# DISTRIBUTION EFFECTS AND THE AGGREGATE CONSUMPTION FUNCTION*

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## 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, based presumably on Keynes' own intuition, that it does -- in particular that equalization of the income distribution would increase consumption.  With the publication of Kuznets' [1942] and Goldsmith's [1955] data, and the ascendancy of the Friedman [1957] and Modigliani-Brumberg [1954] models of consumer 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 that 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 resdistribution 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, total 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 psotwar 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

### II.1.  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 wide-
spread 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:

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

where  $r$  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.  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:

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

with the further stipulation that $\delta = \beta$ .[1] Thus (2) represents a minor
generalization of the Modigliani-Brumberg-Friedman (henceforth MBF) model
-- a generalization with important consequences for the question of distri-
bution effects.

The maximization of (2) subject to (1) is a well-known problem
which was first studied by Strotz [1955-56]. One simple method of solution
is given in the appendix. For present purposes, it suffices to note that
the optimal plan is given implicitly by the equations:

(3) $\qquad c(t) = c_o e^{gt}$ $\qquad\qquad$ where $\quad g \equiv (r-\rho)/\delta$

(4) $\qquad c_o = \phi(r,\rho,\delta,T) \ [W - K_T e^{-rT}]$

(5) $\qquad K_T = (b \ e^{rT})^{1/\beta} \ c_o^{\ \delta/\beta}$

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 (1) - (3) under two sets of circumstances:

(a) $b = 0$ : This is the strict life-cycle model of
Modigliani-Brumberg-Ando. 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, $c_o$
(and hence all $c(t)$) will be proportional to $W$ . Speci-
fically, $c(t) = \phi e^{gt} W$ .

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

(b) $\delta = \beta$ : This is a modification mentioned by Ando and Modigliani [1960]. In this case $K_T$ is proportional to $c_o$ by (5), so that (3) - (5) can be solved to yield:

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

It is important to note that these are the <u>only</u> two cases which give rise to strict proportionality, i.e., a constant lifetime 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:

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

where $c^* \equiv \int_o^T c(t) e^{-rt} dt$ , 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 $\beta$ , the elasticity of the marginal utility of bequests, exceeds $\delta$ , the elasticity of the marginal utility of consumption, then consumption is the luxury good, i.e., 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; but, in the absence of any convincing empirical evidence, each reader is free to make up his own mind. My only purpose is to establish that it is possible, within the basic MBF model, to have an MPC which either rises or falls with income.

## II.2.  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:

$$c = c(y)$$

$$1 > c'(y) > 0$$

$$c''(y) \gtreqless 0 \qquad \text{according as} \quad \delta \lesseqgtr \beta \; .$$

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  $\mu$ ,  a  and  b  denote respectively the average, lowest and highest permanent income in the population.

It is easily established that:[2]

$$(6) \qquad \mu = b - \int_a^b F(y,d) \, dy \; .$$

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[2]Proof:  By definition  $\mu = \int_a^b yf(y,d)dy$ .  Integrating this by parts

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

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

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") 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:

(7a) $\qquad \dfrac{\partial a}{\partial d} = \dfrac{\partial b}{\partial d} = 0$ ;

(7b) $\qquad F_d(y,d)$  is continuous on the interval  $a \leq y \leq b$ ;

(7c) $\qquad$ there is some  y*  in the interval  (a,b)  such that:

$$F_d(y,d) \geq 0 \qquad\qquad \text{for} \qquad\qquad a \leq y \leq y^* \text{ , and}$$

$$F_d(y,d) \leq 0 \qquad\qquad \text{for} \qquad\qquad y^* \leq y \leq b \text{ ;}$$

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

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

With the preliminaries thus established, the proof is quite direct.[3]  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:

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

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

Integrating this by parts yields:[4]

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

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 $\frac{\partial C}{\partial d}$ must be negative. Conversely, if $c'(y)$ were a rising function of $y$, $\frac{\partial C}{\partial d}$ would be positive. Of course, if $c'(y)$ is constant, (7d) immediately implies that $\frac{\partial C}{\partial d} = 0$. In words, I have established:

> PROPOSITION D: A mean-preserving spread in the income distribution will decrease, leave unchanged, or increase aggregate consumption according as $\delta$ is greater than, equal to, or less than $\beta$.

