assumption that θ is known, these ratios are indicative only; θ is the weight given to actual inflation in forming inflationary expectations (see eq. [15]).

* The equation with θ = 0.5 had virtually the same standard error.

quintile has the same MPC as the fourth. For want of a better criterion, I omitted the quintile whose estimated MPC was closest to that of F_4; this led to the regression reported in column 3 of table 2. The resulting standard error of estimate is still larger than in column 1 and autocorrelation is no less severe. As before, computed standard errors for individual shares are very high, suggesting multicollinearity.[20] I then imposed the further constraint that the first and third quintiles have identical MPCs to arrive at the regression in column 4. Here at last the standard error of the regression is reduced below that of the no-distribution-effects case, but the distribution variables are not very important.[21] Finally, using the same criterion, I omitted still another

[20] In view of the way the hypothesis is obtained from the data, these standard errors are purely indicative and cannot be used for hypothesis testing. In particular, the ratio of a coefficient to its "standard error" certainly does not have a t-distribution.

[21] An "F-test" of the null hypothesis that both distribution variables have zero coefficients fails to reject the null hypothesis at the 5 percent level, but just barely. The computed F-value is 3.04, as compared to a critical 5 percent point in the $F(2, 17)$ distribution of 3.59. As noted in n. 20 above, the test statistic does not really have an F-distribution because the restrictions $a_2 = 0$ and $a_1 = a_3$ were obtained from the data. The "test" is meant to be heuristic only.

variable, leaving a single distributional variable in the regression. Column 5 is equation (16) subject to the constraints $a_2 = 0$, $a_1 = a_3 = a_5$.

Of all the regressions reported in table 2, only column 5 can really be said to be an improvement over column 1. The lone distribution variable, $F_1 + F_3 + F_5$, has a coefficient nearly twice as large as its standard error.[22] This equation implies that the short-run MPCs (evaluated at the mean interest rate) are 0.26 for the second and fourth quintiles and 0.78 for the first, third, and fifth. The corresponding long-run MPCs are 0.36 and 1.09. The pattern is not entirely believable but is probably the best job of estimating quintile-specific MPCs that can be done in the absence of quintile-specific consumption data. In any event, there is certainly no indication that MPCs decline in higher income brackets, as is commonly assumed. Precisely what they do is not illuminated very well by table 2.

C. The First Compromise Model

It is clear from these results that some compromise with the theory must be made if any estimation is to be done. And it is not clear that omitting variables (i.e., constraining certain MPCs to be equal) is the ideal compromise. I therefore tried two other approaches which at least have the virtue of allowing every distributional shift to affect aggregate consumption. The first involves constraining the way the MPC varies by income class and is explained in this subsection. The second entails replacing the quintile shares by one or another aggregate index of inequality and is discussed in the following subsection.

The basic model which I would like to estimate is essentially

$$c_{it} = \gamma_i + (k_0^i + k_1^i r_t) y_{it} + \lambda c_{i,t-1} + u_{it}. \tag{17}$$

The problem is that collinearity among the y_{it} precludes accurate estimation of the k_0^i and the k_1^i. Taking a cue from the technique introduced by Almon (1965) to cope with a similar problem in the case of distributed lags, one possibility is to assume a functional form for the dependence of k_0^i and/or k_1^i on i. I report below regressions based on the assumption that both of these coefficients are linear functions of i, but I also ran equations with k's assumed to be either quadratic or logarithmic functions of i. The results were essentially identical. Appending to equation (17) the constraints $k_0^i = m_0 + m_1 i$ and $k_1^i = n_0 + n_1 i$, summing over i, and simplifying leads to

$$C_t = \gamma + m_0 Y_t + n_0 r_t Y_t + m_1 \sum_{i=1}^{5} i y_{it} + n_1 r_t \sum_{i=1}^{5} i y_{it} + \lambda C_{t-1} + V_t,$$

[22] The reader is again reminded of the caveat raised in n. 20 above.

where

$$\gamma \equiv \sum_{i=1}^{5} \gamma_i, \qquad V_t = \sum_{i=1}^{5} u_{it}, \qquad Y_t = \sum_{i=1}^{5} y_{it}.$$