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

It is worth pausing at this junction to consider what has <u>not</u> been proved. Proposition D refers only to transfers within a group which is identical in all relevant respects save income. It is not applicable to transfers where the age distribution of the givers differs from the age

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

distribution of the recipients. Nor is it 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. The practical implication of this, of course, is that Proposition D gives no basis for predicting the effect on aggregate consumption of most real-world redistributions.[5] That is, it certainly does _not_ establish Proposition B.

## III. Testing for Distribution Effects

### III.1. 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:

$$(9) \qquad C_t = aY_t + bC_{t-1} + u_t \;,$$

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:

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[5] 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).

$$C_t = a_1 L_t + a_2 P_t + b C_{t-1} + u_t ,$$

and test $a_1 = a_2$ against the alternative $a_1 > a_2$ . Generally the null hypothesis of no distribution effects cannot be rejected.[6]

What is the rationale for this test? The theory outlined in Section II suggests that MPC's 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. The distributions

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[6]The most recent example of this is Taylor [1971], who cannot reject $a_1 = a_2$ but does find significantly different MPC's out of transfers and other types of income.

of "Wages and Salaries" and "Business and Property Income" are not radi-
cally different in the sense that knowing whether a given dollar went to
"labor" or to "capital" conveys relatively little information about where
that dollar landed in the size distribution.  In fact, it is not unambigu-
ously 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 or not 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.

## Table 1

### U.S. Income Distribution in 1962 by Components

Percentage Share in:

| Decile Group | Wages and Salaries | Business and Property Income | Pensions & 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).

### III.2. The Definition of "Consumption"

The second error is the utilization of a theory of consumption to explain the behavior of consumer expenditures.[7] 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

UD = use value of the stock of consumer durables, defined as the sum of depreciation plus imputed income

Then the following relation holds:

$$CE = C + CD - UD .$$

So long as the use value of consumer durables is not affected by current disposable income, the marginal propensity to spend on consumer goods and services is related to the theoretical marginal propensity to consume by:

$$\frac{\partial CE}{\partial Y} = \frac{\partial C}{\partial Y} + \frac{\partial CD}{\partial Y} ,$$

from which it follows that:

$$(10) \qquad \frac{\partial^2 CE}{\partial Y^2} = \frac{\partial^2 C}{\partial Y^2} + \frac{\partial^2 CD}{\partial Y^2} .$$

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

It is clear from this equation that a theoretical model which

implies $\frac{\partial^2 C}{\partial Y^2} < 0$ carries no necessary prediction about the sign of $\frac{\partial^2 CE}{\partial Y^2}$ ;

in particular, it is possible for $\frac{\partial^2 CE}{\partial Y^2}$ 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)[8] has been constructed by the builders of the Federal Reserve-MIT-

Penn (FMP) econometric model.[9]


### III.3. A Statistical Model Deduced from the Theory

While the theoretical bases of Friedman's permanent income model

and Modigliani-Brumberg life-cycle model are identical, the empirical

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

od you decompose income of the self-employed into "labor" and "property"

components?).

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[8] 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.

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

The model of section II.1 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,

$$(11) \qquad c^*_{it} = \gamma_i + k_i y^*_{it} \, ,$$

where the $\gamma_i$'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 $i^{th}$ income group is:

$$c_{it} = \gamma_i + k_i y^*_{it} + u_{it}$$

where $u_{it}$ is the transitory component.

Summing over $i$ to obtain aggregate consumption yields:

$$(12) \qquad c_t \equiv \sum_i c_{it} = \gamma + \sum_i k_i y^*_{it} + v_t$$

where
$$\gamma \equiv \sum_i \gamma_i \quad \text{and} \quad v_t \equiv \sum_i u_{it} \, .$$

To implement (12) empirically, it is necessary to have proxies for the permanent income of each group. Friedman suggests using:

$$(13) \qquad y^*_{it} = (1-\lambda)[y_{it} + (1+m)\lambda \, y_{i,t-1} + (1+m)^2\lambda^2 y_{i,t-2} + \ldots]$$

where the m is an exogenously estimated growth rate and $0 \leq \lambda \leq 1$.