Dividing through by Y_t, and denoting the distributional variable,

$$\sum_{i=1}^{5} i\left(\frac{y_{it}}{Y_t}\right),$$

by D_t, leads to the estimating equation

$$\frac{C_t}{Y_t} = \frac{\gamma}{Y_t} + m_0 + n_0 r_t + m_1 D_t + n_1 r_t D_t + \lambda \frac{C_{t-1}}{Y_t} + v_t, \qquad (18)$$

where $v_t = V_t/Y_t$. Unfortunately, multicollinearity has still not been purged from the equation. Since D_t is relatively constant, r and rD are almost perfectly correlated, so I had to estimate one of two alternative models:

$$\frac{C_t}{Y_t} = \frac{\gamma}{Y_t} + m_0 + n_0 r_t + m_1 D_t + \lambda \frac{C_{t-1}}{Y_t} + v_t \quad (n_1 = 0); \quad (19a)$$

$$\frac{C_t}{Y_t} = \frac{\gamma}{Y_t} + m_0 + n_1 r_t D_t + m_1 D_t + \lambda \frac{C_{t-1}}{Y_t} + v_t \quad (n_0 = 0). \quad (19b)$$

Both variants are reported in table 3. Once again, the expectations parameter in the definition of the real interest rate (θ) was chosen to minimize the standard error; only the results with the optimal choice of θ are given in the table.[23]

It is obvious from table 3 that the choice between (19a) and (19b) is a matter of indifference. The regressions tell a story which is rather similar to that of table 2. Even if the value of θ were known a priori (so that the t-ratios reported in the table were valid), the null hypothesis of no distribution effects (i.e., the null hypothesis that the coefficient of D or rD is zero) could not be rejected at the 10 percent level (two-tail test). But the point estimate suggests that increasing inequality actually increases consumption. To give the reader some feeling for magnitudes, when both r_t and D_t are at their mean values, the predicted short- and long-run MPCs are 0.63 and 0.82, respectively. If D should then rise by 10 percent, these figures would increase to 0.67 and 0.87. While these are not dramatic changes, they are substantial relative to typical year-to-year fluctuations in the observed APC. Were the point estimate of the distributional coefficient more precise, I would be tempted to conclude that

[23] Other results are available on request. In view of the interest in money illusion elicited by Branson and Klevorick's paper (1969), I experimented with an alternative specification using the inverse of *nominal* disposable income in place of the inverse of *real* disposable income. In every case, the real specification gave a better fit, indicating an absence of money illusion.

TABLE 3

REGRESSIONS WITH CONSTRAINED MPCs, 1949–72

VARIABLE	COEFFICIENTS (*t*-RATIOS) (1)	(2)
$1/Y$	49.64	49.51
	(2.6)	(2.6)
Constant	.283	.279
	(1.2)	(1.2)
C_{-1}/Y	.240	.240
	(3.7)	(3.7)
r	−.0028	...
	(−2.8)	
D	.092	.093
	(1.5)	(1.5)
rD	...	−.0007
		(−2.8)
$\hat{\rho}$	.92	.92
	(11.6)	(11.6)
$\hat{\theta}$	0.7	0.7
R^2	.947	.947
SE	.00352	.00352

NOTE.—See general note to table 2.

there are moderate distributional effects which are opposite in direction to those normally assumed. However the large standard error makes this temptation easy to resist.

D. The Second Compromise Model

My second compromise approach is to give up on estimating separate MPCs by income class in favor of estimating the effect of income inequality on aggregate consumption directly, using some conventional measure of inequality such as the Gini ratio or the variance of the logarithms. The great advantage of this approach is, of course, that it saves on degrees of freedom without constraining any particular MPCs to be equal. The disadvantage is that it embodies a weaker "no distribution effects" assumption of its own. For example, by employing the Gini ratio as the measure of inequality, I essentially impose the constraint that all possible redistributions which raise the Gini ratio by 0.01 have the same effect on consumption. Obviously, this need not be true, but it seems more innocuous than assuming away any distribution effects whatever. In any case, I prefer to view this model as a crude approximation to the true model—an approximation dictated by the weakness of the data.

The permanent income hypothesis is generally specified for regression purposes as

$$C_t = \alpha_0 Y_t + \alpha_4 C_{t-1} + v_t. \tag{20}$$

Friedman derived this by assuming that consumption is proportional to permanent income and that permanent income is a Koyck lag on measured

TABLE 4

REGRESSIONS WITH CURRENT INEQUALITY MEASURES

	INEQUALITY MEASURE					
VARIABLE	σ^2 (1)	σ^2_M (2)	σ^2_F (3)	σ^2_{25} (4)	σ^2_{64} (5)	G (6)
$1/Y$	10.18	52.35	47.13	51.30	57.38	−9.48
	(1.3)	(2.2)	(2.0)	(2.1)	(2.5)	(−3.0)
Constant...	.347	.567	.594	.596	.630	.371
	(7.4)	(11.9)	(9.4)	(12.8)	(17.3)	(4.1)
r..........	−.0030	−.0022	−.0022	−.0021	−.0020	−.0035
	(−3.5)	(−2.6)	(−2.6)	(−2.6)	(−2.51)	(−3.5)
C_{-1}/Y	.509	.290	.307	.293	.244	.631
	(7.0)	(3.4)	(3.7)	(3.6)	(3.4)	(10.9)
d	.083	.012	−.014	−.004	.012	.072
	(2.5)	(0.5)	(−0.4)	(−0.1)	(2.1)	(0.3)
$\hat{\rho}$	.19	.92	.92	.93	.94	.14
	(0.9)	(11.0)	(11.0)	(11.4)	(12.8)	(0.6)
$\hat{\theta}$	1.0	0.9	0.9	0.9	0.8	1.0
R^2	.942	.946	.946	.956	.958	.947
SE........	.00381	.00366	.00367	.00338	.00313	.00392

NOTE.—See general note to table 2. The periods of estimation are: for σ^2, σ^2_M, and σ^2_F, 1949–70; for σ^2_{25} and σ^2_{64}, 1949–69; for G, 1949–52, 1954–68.

income. But since this is the general framework in which I have tested for distribution effects throughout, it is worth noting that (20) can arise in other models as well. For example, (20) could represent Brown's (1952) habit-persistence model. Also, a regression very close to (20) could represent Duesenberry's (1949) relative income hypothesis, since, in annual data for the postwar era, "previous peak income" and lagged income are almost always identical.

The pure theory of consumer behavior implies that α_0 should be a function of the rate of interest, r_t, and I simply propose to add the inequality in the size distribution of income, d_t, to the list of arguments. That is, $\alpha_0 = \alpha_1 + \alpha_2 d_t + \alpha_3 r_t$. The null hypothesis to be tested is $\alpha_2 = 0$ against the two-tailed alternative: $\alpha_2 \neq 0$. The regression to be estimated, then, is[24]

$$C_t = \gamma + (\alpha_1 + \alpha_2 d_t + \alpha_3 r_t) Y_t + \alpha_4 C_{t-1} + v_t. \tag{21}$$

[24] Equation (21) cannot be offered as an accurate representation of Friedman's model for the following reason. Adding distribution effects to Friedman's model in the way I have suggested gives $C_t = (k_0 + k_1 d_t + k_2 r_t) Y_t^*$, where Y^* is permanent income. Adopting the Koyck lag for permanent income, as Friedman suggested, gives $Y_t^* - \lambda Y_{t-1}^* = (1 - \lambda) Y_t$, but applying the same Koyck transformation to C_t gives

$$\begin{aligned} C_t - \lambda C_{t-1} &= k_t Y_t^* - \lambda k_{t-1} Y_{t-1}^* \\ &= k_t (Y_t^* - \lambda Y_{t-1}^*) + \lambda Y_{t-1}^* (k_t - k_{t-1}) \\ &= k_{t(1-\lambda)} Y_t + \lambda Y_{t-1}^* (k_1 d_t - k_1 d_{t-1} + k_2 r_t - k_2 r_{t-1}). \end{aligned}$$

Since Y_{t-1}^* is not observable, this equation is not suitable for empirical analysis. Thus eq. (21) was adopted instead. Note that this difficulty would arise even without distribution effects, as long as the rate of interest is allowed to affect the MPC.