Note that $\lambda$ is assumed to be the same for each income group; this may be false, but the existing data make it impossible to test for different $\lambda$'s. While m could easily be made different for each income class, the actual differences are so trivial that I ignore them. Applying the Koyck transformation to (12) and (13) 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

$$\eta_t \equiv v_t - (1+m)\lambda \, v_{t-1} .$$

Note that if the model is well-specified, $v_t$ will be serially uncorrelated, but $\eta_t$ will be serially correlated. Also, it seems likely that $v_t$ (and hence $\eta_t$) would be heteroskedastic. If the standard deviation of $v_t$ grows proportionately with aggregate disposable income, a more efficient estimating equation would be:

$$(14) \qquad \frac{C_t}{Y_t} = \frac{\gamma^*}{Y_t} + (1-\lambda) \sum_i k_i (\frac{Y_{it}}{Y_t}) + (1+m)\lambda \frac{C_{t-1}}{Y_t} + \epsilon_t$$

where $\gamma^* \equiv \gamma(1 - (1+m)\lambda)$, $\epsilon_t \equiv \eta_t/Y_t$, and $Y_t$ is aggregate disposable income. In words, (14) requires regressing the average propensity to consume on the shares of each income group and the lagged APC adjusted for growth, $\frac{C_{t-1}}{Y_t} = (\frac{C_{t-1}}{Y_{t-1}}) (\frac{Y_{t-1}}{Y_t})$. 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

$k_i = w_i + \xi r_t$ was imposed. That is, $w$, but not $\xi$, is permitted

to vary across income classes.

## IV. Empirical Results

### IV.1. The Data

The average propensity to consume was obtained by dividing

consumption (FMP definition) by disposable income (also FMP definition).

Continuous time series (annually from 1947 to 1972) on the shares received

by each quintile of families were obtained from the Bureau of the Census.[10]

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 commericial banks on time deposits, by savings and loan associations,

and by mutual savings banks.[11] 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:

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

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[10] 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).

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

where $\Pi_t$ 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 unduly influence the series, I

started the series by assuming $\Pi_t = \dfrac{\Delta P_t}{P_{t-1}}$ for $t = 1930$, used the recursion

formula (15) to generate $\Pi_t$ from 1931 forward, and then discarded all the

data prior to 1949. The value of $\theta$ was chosen from the set $\{0, 0.1,$

$0.2, \ldots, 1.0\}$ so as to minimize the standard error of each regression.

This is approximately equivalent to maximizing the likelihood function

over $\theta$ , and the optimal value of $\theta$ is reported (without a standard

error) with each regression.

### IV.2. Discussion of a Failure

In brief, the regression is of the form:

$$(16) \qquad \frac{C_t}{Y_t} = \frac{a_o}{Y_t} + a_4 + a_1 F_{1t} + a_2 F_{2t} + a_3 F_{3t} + a_5 F_{5t} + a_6 r_t + a_7 \frac{C_{t-1}}{Y_t} + \epsilon_t$$

where $F_{it}$ is the share of total income received by the i-th quintile

(counting from the bottom) of families. Note that since $\sum\limits_{i=1}^{5} F_{it} = 1$ 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.[12]

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

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[12] 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.

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  $F_{1t} + F_{2t} + F_{3t} + F_{5t} = F_{4t}$ ,
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 1.  A constrained regression, which assumes equal MPC's 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) is significant at the
2% level (by a two-tail test), we can reject strict proportionality.
Based on an extraneous estimate of the growth rate of disposable income
(FMP concept) of 5.68% per annum (obtained by regressing the logarithm  of
Y  on time), it is possible to identify the underlying parameters.  The
implied estimate of  $\lambda$  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 0.82 evaluated at the mean value of  r ;[13] 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

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[13]The estimated short and long-run MPC is actually quite insensitive to  r .

TABLE 1

Regressions with Quintile Shares, 1949-1972[a]

Coefficients
(t ratios)

| Variable | (1) | (2) | (3)[b] | (4) | (5) |
|---|---|---|---|---|---|
| $1/Y$ | 57.51 | 46.28 | 42.80 | 46.36 | 44.77 |
| | (2.6) | (2.1) | (2.1) | (2.2) | (2.3) |
| Constant | .616 | .066 | .113 | .130 | .264 |
| | (16.0) | (.2) | (.3) | (.4) | (1.5) |
| $r$ | −.0019 | −.0022 | −.0021 | −.0018 | −.0022 |
| | (−2.2) | (−2.1) | (−2.5) | (−2.4) | (−2.9) |
| $C_{-1}/Y$ | .253 | .290 | .312 | .312 | .295 |
| | (3.4) | (3.8) | (4.1) | (3.9) | (4.1) |
| $F_1$ | − | .692 | .796 | .837 | .515 |
| | | (.8) | (1.1) | (1.2) | (1.9) |
| $F_2$ | − | .168 | − | − | − |
| | | (.1) | | | |
| $F_3$ | − | .997 | .908 | .837 | .515 |
| | | (1.2) | (1.1) | (1.2) | (1.9) |
| $F_5$ | − | .721 | .645 | .621 | .515 |
| | | (1.5) | (1.7) | (1.8) | (1.9) |
| $\hat{\rho}$ | .93 | .92 | .92 | .92 | .92 |
| | (12.1) | (11.5) | (11.4) | (11.6) | (11.5) |
| $\hat{\theta}$ | 0.8 | 0.8 | 0.9 | 1.0 | 0.9 |
| $R^2$ | .944 | .953 | .952 | .952 | .952 |
| Standard error | .00351 | .00364 | .00356 | .00344 | .00336 |
| Durbin-Watson | 1.57 | 1.52 | 1.50 | 1.51 | 1.55 |