TABLE 5

REGRESSIONS WITH LAGGED INEQUALITY MEASURES

	INEQUALITY MEASURE					
VARIABLE	σ^2 (1)	σ^2_M (2)	σ^2_F (3)	σ^2_{25} (4)	σ^2_{64} (5)	G (6)
$1/Y$	16.98	35.27	47.28	−8.94	−9.70	53.24
	(2.0)	(2.2)	(2.2)	(−3.1)	(−4.2)	(2.3)
Constant...	.353	.521	.562	.395	.393	.641
	(8.5)	(10.7)	(8.6)	(7.3)	(8.5)	(7.8)
r..........	−.0021	−.0026	−.0022	−.0032	−.0034	−.0070
	(−2.2)	(−3.2)	(−2.7)	(−3.4)	(−3.8)	(−3.2)
C_{-1}/Y	.457	.340	.311	.580	.588	.220
	(5.6)	(5.5)	(4.0)	(8.0)	(9.3)	(3.4)
d_{-1}	.099	.034	.010	.060	.071	.066
	(3.1)	(1.5)	(0.3)	(1.8)	(2.8)	(0.4)
$\hat{\rho}$	.07	.91	.92	.29	.19	.93
	(0.4)	(10.2)	(11.4)	(1.4)	(0.8)	(10.6)
$\hat{\theta}$	1.0	0.8	0.9	1.0	1.0	0.3
R^2	.940	.948	.947	.928	.940	.961
SE........	.00382	.00353	.00359	.00414	.00376	.00334

NOTE.—See general note to table 2. The periods of estimation are: for σ^2, σ^2_M, and σ^2_F, 1949–71; for σ^2_{25} and σ^2_{64}, 1950–70; for G, 1949–53, 1955–69.

I again correct for heteroskedasticity by dividing (21) through by Y_t and estimate

$$\frac{C_t}{Y_t} = \frac{\gamma}{Y_t} + \alpha_1 + \alpha_2 d_t + \alpha_3 r_t + \alpha_4 \frac{C_{t-1}}{Y_t} + e_t, \tag{22}$$

where $e_t = v_t/Y_t$.

Several time series on overall income inequality, d_t, are available. Since it is by no means clear which measure best captures the relevant distributional shifts, I have run regressions with each of them. The measures are:

G = the Gini concentration ratio of the distribution of money income (CPS concept) among families and unrelated individuals. This series is available for 1948–68 (with 1953 missing) in Budd (1970).

σ^2 = the variance of the logarithms of CPS money income among all persons with income over 14 years of age. This is available over 1947–70, as computed by Schultz (1971).

σ^2_M = same as σ^2, but restricted to males.

σ^2_F = same as σ^2, but restricted to females.

σ^2_{25} = same as σ^2_M, but restricted to men at least 25 years of age. This has been calculated over 1949–69 by Chiswick and Mincer (1972).

σ^2_{64} = same as σ^2_{25}, but excluding men over 65.

Following a suggestion made by Lubell (1947), I tried each variant of d in both current and lagged form. Regressions using current d are reported in table 4, and regressions using lagged d are reported in table 5.

Note that the period of estimation differs somewhat, depending on which variant of d is employed. Because of the paucity of data, I used every available data point rather than confine myself to a common sample period (which would have been 1949–52, 1954–68).

If σ^2, the variance of logarithms over the entire adult population, is used as the inequality measure, it does not matter much whether d_t or d_{t-1} enters the regression. In either the current or the lagged version the impact of inequality is apparently measured with some precision, and an increase of 0.03 in σ^2 (which is a fairly typical year-to-year change) would increase the average propensity to consume by about 0.3 of a percentage point in the short run and about 0.5 of a percentage point in the long run. These two equations are also notable for the absence of autocorrelation (a rare finding in this study) and for the slow speed of adjustment. In fact, inspection of the tables reveals a systematic relationship: the equations with slow adjustment speeds do not have autocorrelated residuals.

The only other measure which is "significant"[25] in both the current and the lagged specification is σ_{64}^2, the log variance among males aged 25–64. The current version exhibits trivially small distribution effects, and the lagged version has much larger ones. Other than these, only the regression with lagged σ_{25}^2 indicates significant distribution effects (at the 10 percent level in a two-tailed test). However, the persistent sign pattern is suggestive. Except for two trivially small coefficients, the point estimates all say that a rise in an inequality index leads to higher consumption. The next section is devoted to interpreting this conclusion and convincing the reader that it is not quite so outlandish as it may seem.

V. Can the Results Be Right?

I began this paper by contrasting what might be called the educated layman's view (that more equal income distributions give rise to more consumption) with the view that is now dominant among macroeconomists (that the income distribution does not matter). The empirical results certainly contradict the layman's view. Instead, they suggest either that consumption is independent of the income distribution or that distributions with less measured inequality give rise to somewhat less consumption. Is the latter possibility believable?[26]

To begin with pure theory, I showed in Section II*A* that—in the optimal life-cycle consumption model—transfers from poor to rich will actually increase consumption if the elasticity of the marginal utility of bequests, β, exceeds the elasticity of the marginal utility of consumption,

[25] As noted, significance tests are valid only on the assumption that θ is known a priori.

[26] To be sure, the arguments I am about to give are not terribly convincing on a priori grounds. They are offered as conceivable explanations of a counter-intuitive result which gets at least mild support from the data.

TABLE 6

POSTWAR CHANGES IN INCOME INEQUALITY

Measure	Initial Value*	Final Value†
σ^2	0.785	1.406
σ^2_M	0.668	1.187
σ^2_F	0.670	1.169
σ^2_{25}	0.742	0.729
σ^2_{64}	0.653	0.581
G	0.424	0.406

* 1947 for σ^2, σ^2_M, and σ^2_F; 1948 for G; and 1949 for σ^2_{25} and σ^2_{64}.
† 1968 for G; 1969 for σ^2_{25} and σ^2_{64}; 1970 for σ^2, σ^2_M, and σ^2_F.

δ. So one interpretation of the data—an interpretation which I do not find particularly appealing—is that the rich actually consume a larger fraction of their lifetime resources because bequests have a wealth elasticity less than unity.[27]

Duesenberry's relative income hypothesis gives an alternative theoretical rationale for the empirical findings. In his model, utility attaches not to consumption but to the *ratio* of own consumption to a weighted average of consumption of others. The weights reflect the frequency of contact with individuals in other consumption classes, and Duesenberry hypothesizes that more contacts with individuals with higher consumption will increase the fraction of income that is consumed. Thus, *it is possible* that an equalization of the income distribution could reduce the number of contacts which most people have with persons much better off than themselves, and therefore reduce aggregate consumption (Duesenberry 1949, pp. 44–45).[28]

A third explanation of the findings, and the one I find most satisfying, rests on the distinction between the kind of "ideal" redistributions that pure theory envisions and the actual redistributions that are reflected in postwar United States data. Table 6 shows the net change over the entire sample period in each of the six measures of income inequality. The variable G is conceptually different from the other variables in two ways. First, it is a Gini ratio, not a log variance. But, more important, it uses families and unrelated individuals (pooled) as the recipient unit, whereas all the others are based on individuals. Thus, the six measures tell the following story. The distributions of income among *families*, and among males over 24 years old, hardly changed in the postwar period; the decline in inequality was very slight. However, inequality fell much more

[27] Becker's (1974) analysis shows that $\delta \leq \beta$ implies that the intergenerational distribution of consumption shifts in favor of the current generation as the economy gets richer. See n. 3 above.

[28] While this is possible, the reverse could also happen, as was pointed out by Johnson (1951).

TABLE 7

FRACTIONS OF ALL PERSONS RECEIVING INCOME, BY AGE-SEX GROUP

Age Group	Males 1948	Males 1972	Females 1948	Females 1972
14–19	3.8	5.7	3.0	4.9
20–24	7.2	6.6	4.3	5.8
25–34	15.4	11.0	6.6	7.4
35–44	14.0	8.9	6.1	6.1
45–54	11.7	9.1	4.8	6.5
55–64	8.9	7.2	3.5	5.8
65 and up	6.6	6.8	4.1	8.3

SOURCE.—Computed by the author from data in U.S. Bureau of the Census (1967, table 14; 1973, table 47).
NOTE.—Totals may not add to 100% as a result of rounding.

noticeably among prime-age males (25–64 years old)[29]—suggesting that the gap between old and prime-age men widened. Further, among all individuals above 14 years of age, inequality rose substantially in the total population, among males, and among females. This suggests that substantial increases in the labor force participation of young people and women may have raised inequality by adding many new income recipients to the lower tail of the distribution.