[a] Estimation was by the Conchrane-Orcutt iterative technique. $\hat{\rho}$ is the estimated autocorrelation coefficient. t ratios are in parentheses. $\theta$ is the weight given to actual inflation in forming inflationary expectations.

[b] the equation with $\theta = .5$ had virtually the same standard error.

error of the regression exceeds that for column (1), and the estimated

short-run MPC's 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 MPC's.

One obvious approach is to omit one or more of the shares.

This already compromises the theoretical model by imposing the constraint

that some 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 1.

The resulting standard error is still larger than in column (1) and

autocorrelation is no less severe.  As before, t-ratios on the individual

shares are very low, suggesting multicollinearity.  I then imposed the

further constraint that the first and third quintiles have identical

MPC's 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 significant ($F_5$

is just significant at the 10% level by a two-tailed test).  An F-test

of the null hypothesis that both distribution variables have zero

coefficients fails to reject the null hypothesis at the 5% level, but

just barely.[14]  Finally, using the same criterion, I omitted still another

variable, leaving a single distributional variable in the regression.

Column (5) is equation (16) subject to the constraints $a_2 = 0$ , $a_1 = a_3 = a_5$ .

_____

[14]The computed F value is 3.04, as compared to a critical 5% point in the
F(2,17) distribution of 3.59.

Of all the regressions reported in Table 1, only column (5) can really be said to be an improvement over column (1). The lone distribution variable, $F_1 + F_3 + F_5$, is significant at the 10% level by a two-tailed test. This equation implies that the short-run MPC's (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 MPC's are 0.37 and 1.16. The pattern is not entirely believable, but is probably the best job of estimating quintile-specific MPC's that can be done in the absence of quintile-specific consumption data. In any event, there is certainly no indication that MPC's decline in higher income brackets as is commonly assumed. Precisely what they do is not illuminated very well by Table 1.

### IV.3. 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 MPC's 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:

$$(17) \qquad C_{it} = \gamma_i + (k_o^i + k_1^i r_t) Y_{it} + \lambda C_{i, t-1} + u_{it} \ .$$

The problem is that collinearity among the $Y_{it}$ precludes accurate estimation of the $k_o^i$ 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 $k_o^i$ 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 the $k$'s assumed to be either quadratic or logarithmic functions of $i$. The results were essentially identical. Appending to equation (17) the constraints:

$$k_o^i = m_o + m_1 i$$

$$k_1^i = n_o + n_1 i ,$$

summing over $i$, and simplifying leads to:

$$C_t = \gamma + m_o Y_t + n_o r_t Y_t + m_1 \sum_{i=1}^{5} i Y_{it} + n_1 r_t \sum_{i=1}^{5} i Y_{it} + \lambda C_{t-1} + V_t$$

Where $\gamma \equiv \sum_{i=1}^{5} \gamma_i$ , $V_t \equiv \sum_{i=1}^{5} u_{it}$ , $Y_t \equiv \sum_{i=1}^{5} Y_{it}$ . Dividing through by $Y_t$ ,

and denoting the distributional variable, $\sum_{i=1}^{5} i(\frac{Y_{it}}{Y_t})$ , by $D_t$ , leads

to the estimating equation:

$$(18) \quad \frac{C_t}{Y_t} = \frac{\gamma}{Y_t} + m_o + n_o r_t + m_1 D_t + n_1 r_t D_t + \lambda \frac{C_{t-1}}{Y_t} + v_t .$$

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:

(19a)
$$\frac{C_t}{Y_t} = \frac{Y}{Yt} + m_o + n_o r_t + m_1 D_t + \lambda \frac{C_{t-1}}{Y_t} + v_t \qquad (n_1 = 0)$$

(19b)
$$\frac{C_t}{Y_t} = \frac{Y}{Y_t} + m_o + n_1 r_t D_t + m_1 D_t + \lambda \frac{C_{t-1}}{Y_t} + v_t \qquad (n_o = 0) \quad .$$

Both variants are reported in Table 2 below. Once again, the expectations parameter in the definition of the real interest rate $(\theta)$ was chosen to minimize the standard error; only the results with the optimal choice of $\theta$ are given in the table.[15]

It is obvious from Table 2 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 1. The null hypothesis of no distribution effects (i.e., the null hypothesis that the coefficient of $D$ or $rD$ is zero) cannot be rejected at the 10% level (two-tail test), but there is a hint 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 MPC's are 0.63 and 0.83 respectively. If $D$ should then rise by 10%, these figures would increase

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[15] 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 <u>nominal</u> disposable income in place of the inverse of <u>real</u> disposable income. In every case, the real specification gave a better fit, indicating an absence of money illusion.

## TABLE 2

### Regressions with Constrained MPC's, 1949-1972

Coefficients[a]

(t-ratios)

| Variable | (1) | (2) |
|---|---|---|
| $1/Y$ | 49.64 | 49.51 |
| | (2.6) | (2.3) |
| Constant | .283 | .279 |
| | (1.2) | (1.2) |
| $C_{-1}/Y$ | .240 | .240 |
| | (3.7) | (3.7) |
| $r$ | -.0028 | -.0007 |
| | (-2.8) | (-2.8) |
| $D$ | .092 | - |
| | (1.5) | |
| $rD$ | - | .093 |
| | | (1.5) |
| $\hat{\rho}$ | .92 | .92 |
| | (11.6) | (11.6) |
| $\hat{\theta}$ | 0.7 | 0.7 |
| $R^2$ | .947 | .947 |
| Standard error | .00352 | .00352 |
| Durbin-Watson | 1.62 | 1.62 |

[a]The fact that the coefficient of D in column (1) and
of rD in column (2) are virtually the same is a
coincidence attributable to the fact that the mean
value of r is nearly 1.0.

to 0.67 and 0.88.  While these are not dramatic changes, they are rather

large compared 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 there are substantial distributional

effects which are opposite in direction from those normally assumed.  How-

ever the low  t  ratio makes this temptation easy to resist.


### IV.4.  The Second Compromise Model

My second compromise approach is to give up on estimating separate

MPC's  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  MPC's  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 innocu-

ous 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:

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

Friedman derived this by assuming that consumption is proportional to permanent income and that permanent income is a Koyck lag on measured 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 rperesent 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 $\alpha_0$ 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:[16]

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[16] 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:

$$Y^*_t - \gamma \, Y^*_{t-1} = (1-\lambda) Y_t ,$$

but applying the same Koyck transformation to $C_t$ gives:

$$
\begin{aligned}
C_t - \gamma C_{t-1} &= k_t Y^*_t - \gamma \, k_{t-1} \, Y^*_{t-1} \\
&= k_t [Y^*_t - \gamma \, Y^*_{t-1}] + \gamma \, Y^*_{t-1} [k_t - k_{t-1}] \\
&= k_t Y_t + \gamma \, Y^*_{t-1} [k_1 d_t - k_1 d_{t-1} + k_2 r_t - k_2 r_{t-1}] .
\end{aligned}
$$

Since $Y^*_{t-1}$ is not observable, this equation is not suitable for empirical analysis. Thus equation (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 .

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

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

estimate:

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

where  $e_t = \dfrac{v_t}{Y_t}$ .

Several time series on overall income inequality, $d_t$ , are avail-

able. Since it is by no means clear which measure best captures the rele-

vant 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 from 1948-1968 (with 1953 missing) in Budd [1970].

$\sigma^2$: The variance of the logarithms of CPS money income among all persons with income over 14 years of wage. This is available over 1947-1970, as computed by Schultz [1971].

$\sigma^2_M$: Same as $\sigma^2$ , but restricted to males.

$\sigma^2_F$: Same as $\sigma^2$ , but restricted to females.

$\sigma^2_{25}$: Same as $\sigma^2_M$ , but restricted to men at least 25 years of age. This has been calculated over 1949-69 by Chiswick and Mincer [1972].