A more detailed look at these phenomena can be obtained by consulting tables 7–10. Tables 7 and 8 analyze the age-sex composition of the population which underlies each of the income distribution measures other than G, that is, all persons with income. It is clear from table 7 that the distribution included many more teenage boys, and females of all ages except 35–44, in 1972 than it did in 1948. There are two principal reasons for such changes over time: demographic shifts in the age-sex composition of the population as a whole, and changes in labor force participation rates.[30] Table 8 shows that it is the former that accounts for the increased importance of teenage boys; their participation rate actually fell. Similarly, exogenous demographic factors appear to play the major role in the increased numbers of very young and very old women having income. However, for women aged 20–64, it appears that their increased importance in the income distribution is largely attributable to higher labor force participation.

Tables 9 and 10 contain similar information for shares in *income*, rather than in population. Table 9 highlights the increased importance of women over 20 in the overall income distribution. While they received only 17.2 percent of all income in 1948, this share rose to 24.4 percent by

[29] Shultz (1971, p. 11) actually finds a 17 percent *rise* in the log variance of income among males aged 25–64 over the 1947–70 period. But most of this occurs between 1947 and 1949.

[30] This ignores any changes in the fraction of each age-sex group receiving property (but not labor) income.

TABLE 8

LABOR FORCE PARTICIPATION RATES, BY AGE-SEX GROUP

Age Group	Males 1948	Males 1972	Females 1948	Females 1972
14–19	54.2	46.9	32.7	35.8
20–24	85.7	85.9	45.3	59.1
25–34	96.1	95.9	33.2	47.6
35–44	98.0	96.5	36.9	52.0
45–54	95.8	93.3	35.0	53.9
55–64	89.5	80.5	24.3	42.1
65 and up	46.8	24.4	9.1	9.3

SOURCE.—Figures for 14–19 age group were computed by the author from data in U.S. Office of the President (1974, table A-1). Other figures came directly from that source.

TABLE 9

SHARES IN TOTAL MONEY INCOME, BY AGE AND SEX

Age Group	Males 1948	Males 1972	Females 1948	Females 1972
14–19	1.1	1.3	0.9	0.8
20–24	6.0	5.2	2.5	3.0
25–34	19.4	16.7	4.2	5.0
35–44	21.4	17.0	4.0	4.3
45–54	17.2	17.4	3.3	4.8
55–64	11.5	11.7	1.8	3.7
65 and up	5.1	5.7	1.4	3.6

SOURCE.—Computed by the author from data in U.S. Bureau of the Census (1967, tables 14, 36; 1973, table 47).
NOTE.—Totals do not add to 100% as a result of rounding.

TABLE 10

RELATIVE MEAN INCOMES, BY AGE AND SEX*

Age Group	Males 1948	Males 1972	Females 1948	Females 1972
14–19	0.20	0.12	0.20	0.09
20–24	0.54	0.42	0.38	0.27
25–34	0.83	0.79	0.41	0.35
35–44	1.00	1.00	0.42	0.37
45–54	0.95	0.99	0.44	0.39
55–64	0.84	0.85	0.35	0.23
65 and up	0.51	0.44	0.23	0.29

SOURCE.—Same as for table 7.
* Ratio of mean income in each group to mean income of males aged 35–44.

1972. And table 10 shows that this was accomplished despite a widening in the relative income gap between men and women. In fact, it is apparent from table 10 that all groups except elderly women lost ground relative to prime-age males during this period.

In a word, these tables suggest the following anatomy of the rise in *measured* inequality *among persons with income* over the postwar period:

1. Purely demographic forces[31] led to a substantial increase in the number of teenagers and elderly women receiving income. Since these groups generally have low mean incomes (and, in the case of teenagers, also relatively high within-group variance), this alone would tend to raise inequality. Compounding this, the mean income of teenagers relative to prime-age males fell so sharply that their share in total income actually declined.[32]

2. Marked increases in labor force participation of women between 20 and 64 years of age added a great number of relatively low incomes to the distribution, thus raising inequality as conventionally measured. This alleged rise in inequality strikes me as particularly illusory, since it is attributable to an artifact in construction of the data, that is, to defining the population as "all persons with income" rather than "all persons." Presumably, if all the "zeros" in both years had been included in the income distribution, the growth of the female labor force would have *reduced* measured inequality.

3. Incomes of males aged 20–24 and over 64, as well as all but the oldest women, declined relative to prime-age males. This represents a bona fide increase in inequality by most reasonable criteria, but it is very far afield from the idealized "regressive transfers" discussed in Section II*B*.

If these were the underlying changes in the income distribution, the regressions in Section IV*D* could be associating greater inequality with greater consumption if any of the following are true: (*a*) elderly women have higher-than-average marginal propensities to consume, (*b*) women have greater MPCs than men, or (*c*) prime-age males have lower-than-average MPCs.

Of these possibilities, the theory of optimal life-cycle consumption gives every reason to believe that (*a*) and (*c*) would be true. Given a typical "humped" income profile, consumer units will dissave while very young and very old, and save during the prime earning years. However, there seems to be no theoretical reason to believe in the veracity of (*b*).[33] Indeed, the theory of the household gives every reason to believe that the consumption behavior of married women and married men should be

[31] This phrase is not meant to deny that there may have been economic reasons behind these phenomena. It simply connotes that the variable itself (age-sex distribution) is a demographic one rather than an economic one (like labor-force participation rates).

[32] The drop in relative incomes may well have been caused by the increase in relative labor supplies, but that is beyond the scope of this paper.

[33] This, of course, does not prove that (*b*) is false. It would be true, for example, if wives typically entered the labor force to finance the acquisition of specific consumer durables which the family wished to purchase.

essentially identical.[34] And it is the married women (in particular, those married with spouse present) who have registered the greatest gains in labor force participation over this period.[35]

Finally, the only other study known to me which included the *size* distribution of income in the consumption function also obtained the "odd" result that increased inequality led to increased consumption. Metcalf (1972) was rather puzzled by the finding and noted that "while a number of significant relationships were uncovered, it is not yet clear how the results should be interpreted" (p. 148–49). The distributional variable in his preferred consumption function is the ratio of income at the ninetieth percentile to mean income, and he concluded that "the higher the top decile income relative to the mean, the higher the marginal propensity to consume" (p. 152).[36]

VI. Summary

In this paper I have shown that the established theory of consumer behavior carries definite implications as to the effect of a sequence of regressive transfers on aggregate consumption. Such an increase in inequality must reduce *consumption* if bequests are a luxury good (or increase consumption if own consumption is the luxury good). However, this does *not* say that such a redistribution would necessarily reduce *consumer expenditures*. Nor does it say that aggregate consumption must fall as a result of the kinds of redistributions that have taken place in the post-war United States. Finally, the theory (and the facts) give no reason to believe that a shift in the *factor share* distribution will have any particular effect upon consumption.

The only rigorously correct way to test for the existence of distribution

[34] In March 1972, 58.4 percent of the females in the labor force were married with a husband present. The remainder included single, widowed, divorced, and separated women (see U.S. Office of the President 1974, table B-1).

[35] Labor force participation among "married, spouse present" women rose from 22 percent in 1948 to 41.5 percent in 1972, while participation rates were virtually trendless for other categories of women. (See U.S. Office of the President 1974, table B-2.)

[36] A further possibility, which I regard more as an intellectual curiosum than as a practical explanation of the results, is that the log variance could conceivably *increase* while inequality is actually falling by the mean-preserving spread criterion. To show this, I follow Atkinson's (1970) approach to inequality measurement, which assumes an additive social welfare function, $W = \int u(y) f(y)\, dy$, where $u(y)$ is the social welfare significance of a person's receiving income y, and $f(y)$ is the income density function. The specific utility function implicit in using the variance of logarithms, $V = \int (\log y - k)^2 f(y)\, dy$, where $k \equiv E(\log y)$, as the inequality measure is clearly $u(y) = (\log y - k)^2$. As first noted by Atkinson (1970, p. 13), this function is not concave over its entire range; therefore a sequence of *regressive* transfers might actually *raise* social welfare (Rothschild and Stiglitz 1973). Since five of the six inequality measures are variances of logarithms, they may not be correct indicators of the direction of change in inequality.

effects in the aggregate consumption function is to estimate directly separate marginal propensities to consume by income class. Unfortunately, the data are too weak to allow this, so several "second best" procedures were adopted. First, various MPCs were constrained to be equal to one another; then a method similar to the Almon lag technique was used to constrain the variation in MPCs; finally, the MPCs were ignored and an aggregate measure of inequality was inserted in the consumption function. The upshot of all this appears to be that equalizing the income distribution will either have no bearing on or (slightly) reduce aggregate consumption.