$\sigma^2_{64}$: Same as $\sigma^2_{25}$ , but excluding men over 65.

Following a suggestion made by Lubell [1947], each variant of  d  was tried

in both current and lagged form. Regressions using current  d  are reported

in Table 3; and regressions using lagged  d  are reported in Table 4.

TABLE 3

Regressions with Current Inequality Measures[a]

| Variable | Inequality Measure | (1) $\sigma^2$ | (2) $\sigma^2_M$ | (3) $\sigma^2_F$ | (4) $\sigma^2_{25}$ | (5) $\sigma^2_{64}$ | (6) G |
|---|---|---|---|---|---|---|---|
| $1/Y$ | | 10.18 (1.3) | 52.35 (2.2) | 47.13 (2.0) | 51.30 (2.1) | 57.38 (2.5) | −9.48 (−3.0) |
| Constant | | .347 (7.4) | .567 (11.9) | .594 (9.4) | .596 (12.8) | .630 (17.3) | .371 (4.1) |
| $r$ | | −.0030 (−3.5) | −.0022 (−2.6) | −.0022 (−2.6) | −.0021 (−2.6) | −.0020 (−2.51) | −.0035 (−3.5) |
| $C_{-1}/Y$ | | .509 (7.0) | .290 (3.4) | .307 (3.7) | .293 (3.6) | .244 (3.4) | .631 (10.9) |
| $d$ | | .083 (2.5) | .012 (0.5) | −.014 (−0.4) | −.004 (−0.1) | .012 (2.1) | .072 (0.3) |
| $\hat{\rho}$ | | .19 (.9) | .92 (11.0) | .92 (11.0) | .93 (11.4) | .94 (12.8) | .14 (0.6) |
| $\hat{\theta}$ | | 1.0 | 0.9 | 0.9 | 0.9 | 0.8 | 1.0 |
| $R^2$ | | .942 | .946 | .946 | .956 | .958 | .947 |
| Std. error | | .00381 | .00366 | .00367 | .00338 | .00313 | .00392 |
| Durbin Watson | | 1.69 | 1.47 | 1.41 | 1.22 | 1.21 | 1.83 |

[a]The periods of estimation are: for $\sigma^2$, $\sigma^2_M$ and $\sigma^2_F$, 1949-1970; for $\sigma^2_{25}$ and $\sigma^2_{64}$, 1949-1969; for G, 1949-1952, 1954-1968.

TABLE 4

Regressions with Lagged Inequality Measures[a]

| Variable \ Inequality Measure | (1) $\sigma^2$ | (2) $\sigma^2_M$ | (3) $\sigma^2_F$ | (4) $\sigma^2_{25}$ | (5) $\sigma^2_{64}$ | (6) $G$ |
|---|---|---|---|---|---|---|
| $1/Y$ | 16.98 (2.0) | 35.27 (2.2) | 47.28 (2.2) | -8.94 (-3.1) | -9.70 (-4.2) | 53.24 (2.3) |
| Constant | .353 (8.5) | .521 (10.7) | .562 (8.6) | .395 (7.3) | .393 (8.5) | .641 (7.8) |
| $r$ | -.0021 (-2.2) | -.0026 (-3.2) | -.0022 (-2.7) | -.0032 (-3.4) | -.0034 (-3.8) | -.0070 (-3.2) |
| $C_{-1}/Y$ | .457 (5.6) | .340 (5.5) | .311 (4.0) | .580 (8.0) | .588 (9.3) | .220 (3.4) |
| $d_{-1}$ | .099 (3.1) | .034 (1.5) | .010 (0.3) | .060 (1.8) | .071 (2.8) | .066 (0.4) |
| $\hat{\rho}$ | .07 (0.4) | .91 (10.2) | .92 (11.4) | .29 (1.4) | .19 (0.8) | .93 (10.6) |
| $\hat{\theta}$ | 1.0 | 0.8 | 0.9 | 1.0 | 1.0 | 0.3 |
| $R^2$ | .940 | .948 | .947 | .928 | .940 | .961 |
| Std. error | .00382 | .00353 | .00359 | .00414 | .00376 | .00334 |
| Durbin-Watson | 1.82 | 1.79 | 1.67 | 1.57 | 1.56 | 0.96 |

[a] Periods of estimation are: for $\sigma^2$, $\sigma^2_M$ and $\sigma^2_F$, 1949-1971; for $\sigma^2_{25}$ and $\sigma^2_{64}$, 1950-1970; for $G$, 1949-1953, 1955-1969.

Note that the period of estimation differs somewhat, depending on which variant of d is employed. Due to the paucity of data, I decided to use every available data point rather than confine myself to a common sample period (which would have been 1949-1952, 1954-1968).

If $\sigma^2$, 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 inequality measure is significant, and an increase of .03 in $\sigma^2$ (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 in both the current and the lagged specification is $\sigma^2_{64}$, the log variance among males aged 25-64. The current version exhibits trivially small distribution effects, and the lagged version has much larger areas. Other than these, only the regression with lagged $\sigma^2_{25}$ indicates significant distribution effects (at the 10% 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 greater inequality leads to higher consumption. The next section is devoted to convincing the reader that this result 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 will 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 more equal distributions give rise to somewhat less consumption. Is the latter possibility believable?[17]

To begin with pure theory, I showed in section II.1. 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, $\beta$ , exceeds the elasticity of the marginal utility of consumption, $\delta$ . 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.

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

---

[17]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.

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.[18]

A third explanation of the findings, and the one I find most

satisfying, rests on the distinction between the kind of "ideal" redis-

tributions that pure theory envisions and the actual redistributions that

have taken place in the postwar United States.  Table 5 shows the net

change over the entire sample period in each of the six measures of income

Table 5

Postwar Changes in Income Inequality

| Measure | Initial Value[a] | Final Value[b] |
|---------|------------------|----------------|
| $\sigma^2$ | .785 | 1.406 |
| $\sigma^2_M$ | .668 | 1.187 |
| $\sigma^2_F$ | .670 | 1.169 |
| $\sigma^2_{25}$ | .742 | .729 |
| $\sigma^2_{64}$ | .653 | .581 |
| $G$ | .424 | .406 |

[a]1947 for $\sigma^2$ , $\sigma^2_M$ and $\sigma^2_F$ ; 1948 for $G$ ; 1949 for $\sigma^2_{25}$ and $\sigma^2_{64}$ .

[b]1968 for $G$ ; 1969 for $\sigma^2_{25}$ and $\sigma^2_{64}$ ; 1970 for $\sigma^2$ , $\sigma^2_M$ and $\sigma^2_F$ .

[18]Duesenberry ([1949], pp. 44-45).

inequality.  G  is conceptually different from the other variables in two ways.  First, it is a Gini ratio, not a log variance.  But, more importantly, 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 25 years old, hardly changed in the postwar period; the decline in inequality was very slight.  However, inequality fell much more noticeably among prime-age males (25-64 years old) -- 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.  Thus substantial increases in the labor force participation of young people and women appear to have raised inequality by adding many new income recipients to the lower tail of the distribution.

Further information on postwar changes in income distribution is provided by Tables 6 and 7.  Table 6 shows the share in total income received by each age-sex group.  It is clear that the major shift was the decreased importance of men aged 20-44 and the increased importance of women over 45.  Moreover, Table 7 shows that this occurred despite the fact that both young and old men and almost all women lost ground to prime age males in terms of relative average incomes.

In a word, overall inequality among persons with income rose since (a) an increasing share of total income went to groups other than prime age males, while (b) the average incomes received by these groups fell relative to prime age males.  This is very different from the sorts of

"regressive transfers" discussed in Section II.2. above. What the results
of subsection IV.4. seem to say, then, is that prime-age males have rather
higher average propensities to save than the other groups -- not an unrea-
sonable proposition.

## Table 6

### Shares in Total Money Income by Age and Sex

| Age Group | Males 1948 | 1969 | Females 1948 | 1969 |
|-----------|------|------|------|------|
| 14-19 | 1.1 | 1.2 | 0.9 | 0.8 |
| 20-24 | 6.0 | 4.6 | 2.5 | 3.0 |
| 25-34 | 19.4 | 16.5 | 4.2 | 4.1 |
| 35-44 | 21.4 | 18.2 | 4.0 | 4.3 |
| 45-54 | 17.2 | 17.8 | 3.3 | 5.1 |
| 55-64 | 11.5 | 11.7 | 1.8 | 3.8 |
| 65 and up | 5.1 | 5.6 | 1.4 | 3.2 |

Source: Computed by the author from data in U.S. Bureau of the Census
([1967], Tables 14, 36) and U.S. Bureau of the Census ([1970],
Table 45).

Note: Totals do not add to 100% due to rounding.