Several reasons for the latter possibility were suggested. Of these, I find two most appealing. First, if the kinds of "demonstration effects" stressed by Duesenberry are at all important, disequalization can conceivably lead to more rather than less consumption. Second, income inequality in the postwar United States increased largely because of demographic shifts and increased labor force participation of women. Had these women been counted as "zeros" in the income distribution while out of the labor force, rather than omitted, income inequality might well have fallen. Thus the observed positive effect of certain inequality measures on consumption may be misleading.

As a by-product, this study has shed some additional light on other properties of the consumption function. In all specifications, I find an absence of money illusion, a very small negative interest elasticity, and a rather fast adjustment of inflationary expectations to actual inflation.

Appendix

Solution of the Optimal Consumption Problem with a Bequest Motive

The problem is to pick a time pattern of consumption, $c(t)$, and a level of terminal assets, $K(T)$, so as to maximize

$$\int_0^T U[c(t)]e^{-\rho t}\,dt + B[K(T)], \tag{A1}$$

subject to the lifetime budget constraint

$$\int_0^T c(t)e^{-rt}\,dt + K(T)e^{-rT} = W. \tag{A2}$$

Defining the functional,

$$L[c(t), K(T)] = \int_0^T U[c(t)]e^{-\rho t}\,dt + B[K(T)] + \lambda\left[W - \int_0^T c(t)e^{-rt}\,dt - K(T)e^{-rT}\right],$$

the first-order conditions are[37]

$$\frac{\partial L}{\partial c(t)} = U'[c(t)]e^{-\rho t} - \lambda e^{-rt} = 0 \text{ for all } t; \tag{A3}$$

$$\frac{\partial L}{\partial K(T)} = B'[K(T)] - \lambda e^{-rt} = 0. \tag{A4}$$

Equation (A4), of course, is simply the transversality condition, since $\lambda e^{-rT} = U'[c(T)]e^{-\rho T}$ by (A3). Solving (A3) under the specific functional form $U[c(t)] = [c(t)^{1-\delta}]/(1 - \delta)$ gives

$$c(t)^{-\delta} = \lambda e^{-(r-\rho)t}, \tag{A5}$$

from which it follows that λ is related to the initial level of consumption by

$$c(0)^{-\delta} = \lambda, \qquad \text{or } c(0) = \lambda^{-1/\delta}. \tag{A6}$$

Therefore (A5) becomes

$$c(t) = \lambda^{-1/\delta}e^{[(r-\rho)/\delta]t} = c(0)e^{[r-\rho)/\delta]t}$$

by (A6), which is equation (3) in the text.

Now use (A6) and the specific functional form

$$B[K(T)] = \frac{bK(T)^{1-\beta}}{1-\beta}$$

to express (A4) as $bK_T^{-\beta} = c(0)^{-\delta}e^{-rT}$, or $K_T = (be^{rT})^{1/\beta}c(0)^{\delta/\beta}$, which is equation (5) in the text.

Finally, return to the budget constraint, equation (A2), to write

$$\begin{aligned} W - K(T)e^{-rT} &= \int_0^T c(t)e^{-rt}\,dt \\ &= c(0)\int_0^T e^{[(r-\rho)/\delta]t}e^{-rt}\,dt \\ &= c(0)\int_0^T e^{\{[r(1-\delta)-\rho]/\delta\}t}\,dt, \end{aligned}$$

which is equation (4) in the text with

$$\phi(r, \rho, \delta, T) \equiv \left[\int_0^T e^{\{[r(1-\delta)-\rho]/\delta\}t}\,dt\right]^{-1}.$$

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[37] Assuming $U(\cdot)$ and $B(\cdot)$ are strictly concave, these are also sufficient. Also the assumptions

$$\lim_{c(t)\to 0} U'[c(t)] = \infty \qquad \text{and} \qquad \lim_{K(T)\to 0} B'[K(t)] = \infty$$

rule out the possibility of corner solutions.

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