## Table 7

### Relative Mean Incomes by Age and Sex[a]

| Age Group | Males 1948 | Males 1969 | Females 1948 | Females 1969 |
|-----------|------|------|------|------|
| 14-19 | .20 | .11 | .20 | .09 |
| 20-24 | .54 | .41 | .38 | .28 |
| 25-34 | .83 | .83 | .41 | .33 |
| 35-44 | 1.00 | 1.00 | .42 | .36 |
| 45-54 | .95 | .98 | .44 | .39 |
| 55-64 | .84 | .84 | .35 | .22 |
| 65 and up | .51 | .43 | .23 | .29 |

[a]Ratio of mean income in each group to mean income of males age 35-44.

Source: Same as Table 6.

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 ineqaulity 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. 148-9). The distributional variable in his preferred consumption function is the ratio of income at the 90th

percentile to mean income, and he concluded that "the higher the top decile income relative to the mean, the higher the marginal propensity ot consume" (p. 152).[19]

## 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. Assuming that bequests are a luxury good, such an increase in inequality must reduce consumption. 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 disequalizing redistributions that have taken place in the postwar United States. Finally, the theory (and

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[19]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 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.

the facts) give no particular 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 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 MPC's where constrained to be equal to one another; then a method similar to the Almon lag technique was used to constrain the variation in MPC's; finally, the MPC's 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 increased labor force participation by low-paid demographic groups. And these groups -- teenagers, women and old men -- may well have higher propensities to consume than prime-age males.

As a by-product, this study has shed some additional light on two other properties of the consumption fucntion. In all specifications, I find an absence of money illusion and a negative interest elasticity.

# 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:

$$\int_0^T U[c(t)]e^{-\rho t}dt + B[K(T)] \tag{A.1}$$

subject to the lifetime budget constraint:

$$\int_0^T c(t) \, e^{-rt}dt + K(T) \, e^{-rT} = W \tag{A.2}$$

Defining the functional:

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

the first-order conditions are:[20]

$$\frac{\partial L}{\partial c(t)} = U'[c(t)]e^{-\rho t} - \lambda \, e^{-rt} = 0 \quad \text{for all} \quad t \tag{A.3}$$

$$\frac{\partial L}{\partial K(T)} = B'[K(T)] = \lambda \, e^{-rt} = 0 \; . \tag{A.4}$$

(A.4), of course, is simply the transversality condition since

$$\lambda \, e^{-rT} = U'[c(T)]e^{-\rho T} \quad \text{by} \quad (A.3).$$

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[20] Assuming $U(\cdot)$ and $B(\cdot)$ are strictly concave, these are also sufficient. Also the assumptions $\lim_{c(t)\to 0} U'[c(t)] = \infty$ and $\lim_{K(T)\to 0} B'[K(T)] = \infty$ rule out the possibility of corner solutions.

Solving (A.3) under the specific functional form

$$U[c(t)] = \frac{c(t)^{1-\delta}}{1-\delta} ,$$

gives:

$$c(t)^{-\delta} = \lambda \ e^{-(r-\rho)t} , \tag{A.5}$$

from which it follows that $\lambda$ is related to the initial level of consumption by:

$$c(0)^{-\delta} = \lambda , \quad \text{or} \quad (0) = \lambda^{-\frac{1}{\delta}} . \tag{A.6}$$

Therefore (A.5) becomes:

$$c(t) = \lambda^{-\frac{1}{\delta}} e^{(\frac{r-\rho}{\delta})t} = c(0) \ e^{(\frac{r-\rho}{\delta})t} \quad \text{by} \quad (A.6) ,$$

which is equation (3) in the text.

Now use (A.6) and the specific functional form

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

to express (A.4) as:

$$b \ K_T^{-\beta} = c(0)^{-\delta} e^{-rT} ,$$

or

$$K_T = (b \ e^{rT})^{\frac{1}{\beta}} c(0)^{\frac{\delta}{\beta}} ,$$

which is equation (5) in the text.

Finally, return to the budget constraint, equation (A.2), to write:

$$W - K(T)\ e^{-rT}\ =\ \int_0^T c(t)\ e^{-rt}dt$$

$$=\ c(0)\ \int_0^T e^{(\frac{r-\rho}{\delta})t}\ e^{-rt}dt$$

$$=\ c(0)\ \int_0^T e^{\{\frac{r(1-\delta)-\rho}{\delta}\}t}\ dt$$

which is equation (4) in the text with:

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

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